EUROPEAN CENTRAL BANK

WORKING PAPER SERIES

NO. 541 / NOVEMBER 2005

EUROSYSTEM INFLATION PERSISTENCE NETWORK

CONSUMER PRICE BEHAVIOUR IN LUXEMBOURG

EVIDENCE FROM MICRO CPI DATA

by Patrick Lünnemann

and Thomas Y. Mathä

EUROPEAN CENTRAL BANK

WORKING PAPER SERIES
NO. 541 / NOVEMBER 2005

EUROSYSTEM INFLATION PERSISTENCE NETWORK

CONSUMER PRICE BEHAVIOUR IN LUXEMBOURG

EVIDENCE FROM MICRO CPI DATA [1]

by Patrick Lünnemann [2]
and Thomas Y. Mathä [2]

In 2005 all ECB publications will feature a motif taken from the €50 banknote.

This paper can be downloaded without charge from http://www.ecb.int or from the Social Science Research Network electronic library at http://ssrn.com/abstract_id=828986.

1 This paper was written in the context of the Eurosystem Inflation Persistence Network (IPN), a research network between the European Central Bank and the National Central Banks of the Eurosystem. We would like to thank STATEC for granting us access to their data, Roland Kerschenmeyer and Marco Schockmel for explaining the methodological issues underlying the price collection to us, the IPN members and in particular Luis Álvarez, Josef Baumgartner, Emmanuel Dhyne, Alfred Stiglbauer, an anonymous referee for the constructive criticism and helpful comments. The views expressed in this paper are those of the authors and do neither necessarily reflect the views of the Banque centrale du Luxembourg nor those of the Eurosystem.

2 Banque centrale du Luxembourg, Département Monétaire, Economique et Statistique, 2, Boulevard Royal, L-2983 Luxembourg, Luxembourg; e-mail: Patrick.Lunnemann@bcl.lu; Thomas.Mathae@bcl.lu.

The Eurosystem Inflation Persistence Network

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

Since June 2005 the IPN is chaired by Frank Smets; Stephen Cecchetti (Brandeis University), Jordi Galí (CREI, Universitat Pompeu Fabra) and Andrew Levin (Board of Governors of the Federal Reserve System) act as external consultants and Gonzalo Camba-Méndez as Secretary.

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

Address
Kaiserstrasse 29
60311 Frankfurt am Main, Germany

Postal address
Postfach 16 03 19
60066 Frankfurt am Main, Germany

Telephone
+49 69 1344 0

Internet
http://www.ecb.int

Fax
+49 69 1344 6000

Telex
411 144 ecb d

ISSN 1561-0810 (print)
ISSN 1725-2806 (online)

CONTENTS

Abstract:

This paper uses micro-level price data and analyses the behaviour of consumer prices in Luxembourg. We find that the median duration of consumer prices is roughly 8 months. The median durations of energy and unprocessed food are about 1.5 and 5 months, while prices of services typically change fewer than once a year. For some product types, such as non-energy industrial goods and processed food, a relatively large share of the observed price changes is reverted afterwards. With the exception of services, individual prices do not show signs of downward rigidity. On average, price decreases are as large as price increases. Price changes are determined both by state- and time-dependent factors. Accumulated price and wage inflation, wage adjustment due to indexation, the cash changeover and a larger number of competitors increase the probability of a price change, while pricing at attractive pricing points and price regulation have the opposite effect.

Keywords: Price setting, consumer prices, rigidity, wage indexation, sales

JEL-Codes: E31, C23, C41

Non-technical summary

This paper analyses behavioural patterns of micro consumer prices in Luxembourg and investigates into the explanatory factors of price changes. We compute different measures of price stickiness for more than 230 product categories.

A. We find that the implied median and mean durations are 8 and 12 months. The frequency of price change varies substantially across product types. Whereas energy prices, on average, change every 1.5 months, services' prices change less than once per year. Overall there are no strong signs of downward price rigidity as price decreases represent almost 40 percent of all price changes. In contrast, the corresponding share is less than 20 percent for services.
B. The average size of price changes is relatively large compared to the dynamics in aggregate inflation. The average sizes of both price increases and price reductions are similar.
C. Our results suggest that the overall frequency of price change is substantially affected by price change reversion, which, in turn, is frequently observed during end–of–season sales periods. Price reversions may contribute as much as roughly 40 percent to the frequency of price change in the case of non-energy industrial goods.
D. The analysis reveals a relatively high degree of synchronisation with respect to price changes compared to other euro area countries. This may reflect the compact size of the Luxembourg economy. The aggregate unconditional hazard rate reveals mass points at durations of 1 and 5 months and to a smaller extent at 12 and 24 months. The mass point at 5 months is due to the influence of end–of–season sales.

With regard to the factors explaining price rigidity, our results suggest that both state- and time-dependent factors contribute to the observed pattern of price changes. This applies to price increases as well as to price reductions.

A. With respect to time-dependent factors, the probability of a price change increases at specific truncation lengths (in particular at 1, 5, 6 and 12 months). Interestingly, the probability of a price change is particularly strong in January, but decreases after 2002.
B. Cumulated price and wage inflation significantly increase the probability of a price change. While this extends to price increases in general, no such effects were found for price decreases. In contrast, for services, the rareness of observed price reductions may have to do with fact that wages rarely decline. Hence, services are not necessarily intrinsically downward rigid. Consistent with widespread commercial practices of temporary promotions and sales, the magnitude and the direction of the preceding price change affect the probability of a price adjustment. In particular, the probability of a price change or of a price increase is higher when subsequent to a price reduction. In addition, the probability of a price change and price increase is an increasing function of the size of the former price reduction.
C. Specific events tend to increase the probability of a price adjustment. For example, the event of an automatic wage adjustment due to indexation tends to increase the probability of a price change and of a price increase, and may thus cause inflationary effects. The cash changeover increased the probability of a price change, both upward and downward. However, the former clearly dominated the latter, in particular for services. Attractive pricing policies generally reduce the probability of a price change. Interestingly, they are found to contribute to upward rigidity but less so to downward rigidity. Moreover, the impact of attractive pricing may depend on the currency in use. Similarly, the evidence clearly suggests that product prices considered subject to price regulation are more rigid; this being the case in either direction.
D. The number of competitors is found to add to the probability of price adjustments, in particular downwards. In contrast, the market share reduces the probability of a price change. These results are very much in line with recent evidence from surveys and stress the importance of the competitive environment for the observed degree of price rigidity.

1 Introduction

Nominal rigidities are one commonly cited reason for non-neutrality of money in the short run. They arise because prices of most goods and services do not change instantaneously following shocks, but rather remain constant for a certain period of time. The reasons prices do not adapt instantaneously to new economic conditions are numerous. Firms may prefer not to adjust prices immediately, if the costs involved exceed the benefit of an instantaneous adjustment. Alternatively, firms may not be in a position to adjust their prices instantaneously due to prices being subject to regulation or due to the existence of either written formal (explicit) contracts or informal (implicit) contracts with customers. In line with recent empirical evidence based on surveys for other countries, the latter two are found to be among the most prominent reasons for observing rigid prices in Luxembourg (Lünnemann & Mathä, 2005a).[1] Other possible factors causing price rigidity include constant marginal costs, procyclical elasticity of demand, non-price factors (e.g. adjustment in delivery time or after-sales services), psychological pricing points, coordination failure or kinked demand curves.[2]

Despite progress in modelling nominal rigidities, either as time-dependent price adjustment processes (e.g. Calvo, 1983; Taylor, 1980, 1999), state-dependent price adjustment processes (Caplin & Spulber, 1987), or processes that combine elements of both (e.g. Dotsey et al. 1999), comprehensive analyses based on micro level data have until recently been very scarce. This is due to the lack of individual data and, in particular, due to the restricted access to the national statistical institutes' price statistics. Hence, most available evidence relates to very specific products and markets and, in most cases, refers to the U.S. Specific analyses include for example newsstand prices of magazines (Cecchetti, 1986), retail catalogues (Kashyap, 1995), the refrigerated or frozen orange juices market (Dutta et al., 2002), city-level retail and wholesale gasoline prices (Borenstein et al., 1997) and processed meat products (Ratfai, 2003). Food and retail store prices have among others been analysed by Lach & Tsiddon (1992, 1996) and Kackmeister (2005). Due to the limited number of products and markets these studies can barely be used to derive economy wide generalisations and implications for monetary policy.

In contrast, the use of comprehensive CPI micro data sets is particular rewarding, as it allows firstly analysing the (in-)frequency of price adjustments, the size of such adjustments and the degree to which price changes are synchronized across a broad range of products, sectors as well as types of outlets. Secondly, it allows identifying the areas with relatively strong price rigidities within a given economy. Thirdly, such micro data analysis may contribute to a better understanding of the explanatory factors

[1] The pricing behaviour of firms has recently been analysed in several euro area countries. See Fabiani et al. (2005) for a summary of these studies and Blinder et al. (1998), Hall et al. (2000) and Apel et al. (2005) for similar studies for the U.S., the UK and Sweden.

[2] See for example Blinder et al. (1994) for a comprehensive list of theories.

of price rigidity. Finally, individual price data are essential for understanding the short-term impact of monetary policy and for the development of macro models consistent with micro data evidence.

The most prominent recent empirical contributions in this respect are those by Bils & Klenow (2004) and Klenow & Kryvtsov (2004) who study retail price stickiness in the U.S. using monthly BLS consumer price data. For the euro area, the empirical evidence has until recently been particularly scarce. During 2004 and 2005, however, several papers analysing consumer price behaviour in individual EMU countries have been released within the Eurosystem Inflation Persistence Network (IPN). The present paper is part of this initiative and analyses the behaviour of consumer prices in Luxembourg.[3]

Apart from providing descriptive statistics for some key indicators, such as the price change frequency, the size of price changes and their synchronisation, and analysing commonly used time- and state dependent factors in explaining the probability of a price change, this paper emphasises following issues and results: First, a large fraction of the observed price changes are price change reversals caused by temporary or end–of–season sales. It is questionable whether these price changes reflect true underlying price flexibility. Second, we analyse the effect of the cash changeover. The results suggest that firms strategically brought forward or postponed price adjustments. Third, the event of an automatic wage adjustment due to indexation, which is in place in Luxembourg to date, increases the probability of a subsequent price increase, and may thus cause inflationary effects. Fourth, the currency in use may systematically influence the observed rigidity of attractive prices. Fifth, services are found to be more rigid than other sectors. However, the results suggest that the rareness of observed price reductions in this sector may be due to the fact that wages rarely decline, and thus that services prices are not intrinsically downward rigid. Sixth, the price change flexibility is larger, the larger the number of competitors. This stresses the importance of the competitive environment for the observed degree of price rigidity.

This paper is organised as follows. Section 2 gives an overview of the data used in the Luxembourg national index of consumer prices (NICP) and describes the definition of variables and of methods used. Section 3 presents descriptive evidence for price durations, the size of price changes and the degree to which price changes are synchronized. Section 4 summarizes the time- and state-dependent, as well as other factors thought to determine the probability of observing a price change before proceeding with an econometric analysis thereof. Section 5 concludes.

---

[3] Companion papers, for which similar country-specific studies have been conducted within the network, are those by Baumgartner et al. (2005) for Austria, Aucremanne & Dhyne (2004, 2005) for Belgium, Baudry et al. (2004) and Fougère et al. (2005) for France, Vilmunen & Laakonen (2005) for Finland, Hoffmann & Kurz-Kim (2005) for Germany, Veronese et al. (2005) for Italy, Jonker et al. (2004) for the Netherlands, Dias et al. (2004, 2005) for Portugal and Álvarez et al. (2004, 2005) for Spain. Dhyne et al. (2005) provide cross-country information for a small selection of consumer goods.

2 Data, variables and methods

2.1 Data source

The data used in this study refer to individual consumer prices, as collected by the national statistical institute in Luxembourg, the *Service Centrale des Statistiques et des Etudes Economiques* (STATEC), for the purpose of compiling the NICP. The construction of the index of consumer prices in Luxembourg encompasses all COICOP 2-digit categories and is based on approximately 7.500 individual price quotes per month. In Luxembourg, the prices of goods and services are centrally collected by STATEC from a group of outlets that represents as closely as possible the structure of consumer expenditure. The sample of goods and services is defined on the basis of consumption surveys as well as on other information sources, such as retail trade sales figures and, ultimately, expert opinion. The weights reflect the shares in consumption expenditures used for the compilation of the NICP and are revised annually. Hereafter, a product category is generally defined as a 10-digit COICOP code. Overall, we distinguish more than 230 product categories.[4]

2.2 Sample period selection

Three major events stand out that may have affected the compilation of consumer prices in Luxembourg and/or the price setting practices of Luxembourg firms in recent years. First, since January 1999, prices quotes reported by STATEC take account of sales prices. The integration of sales prices is due to a methodological revision with regard to the collection of consumer prices underlying the compilation of the NICP.[5] In Luxembourg, end–of–season sales take place in January and in July. As illustrated in Lünnemann & Mathä (2004), the introduction of sales prices may fundamentally change the time series properties of price indices and may have strong implications for the degree of measured inflation persistence. Second, the inception of the single monetary policy in January 1999 characterises a general change to the monetary regime that may have affected the aggregate inflation process as well as the individual firm's price setting behaviour. Third, and similar to all other euro area countries, the price setting behaviour of firms may have been affected by the introduction of euro banknotes and coins in the run-up to the euro cash changeover on 1 January 2002 (e.g. BCL, 2002, 2003).

In order to avoid structural breaks due to the new monetary policy regime or the introduction of sales and in order to leave a sufficient number of observations both before and after the cash changeover, the sample period considered ranges from January 1999 to December 2004.

[4] For some product categories, such as fresh fruit and fresh vegetables, product categories are defined at higher levels of aggregation (e.g. 6-digit COICOP).

[5] This holds for both the HICP and the NICP. The latter is used throughout this paper

2.3 Product and price information contained in sample

For each individual price collected, the data set entails the following information: an item code, a price code, a point of sale code (includes a department (within store) code), a price type code (normal/estimated/missing etc.), a product description and a quality code (quality adjustment). All prices included in the dataset are given in EUR. The information set entails essentially the raw data collected by STATEC. Initially, the dataset encompasses all price quotes collected over entire sample period.

For the large majority of products included in the NICP, prices are collected at monthly frequency. Exceptions apply to selected product categories, for which prices are collected at quarterly frequency (e.g. package holidays, selected clothing articles) or twice a year (e.g. housing rents). The treatment of prices of seasonal products (e.g. cherries, skis) while not available follows specific particularities. Hereafter, these items are excluded from the analysis and – given their small weight in the NICP – their exclusion does not substantially affect the overall results. In addition, some price quotes were removed from the sample due to the way data have been coded and stored by STATEC.[6] In total, the remaining sample contains more than 380.000 price quotes representing approximately 85% of the total NICP (see Table 1). The removal of selected prices slightly affects the product type weights relative to the full sample. Relative to the full NICP, services are slightly under-represented, but nevertheless account for approximately 30 percent of all price quotes.[7]

2.4 Definition of price trajectories and price spells

A price *trajectory* refers to a series of price quotes for a specific article of a specific brand in a specific outlet. Price trajectories can be divided into *price spells*, i.e. time periods of equal prices for a specific article of a specific brand in a specific outlet. Price spells may be *uncensored* (i.e. the price spell starts and ends with a price change), *left-censored* (i.e. the price spell ends with a price change and its starting date is unknown), *right-censored* (i.e. the price spell starts with a price change and its end date is unknown) or *double-censored* (i.e. both start and end date of the price spell are unknown). Censoring may lead to a downward bias in the estimation of the duration of price spells, as spells with a long duration tend to be overrepresented in the class of censored price spells. Put differently, censoring is likely to lead to an overrepresentation of "short duration" spells. Censoring may result from a number of reasons. In particular, price spells may become truncated as a given product or shop disappears. Hereafter, a product replacement within a given shop is not considered as a source of censoring. In-

[6] An unambiguous identification of the sales point was impossible for shops that do not exist anymore and for shops in which no more than one price is recorded. Price observations made at shops not existing in December 2004 are not considered.

[7] The number of prices collected has no immediate implications for the weight given to that product category within the overall index. Prices subject to regulation, such as liquid fuels, may be identical across all sales points and recorded once at a time only. Yet their weight in the full CPI may be relatively large.

stead, in-shop product replacement is considered equivalent to a price change.[8] In contrast, truncated price spells due to a store replacement are considered as censored price spells. This as the end of a price spell due to the disappearance of an outlet is clearly unaffected by the outlet's pricing policy.

2.5 Measures of interest

2.5.1 Frequency of price changes

We compute the frequency of price change (or the implied duration), the size of price changes (when they occur) and the degree to which price changes are synchronized. In order to take particular account of the assumed phenomenon of downward rigidity, we distinguish at any stage between price decreases and price increases. In general, each price quote p_{ijt} is characterized by the product characteristics (article code i, where $i = 1$ to n_j), its selling (outlet code), the date of the quote and the corresponding product category j. F_j denotes the frequency of price change for product category j. It is defined as the fraction of price changes between period t and period $t+1$ relative to the potential number of price changes. A distinction into price decreases and price increases yields F_j^+ (frequency of upward price adjustment) and F_j^- (frequency of downward price adjustment).

In what follows, we adopt the frequency approach of measuring price durations. This is the approach adopted by Bils & Klenow (2004) and other papers within the IPN. It offers a number of advantages over the duration approach. The main advantage of the frequency statistic over the average duration of price spell is that it uses all relevant statistical information at hand.[9] Importantly, the frequency approach does not require an explicit treatment of censoring, and provided that censoring is independent of the duration process, the inverse of the frequency is a consistent estimator of the average duration (e.g. Dias et al., 2004). In addition, computing the frequency of price changes does not require long time series (as long as the assumptions of homogeneity and stationarity remain valid), as it is possible to estimate price durations even if the number of price quotes is smaller than the average duration of a price spell, while still being less sensitive to specific events, such as VAT rate changes.

In contrast, the direct computation of the average duration of price spells must take into account uncensored price spells only (i.e. price spells that both start and end with a price change). The restriction to uncensored spells, however, implies discarding relevant information when computing the average duration. Furthermore, the obtained values may be subject to a selection bias, as longer price spells are more likely to be censored and, hence more likely to be discarded (e.g. Álvarez & Hernando, 2004; Aucremanne & Dhyne, 2004; Dias et al., 2004).

---

[8] Note though that if the in-shop product replacement involves a quality change, the size of the price change is ignored in the calculation of the average size of price changes.

[9] If a price spell is censored the computation of the price change frequency only discards the price changes for which the price is not observed (e.g. Dias et al. 2004).

Assuming stationarity and homogeneity of price change behaviour in the cross-sectional dimension, a convenient property is that the inverse of the frequency of price changes converges asymptotically to the mean duration in a large sample (e.g. Baudry et al., 2004). In continuous time representation and under the assumption of a constant probability of a price change throughout a given month, the average duration $\bar{S}_j$ and median duration $S_{50,j}$ can respectively be written as:

$$\bar{S}_j = -\frac{1}{\ln(1-F_j)} \qquad \text{and} \qquad S_{50,j} = \frac{\ln(0.5)}{\ln(1-F_j)}.$$

2.5.2 *Hazard rate*

The hazard rate allows a more detailed description of the distribution of the length of price spells, as it is able to capture duration dependence and changes in the probability of a price change as the elapsed duration of the spell is exceeded. The hazard rate *h(s)* reflects the conditional probability that a price spell, having lasted until duration *s*, will end instantaneously after duration *s*. The hazard function can be written as:[10]

$$h(s) = \lim_{ds \to 0} \frac{P(S < s + ds \mid S \geq s)}{ds} = \frac{f(s)}{1 - F(s)},$$

where F is the cumulative density function of the completed spell duration S and f is the corresponding distribution function.

2.5.3 *The size of price changes*

The size of a price change is defined as the log difference between prices observed in consecutive periods for identical products in identical shops. This is to obtain identically sized price increases and decreases in case of price changes that are reversed afterwards. While the use of log differences may overstate the true (percentage) difference in case of extreme price adjustments, the distribution of the size of percentage price changes is necessarily asymmetric; bounded by -1 but unbounded from above. As for the frequency of price changes, we assume prices to change once per month at most.

2.5.4 *The synchronisation of price changes*

The degree to which price changes are or are not synchronised is measured by means of the synchronisation ratio. In case of perfect synchronisation of price changes, the proportion of price changes at time *t* is either equal to 1 or to 0. With the average frequency of price changes over the sample period being equal to F_j, in the case of perfect synchronisation, all the firms change their price simultaneously in F_j percent of the cases, whereas they do not change their price in 1- F_j percent of the cases. Follow-

[10] See for example Kiefer (1998) for a survey on duration data and hazard functions.

ing Fisher & Konieczny (2000) the theoretical value of the standard deviation of the proportion of price changes over time in case of perfect synchronisation may be calculated as

$$SDMAX_j = \sqrt{F_j(1-F_j)^2 + (1-F_j)(0-F_j)^2} = \sqrt{F_j(1-F_j)}.$$

This expression gives the upper limit for the standard deviation of the proportion of price changes. Similarly, in the case of perfect staggering, a constant proportion F_p of firms changes its price each month and the standard deviation of the proportion of price changes over time is equal to 0. The effective standard deviation of price changes for product category p is given by

$$SD_j = \sqrt{\frac{1}{\tau - 1}\sum_{t=2}^{\tau}(F_{jt} - F_j)^2},$$

with τ being equivalent to the number of months for which prices are observed. The synchronisation ratio of product classification j is defined as the ratio of effective standard deviation to the theoretical maximum standard deviation of price changes

$$SYNC_j = \frac{SD_j}{SDMAX_j}.$$

The degree of synchronisation is an increasing function of $SYNC_j$. Similar expressions can be derived for $SYNC_j^+$ and $SYNC_j^-$, the synchronisation ratio of price increases and price decreases.

2.5.5 Aggregation of indicators

Aggregate indicators are obtained across weighting figures for the corresponding sub-aggregates. In general, weighting is done at the product category level, as no weights are available for single products. In general, product category specific estimates are unweighted averages obtained at the 10-digit COICOP level by using all price quotes. The weights refer to the NICP. Thus, the estimate for price duration at the aggregate level is given by the expression

$$\bar{S} = \sum_j w_j \cdot \frac{1}{F_j} = \sum_j w_j \cdot \bar{S}_j,$$

where w_j denotes the weight of product category j in the Luxembourg NICP in 2000. Note that, for some product categories, the frequency of price change (in particular that of price decreases) is close to zero, thereby implying very long durations. A relatively small number of indices may therefore substantially affect the estimate of aggregate duration. Hence, we use weighted medians when computing aggregate figures. This measure is also used by Bils & Klenow (2004) and is less sensitive to the exis-

tence of few items with very low frequencies of price change. Nevertheless, this measure cannot be interpreted as an estimator for the average price duration.[11]

3 Main findings for key indicators

3.1 Aggregate developments of consumer prices between 1999 to 2004

The sample period is characterized by moderate inflation rates. Overall, the annual increase in the price level was on average approximately 2.2% (see Figure 1). During the period 2002 – 2004, the increase in the price level (2.1% – 2.2%) remained close to the upper level of the definition of price stability adopted by the Eurosystem. The measured aggregate inflation rate remained very low in 1999 (1.0%), whereas it exceeded 3% in 2000.[12] In addition, there is a strong degree of heterogeneity across different product categories (see Table 2). At the 2-digit COICOP level, the highest average annual inflation is recorded for '*education*' (4.1%), whereas the lowest average is obtained for '*communications*' (-6.1%). The annual inflation rates are always positive except for '*communications*' and '*health*'. Similarly, m-o-m inflation rates tend to be positive, only in the case of '*communications*', '*housing, water, electricity, gas and other fuel*' and '*recreation and culture*', we observe a m-o-m decrease in the price level in 1 out of 3 months or more for the period under investigation.

As there are substantial differences in the relative importance of price decreases and the magnitude of price changes across consumer price categories in Luxembourg, we expect to find strong differences with respect for the previously described key indicators.

3.2 Frequency of price change

Table 3 reports a choice of key indicators for the 12 COICOP-2 groups and product types, while Table 4 sketches the frequency distribution by presenting weighted percentiles of the frequency of price changes across the total NICP and the five different product types. The weighted overall frequency of price change is 17%, of which approximately 63% and 37% are price increases and decreases. Overall, the frequency of price change may range from values close to 0% to values larger than 90% (mostly energy-related products).[13] As shown in Table 4, the frequency of price change is 12% at the weighted median, while the weighted 25 and 75 percentiles are approximately 7% and 18%. In weighted terms, the median frequencies of price increases and decreases are 8% and 4%, respectively.

Despite the relatively high overall frequency of price change, for roughly 1 in 3 product categories in the NICP (making up 30% of the sample in terms of weight), prices change once a year at most. A

[11] See for example Baudry et al. (2004).
[12] The low value for 1999 may be the result a base effect originating from the inclusion of sales prices from January 1999 onwards.
[13] Note that the number of observations may differ substantially across product categories.

large majority of products categories exhibit price changes fewer than four times a year. Only 10% of the 234 product categories (in terms of weight) reveal 5 or more price changes a year, while for 70% of the product categories, prices change on average twice or less often per year.

Consistent with evidence for other euro area countries (e.g. Baumgartner et. al. 2005), energy prices change by far most frequently (52%). A distant second place goes to unprocessed food with 26%. Non-energy industrial goods and processed food trail further behind with averages of 16% and 12%. Lastly, services' prices change rarely (7%), which again is also reported for other euro area countries (e.g. Álvarez & Hernando, 2004). Both energy and unprocessed food exhibit are very skewed distributions, as can be seen in Table 4. First, the weighted 25 and 50 percentiles of unprocessed food are below those of processed food and non-energy industrial products, while the weighted 75 percentile is close to energy products. Second, for both energy and unprocessed food, the difference in the frequency of price change between the weighted 25 and 75 percentiles is 44 and 52 percentage points, which is in startling contrast to the other three product types.

The same ranking holds, in absolute terms, for price increases and price decreases. The relative share of price decreases in price changes differs across product categories. For services the share of price decreases is less than 20%, while for unprocessed food and non-energy industrial goods it is more than 40% (see Table 3). Whereas for a dozen categories, all price changes are entirely characterised by a rise in prices, for another dozen categories price decreases were observed more frequently than price increases. For a large fraction of the NICP basket, though, the share of price decreases in price changes is between 0.3 and 0.5 (see Table 3). Strikingly, processed food and services are not only characterized by the lowest frequency of price changes, but also reveal the highest ratio of price increases to decreases.

3.3 The role of price change reversals

Whereas the above measures of price frequency apply to all price changes recorded over the sample period, the nature of the price changes may differ. On the one hand, firms may choose to adjust prices based on a thorough price review taking stock of the fundamentals of their pricing polices. On the other hand, firms may decide to change prices for other reasons (e.g. institutional habits, stock clearance, etc…). Examples for such price changes are typically found during periods of promotional or end–of–season sales. During sales periods, firms often reduce prices and revert to the pre-sales or regular prices afterwards.[14]

[14] To give a specific example, consider a price that is reduced from EUR 9.99 in December to 8.99 in January and with the end of the promotional activity is raised again to EUR 9.99 in February.

Promotional sales and end–of–season sales are a common phenomenon, in particular for retail and food products.[15] Baumgartner et al. (2005) report for Austrian micro CPI data, that price changes due to temporary promotions and end-of-season sales account for 5.5% of all price quotes and 2.3% of the weighted aggregate price change frequency. They are typically encountered in the product categories '*food and alcoholic beverages*' and '*clothes and footwear*', respectively. Using prices quotes collected for the compilation of the U.S. CPI, Klenow & Kryvstov (2004) report that roughly 11% of price quotes are sales prices. Sales prices are particularly frequently encountered for food items, where they comprise 15% of all price quotes and 2/3 of the sales prices revert to their regular prices in the next period. Similarly, Warner & Barsky (1995) report for specific goods in selected U.S. retail stores that prices of products in temporary sales are most often reverted afterwards.

In order to single out the frequency of price change due to price change reversals, we split the overall frequency of price change into three components, namely a) the frequency of price changes due to reverting price changes during the end–of–season sales in January/February and July/August, b) the frequency of price changes due to reverting price changes during periods other than end–of–season sales and c) the frequency of price changes resulting from price changes that are not reverted afterwards. Hereafter, price change reversals are considered equivalent to price changes that are exactly offset by the following price change, regardless of the time elapsed until reversal.

Table 3 and Figure 2 illustrate the results from the decomposition exercise at the COICOP 2-digit level and for all product types. Overall, approximately one out of four price changes is offset afterwards. In other words, when eliminating price changes due to price change reversals, the weighted overall frequency of price change declines from 17% to 12%. Most price change reversals take place during end–of–season sales periods. In particular, prices of non-energy industrial goods switch particularly often between two levels during end–of–season sales periods. Indeed, for some COICOP-2 product categories (e.g. '*cp02 clothing and footwear*' and '*cp05 furnishings, household equipment & maintenance of the house*'), switching between regular and sales prices is the main source of price changes (i.e., up to 70% of all price changes). As expected, the relative importance of price change reversals differs substantially across product types. Energy products, unprocessed food and services are by and large unaffected by price change reversals due to end–of–season sales.

Price change reversals in periods other than end–of–season sales are relatively frequent for food products (more than 1 in 6 price changes), but less so for energy products and non-energy industrial goods

[15] For theoretical explanations for temporary promotions and end–of–season sales, see for example Varian (1980), Salop & Stiglitz (1982), Lazear (1986) and Pashigian (1988).

(1 in 10 price changes or fewer).[16] For the product categories '*food and & non alcoholic beverages*' and '*alcohol, tobacco & narcotics*', these price change reversals are likely to characterise promotional sales activity and contribute significantly to their overall frequency of price changes. This particular pricing behaviour is particularly frequently encountered in supermarkets and retail stores, and is in line with results reported elsewhere (e.g. Baumgartner et al. 2005).

Hence, price change reversals are a common phenomenon. They contribute substantially to the overall frequency of price changes, although they may only be of limited use for the inference of the degree of price rigidity. In particular, from a monetary policy perspective, the relevance of such price changes remains questionable.

3.4 The average size of price changes

In weighted terms, the average size of price increases and price decreases has been roughly equivalent (both around 8%). Hence, price increases and price decreases are quite sizable relative to the overall inflation rate. There are substantial differences in the size of prices changes across product types. With both average price reductions and increases of 4%, they tend to be smallest for energy products (see Table 3). The largest average price changes are observed for unprocessed food products (+13% and –17% for price increases and reductions). For the remaining three categories (processed food, services and non-energy industrial goods), the average price increases are relatively close to the overall 8%. They differ with respect to the average size of price reductions though. At 8%, price reductions and increases are relatively similar for non-energy industrial goods. For processed food, the price reductions (–11%) are on average more sizeable than price increases (+8%), while price reductions (–5%) tend to be smaller than the price increases (+7%) for services.

3.5 Synchronisation ratios

Overall, the weighted synchronization ratio of price changes is 0.5 (see Table 5). These results suggest a relatively high degree of price change synchronisation in Luxembourg compared to other euro area countries (see for example Dhyne et al., 2005). This may, however, relate to the relatively compact geography and the relatively small distances between outlets. As indicated by Veronese et al. (2005), in larger countries, synchronisation ratios tend to increase when computed at the regional level (rather than at the national level). Again, substantial differences occur across product types. Energy related products reveal the strongest degree of price change synchronisation. Due to the fact that a number of energy related products are subject to some form of price regulation, the synchronisation ratio for price changes, price reductions and price increases are approximately 0.9.[17] Second and third to energy, the

[16] For some categories at the finer COICOP-10 level, non-sales driven price change reversals may contribute up to 40% to the total frequency of price changes.

[17] In fact, using a more detailed product-specific categorisation the synchronisation ratio is 1.

synchronisation ratios for service prices and for non-energy-industrial goods are 0.56 and 0.49. In the case of non-energy industrial goods, the relatively high ratio of synchronisation may be due to time-dependent phenomenon of end–of–season sales. Below average price synchronisation is obtained for processed food (0.32) and unprocessed food (0.25). The degree of synchronisation is broadly comparable for price increases and price decreases for energy, non-energy industrial goods and unprocessed food. Services prices tend to change in a more synchronised manner upwards than downwards with the difference in the synchronisation ratios exceeding 15 percentage points. Here, the high degree of synchronisation with respect to price increases may to some extent be related to the periodic wage adjustment due to indexation.

4 The probability of a price change: an econometric investigation

In structural macroeconomic models, nominal price rigidities are commonly implemented by incorporating time- or state-dependent price adjustment processes. In this context, time dependency is typically obtained by assuming that prices are set for a fixed period of time (e.g. Taylor, 1980 & 1999) or that each period a certain fraction of firms adjusts its prices with the distribution of price adjustments following a Poisson process (e.g. Calvo, 1983) or by combining properties of both the Taylor and the Calvo model ("truncated Calvo model"). A Calvo price adjustment process implies a constant probability of price change, whereas Taylor type price adjustment process implies that a firm only changes its price when the fixed contract duration is reached.

The above named models have been criticised for the exogeneity of the price setting intervals, as in reality the price setting mechanism is likely to depend on the state of the economy. State-dependent models endogenise the price setting mechanism. Faced with a (fixed) cost of price adjustment, firms decide each period dependent on the state of the economy whether or not to change their price and by how much (e.g. Caplin & Spulber, 1987). Dotsey et al. (1999) present a model that combines properties of Calvo price adjustment processes with state-dependent pricing features. In particular, they allow the fraction of firms that changes prices to increase with increasing inflation rates; a feature that is generally supported by the empirical evidence (e.g. Dhyne et al., 2005). These models bear important implications with regard to the probability of observing a price change at a given time *t*.

4.1 Factors determining the behaviour of individual consumer prices

In this section, we identify explanatory factors that are thought to affect whether or not prices are adjusted at a given point in time. We distinguish between time- and state-dependent factors, as well as other factors such as attractive pricing policies. Hereafter, we briefly discuss these factors before proceeding to more formal estimation techniques.

4.1.1 Time-dependent factors

4.1.1.1 UNCONDITIONAL HAZARD RATES

Overall, the hazard plot, based on non-weighted survival estimates presented in Figure 3, demonstrates mass points at durations of 1 and 5 months and – though less prominent – 12 and 24 months (see Figure 4). The overall hazard rate is slightly downward sloping. These mass points reveal the existence of a substantial degree of time dependency, consistent with truncated Calvo price adjustment processes. Mass points at 1, 12 and 24 months but also the negative slope of the hazard function are similarly found for other euro area countries (see e.g. Álvarez et al., 2005, Dhyne et al., 2005), whereas the mass point at d = 5 months seems to be specific to Luxembourg. This may to some extent be explained by the fact that some of the datasets used in other euro area country studies do not include sales prices.[18]

The product type specific charts reveal that the 5-month peak in the hazard rate is primarily due to non-energy industrial goods. More than 70% of the price spells non-prevailing after 5 months originate from this category (see Figure 3). This spike highlights the importance of end–of–season sales for price setting of non-energy industrial goods in Luxembourg. End–of–season sales apply to a very large range of products in Luxembourg and it seems that prices often switch between sales prices (in January and July) and regular prices (from February to June and from August to December). Hazard rates for duration d = 1 – to a lesser extent also for d = 2 – are most important for those categories characterized by a high frequency of price change (i.e. energy and unprocessed food). Energy products reveal a particularly strong mass points at 11 months, thereby illustrating the high degree of heterogeneity in this category. Contrary to all other product types, services have their peak hazard rate at d=14. At the same time, and contrary to all other product types, the hazard rate for services reveals a mass point at d = 6 to 7 months. At the same time, the hazard rates at durations d = 1 and d = 2 are relatively small.

4.1.1.2 INTRA-YEAR PRICE CHANGE FREQUENCIES AND SEASONALITY

As reported in Lünnemann & Mathä (2004) at the level of disaggregate consumer price indices, the frequency of price changes (as well as the other indicators analysed in section 2) may display seasonal patterns. Overall, the weighted frequency of price change is highest in January. Approximately 13% of all price changes occur in January (see Figure 5). In weighted terms, the frequency of price change is relatively important in February, July and August, too. It seems that even at the aggregate level, sales are an important factor in driving price changes. Each of these months subsumes more than 10% of all

[18] This is for example the case for the Belgian and Spanish dataset analysed by Aucremanne & Dhyne (2004, 2005) and Álvarez & Hernando (2004) and Álvarez et al. (2005), while Baumgartner et al. (2005), using existing information (available for sub-period of the dataset only) on whether or not a price constitutes a sales price, decide to omit the latter in the analytic part, by arguing that they may not be relevant from a monetary policy analysis point of view.

price changes. February and August reveal the highest frequency of price increases (approximately 17%) whereas in January (16%) and July (14%) the highest frequency of price reductions is observed. The share of price cuts in price changes is equal to and smaller than 20% in February and August. Only in January (55%) and July (65%) price cuts are more frequent than price increases.

The relevance of end–of–season sales becomes particularly prominent for non-energy industrial goods. More than 1 out of 4 prices decline in January and July and the same fraction increases in the subsequent months February and August. The share of price cuts in price changes is larger than 75% in January and July, whereas in August and in February price increases account for more than 90% of all price changes. The frequency of price change ranges from 4% in December to almost 40% in January. At the same time, the frequency of price changes is higher for non-energy industrial goods than for processed food during end–of–season sales only. In contrast, unprocessed food prices are relatively unaffected by seasonality. This holds both in relative and in absolute terms. In this product category, the frequency of price change ranges from 24% in December to 29% in April. The fraction of price decreases varies from 38% to 50% (note that the corresponding bandwidth for non-energy industrial goods is from less than 10% to more than 80%).

The lower left hand side panel of Figure 5 illustrates the special character of the seasonal pattern in price change frequency for service related products. In January, 1 out of 7 services' prices change (a price frequency higher than that obtained for processed food in the same month). Approximately 1 in 6 price changes occurs in January. Almost 80% of all price changes in January constitute price increases. The share of price cuts in price changes is always smaller than 30%. In December, the share of price cuts is as small as 5%. In absolute terms, the frequency of price cuts in December is 0.3%.

The role of the calendar month in determining the frequency of price change is confirmed by simple categorical regressions across the set of product categories (not shown). From February to December, the frequency of price change is significantly lower than in January. The size of the frequency differential is approximately 20 percentage points for all months not concerned by end–of–season sales. The remaining three months affected by end–of–season sales reveal frequencies of price change significantly lower than January, the differential being substantially smaller though (approximately 6 to 8 percentage points).[19]

[19] For more details on the seasonal patterns of price changes, refer to section 4.3.

4.1.2 State-dependent factors

4.1.2.1 The level of inflation

Recent research on micro consumer prices has shown that the frequency of price change positively depends on the level of inflation prevailing over the sample period (e.g. Woodford, 1999). Whereas some papers find a higher frequency of price change for periods of higher (aggregate) inflation, others find that for a given product category the frequency of price change increases with higher inflation rate specific to the product category. For example, Cecchetti (1986) reports that the average duration of US magazine prices was 7 years during the 1950s, a period of low inflation, while it was much lower at 3 years during the 1970s, a period of high inflation. Similarly, Lach & Tsiddon (1992) report that the average duration between price changes is shorter in periods of high inflation

This dependency of the price adjustment frequency on the (cumulated) inflation rate is also theoretically founded. Cecchetti (1986), for example, provides a theoretical pricing model, where the probability of price change depends on the size of the difference between the actual and the desired price. If this difference exceeds a certain threshold h firms decide to adjust prices, as the gain from adjusting outweighs the costs of the price adjustment. He derives an estimable function where the probability of price change is related to the accumulated inflation and the elapsed time since the last price change and the size of the previous price change. Aucremanne & Dhyne (2005) and Baumgartner et al. (2005) provide empirical support for the existence of such a relationship.

4.1.2.2 Characteristics of last price change

The direction and the size of the previous price change may contain information about the next price change to be expected (e.g. Cecchetti, 1986). On the one hand, a large previous price change may indicate that the adjustment costs of changing prices are rather high. Hence, the firm is expected to change prices infrequently and by large amounts. A small previous price change, in turn, may indicate that the adjustment costs are rather small or convex, as in Rotemberg (1982), and correspondingly we would expect prices to change relatively frequently. The probability of a price change in the next period would be expected to negatively depend on the size of the previous price change. Furthermore, in a state of positive aggregate inflation, we would successive price changes expect to exhibit a positive sign. Carlton (1986) for example reports a positive correlation between price rigidity and the average absolute price change for selected intermediate products in manufacturing. The less frequently prices are adjusted, the greater is the price change when prices change.

On the other hand, the size of the previous price change may signal that large price changes follow each other very frequently. A typical feature of such a price change sequence would also entail switching of signs in price changes. Recent micro consumer price studies for Belgium and Austria have, indeed, confirmed that the probability of a price change is larger if the previous price change was a price

reduction. This is consistent with widespread commercial practices, in particular with temporary promotions or end–of–season sales. Furthermore, the size of a prior price reduction increases the probability of a price change by more than the size of a prior price increase (e.g. Aucremanne & Dhyne, 2005; Baumgartner et al., 2005). Sizeable markdowns are a typical feature of sales and promotions in general.

4.1.2.3 THE AUTOMATIC WAGE INDEXATION MECHANISM

Contrary to all other European countries, in Luxembourg, wages and pensions are integrally indexed to the national index of consumer prices.[20] Whenever the 6-month moving average of the national consumer price index exceeds the level at which the former "wage indexation" took place by more than 2.5%, wages and pensions are adjusted.[21] The adjustment takes place as of the beginning of the month following the breach of the 2.5% threshold. Thus, the non-negotiated wage indexation mechanism is completely backward looking.

In general, the wage indexation mechanism in Luxembourg does not distinguish between different sources of inflation (in particular, it does not depend on whether there has been a shock to some domestic variable, such as higher consumer prices due to higher VAT rates, or to some exogenous variable, such as higher oil prices). However, the Luxembourg Government may temporarily choose to deviate from a purely mechanical implementation of the automatic adjustment mechanism, as for example in the wake of the devaluation of the Belgian and the Luxembourg Franc in 1982 (e.g. Fontagné, 2004).

The wage adjustment, in turn, may have a direct and an indirect impact on the behaviour of consumer prices. The adjustment is directly passed through to consumer prices that take the form of wages, in particular in domestic services. Higher adjusted wages may – depending on the degree of competition (in particular from abroad) – be indirectly passed on to consumer prices. Nevertheless, as suggested by the 2004 survey among Luxembourg firms, firms may be unable to instantly adjust prices due to, for example, explicit or implicit contractual agreements with their customers (Lünnemann & Mathä, 2005a). In practice, the impact of wage indexation on the frequency and on the size of price adjustment in Luxembourg is not obvious. Besides, wage indexation and its effect on price adjustments have so far rarely been studied and empirical evidence on this issue is scarce.

In theory, and similar to the euro cash changeover or VAT rate changes, automatic wage adjustments are exogenous shocks to the individual firm. These adjustments do not come as a surprise to Luxem-

[20] Prior to the law of 27 May 1975, and since 1921, indexation was applied to salaries of civil servants and railway staff only (e.g. Adam and da Costa, 2002).

[21] Strictly speaking, a small number of factors affecting the NICP are not considered when computing the threshold consumer price level at which the new wage adjustment takes place.

bourg firms, however. While some degree of uncertainty may exist with respect to the exact date of the next adjustment, firms have the possibility to broadly anticipate the next wage adjustment, as the present level of inflation and inflation forecasts by the national statistical institute and the Central Bank of Luxembourg are publicly accessible.[22]

We expect the automatic wage indexation mechanism to have an influence on firms' price setting behaviour. Besides, indexation may also apply to contracts other than explicit wage agreements, such as insurance contracts, rents and prices underlying construction projects. Throughout the sample period, 6 automatic wage adjustments took place, namely on 1 August 1999, 1 July 2000, 1 April 2001, 1 June 2002, 1 August 2003 and 1 October 2004. We expect wage indexation to be of particular importance in the services sector, as the labour share in services is particularly high. Furthermore, services included in the NICP are mostly domestic services and are not subject to international competition as is the case of industrial activity.[23] Moreover, as reported in Lünnemann & Mathä (2005a), Luxembourg firms in the services sector are more likely to make use of rules of thumb when recalculating prices and are less forward looking than the average Luxembourg firm, in particular in the case of medium-sized and large firms. In addition, firms in the services sector judge wage indexation to be the second most important reason for price increases, while wage indexation seems of lower importance in other sectors (except for construction).

4.1.2.4 EURO CASH CHANGEOVER

The euro cash changeover at the beginning of 2002 can be thought of an economy-wide shock potentially affecting all firms' prices setting behaviour. As firms had to adjust the nominally displayed prices anyway, the costs of genuine price changes seem trivial. Some Eurosystem NCBs and NSIs analysed the inflationary effects of the euro cash changeover as media attention increased consumers' fears of firms unduly increasing prices.[24] These studies showed firstly that a large share of prices in national currency consisted of attractive prices, and secondly that large price increases were primarily confined to the services sector (e.g. Folkertsma et al. 2002; BCL, 2003; Cornille, 2003).

Jonker et al. (2004) report for Dutch CPI data that the hazard ratio for a price spell ending in December 2001 relative to the baseline hazard was double as high, clearly indicating the impact of the conversion to euro on the pricing behaviour of firms. Baumgartner et al. (2005) report for Austrian CPI data that the probability of a price change in January 2002 was 1.8 percentage points higher. In general, the cash changeover period increased the probability of a price spell completion, thus supporting the idea of the importance of state dependent elements in firms' pricing behaviour.

---

[22] For example the national statistical institute informed the general public on 21 September 2004 about the next automatic wage indexation with effect of 1 October 2004 (Statec, 2004).

[23] On the importance of the wage indexation mechanism for services see also BCL (2001).

[24] For a review of the inflationary effects of the cash changeover see Mathä (2002).

In Luxembourg, the euro cash changeover also sparked a great interest in the public, mainly due to consumer fears that prices would be rounded upwards. A joint study by STATEC and the BCL reports an increase in the price change frequency in January 2002. In January 2002, about almost 13% of all LUF prices considered attractive in the Luxembourg CPI were rounded to attractive EUR prices, implying a substantial increase the price change frequency (e.g. BCL, 2003). In January 2002, a significantly higher proportion of prices were changed than in other years. Figure 6 presents the price increases and price decreases for each month for different product types. The cash changeover effect cannot be overlooked, in particularly not for services.

4.1.3 Other factors

4.1.3.1 ATTRACTIVE PRICES

A large number of consumer prices are set at levels such as EUR 9.99 commonly referred to as pricing points or threshold prices. The reason for this type of pricing behaviour is that such prices appear attractive to consumers and that the price label contains information that transcends the pure price signal. Firms use these signalling properties, as doing so allows them to influence consumers' purchasing decisions.[25] Bergen et al. (2003) show that pricing at psychological pricing points is very common and that price changes often occur in multiples of 10 cent. They conclude that threshold pricing contributes to price rigidity, as many prices seem to be stuck at 9-ending points. Several recent micro CPI price studies corroborate these findings; attractive prices are reported to be more rigid than others (e.g. Álvarez & Hernando, 2004; Aucremanne & Dhyne, 2005; Baumgartner et al., 2005; Veronese et al., 2005).

In close correspondence to previous studies analysing the cash changeover, we distinguish between three different types of attractive prices – round, psychological and fractional prices. Psychological prices are generally defined as prices ending with the digits "9", "95" or "99". Fractional prices are prices convenient to pay and are generally defined as prices with the last digit "5" or "0". Round prices are defined as integers. Moreover, the exact definition of psychological, fractional or round prices varies as prices exceed different thresholds. As our data set spans from 1999 to 2004, we had to consider varying definitions for attractive prices both in LUF and EUR.[26] The distribution of the last two digits in euro prices is presented in Figure 7.

The consequences of the introduction of the euro are twofold: Firstly, the unit increments at which prices can be changed differ according to the currency in use. Secondly, attractive pricing points differ

[25] See for example Schindler & Kibarian (1996) who report that psychological pricing increases consumer spending.
[26] In LUF, decimal prices are for example never considered as attractive prices; such prices were not common in retailing and could not be paid for.

according to the currency in use, and so does the increment to reach the next attractive threshold.[27] We, thus, decide to distinguish between attractive pricing points in LUF and EUR and analyse whether currency denomination differences affect firms' price setting behaviour.

4.1.3.2 PRODUCTS SUPPLIED BY PUBLIC INSTITUTIONS OR SUBJECT TO PRICE REGULATION

One of the more common findings from recent studies on consumer price behaviour suggests that prices subject to regulation display more price rigidity than freely determined prices. Dhyne et al. (2004) report for a selection of 50 products that regulated prices exhibit a roughly 10% lower probability of price change than freely determined prices. Lünnemann & Mathä (2005b) use index data for 15 EU countries and report that price indices considered subject to price regulation exhibit a 50% longer median price duration than non-regulated indices. Similarly, Dexter et al. (2004) report that significant inertia in aggregate price adjustments is due to the presence of price regulation.

In Luxembourg, as well as in other countries, an exact account of what is considered a regulated price is very cumbersome, as the measures undertaken and their scope may vary substantially (e.g. ECB, 2003). According to official sources at the national level, it is not exactly known which prices are regulated, and the degree to which prices are not freely determined may differ substantially. Hereafter, we consider two types of products subject to price regulation: 1) products that are typically supplied by (quasi-)public sector institutions, such as hospital services, theatres and combined passenger transport by road, 2) products, the prices of which are fixed (as point numbers or ranges) by public authorities, such as liquid fuel and transportation by taxi. These products are typically supplied by private enterprises but, in cases of thorough regulation, the price can be identical across all supplying firms. The product prices, which are not considered freely determined, are henceforth referred to as 'regulated'. A full list is presented in Table A1 the appendix.

In general and in analogy to the empirical findings in other studies we expect regulated prices to be more rigid than freely determined prices. However, price regulation per se must not necessarily render prices stickier. In Luxembourg, for example, fuel prices are subject to ceilings but nevertheless reveal a high frequency of price change.

4.2 The econometric model

In this section, we analyse whether the aforementioned determinants affect the probability of price changes. The econometric analysis follows Cecchetti (1986), Aucremanne & Dhyne (2005) and Baumgartner et al. (2005) in that we use a LOGIT model to model the probability of observing a price

[27] As example, if we simply defined psychological prices as prices ending with the last digit "9", then there would 12 psychological pricing points in 1 EUR, while expressed in LUF, there would only be 4.

change. First, we define a binary variable $y_{i,j,t}$ that characterises whether the price of product $p_{i,j,t}$ has changed between time t and $t+1$.

$$y_{ijt} = \begin{cases} 1 \; if \; p_{ijt+1} \neq p_{ijt} \\ 0 \; otherwise \end{cases} \tag{1a}$$

Using a LOGIT representation and allowing for product-specific random effects, the probability of observing a price change of product i in period $t+1$ can be written as

$$\text{Prob}\left[y_{ijt} = 1\right] = \frac{\exp(x_{ijt}\beta + u_i + \varepsilon_{ijt})}{1 + \exp(x_{ijt}\beta + u_i + \varepsilon_{ijt})}, \text{ with} \tag{2a}$$

$$\begin{aligned} x_{ijt} = {} & \chi_1 \sum_{t=T}^{t} \left| p\pi_{j,t-T} \right| + \varphi_1 \sum_{t=T}^{t} \left| w\pi_{t-T} \right| \\ & + \delta_1 \Delta^{+}_{ij,t-T} + \delta_2 \Delta^{-}_{ij,t-T} + \delta_3 lpcdw_{ij,t-T} \\ & + \xi_1 cco1 \\ & + \phi_1 itr1\_2 \\ & + \vartheta_1 attr^{ALL}_{ijt} \\ & + \gamma_1 regulated_{ijt} \\ & + \alpha_1 leng_{ij,t-T} \\ & + \beta_1 dur1_{ijt} + \beta_2 dur5_{ijt} + \beta_3 dur6_{ijt} + \beta_4 dur12_{ijt} + \beta_5 dur24_{ijt} \\ & + \sum_{m=2}^{12} \gamma_m month\_m + \sum_{y=2}^{6} \delta_y year\_y \end{aligned}$$

u_i reflects the product-specific random effects and ε_{ijt} reflects the independently distributed error term. The random effects LOGIT model accounts for unobserved heterogeneity across products i, which has been found to be one major reason for downward sloping hazard functions (e.g. Álvarez et al., 2005; Dhyne et al., 2005).

leng characterises the length of an individual price spell since its start in *t-T* until date *t*. If heterogeneity is appropriately accounted for, we expect the coefficient of *leng* to be close to unity or to be insignificant. Δ^+ and Δ^- refer to the size of the price change completing the preceding price spell. Separating the size of positive and negative preceding price adjustments allows analysing asymmetric effects. *lpcdw* represents a dummy variable and takes the value of 1 if the preceding price change for product *i* was negative. $\Sigma|p\pi|_{j,t-T,t}$ and $\Sigma|w\pi|_{t-T,t}$ represent the accumulated absolute price index inflation at the 10-digit COICOP level and the accumulated absolute aggregate wage inflation in the economy since the completion of the last price spell.[28] The inclusion of the above named variables is motivated by the state-dependent pricing model derived in Cecchetti (1986).

[28] The accumulation period is t-T, where t is the date of the current price and T is the period of the price spell beginning, while the estimated probability of price change is Prob(y_{ijt+1}) and hence refers to a the probability of price change in the period t+1. Hence these measures are purely backward looking. There is no simultaneity involved with respect to a price index change and an individual price change.

Other state-dependent variables included are those capturing the dates when wages are newly adjusted due to indexation (*itr1_2*) and the cash changeover period, where *cco1* captures January 2002. Attractive prices are denoted by the variable $attr^{ALL}$.[29] *regulated* refers to a dummy variable taking the value 1 if the product is subject to price regulation. *dur1-dur24* characterise dummy variables for spell lengths of 1, 5, 6, 12 and 24 months. These dummies are included in order to capture the spikes of the hazard rates as presented in Figure 4. Similarly, we include dummy variables for each month and year to account for unobserved seasonality and economic conditions not captured otherwise.

As we are also interested in differences and asymmetries between patterns of price increases and price reductions we correspondingly specify the two binary variables

$$y_{ijt}^{+} = \begin{cases} 1 \; if \; p_{ijt+1} > p_{ijt} \\ 0 \; otherwise \end{cases} \tag{1b}$$

$$y_{ijt}^{-} = \begin{cases} 1 \; if \; p_{ijt+1} < p_{ijt} \\ 0 \; otherwise \end{cases} \tag{1c}$$

Again, we assume that the probability of a price increase and price decrease can be modelled by the respective LOGIT representations

$$\text{Prob}\left[y_{ijt}^{+} = 1 \right] = \frac{\exp(x_{ijt}\beta + u_i + \varepsilon_{ijt})}{1 + \exp(x_{ijt}\beta + u_i + \varepsilon_{ijt})} \tag{2b}$$

$$\text{Prob}\left[y_{ijt}^{-} = 1 \right] = \frac{\exp(x_{ijt}\beta + u_i + \varepsilon_{ijt})}{1 + \exp(x_{ijt}\beta + u_i + \varepsilon_{ijt})} \tag{2c}$$

with

$$\begin{aligned} x_{ijt} = {} & \chi_1 \sum_{t=T}^{t} (\pi_{j,t-T}^{p} \mid \pi_{j,t-T}^{p} > 0) + \chi_2 \sum_{t=T}^{t} (\pi_{j,t-T}^{p} \mid \pi_{j,t-T}^{p} < 0) \\ & + \varphi_1 \sum_{t=T}^{t} (\pi_{t-T}^{w} \mid \pi_{t-T}^{w} > 0) + \varphi_2 \sum_{t=T}^{t} (\pi_{t-T}^{w} \mid \pi_{t-T}^{w} < 0) \\ & + \delta_1 \Delta_{ij,t-T}^{+} + \delta_2 \Delta_{ij,t-T}^{-} + \delta_3 lpcdw_{ij,t-T} \\ & + \xi_1 cco1 \\ & + \phi_1 itr1\_2 \\ & + \vartheta_1 attr_{ijt}^{ALL} \\ & + \gamma_1 regulated_{ijt} \\ & + \alpha_1 leng_{ij,t-T} \\ & + \beta_1 dur1_{ijt} + \beta_2 dur5_{ijt} + \beta_3 dur6_{ijt} + \beta_4 dur12_{ijt} + \beta_5 dur24_{ijt} \\ & + \sum_{m=2}^{12} \gamma_m month\_m + \sum_{y=2}^{6} \delta_y year\_y \end{aligned}$$

[29] These three issues will be more thoroughly explored in Table 7.

The only difference with respect to the model in equation (2a) is that we now allow for asymmetric effects between positively and negatively accumulated price and wage inflation. Note that the dummy variable for a duration of 5 months (*dur5*) is not included when separately estimating y_{ijt}^{+} and y_{ijt}^{-}; in these cases, a sales effect is not discernible at these fixed durations.

As indicated in section 2, censoring may adversely affect estimates of price spell duration. Whereas right-hand censoring can relatively easily be overcome, left-hand censoring is particularly worrying, as no information is available on the exact length of the price spell. In order to overcome this sample selection problem, we omit left-hand censored and double-censored price spells in the econometric analysis.

As discussed in section 3, a significant share of price changes is reversed at a later point in time. As switching between two prices may both not reflect true price flexibility and may not relate to the fundamentals of firms' pricing policies, we provide estimation results for both the original price trajectories (referred to as *baseline sample*) as well as for constructed trajectories where price reversals are not considered as price changes (referred to as *non-reversals - NR*). The results are presented in Table 6. We present both the odds ratio and marginal effects. Specifications referring to price changes y_{ijt}, price increases y_{ijt}^{+} and price decreases y_{ijt}^{-} are denoted ±, + and –, respectively.

4.3 The estimation results

4.3.1 Baseline results

The baseline results presented in Table 6 (specification *I*) clearly suggest that both time-dependent and state-dependent factors significantly contribute to the determination of the probability of a price change, a price increase and decreases.

4.3.1.1 THE SHAPE OF THE HAZARD

Consistent with the results from the unconditional hazard rates, the results suggest that time-dependent features contribute significantly to the observed price change probabilities. The dummies representing truncations at 1, 5, 6 12, 24 contribute positively to the probability of a price change. The probability of a price increase is significantly larger for truncations at 1, 6, 12, and 24 months, while the truncations at 1 and 6 months significantly reduce the probability of a price decrease. The odds ratio for the duration without price change *leng* is mostly below unity (specification *I*± and *I*±*NR*), suggesting that the hazard is downward sloping. This may indicate that despite trying to account for unobserved het-

erogeneity, we are not able to completely remove its effect of aggregating over different products.[30] The size of the effect is however very small, suggesting that an increase by one month reduces the probability of a price change, a price increase and a decrease by 0.1, 0.2 and 0.04 percentage points, respectively. Furthermore, in the case of price changes (see specification $I\pm$), the coefficient estimate is not significant at the 10 % level, while it is significant at the 1% and 10% level in case of price increases and decreases.

4.3.1.2 SEASONALITY

The coefficient estimates for the month dummies suggest the presence of a considerable intra-year pattern. The probability of a price change is largest in January, followed by July, thus highlighting the impact of seasonal sales. This increased price change probability is explained by the high probability of price reductions. Correspondingly, the relatively high probability of a price change in February and August is mainly explained by end–of–season sales, thereby implying prices increases in these months. This explanation for the obtained results is further corroborated by the separate estimations for price increases and decreases (see Table 6, specifications $I+$ and $I–$).

4.3.1.3 YEAR DUMMIES

The coefficient estimates for the year dummies indicate a less pronounced price change pattern across the sample years. The estimates suggest that, relative to 1999, the probability of price change is significantly higher in 2002, but lower in 2003 and 2004. These results generally extend to the probability of a price increase and decrease. In addition, for the probability of a price increase, the year 2000 is not significantly different from 1999, while the probability of a price increase is significantly lower in 2001. For a price decrease, a significantly lower probability is reported for 2000.

4.3.1.4 CUMULATED PRICE AND WAGE INFLATION

The probability of observing a price adjustment hinges on the product category specific accumulated price inflation and the aggregate wage inflation since the last price change took place. A 1-percentage point increase in cumulated price and wage inflation increases the probability of observing a price change by 0.5 and 0.1 percentage points, respectively. The results further suggest the presence of asymmetric effects in the accumulated price and wage inflation around zero. A 1 percentage point increase in positive (negative) accumulated price inflation increases (reduces) of the odds of observing a price increase by 1.7% (20%). In contrast, accumulated price inflation, whether positive or negative, has not any significant effect on the probability of a price decrease. The accumulated wage inflation increases the odds of observing a price increase, which is increased by 1.9%, but not the odds of ob-

[30] This interpretation is further strengthened by estimation results obtained from a random sample, where 1 observation per product is drawn. Doing so reduces the overrepresentation problem of short price spells in the dataset (see Dias et al. 2005; Fougère et al. 2005 on this issue) but comes at the expense of lower efficiency.

serving a price decrease. In the other cases, no significant effect is found. Hence, neither negative sector-specific accumulated price inflation nor negative aggregate accumulated wage inflation has an effect on the probability of observing a price decrease.

4.3.1.5 CHARACTERISTICS OF LAST PRICE CHANGE

With regard to the size and the sign of the preceding price change, the results indicate that the odds of observing a price change and a price increase are larger if the preceding price change was a price reduction, while the odds to observe a second successive price reduction are significantly lower. Furthermore, the larger the size of the negative preceding price change, the larger the odds of observing a price change, a price increase and a price decrease. A 1-percentage point increase in the size of the preceding price reduction raises the odds of observing a price change, a price increase and a price reduction by 45%, 46% and by 32%, respectively. Conversely, the odds of observing a price rise (reduction) diminish (increase), as the size of a preceding price rise increases. A 1-percentage point increase in the size of the preceding price rise does not significantly raise the odds of a price change, while it reduces the odds of observing a second consecutive price rise by 37% and increases the odds of observing a subsequent price reduction by 29%. On average, a price rise is followed by a price reduction and vice versa, which is generally consistent with commercial practices, and in particular temporary price promotions and the end−of−season sales. The positive coefficient for the size adds to the interpretation, which is further corroborated by the results in presented in section 4.3.2.

4.3.1.6 THE EURO CASH CHANGEOVER

As expected, the euro cash changeover has resulted in increasing the frequency of firms adapting their prices within a narrow time interval. The odds of a price change and a price increase were almost 120 times larger in January 2002 than in January 1999, implying an increase in the probability of a price change and an increase by 10 and 6 percentage points, respectively! Having said this, the cash changeover did not only increase the odds of observing a price rise, but also the odds to observe a price reduction. While not being nearly as sizeable, the odds of observing a price reduction in January 2002 were almost 45% higher than otherwise the case, implying an increased in the probability of a price reduction by 1.2 percentage points. However, in order to analyse strategic price setting behaviour, we include two more dummies in a separate regression capturing the ex-ante and ex-post period. The results are discussed in section 4.3.3.

4.3.1.7 AUTOMATIC WAGE INDEXATION

The event of a new adjustment of wages due to indexation contributes positively to the probability of observing a price change and a price increase, but reduces the probability of a price decrease. The probability of observing a price change and a price increase in the month and the month following a wage adjustment increase by 0.6 and 0.7 percentage points, while the probability of a price decrease

shrinks by 0.2 percentage points. As the adjustment of wages is caused by the accumulated rise in the aggregate price level, the automatic wage indexation mechanism has indeed important implications for the inflation process in Luxembourg. Therefore, we decide to explore this issue in more detail (see section 4.3.4).

4.3.1.8 ATTRACTIVE PRICES

As expected, pricing at attractive pricing points contributes to the degree of observed price rigidity. This holds for price changes and price increases, while attractive prices do not seem to contribute to downward rigidity of prices. The latter result may be related to sales and promotional activities, as attractive prices are more likely to be temporarily marked down. Differences between attractive prices in different currencies and different types of attractive prices will be explored in section 4.3.5.

4.3.1.9 REGULATED PRICES

As expected, regulated prices reveal unambiguous signs of increased price rigidity, both upward and downward. Compared to freely determined prices, the probability of a price change, increase and decrease are 7.6, 4.2 and 1.2 percentage points smaller.

4.3.1.10 RESULTS FOR DIFFERENT PRODUCT TYPES

The dummies for different product types are as reported in the descriptive part of the paper. Energy products display the highest probability of a price change, a price increase and a decrease. On the other hand of the scale, services reveal the lowest probability of a price change, a price increase and a price decrease.

4.3.2 Non-reverted price changes

In contrast to the baseline estimation results, the results presented hereafter refer to those price changes that are not offset by the following price change (see Table 6, specification *NR*). As a general observation, the results obtained for the baseline specification also apply to the sample including non-reverting price changes only. Time-dependent dummies for durations d = 1, 12 and 24 months significantly increase the probability of a price change. Contrary to the baseline specification, however, these dummies tend to enhance not only the probability of a price change and a price increase, but also the probability of a price decrease, while the coefficient estimates for durations 6, 12 and 24 months are very similar in explaining the probabilities of a price change and of a price increase. Differences between the *baseline* and the *non-reversals* exist with regard to the coefficient estimates for a duration of 1 month, which is substantially smaller in the latter specification, reflecting the importance of short-lived promotions in general. Lastly, the length of the price spell continues to contribute negatively to the probability of price adjustments.

In general, the probability of observing a price change, a price decrease and a price increase still reveal a marked intra-year pattern. However, the January dummy becomes weaker in the case of price changes and price decreases. The coefficient estimates suggest that the probability of a non-reverted price change is strongest in October. The coefficient estimate for the February and the August dummies decline substantially, reflecting the disappearance of (reverting) sales price changes in these months. In contrast, the coefficient estimates for all monthly dummies decline substantially, again reflecting the smaller impact of sales related price reductions in January.

In general, the yearly differences in the odds ratios become more pronounced when compared to the baseline specification. For all price changes, the odds ratios are significantly lower in 2000, but significantly higher in 2002. While for price decreases, the odds ratios were not considered significantly different from 1 in 2004, they were significantly smaller in 2000 and 2003.

In comparison to the baseline specification, the state-dependent elements remain broadly unchanged. The odds ratio for the absolute size of disaggregate accumulated inflation is significantly larger than 1 in the case of price changes, and so is the odds ratio for the accumulated wage inflation.

With regard to the characteristics of the last prices change, we note that the magnitudes of the estimated coefficients change in the expected direction, as they are generally smaller than in the baseline specification, as a large fraction of reverse sign price changes are eliminated. This further strengthens the interpretation that the negative correlation between the preceding price change and the subsequent price change, and in particular the positive size effect, are largely due to commercial practices. The odds ratio for the size of the last price change is smaller in the case of a price increase, whereas the opposite holds in the case of a price decrease. The size of the last price decrease reveals a significantly larger odds ratio for both price increases and price decreases (though the magnitude of the impact is smaller in size when compared to the baseline specifications). Contrary to the baseline specification, the odds are not significantly different in the case of price changes.

Similar to the result for the baseline specification, the probability of a price change is generally larger in January 2002. Quantitatively, the cash-changeover impact becomes even more pronounced for price changes and decreases than in the baseline specification. Wage adjustments due to indexation continue to increase (decrease) the probability of observing price increases (decreases). Similarly, attractive prices continue to be more rigid than non-attractive prices. In analogy to the baseline scenario, product price considered subject to price regulation reveal a smaller probability of price adjustment, both upward and downward.

With respect to the product type, the ranking remains almost identical (i.e. all odds ratios significantly smaller than energy products as well as the smallest coefficient for service products), with processed food and non-energy industrial goods switching their relative position (see also summary statistics in Table 3). With respect to price increases, again, the odds ratios for all product types are significantly smaller than the odds ratio for energy. Contrary to non-energy industrial goods and processed food, the odds ratio of the services dummy is hardly affected (relative to the baseline specification).

4.3.3 The euro cash changeover in more detail

While most firms may have decided to adjust prices at the actual date of the changeover, the 1 January 2002, others may have decided to either adjust their prices long before the actual cash changeover, when consumers were not yet sensitive to euro induced price changes, or after the period of dual price display, when consumers could not easily compare euro prices with former national prices any more (e.g. Mathä, 2002). In other words, some firms may have behaved strategically, which would also be in line with increased media attention and news in the second half of 2001 reporting that firms were unduly increasing prices. This may have contributed to increased caution about the timing and the size of price adjustments.

The additional variables *cco2* and *cco3* included in specifications *II* and *III* presented in Table 7 reflect dummy variables capturing the periods 1-3 months before and after and 4-6 months before and after January 2002 (*cco1*), respectively. So, taken together a full year around the cash changeover is considered. While the probability of a price adjustment is significantly larger in the 4-6 months (*cco3*) prior and after as well as the actual date of the cash changeover (*cco1*), the reverse holds for the 1-3 months preceding and succeeding the changeover (*cco2*). Particularly interesting is the size of the estimated coefficients. The increase in the probability of a price change in January 2002 is estimated to be 14 percentage points!

While the cash changeover also led to a significantly higher probability of observing a price reduction, it is in particular its effect on price increases that stands out. The probability of a price increase in January 2002 is 10 percentage points higher, while the probability of a price decrease is less than 1 percentage point higher. While the results for price increases generally mirror those for price changes in general, there are some noteworthy differences with regard to price decreases. The probability of observing a price reduction is larger within 1-3 months and 4-6 months prior and after the cash changeover. These results are consistent with previous simulation results and ex post evaluation of the cash changeover (BCL 2002, 2003). Taken together, they indeed indicate that firms behaved strategically; as during the 1-3 months prior and after January 2002 increased caution was exerted with respect to price adjustments; if prices were changed then there was a tendency to change them downward.

A possible explanation for the lower odds of observing a price change and a price increase in 2003 and 2004 is provided by Hobijn et al. (2004). Firstly, the cash changeover period incites all firms to review (raise) prices around the same time, implying that a disproportionate number of firms change (raise) their prices and resulting in a temporarily higher inflation rate. Secondly, firms are cognisant about the future need to convert prices at the euro cash changeover. Hence, firms' price changes prior to the cash changeover do not reflect the future expected marginal cost changes after January 2002, while the prices changed at the date of the cash changeover fully account for future expected marginal cost changes. As a result, this change in the time horizon implies a reduction in the price change frequency once firms have completed their price reviews and changes.

4.3.4 The automatic wage indexation mechanism and price increases

In Table 7, we allow for the automatic wage indexation at time t to have two separate effects in t and in $t+1$. The coefficients of both dummy variables are statistically significant. The probability of a price adjustment increases by 0.8 percentage points in the same month t and by 0.4 percentage points in the subsequent month $t+1$ if wages are newly indexed in t. This effect is clearly asymmetric, as the probability of a price increase rises by 0.5 and 0.9 percentage points, respectively, while the probability of a price decrease is 0.4 percentage points smaller.

At the current juncture, it is, however, unknown whether and to what extent prices of a given product category are affected by the wage indexation mechanism. For this purpose, we proceed in two ways. First, we identify which product categories are affected by the wage indexation by means of a multivariate regression. This regression includes the dependent variables' own lags plus a dummy variable indicating the lags of a new wage adjustment due to indexation. Second, we estimate a LOGIT regression for selected product categories to vindicate the results obtained by the inflation regressions.[31]

First, we estimate the following specification for 255 10-digit COICOP categories:

$$\pi^{p}_{i,t} = \alpha + \beta_1 \pi^{p}_{i,t-1} + \sum_{l=1}^{L-1} \gamma_l \Delta \pi^{p}_{i,t-l} + \beta_2 I_{t-1} + \sum_{l=1}^{L-1} \delta_l \Delta I_{t-l} + \sum_{m=1}^{11} \varepsilon_m M \,,$$

where π denotes the inflation rate of the price index p_i at the 10-digit COICOP level at time t. I denotes a 2.5% increase in the wages whenever wages are newly indexed to the price level. Δ, L and M represent the difference operator, the lag operator (with L being the total number of lags) and monthly dummies, respectively. For the purpose of robustness, we estimate the equation above separately for

[31] While the automatic wage indexation certainly deserves more attention and a more rigorous analysis, we have to content ourselves with providing some specific examples here. A detailed study is beyond the scope of this paper and has to be referred to a later stage.

lag lengths from to 1 to 6 months. Our estimates suggest a significant impact of changes in the relevant index on changes of the disaggregate price index for 10 product categories (see Table 8).[32]

In general, the results are well in line with expectations. In particular, they suggest that inflationary effects are present for those COICOP-10 categories, where prices seem directly or closely linked to wage costs, i.e. '*services for the maintenance of the dwelling*', '*repair of household appliances*', '*domestic services and household services*', '*maintenance and repair of personal transport equipment*' and '*hairdressing salons*'.[33] For a few COICOP categories, the suggested link between wage indexation and price level comes at a surprise, however, and we cannot rule out spurious effects not accounted for otherwise (e.g. in the case of '*mineral and spring water*', '*furniture*' and '*information processing equipment*').

Second, what are the implications of a wage adjustment due to indexation on the probability of observing a price change? We estimate a LOGIT specification similar to the baseline model on individual price data for selected COICOP-10 sub-categories (see Table 9).[34] We allow for different lags between the wage adjustment and the subsequent price change. In Figure 8, we see that for some indices, up to four months may lie between the wage adjustment due to indexation and the subsequent price rise. For all individual COICOP-10 categories presented, the results show a significant effect of wage adjustment on the probability of a subsequent price rise, indicating that for the selected indices, wage indexation may cause inflationary effects.

4.3.5 Attractive prices and price rigidity

As argued, differences in pricing points in different currencies may have implications for price rigidity. Thus, we separate $attr^{ALL}$ into $attr^{EUR}$ and $attr^{LUF}$, with respective superscripts referring to prices displayed in *EUR* and *LUF*. Within the set of attractive prices, we further distinguish between psychological, round and fractional prices in the respective currency (i.e. $round^{EUR}$, $round^{LUF}$, $psycho^{EUR}$, $psycho^{LUF}$, $fract^{EUR}$ *and* $fract^{LUF}$).

The estimations in Table 7 show firstly that psychological, fractional and round prices reduce the probability of observing a price change and price increase. This is irrespective of the currency in use. A formal coefficient equality test is rejected, confirming the choice to include these regressors separately. Furthermore, attractive prices in LUF are stickier than attractive prices in EUR. This is in particular due to fractional and psychological prices in LUF being stickier than their EUR counterparts.

---

[32] More specifically, for these 10 product categories, a significantly positive index change coefficient is found for at least 3 different lag lengths (among the six specifications tested) at the 5 percent significance level.

[33] For some of these indices, notably hairdressing salons, a significant role of the indexation mechanism had been found in former studies as well (see for example Adam & da Costa, 2002).

[34] Notation as described in section 4.2.

No such effect can be discerned for the round prices. With regard to price reductions, only psychological prices in LUF seem to result in more downward rigidity. In contrast, round and fractional prices in LUF, as well as psychological prices in EUR seem to result in a higher probability of price reductions.

Taken together, these results indicate very clearly, as posited that firstly, pricing at pricing points contribute to the observed price rigidity, secondly that attractive prices tend to contribute to an increased upward rather than downward rigidity, and finally that the observed rigidity may be related to the currency in use. Concerning the last point, it seems that the number of possible pricing points within a certain range of values is of importance for the observed rigidity.

4.3.6 Services

New evidence for individual euro area countries suggests that, compared to other product types, services' prices change less frequently, and if so then rarely downwards (e.g. Dhyne et al., 2005; Fabiani et al., 2005; Lünnemann & Mathä, 2005b). As services seem to be a major contributing factor to the alleged price rigidity in the euro area countries, we estimate the LOGIT regression in slightly different form for services only. The results are presented in Table 10.

The results for services reveal that a rise in the accumulated absolute wage inflation increases the probability of a price change.[35] Similarly, increases in the accumulated positive (negative) wage inflation increase the probability of observing a price increase (decrease). The marginal effect of negative accumulated inflation is actually larger in specification $IV-$ than in the specifications $IV\pm$ and $IV+$. Thus, contrary to the results in the baseline regression in Table 6, services prices seem to respond well to negative wage inflation. This gives rise to two different interpretations. Firstly, overall price reductions are governed by factors other than price and negative wage inflation. Secondly, services rarely adjust downwards, as wages rarely adjust downwards. Hence, the low frequency of price reductions does not reflect intrinsic downward price rigidity.

With regard to the other determinants, we note that the results are by and large as previously reported. The marginal effect of the cash changeover is much larger (both upwards and downwards) than reported for the baseline specification, confirming the euro area NCBs' previous results pointing to particularly strong effects in the services sector. The automatic wage adjustment is also reported to lead to a higher probability of a price change and a price increase[36], while the probability of a price reduction is reduced. Attractive prices add to price stickiness, as the probability of a price change and a price in-

[35] For services, the accumulated price inflation did not matter, and was thus excluded from the regression. The insignificance should, however, not come as surprise, as services price are to large extent driven by wage developments.

[36] This is in line with evidence from former studies (see BCL, 2001).

crease are reduced, while the effect for price decreases is weaker. Prices of products subject to regulation are generally stickier, no matter in what direction, than freely determined prices.

4.3.7 Competition among supermarkets

A key element for understanding the flexibility of prices relates to competition.[37] Several theoretical and empirical contributions suggest a negative relationship between competition and price rigidity. Rotemberg & Saloner (1987) show that monopolists change prices less frequently than oligopolists, while Dornbusch (1987) shows that the extent of price adjustment generally depends on the market structure and on the number of firms. On the empirical front, Carlton (1986) has shown that price rigidity is strongly correlated with industry level concentration, while Dutta et al. (2002) argue that price flexibility in the U.S. orange juices market is related to fierce competition in this industry. Recent survey evidence for several euro area countries as well as for Luxembourg suggested that competitors' prices are an important factor for firms to reduce their prices (Fabiani et al., 2005; Lünnemann & Mathä, 2005a).

To analyse the effect of competition on the price setting behaviour, we focus on observations collected from supermarkets. When analysing the effect of competition on price flexibility, we need to make sure that the firms are competing amongst each other. Supermarkets serve as a good case in point. Secondly, we do not want to restrict ourselves to a case study of a single product, such as petroleum prices, but rather aim at still analysing a non-negligible part of the Luxembourg NICP.

We introduce two new variables considered as proxies for the degree of competition. The first variable captures how many supermarkets sell products in a particular COICOP-10 product category (*Σsupermkts*). For services, this variable's mean is close to 0, indicating that not many supermarkets offer services at all, while unprocessed and processed food are supplied by a large number of supermarkets. The second variable captures the overall market share of a particular supermarket. The market share of a given supermarket is simply taken to reflect the total number of observations from supermarket *s* divided by the total number of observations from all supermarkets (i.e. $\Sigma p_s / \Sigma p$).[38] Based on our reduced sample, the largest supermarket has a share of more than 25% of all supermarket observations, while the smallest retains a share of less than 5%.

The results presented in Table 11 show clearly that both the number of supermarkets selling products in a given COICOP-10 product category and the market share are important determinants for supermarkets' price setting behaviour. The higher the number of competitors, the more often prices change.

---

[37] For a survey on the link between competition and inflation see for example Asplund & Friberg (1998).

[38] Clearly, this is only an imperfect measure. However, given the representativeness of the price quotes collected by STATEC, this measure is expected to be a rather accurate reflection of market share.

This is the case for price increases and price decreases. In contrast, the larger the market share of a supermarket, the less frequent price adjustments are. This is particularly the case for price reductions. Thus, these results seem to suggest that the supermarkets with more market power adjust prices downward more rarely. This may be due to the fact that they can act as price leaders.

5 Conclusion

This paper analyses behavioural patterns of micro consumer prices in Luxembourg and investigates into the explanatory factors of price changes. We compute different measures of price stickiness for more than 230 product categories.

1. We find that the implied median and mean durations are 8 and 12 months. The frequency of price change varies substantially across product types. Whereas energy prices, on average, change every 1.5 months, services' prices change less than once per year. Overall there are no strong signs of downward price rigidity as price decreases represent almost 40 percent of all price changes. In contrast, the corresponding share is less than 20 percent for services.
2. The average size of price changes is relatively large compared to the dynamics in aggregate inflation. The average sizes of both price increases and price reductions are similar.
3. Our results suggest that the overall frequency of price change is substantially affected by price change reversion, which, in turn, is frequently observed during end–of–season sales periods. Price reversions may contribute as much as roughly 40 percent to the frequency of price change in the case of non-energy industrial goods.
4. The analysis reveals a relatively high degree of synchronisation with respect to price changes compared to other euro area countries. This may reflect the compact size of the Luxembourg economy. The aggregate unconditional hazard rate reveals mass points at durations of 1 and 5 months and to a smaller extent at 12 and 24 months. The mass point at 5 months is due to the influence of end–of–season sales.

Results based on panel estimates indicate that both state- and time-dependent factors contribute to the observed pattern of price changes. This applies to price increases as well as to price reductions.

1. With respect to time-dependent factors, the probability of a price change increases at specific truncation lengths (in particular at 1, 5, 6 and 12 months). Interestingly, the probability of a price change is particularly strong in January, but decreases after 2002.
2. Cumulated price and wage inflation significantly increase the probability of a price change. While this extends to price increases in general, no such effects were found for price decreases. In contrast, for services, negative accumulated wage inflation leads to an increased probability of observing price reductions, which suggests that the rareness of observed price reductions in this sector may have to do with fact that wages rarely decline, and that services prices are not intrinsically downward rigid.

3. Consistent with widespread commercial practices of temporary promotions and sales, our results suggest that the magnitude and the direction of the preceding price change affect the probability of a price adjustment. In particular, the probability of a price change or of a price increase is higher when subsequent to a price reduction. In addition, the probability of a price change and price increase is an increasing function of the size of the former price reduction.

4. Specific events tend to increase the probability of a price adjustment. For example, the event of an automatic wage adjustment due to indexation tends to increase the probability of a price change and of a price increase, and may, thus, cause inflationary effects. The cash changeover increased the probability of a price change, both upward and downward. However, the former clearly dominated the latter, in particular for services. Attractive pricing policies generally reduce the probability of a price change. Interestingly, they are found to contribute to upward rigidity but less so to downward rigidity. Moreover, the impact of attractive pricing may depend on the currency in use. Similarly, the evidence clearly suggests that product prices considered subject to price regulation are more rigid; this being the case in either direction.

5. The number of competitors is found to add to the probability of price adjustments, in particular downwards. In contrast, the market share reduces the probability of a price change. These results are very much in line with recent evidence from surveys and stress the importance of the competitive environment for the observed degree of price rigidity.

6 References

Adam, F. and Z. da Costa (2002): "Le Phénomène d'Auto-Allumage dans le Contexte de l'Indexation des Salaires", in *Bulletin du Statec*, Volume 49, No. 2, pp. 114-133.

Álvarez, L. and I. Hernando (2004): "Price Setting Behaviour in Spain: Stylised Facts Using Consumer Price Micro Data", *ECB Working Paper*, No. 416.

Álvarez, L., P. Burriel, Pablo and I. Hernando (2005): "Do Decreasing Hazard Functions Make Sense?", *ECB Working Paper* No. 461.

Apel, M., R. Friberg and K. Hallsten (2001): "Micro Foundations of Macroeconomic Price Adjustment: Survey Evidence from Swedish Firms", in *Journal of Money, Credit and Banking*, Vol. 37, No. 2, pp. 313-338.

Asplund, M. and R. Friberg (1998): "Links between Competition and Inflation", in *Sveriges Riksbank Quarterly Review*, Vol. 3, pp. 50-73.

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

Aucremanne, L. and E. Dhyne (2005): "Time-Dependent versus State-Dependent Pricing: A Panel Data Approach to the Determinants of Belgian Consumer Price Changes", *ECB Working Paper* No. 462.

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

Banque centrale du Luxembourg (2001): "*Rapport Annuel 2000*", (BCL, Luxembourg).

Banque centrale du Luxembourg (2002): "*Rapport Annuel 2001*", (BCL, Luxembourg).

Banque centrale du Luxembourg (2003): "*Rapport Annuel 2002*", (BCL, Luxembourg).

Baumgartner, J., E. Glatzer, F. Rumler and A. Stiglbauer (2005): "How Frequently Do Consumer Prices Change in Austria? Evidence from Micro CPI Data", *ECB Working Paper* No. 523.

Bergen, M., H. Chen and D. Levy (2003): "Making Sense of Ignoring Cents: Another Implication of Rational Inattention", mimeo.

Bils, M. and P. Klenow (2004): "Some Evidence on the Importance of Sticky Prices, in *Journal of Political Economy*, Vol. 112, No. 5, pp. 947-985.

Blinder, A.S. (1994): "On sticky prices: Academic Theories Meet the Real World", in Mankiw, G. (ed.), *Monetary Policy*, NBER Studies in Business Cycles, (University Press, Chicago), pp. 117-150.

Blinder, A.S., E. Canetti, D.E. Lebow and J.B. Rudd (1998): *Asking About Prices: A new Approach to Understanding Price Stickiness*, (Russell Sage Foundation, New York).

Borenstein, S., C. Cameron and R. Gilbert (1997): "Do Gasoline Prices Respond Asymmetrically to Crude Oil Price Changes, in *Quarterly Journal of Economics*, Vol. 112, pp. 305-339.

Calvo, G. (1983): "Staggered Prices in a Utility Maximizing Framework", in *Journal of Monetary Economics*, Vol. 12, pp. 383-98.

Caplin, A. and D. Spulber (1987): "Menu Costs and Optimization with State-Dependent Pricing", in *Quarterly Journal of Economics*, Vol. 110, 703-725.

Carlton, D.W. (1986): "The Rigidity of Prices", in *American Economic Review,* Vol. 76, No. 4, pp. 637-658.

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

Cornille, D. (2003): "The Adaptation of Prices to the Changeover to the Euro", *National Bank of Belgium Economic Review*, 4th Quarter.

Dexter, A., M. Levi and B. Nault (2004): "Sticky Prices: The Impact of Regulation", in *Journal of Monetary Economics*, Vol. 49, No. 4, pp. 797-821.

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

Dias, M., C. Robalo Marques and J.M.C. Santos Silva (2005): "Time or State Dependent Price Setting Rules? Evidence from Portuguese Micro Data", *ECB Working Paper* No. 511.

Dhyne, E., L. Álvarez, H. Le Bihan, G. Veronese, D. Dias, J. Hoffmann, N. Jonker, P. Lünnemann, F. Rumler and J. Vilmunen (2005): "Price Setting in the Euro Area: Some Stylised Facts From Micro Consumer Price Data", *ECB Working Paper* No. 524

Dornbusch, R. (1987): "Exchange Rates and Prices", in *American Economic Review*, Vol. 77, No.1, pp. 93-106.

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

Dutta, S., M. Bergen and D. Levy (2002): "Price Flexibility in Channels of Distribution: Evidence from Scanner Data", in *Journal of Economic Dynamics & Control*, Vol. 26, pp. 1845-1900.

ECB (2003): "Direct Impact in Administered Prices and Indirect Taxes on Euro Area HICP Developments" in *ECB Monthly Bulletin*, No. 3, pp. 35-36, (ECB: Frankfurt).

Fabiani, S., M. Druant, I. Hernando, C. Kwapil, B. Landau, C. Loupias, T. Mathä, R. Sabbatini, H. Stahl A. Stokman (2005): "The Pricing Behaviour of Firms in the Euro Area: New Survey Evidence", *ECB Working Paper* No. 535.

Fisher, T. and J. Konieczny (2000): "Synchronization of Price Changes by Multiproduct Firms: Evidence from Canadian Newspaper Prices", in *Economics Letters*, Vol. 68, pp. 271-277.

Folkertsma, C., C. Van Renselaar and A. Stokman (2002): "Smooth Cash Changeover, Higher Prices? Results of a Survey Among Dutch Retailers" in *DNB Research Memorandum*, WO No. 682.e.

Fontagné, L. (2004): "Compétitivité du Luxembourg: Une Paille dans l'Acier", Report to the Minister of Economics and Foreign Affairs of the Grand-Duchy of Luxembourg.

Fougère, D., H. Le Bihan and P. Sevestre (2005): "Heterogeneity in Consumer Price Stickiness: A Microeconometric Investigation", *ECB Working Paper* No. 536.

Hobijn, B., F. Ravenna and A. Tambalotti (2004): "Menu Costs at Work: Restaurant Prices and the Introduction of the Euro", *Federal Reserve Bank of New York* Staff Report No. 195.

Hoffmann, J. and J.-R. Kurz-Kim (2005): "Consumer Price Adjustment under the Microscope: Germany in a Period of Low Inflation", *Deutsche Bundesbank*, mimeo.

Hall, S., M. Walsh and A. Yates (2000): "Are UK companies' Prices Sticky?", in *Oxford Economic Papers*, Vol. 52, pp. 425-46.

Jonker, N., C. Folkertsma and H. Blijenberg (2004): "Empirical Analysis of Price Setting Behaviour in the Netherlands in the Period 1998-2003 using Micro Data", *ECB Working Paper* No. 413.

Kackmeister, A. (2005): "Yesterday's Bad Times are Today's Good Old Times: Retail Price Changes in the 1890s were Smaller, Less Frequent, and More Permanent", *Federal Reserve Board of Governors*¸ Finance and Economics Discussion Series, 2005-18.

Kashyap, A. (1995): "Sticky Prices: New Evidence from Retail Catalogues", in *Quarterly Journal of Economics*, Vol. 105, No. 1, pp. 245-274.

Kiefer, N. (1988): "Economic Duration Data and Hazard Functions", in *Journal of Economic Perspectives*, Vol. 26, No. 2, pp. 646-679.

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

Lach S. and D. Tsiddon (1992): "The Behaviour of Prices and Inflation: An Empirical Analysis of Disaggregated Price Data", in *Journal of Political Economy*, Vol. 100, pp. 349-389.

Lach S. and D. Tsiddon (1996): "Staggering and Synchronization in Price-Setting: Evidence from Multiproduct Firms", in *American Economic Review*, Vol. 86, No. 5, pp. 1175-1196.

Lazear, E.P. (1986): "Retail Pricing and Clearance Sales", in *American Economic Review*, Vol. 76, No. 1, pp. 14-32.

Lünnemann, P. and T.Y. Mathä (2004): "How Persistent is Disaggregate Inflation? An Analysis Across EU Countries and HICP Sub-Indices", *ECB Working Paper* No. 415.

Lünnemann, P. and T.Y. Mathä (2005a): "New Survey Evidence on the Pricing Behaviour of Luxembourg Firms", mimeo, *Banque centrale du Luxembourg*.

Lünnemann, P. and T.Y. Mathä (2005b): "Regulated and Services' Prices and Inflation Persistence", *ECB Working Paper* No. 466.

Mathä, T.Y. (2002): 'Die Eurobargeldeinführung, Regionale Preisentwicklung und die Wahrgenommene Inflation der Verbraucher', in *BCL Bulletin*, No. 3, pp. 60-73, Banque centrale du Luxembourg.

Pashigian, B.P. (1988): "Demand Uncertainty and Sales: A Study of Fashion and Markdown Pricing, in *American Economic Review*, Vol. 78, No. 5, pp. 936-953.

Ratfai, A. (2003): "The Frequency and Size of Price Adjustment: Microeconomic Evidence", mimeo *Central European University*.

Rotemberg, J. (1982): "Monopolistic Price Adjustments and Aggregate Output", in *Review of Economic Studies*, Vol. 49, No. 4, pp. 517-531.

Salop, S. and Stiglitz, J.E. (1982): The Theory of Sales: A Simple Model of Equilibrium Price Dispersion with Identical Agents, in *American Economic Review*, Vol. 71, No.5, pp. 1121-1130.

Schindler, R. and T. Kibarian (1996): "Increased Consumer Sales Response Through Use of 99-Ending Prices", in *Journal of Retailing*, Vol. 72, No. 2, pp. 187-199.

Statec (2004): "Echelle Mobile des Salaires: Adaptations des Salaires et Traitements (+2.5%) au 1er Octobre 2004.", Press Statement No. 39.

Taylor, J.B. (1980): "Aggregate Dynamics and Staggered Contracts", in *Journal of Political Economy*, Vol. 88, No. 1, pp. 1-23.

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

Varian, H.R. (1980): "A Model of Sales", in *American Economic Review*, Vol. 70, No. 4, pp. 651-659.

Veronese G., S. Fabiani, A. Gattulli and R. Sabbatini (2005): "Consumer Price Behaviour in Italy: Evidence From Micro CPI Data", *ECB Working Paper*, No. 449.

Vilmunen, J. and H. Laakkonen (2005): "How Often Do Price Change in Finland? Evidence from Micro CPI Data", *Suomen Pankki*, mimeo.

Warner E.J. and Barsky, R.B. (1995): "The Timing and Magnitude of Retail Store Markdowns: Evidence from Weekends and Holidays", in *Quarterly Journal of Economics*, Vol. 110, No. 2, pp. 321-252.

Tables and Figures

Table 1: Number of price quotes in the Luxembourg CPI data, 1999 – 2004

COICOP2	Price quotes
cp01 Food & non-alcoholic beverages	108,047
cp02 Alcohol, tobacco & narcotics	17,736
cp03 Clothing & footwear	13,286
cp04 Housing, water, elect., gas & other fuels	18,152
cp05 Furnish., househ. equip. & maint. of house	61,231
cp06 Health	6,732
cp07 Transport	24,828
cp08 Communications	10,200
cp09 Recreation & culture	53,192
cp10 Education	2,508
cp11 Restaurants & hotels	34,830
cp12 Miscellaneous goods & services	29,469

Product type	Price quotes
Energy	9,909
Non-energy industrial goods	134,972
Processed food	78,007
Services	109,547
Unprocessed food	47,776

All	380,211

Table 2: Summary statistics m-o-m and y-o-y inflation

	Y-o-y inflation	M-o-m inflation	
	Average	Average	Share of negative changes
cp01 Food & non-alcoholic beverages	2.6	0.2	0.17
cp02 Alcoholic beverages, tobacco & narcotics	3.4	0.3	0.26
cp03 Clothing & footwear	1.1	0.2	0.22
cp04 Hous., water, electr., gas & other fuels	2.8	0.3	0.33
cp05 Furn., househd. eqmt. & maint. of house	1.8	0.2	0.21
cp06 Health	1.5	0.1	0.22
cp07 Transport	2.8	0.3	0.29
cp08 Communications	-6.1	-0.5	0.72
cp09 Recreation & culture	1.6	0.1	0.39
cp10 Education	4.1	0.3	0.00
cp11 Restaurants & hotels	2.9	0.3	0.04
cp12 Miscellaneous goods & services	2.0	0.2	0.25
cp00 All-items	**2.2**	**0.2**	**0.25**

Table 3: Summary statistics on the frequency and the size of price changes

	F±	F+	F-	Implied median duration	Implied mean duration	Share F-	F sales	F reversion (ex sales)	F (excl. reversion)	S+	S-
COICOP 2											
cp01 Food & non-alcoholic beverages	0.19	0.11	0.08	5.61	8.10	0.42	0.01	0.04	0.14	0.11	0.14
cp02 Alcohol, tobacco & narcotics	0.14	0.11	0.03	4.68	6.75	0.23	0.00	0.03	0.11	0.06	0.10
cp03 Clothing & footwear	0.20	0.11	0.09	5.28	7.62	0.44	0.14	0.01	0.05	0.11	0.10
cp04 Housing, water, elect., gas & other fuels	0.29	0.18	0.11	5.46	7.88	0.38	0.01	0.02	0.27	0.06	0.05
cp05 Furnish., househ. equip. & maint. of house	0.18	0.11	0.07	3.94	5.68	0.39	0.08	0.01	0.09	0.08	0.08
cp06 Health	0.03	0.03	0.01	63.10	91.04	0.18	0.00	0.00	0.03	0.10	0.13
cp07 Transport	0.21	0.14	0.07	6.23	8.98	0.32	0.00	0.02	0.18	0.04	0.03
cp08 Communications	0.04	0.02	0.02	17.19	24.80	0.53	0.00	0.00	0.04	0.12	0.26
cp09 Recreation & culture	0.13	0.07	0.06	9.13	13.17	0.43	0.03	0.02	0.08	0.10	0.09
cp10 Education	0.05	0.04	0.01	13.89	20.03	0.16	0.00	0.00	0.05	0.14	0.17
cp11 Restaurants & hotels	0.05	0.04	0.01	14.36	20.72	0.11	0.00	0.00	0.05	0.06	0.05
cp12 Miscellaneous goods & services	0.11	0.08	0.03	9.93	14.33	0.31	0.03	0.01	0.07	0.07	0.05
Product type											
Energy	0.52	0.32	0.20	1.41	2.03	0.38	0.00	0.06	0.46	0.04	0.04
Non-energy ind. goods	0.16	0.10	0.07	7.37	10.63	0.41	0.06	0.01	0.09	0.08	0.08
Processed food	0.12	0.09	0.04	5.95	8.59	0.30	0.00	0.03	0.09	0.08	0.11
Unprocessed food	0.26	0.15	0.12	4.52	6.52	0.44	0.01	0.05	0.21	0.13	0.17
Services	0.07	0.06	0.01	13.00	18.76	0.17	0.00	0.00	0.07	0.07	0.05
All											
Total	0.17	0.11	0.06	8.15	11.76	0.37	0.03	0.02	0.12	0.08	0.08

Note:
F±: frequency of price change,
F+: frequency of price increases,
F-: frequency of price decreases,
Share F-: share of price decreases,
F sales: frequency of price change due to sales (estimated),

F rev (ex sales): frequency of price change due to price reversions other than sales,
F (ex rev): frequency of price change not (fully) reverted afterwards,

S+: average size increase of price changes,
S-: average size decrease of price changes.

Table 4: Weighted percentiles, per product type

	Price change			Price increase			Price decrease		
Weighted Percentiles	25	50	75	25	50	75	25	50	75
Energy	0.27	0.71	0.71	0.16	0.44	0.44	0.11	0.24	0.27
Non-energy industrial goods	0.11	0.14	0.19	0.07	0.10	0.11	0.02	0.06	0.09
Processed food	0.11	0.13	0.15	0.06	0.08	0.09	0.02	0.04	0.05
Unprocessed food	0.10	0.12	0.62	0.07	0.08	0.34	0.03	0.05	0.28
Services	0.04	0.05	0.08	0.03	0.04	0.07	0.00	0.01	0.01
Total	0.07	0.12	0.18	0.05	0.08	0.11	0.01	0.04	0.07

Note: The weighted x^{th} percentiles are calculated by ordering the respective indicator from smallest to largest and cumulating the weights of the ranked values. The weighted x^{th} percentile then is just the first value for which the cumulated weight is greater than or equal to the x^{th} percentile.

Table 5: Synchronisation ratios by product type and COICOP 2-digit

	Sync±	Sync+	Sync-
cp01 Food & non-alcoholic beverages	0.27	0.28	0.23
cp02 Alcohol, tobacco & narcotics	0.35	0.36	0.13
cp03 Clothing & footwear	0.76	0.76	0.81
cp04 Housing, water, elect., gas & other fuels	0.72	0.72	0.67
cp05 Furnish., househ. equip. & maint. of house	0.51	0.51	0.47
cp06 Health	0.44	0.40	0.19
cp07 Transport	0.61	0.56	0.48
cp08 Communications	0.69	0.67	0.43
cp09 Recreation & culture	0.50	0.48	0.37
cp10 Education	0.45	0.42	0.33
cp11 Restaurants & hotels	0.38	0.34	0.31
cp12 Miscellaneous goods & services	0.55	0.52	0.50
Energy	0.89	0.90	0.88
Non-energy ind. goods	0.49	0.47	0.48
Processed food	0.32	0.32	0.19
Unprocessed food	0.25	0.27	0.23
Services	0.56	0.53	0.36

Note: **Sync±**, **Sync+** and **Sync-** denote the Fisher and Konieczny (2000) synchronisation ratio for price changes, price increases and price decreases, respectively.

Table 6: The probability of price change, increase and decrease: Baseline model

Specification	*I±*		*I+*		*I-*		*I±NR*		*I+NR*		*I-NR*	
Estimat. Techn.	RE LOGIT		RE LOGIT		RE LOGIT		RE LOGIT		RE LOGIT		RE LOGIT	
No. of obs.	292354		292354		292354		292354		292354		292354	
No. of groups	6201		6201		6201		6201		6201		6201	
Obs. per grp: Min/Avg/Max	1 / 47.1 / 70		1 / 47.1 / 70		1 / 47.1 / 70		1 / 47.1 / 70		1 / 47.1 / 70		1 / 47.1 / 70	
Dep. Variable	y_{ijt}		y^{+}_{ijt}		y^{-}_{ijt}		y_{ijt}		y^{+}_{ijt}		y^{-}_{ijt}	
Odds Ratio / Marginal Effect	OR	MFX	OR	MFX	OR	MFX	OR	MFX	OR	MFX	OR	MFX
Σ|pπ |	**1.050**	**0.005**					**1.012**	**0.001**				
Σpπ | pπ>0			**1.017**	**0.001**	0.995	0.000			**1.017**	**0.001**	0.992	0.000
Σpπ | pπ<0			**0.806**	**-0.012**	1.002	0.000			0.997	0.000	**0.957**	**-0.001**
Σ|wπ |	**1.011**	**0.001**					**1.012**	**0.001**				
Σwπ | wπ>0			**1.019**	**0.001**	1.002	0.000			**1.016**	**0.001**	**1.006**	**0.000**
Σwπ | wπ<0			0.990	-0.001	0.530	-0.018			1.864	0.027	1.099	0.002
Δ^{+}	1.017	0.002	**0.627**	**-0.026**	**1.291**	**0.007**	**1.082**	**0.005**	**0.857**	**-0.007**	**1.316**	**0.004**
Δ^{-}	**1.453**	**0.037**	**1.464**	**0.021**	**1.320**	**0.008**	1.014	0.001	*1.066*	*0.003*	***1.095***	***0.001***
lpcdw	**2.622**	**0.116**	**3.208**	**0.085**	**0.582**	**-0.013**	**1.471**	**0.026**	**1.850**	**0.031**	**0.825**	**-0.003**
cco1	**2.170**	**0.102**	**2.165**	**0.060**	**1.437**	**0.012**	**3.266**	**0.119**	**1.977**	**0.040**	**4.273**	**0.049**
itr1_2	**1.059**	**0.006**	**1.128**	**0.007**	**0.919**	**-0.002**	1.018	0.001	**1.125**	**0.005**	**0.861**	**-0.002**
attractive	**0.759**	**-0.028**	**0.735**	**-0.018**	0.980	-0.001	**0.797**	**-0.014**	**0.795**	**-0.010**	**0.858**	**-0.003**
regulated	**0.337**	**-0.076**	**0.320**	**-0.042**	**0.577**	**-0.012**	**0.447**	**-0.037**	**0.352**	**-0.031**	**0.639**	**-0.006**
leng	0.993	-0.001	**0.960**	**-0.002**	*0.986*	*0.000*	***0.990***	***-0.001***	**0.970**	**-0.001**	**0.972**	**0.000**
dur 1	**2.266**	**0.099**	**2.620**	**0.070**	**0.860**	**-0.004**	**1.787**	**0.043**	**1.628**	**0.025**	**1.687**	**0.010**
dur 5	**2.439**	**0.119**										
dur 6	**1.270**	**0.026**	**1.382**	**0.020**	**0.769**	**-0.007**	**1.283**	**0.017**	**1.402**	**0.017**	0.928	-0.001
dur 12	**1.713**	**0.065**	**1.842**	**0.044**	1.093	0.003	**1.896**	**0.051**	**1.884**	**0.036**	**1.644**	**0.010**
dur 24	**1.477**	**0.045**	**1.696**	**0.037**	1.140	0.004	**1.516**	**0.031**	**1.506**	**0.021**	**1.626**	**0.010**
February	**0.675**	**-0.034**	**1.301**	**0.016**	**0.197**	**-0.026**	**0.838**	**-0.010**	**0.880**	**-0.005**	**0.792**	**-0.003**
March	**0.351**	**-0.075**	**0.754**	**-0.014**	**0.243**	**-0.024**	0.998	0.000	**0.856**	**-0.006**	**1.260**	**0.004**
April	**0.417**	**-0.066**	**0.879**	**-0.007**	**0.220**	**-0.025**	1.004	0.000	***0.906***	**-0.004**	**1.182**	**0.003**
May	**0.444**	**-0.063**	**0.890**	**-0.006**	**0.240**	**-0.024**	***1.083***	***0.005***	0.940	-0.003	**1.309**	**0.005**
June	**0.290**	**-0.084**	**0.644**	**-0.021**	**0.185**	**-0.027**	**0.727**	**-0.017**	**0.676**	**-0.015**	0.969	0.000
July	**0.807**	**-0.020**	**0.606**	**-0.023**	1.048	0.001	**0.734**	**-0.017**	**0.599**	**-0.018**	**1.207**	**0.003**
August	**0.595**	**-0.044**	***1.102***	***0.006***	**0.212**	**-0.025**	**0.657**	**-0.022**	**0.616**	**-0.018**	0.912	-0.001
September	**0.304**	**-0.082**	**0.683**	**-0.019**	**0.209**	**-0.026**	**0.826**	**-0.011**	**0.762**	**-0.011**	1.067	0.001
October	**0.457**	**-0.061**	0.995	0.000	**0.237**	**-0.024**	***1.088***	***0.005***	0.966	-0.002	**1.316**	**0.005**
November	**0.335**	**-0.078**	**0.692**	**-0.018**	**0.209**	**-0.026**	**0.796**	**-0.013**	**0.719**	**-0.013**	1.077	0.001
December	**0.254**	**-0.090**	**0.589**	**-0.024**	**0.156**	**-0.028**	**0.601**	**-0.026**	**0.598**	**-0.019**	**0.813**	**-0.003**
Year 2000	0.976	-0.002	1.007	0.000	**0.907**	**-0.003**	**0.856**	**-0.009**	***0.926***	***-0.003***	**0.867**	**-0.002**
Year 2001	0.972	-0.003	**0.896**	**-0.006**	1.015	0.000	***0.926***	***-0.005***	**0.857**	**-0.006**	***1.117***	***0.002***
Year 2002	**1.080**	**0.008**	0.970	-0.002	**1.203**	**0.005**	**1.145**	**0.009**	1.029	0.001	**1.343**	**0.005**
Year 2003	**0.883**	**-0.012**	**0.874**	**-0.007**	**0.890**	**-0.003**	**0.802**	**-0.013**	**0.828**	**-0.008**	*0.923*	*-0.001*
Year 2004	**0.862**	**-0.014**	**0.863**	**-0.008**	***0.920***	***-0.002***	**0.821**	**-0.012**	**0.827**	**-0.008**	1.055	0.001
Unproc. food	**0.192**	**-0.107**	**0.287**	**-0.048**	**0.452**	**-0.017**	**0.513**	**-0.033**	**0.249**	**-0.040**	**0.510**	**-0.009**
Proc. food	**0.079**	**-0.160**	**0.188**	**-0.065**	**0.105**	**-0.040**	**0.182**	**-0.073**	**0.130**	**-0.058**	**0.116**	**-0.022**
No.en.ind.good	**0.137**	**-0.177**	**0.167**	**-0.090**	**0.314**	**-0.030**	**0.168**	**-0.099**	**0.099**	**-0.092**	**0.153**	**-0.028**
Services	**0.044**	**-0.206**	**0.164**	**-0.074**	**0.032**	**-0.064**	**0.156**	**-0.084**	**0.147**	**-0.061**	**0.065**	**-0.030**
Wald-Chi^2	**20052**		**24319**		**12638**		**7541**		**6062**		**4518**	
LR-test: rho=0	**18000**		**4344**		**12000**		**16000**		**4356**		**7985**	
Rho-value	0.29		0.17		0.41		0.26		0.17		0.30	

Note: The columns OR and MFX denote the odds ratio and marginal effects. The marginal effects calculate the probability of a positive outcome, e.g. y_{ijt}=1 assuming that the random effect of that observations's panel is zero. Thus, this may not be similar to the proportion of observed outcomes in the group. **Bold**, ***bolditalics*** and *italics* indicates significance at the 1%, 5% and 10% level respectively. Base is energy in January 1999.

Table 7: The cash changeover, wage indexation and attractive prices

Specification	*II*±		*II*+		*II*-		*III*±		*III*+		*III*-	
Estimat. Techn.	RE LOGIT		RE LOGIT		RE LOGIT		RE LOGIT		RE LOGIT		RE LOGIT	
No. of obs.	292354		292354		292354		292354		292354		292354	
No. of groups	6201		6201		6201		6201		6201		6201	
Obs. per grp: Min/Avg/Max	1 / 47.1 / 70		1 / 47.1 / 70		1 / 47.1 / 70		1 / 47.1 / 70		1 / 47.1 / 70		1 / 47.1 / 70	
Dep. Variable	y_{ijt}		y^{+}_{ijt}		y^{-}_{ijt}		y_{ijt}		y^{+}_{ijt}		y^{-}_{ijt}	
Odds Ratio / Marginal Effect	OR	MFX	OR	MFX	OR	MFX	OR	MFX	OR	MFX	OR	MFX
Σ|pπ |	**1.051**	**0.005**					**1.050**	**0.005**				
Σpπ | pπ>0			**1.017**	**0.001**	*0.995*	*0.000*			**1.019**	**0.001**	*0.994*	*0.000*
Σpπ | pπ<0			**0.806**	**-0.012**	1.002	0.000			**0.807**	**-0.012**	1.002	0.000
Σ|wπ |	**1.011**	**0.001**					**1.011**	**0.001**				
Σwπ | wπ>0			**1.021**	**0.001**	1.002	0.000			**1.020**	**0.001**	1.002	0.000
Σwπ | wπ<0			1.140	0.007	0.541	-0.017			1.135	0.007	0.547	-0.017
Δ^{+}	1.007	0.001	**0.626**	**-0.026**	**1.293**	**0.007**	1.008	0.001	**0.629**	**-0.026**	**1.291**	**0.007**
Δ^{-}	**1.463**	**0.038**	**1.467**	**0.021**	**1.317**	**0.008**	**1.452**	**0.036**	**1.468**	**0.021**	**1.331**	**0.008**
lpcdw	**2.618**	**0.116**	**3.223**	**0.085**	**0.580**	**-0.013**	**2.598**	**0.112**	**3.207**	**0.085**	**0.582**	**-0.013**
cco1	**2.741**	**0.144**	**3.105**	**0.102**	**1.210**	**0.006**	**2.743**	**0.141**	**3.201**	**0.106**	***1.158***	***0.004***
cco2	**0.800**	**-0.021**	**0.636**	**-0.021**	**1.235**	**0.006**	**0.799**	**-0.020**	**0.632**	**-0.022**	**1.237**	**0.006**
cco3	***1.055***	***0.005***	***1.068***	***0.004***	1.004	0.000	***1.051***	***0.005***	***1.061***	***0.003***	1.009	0.000
itr1	**1.079**	**0.008**	**1.092**	**0.005**	1.006	0.000	**1.078**	**0.007**	**1.092**	**0.005**	1.004	0.000
itr2	*1.041*	*0.004*	**1.162**	**0.009**	**0.860**	**-0.004**	1.037	0.004	**1.159**	**0.009**	**0.861**	**-0.004**
attractive EUR	**0.808**	**-0.021**	**0.752**	**-0.015**	1.015	0.000						
attractive LUF	**0.686**	**-0.035**	**0.698**	**-0.018**	*0.952*	*-0.001*						
round EUR							**0.657**	**-0.036**	**0.540**	**-0.028**	1.033	0.001
round LUF							**0.684**	**-0.033**	**0.623**	**-0.023**	***1.084***	***0.002***
fractional EUR							**0.814**	**-0.019**	**0.776**	**-0.013**	0.999	0.000
fractional LUF							**0.755**	**-0.025**	**0.693**	**-0.018**	**1.134**	**0.004**
psycho EUR							**0.924**	**-0.007**	**0.851**	**-0.008**	**1.103**	**0.003**
psycho LUF							**0.723**	**-0.028**	**0.773**	**-0.013**	**0.924**	**-0.002**
regulated	**0.351**	**-0.075**	**0.322**	**-0.042**	**0.579**	**-0.012**	**0.344**	**-0.073**	**0.321**	**-0.042**	**0.583**	**-0.012**
leng	***0.989***	***-0.001***	**0.953**	**-0.003**	*0.985*	*0.000*	*0.991*	*-0.001*	**0.953**	**-0.003**	0.987	0.000
dur1	**2.267**	**0.099**	**2.622**	**0.070**	**0.858**	**-0.004**	**2.248**	**0.096**	**2.603**	**0.069**	**0.861**	**-0.004**
dur5	**2.429**	**0.118**					**2.443**	**0.116**				
dur6	**1.270**	**0.026**	**1.387**	**0.021**	**0.771**	**-0.006**	**1.278**	**0.026**	**1.387**	**0.021**	**0.771**	**-0.006**
dur12	**1.719**	**0.066**	**1.842**	**0.044**	1.095	0.003	**1.718**	**0.064**	**1.840**	**0.044**	1.098	0.003
dur24	**1.486**	**0.046**	**1.708**	**0.038**	1.133	0.004	**1.479**	**0.044**	**1.712**	**0.038**	1.133	0.004
February	**0.698**	**-0.032**	**1.403**	**0.021**	**0.188**	**-0.026**	**0.695**	**-0.031**	**1.396**	**0.021**	**0.188**	**-0.026**
March	**0.362**	**-0.074**	**0.803**	**-0.011**	**0.232**	**-0.024**	**0.363**	**-0.071**	**0.808**	**-0.011**	**0.232**	**-0.024**
April	**0.409**	**-0.067**	**0.857**	**-0.008**	**0.216**	**-0.025**	**0.408**	**-0.065**	**0.861**	**-0.008**	**0.216**	**-0.025**
May	**0.438**	**-0.064**	**0.860**	**-0.008**	**0.243**	**-0.024**	**0.439**	**-0.062**	**0.868**	**-0.007**	**0.243**	**-0.024**
June	**0.284**	**-0.085**	**0.630**	**-0.022**	**0.182**	**-0.027**	**0.285**	**-0.083**	**0.636**	**-0.021**	**0.181**	**-0.027**
July	**0.796**	**-0.021**	**0.589**	**-0.024**	1.046	0.001	**0.799**	**-0.020**	**0.595**	**-0.024**	1.045	0.001
August	**0.582**	**-0.046**	***1.079***	*0.004*	**0.208**	**-0.026**	**0.581**	**-0.044**	*1.078*	*0.004*	**0.209**	**-0.025**
September	**0.301**	**-0.083**	**0.659**	**-0.020**	**0.213**	**-0.025**	**0.303**	**-0.080**	**0.667**	**-0.019**	**0.213**	**-0.025**
October	**0.467**	**-0.060**	1.055	0.003	**0.224**	**-0.025**	**0.469**	**-0.058**	*1.067*	0.004	**0.224**	**-0.025**
November	**0.343**	**-0.077**	**0.721**	**-0.016**	**0.203**	**-0.026**	**0.345**	**-0.075**	**0.730**	**-0.016**	**0.202**	**-0.026**
December	**0.260**	**-0.090**	**0.619**	**-0.022**	**0.150**	**-0.029**	**0.261**	**-0.087**	**0.627**	**-0.022**	**0.150**	**-0.028**
Year 2000	0.973	-0.003	0.993	0.000	**0.895**	**-0.003**	0.972	-0.003	0.998	0.000	**0.898**	**-0.003**
Year 2001	0.978	-0.002	***0.936***	***-0.004***	0.962	-0.001	0.986	-0.001	*0.945*	*-0.003*	0.965	-0.001
Year 2002	0.997	0.000	0.982	-0.001	**1.128**	**0.003**	1.006	0.001	0.973	-0.002	**1.181**	**0.005**
Year 2003	**0.792**	**-0.022**	**0.831**	**-0.010**	**0.852**	**-0.004**	**0.807**	**-0.020**	**0.836**	**-0.009**	**0.889**	**-0.003**
Year 2004	**0.775**	**-0.024**	**0.819**	**-0.010**	**0.879**	**-0.003**	**0.792**	**-0.022**	**0.830**	**-0.010**	***0.914***	***-0.002***
Unproc. food	**0.329**	**-0.082**	**0.289**	**-0.048**	**0.448**	**-0.017**	**0.188**	**-0.105**	**0.287**	**-0.048**	**0.441**	**-0.017**
Proc. food	**0.115**	**-0.144**	**0.189**	**-0.065**	**0.104**	**-0.040**	**0.078**	**-0.156**	**0.186**	**-0.065**	**0.103**	**-0.040**
No.en.ind.good	**0.195**	**-0.147**	**0.168**	**-0.089**	**0.311**	**-0.030**	**0.133**	**-0.176**	**0.182**	**-0.085**	**0.301**	**-0.030**
Services	**0.064**	**-0.187**	**0.165**	**-0.074**	**0.031**	**-0.064**	**0.046**	**-0.199**	**0.173**	**-0.072**	**0.031**	**-0.064**
Wald-Chi^2			24673		12638		20150		24991		12677	
LR-test: rho=0	18000		4330		12000		18000		4100		12000	
Rho-value	0.28		0.17		0.41		0.28		0.16		0.41	

Note: The columns OR and MFX denote the odds ratio and marginal effects. The marginal effects calculate the probability of a positive outcome, e.g. y_{ijt}=1 assuming that the random effect of that observations's panel is zero. Thus, this may not be similar to the proportion of observed outcomes in the group. **Bold**, ***bolditalics*** and *italics* indicates significance at the 1%, 5% and 10% level respectively. Base is energy in January 1999.

Table 8: Product categories for which wages indexation leads to an inflation increase

COICOP10	Product category
01.02.02.01.01	Mineral and spring water
04.01.01.01.01	Actual rentals for housing[39]
04.03.02.01.01	Services for the maintenance of the dwelling
05.01.01.01.01	Furniture and furnishings (living room, dining room, office)
05.01.01.01.03	Furniture and furnishings (bedroom)
05.03.03.01.01	Repair of household appliances
05.06.02.01.01	Domestic services and household services
07.02.03.01.01	Maintenance and repair of personal transport equipment
09.01.03.01.01	Information processing equipment
12.01.01.01.01	Hairdressing salons

Table 9: Wage indexation and subsequent price increases – specific examples

Description	Services for the maintenance of dwellings		Domestic and household services		Maintenance and repair of personal transport equipment		Hairdressing salons	
COICOP-10	0403020101		0506020101		0702030101		1201010101	
Estimat. Techn.	RE LOGIT		RE LOGIT		RE LOGIT		RE LOGIT	
No. of obs.	1071		14330		3619		4214	
No. of groups	20		214		61		17	
Obs. per grp: Min/Avg/Max	15 / 53.5 / 66		31 / 67 / 68		10 / 59.3 / 69		17 / 56.2 / 70	
Dep. Variable	y^{+}_{ijt}		y^{+}_{ijt}		y^{+}_{ijt}		y^{+}_{ijt}	
Odds Ratio / Marginal Effect	OR	MFX	OR	MFX	OR	MFX	OR	MFX
$\Sigma p\pi \mid p\pi>0$	0.756	-0.002	**0.897**	**-0.006**	***0.835***	***-0.005***	**1.135**	**0.004**
cco1	**212.62**	***0.618***	**3.945**	***0.129***	**20.270**	***0.328***	**10.411**	***0.203***
attractive ALL	0.561	-0.005	0.937	-0.003	0.956	-0.001	**0.419**	***-0.035***
Leng	**1.334**	***0.002***	**1.064**	**0.003**	**1.160**	**0.004**	1.016	0.000
itr_t	**46.7190**	***0.220***			**3.113**	**0.049**	**6.622**	**0.120**
itr_{t+1}	**194.570**	**0.515**			**12.987**	**0.206**	**16.150**	**0.254**
itr_{t+2}	**52.776**	***0.231***			**5.519**	**0.095**	**8.894**	**0.155**
itr_{t+3}	***18.924***	0.107	**423.819**	**0.904**	**2.542**	*0.037*		
monthly dummy	Yes		Yes		Yes		Yes	
Yearly dummy	Yes		Yes		Yes		Yes	
Wald-Chi^2	**119**		**2297**		**260**		**289**	
LR-test: rho=0	0.78		**27.9**		**11.0**		**7.0**	
Rho-value	0.03		0.05		0.13		0.06	

Note: The columns OR and MFX denote the odds ratio and marginal effects. The marginal effects calculate the probability of a positive outcome, e.g. y_{ijt}=1 assuming that the random effect of that observations's panel is zero. Thus, this may not be similar to the proportion of observed outcomes in the group. **Bold**, ***bolditalics*** and *italics* indicates significance at the 1%, 5% and 10% level respectively. Base is January 1999.

[39] In this paper, actual rentals for housing are not considered due to lower than monthly frequency of price collection.

Table 10: Services

Specification	*IV*±		*IV*+		*IV*-	
Estimat. Techn.	RE LOGIT		RE LOGIT		RE LOGIT	
No. of obs.	74987		74987		74987	
No. of groups	1863		1863		1863	
Obs. per grp: Min/Avg/Max	1 / 40.3 / 70		1 / 40.3 / 70		1 / 40.3 / 70	
Dep. Variable	y_{ijt}		y^{+}_{ijt}		y^{-}_{ijt}	
Odds Ratio / Marginal Effect	OR	MFX	OR	MFX	OR	MFX
$\Sigma\|w\pi\|$	**1.024**	**0.001**				
$\Sigma w\pi \mid w\pi>0$			**1.033**	**0.001**		
$\Sigma w\pi \mid w\pi<0$					**0.005**	**-0.031**
Δ^{+}	**0.656**	**-0.017**	**0.446**	**-0.026**	0.943	0.000
Δ^{-}	**1.190**	**0.007**	**1.302**	**0.008**	0.996	0.000
lpcdw	**1.388**	**0.014**	**2.118**	**0.030**	**0.493**	**-0.003**
cco1	**8.368**	**0.218**	**8.947**	**0.196**	**3.520**	**0.014**
cco2	**0.465**	**-0.023**	**0.458**	**-0.019**	***0.683***	**-0.002**
cco3	**1.690**	**0.024**	**1.336**	**0.010**	**2.093**	**0.006**
itr2	***1.125***	***0.005***	**1.215**	**0.007**	*0.784*	***-0.001***
attractive EUR	**0.581**	**-0.021**	**0.602**	**-0.016**	**0.377**	**-0.006**
attractive LUF	**0.668**	**-0.014**	**0.530**	**-0.017**	1.083	0.000
regulated	**0.411**	**-0.028**	**0.310**	**-0.027**	*0.783*	*-0.001*
leng	*0.981*	*-0.001*	**0.952**	**-0.002**	0.991	0.000
dur1	**0.841**	**-0.006**	**0.792**	**-0.007**	**1.310**	**0.002**
dur6	**2.183**	**0.043**	**2.360**	**0.040**	***1.344***	***0.002***
dur12	**2.076**	**0.040**	**2.008**	**0.031**	**2.052**	**0.006**
dur24	**1.929**	**0.035**	**2.155**	**0.035**	0.576	-0.003
February	*0.838*	***-0.007***	**0.721**	**-0.009**	1.170	0.001
March	1.124	0.005	1.101	0.003	1.171	0.001
April	**0.563**	**-0.018**	**0.631**	**-0.012**	**0.586**	**-0.003**
May	**1.514**	**0.019**	**1.283**	***0.009***	**1.935**	**0.005**
June	**0.529**	**-0.020**	**0.637**	**-0.012**	**0.400**	**-0.004**
July	**0.570**	**-0.018**	**0.612**	**-0.013**	***0.658***	**-0.002**
August	***0.830***	***-0.007***	0.876	-0.004	0.934	0.000
September	0.875	-0.005	0.977	-0.001	*0.751*	*-0.002*
October	**1.414**	**0.016**	**1.782**	**0.023**	**0.554**	**-0.003**
November	1.021	0.001	1.149	0.005	0.772	*-0.001*
December	*1.161*	*0.006*	**1.499**	**0.015**	**0.459**	**-0.003**
Year 2000	*0.886*	*-0.005*	0.892	-0.004	0.867	-0.001
Year 2001	**0.505**	**-0.022**	**0.517**	**-0.017**	**0.621**	**-0.002**
Year 2002	0.897	-0.004	**0.750**	**-0.009**	***1.442***	*0.002*
Year 2003	**0.596**	**-0.018**	**0.504**	**-0.019**	**1.519**	***0.003***
Year 2004	**0.586**	**-0.019**	**0.553**	**-0.017**	1.184	0.001
Wald-Chi^2	**2969**		**2552**		**1052**	
LR-test: rho=0	**2662**		**1307**		**650**	
Rho-value	0.31		0.31		0.35	

Note: The columns OR and MFX denote the odds ratio and marginal effects. The marginal effects calculate the probability of a positive outcome, e.g. y_{ijt}=1 assuming that the random effect of that observations's panel is zero. Thus, this may not be similar to the proportion of observed outcomes in the group. **Bold**, ***bolditalics*** and *italics* indicates significance at the 1%, 5% and 10% level respectively. Base is January 1999.

Table 11: Competition among supermarkets

Specification	$V\pm$		$V+$		$V-$	
Estimat. Techn.	RE LOGIT		RE LOGIT		RE LOGIT	
No. of obs.	101241		101241		101241	
No. of groups	2100		2100		2100	
Obs. per grp: Min/Avg/Max	1 / 48.2 / 70		1 / 48.2 / 70		1 / 48.2 / 70	
Dep. Variable	y_{ijt}		y^{+}_{ijt}		y^{-}_{ijt}	
Odds Ratio / Marginal Effect	OR	MFX	OR	MFX	OR	MFX
$\Sigma\|p\pi\|$	**1.017**	**0.002**				
$\Sigma p\pi \mid p\pi>0$			***1.016***	***0.001***	1.004	0.000
$\Sigma p\pi \mid p\pi<0$			**0.966**	**-0.002**	0.993	0.000
$\Sigma\|w\pi\|$	**1.007**	**0.001**				
$\Sigma w\pi \mid w\pi>0$			**1.010**	**0.001**	1.000	0.000
$\Sigma w\pi \mid w\pi<0$			***0.271***	***-0.094***	3.080	0.054
Δ^{+}	1.029	0.004	**0.709**	**-0.025**	**1.536**	**0.020**
Δ^{-}	*1.123*	*0.015*	0.972	-0.002	**1.881**	**0.030**
lpcdw	**1.793**	**0.079**	**2.858**	**0.091**	**0.634**	**-0.020**
cco1	**1.660**	**0.075**	**1.691**	**0.047**	1.173	0.008
cco2	**0.846**	**-0.020**	**0.733**	**-0.020**	*1.116*	0.005
cco3	0.998	0.000	*1.082*	*0.006*	*0.918*	*-0.004*
attractive EUR	1.030	0.004	1.008	0.001	1.050	0.002
attractive LUF	**0.893**	**-0.014**	0.975	-0.002	0.971	-0.001
leng	*0.984*	*-0.002*	**0.966**	**-0.003**	*0.979*	*-0.001*
dur1	**2.260**	**0.119**	**2.829**	**0.097**	**1.332**	**0.015**
dur2	**1.099**	**0.012**	**1.185**	**0.013**	***1.111***	***0.005***
dur3	0.972	-0.004	0.950	-0.004	1.038	0.002
dur12	**1.296**	**0.035**	**1.465**	**0.032**	1.032	0.002
dur24	*1.326*	*0.039*	1.137	0.010	***1.728***	***0.033***
February	0.952	-0.006	0.942	-0.004	0.982	-0.001
March	1.047	0.006	0.910	*-0.007*	***1.175***	***0.008***
April	***1.133***	***0.016***	1.005	0.000	***1.188***	***0.009***
May	1.003	0.000	0.969	-0.002	0.997	0.000
June	**0.825**	**-0.023**	**0.701**	**-0.023**	*1.120*	0.006
July	**0.876**	**-0.016**	**0.693**	**-0.023**	**1.240**	**0.011**
August	**0.801**	**-0.026**	**0.643**	**-0.027**	***1.185***	***0.009***
September	*0.914*	*-0.011*	**0.768**	**-0.017**	**1.200**	***0.009***
October	1.035	0.004	**0.861**	**-0.010**	**1.266**	**0.012**
November	**0.823**	**-0.023**	**0.731**	**-0.020**	1.073	0.003
December	**0.570**	**-0.060**	**0.598**	**-0.031**	**0.759**	**-0.012**
Year 2000	**0.760**	**-0.032**	*0.923*	*-0.006*	**0.777**	**-0.011**
Year 2001	**0.811**	**-0.025**	0.993	-0.001	**0.786**	**-0.011**
Year 2002	***0.902***	***-0.013***	0.962	-0.003	1.031	0.001
Year 2003	**0.708**	**-0.040**	*0.911*	*-0.007*	**0.764**	**-0.012**
Year 2004	**0.738**	**-0.036**	**0.825**	**-0.013**	0.937	-0.003
Proc. food	**0.294**	**-0.155**	**0.519**	**-0.048**	**0.263**	**-0.067**
No.en.ind.good	**0.305**	**-0.116**	**0.424**	**-0.051**	**0.341**	**-0.040**
Services	**0.221**	**-0.110**	*0.534*	***-0.035***	**0.062**	**-0.047**
Σ supermkts	**1.172**	**0.020**	**1.108**	**0.007**	**1.141**	**0.006**
supermkt share	**0.302**	**-0.150**	0.704	-0.025	**0.179**	**-0.082**
Wald-Chi^2	**3064**		**5428**		**1669**	
LR-test: rho=0	**5288**		**664**		**1757**	
Rho-value	0.22		0.08		0.19	

Note: The columns OR and MFX denote the odds ratio and marginal effects. The marginal effects calculate the probability of a positive outcome, e.g. y_{ijt}=1 assuming that the random effect of that observations's panel is zero. Thus, this may not be similar to the proportion of observed outcomes in the group. **Bold**, ***bolditalics*** and *italics* indicates significance at the 1%, 5% and 10% level respectively. Base is unprocessed food in January 1999.

Figure 1: Consumer price inflation (m-o-m , y-o-y)

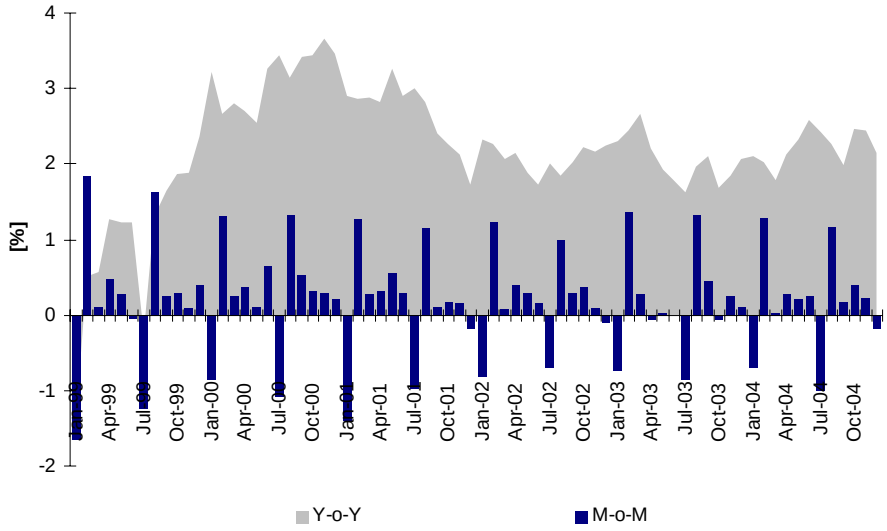


Figure 2: Frequency of price changes: impact of price reversion and sales

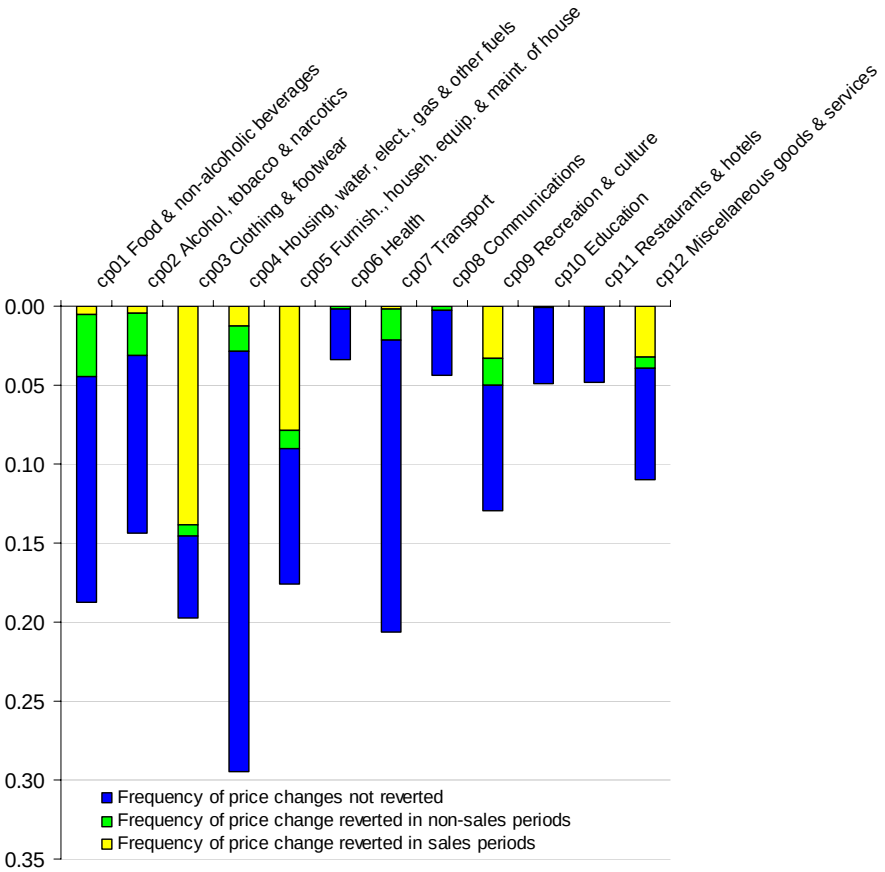
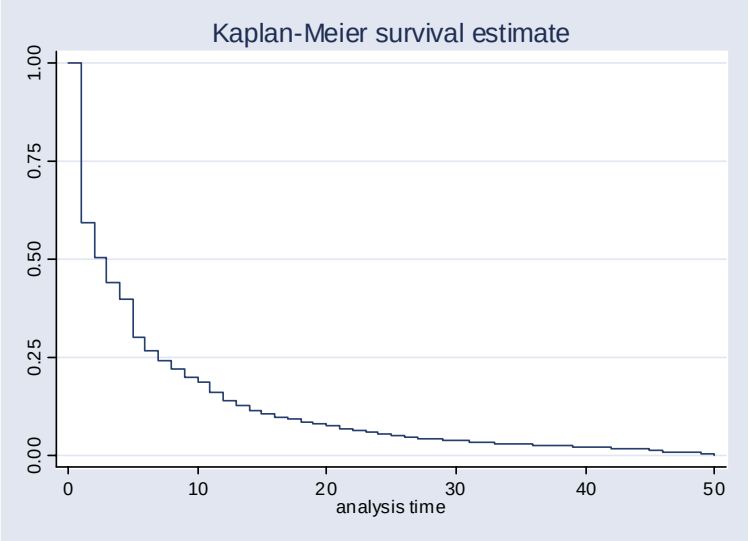


Figure 3: Kaplan-Meier survival estimates

All price spells



Per product type

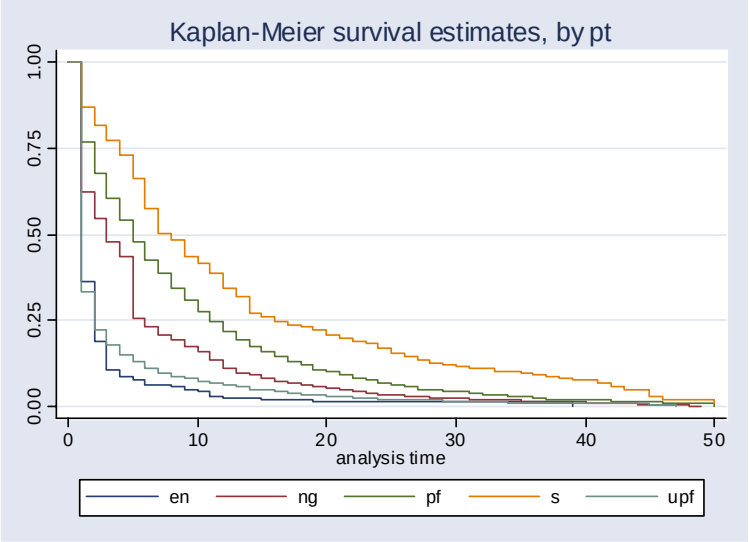


Figure 4: Hazard rates based on Kaplan-Meier survivor estimates

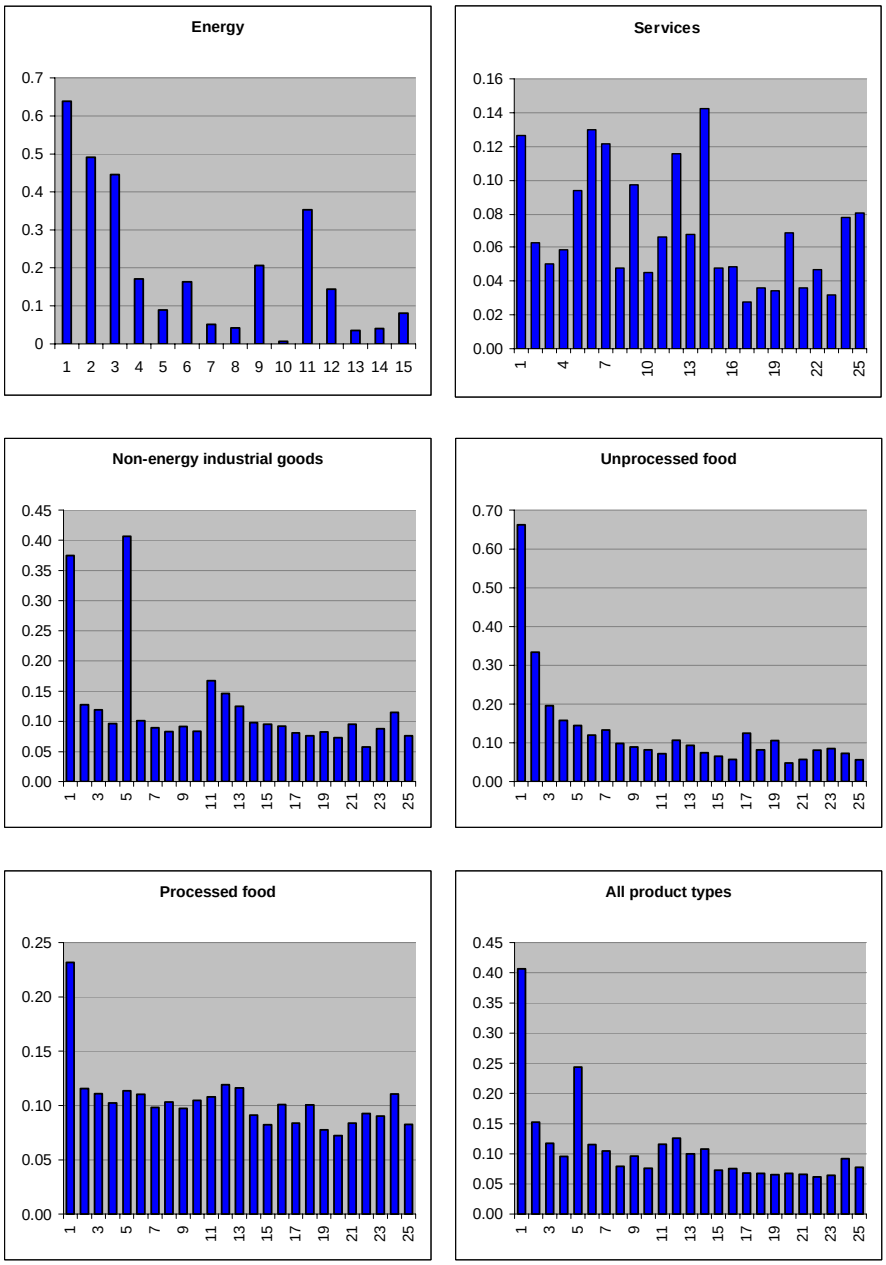


Figure 5: Frequency of price changes per calendar month

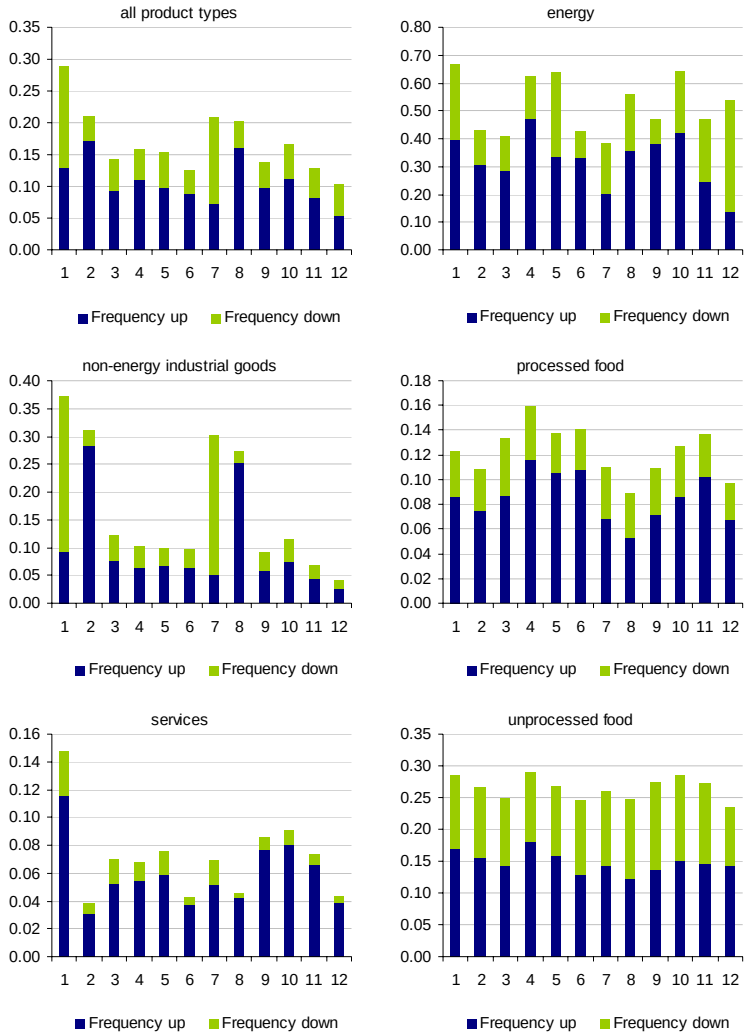


Figure 6: Frequency of price changes for each month

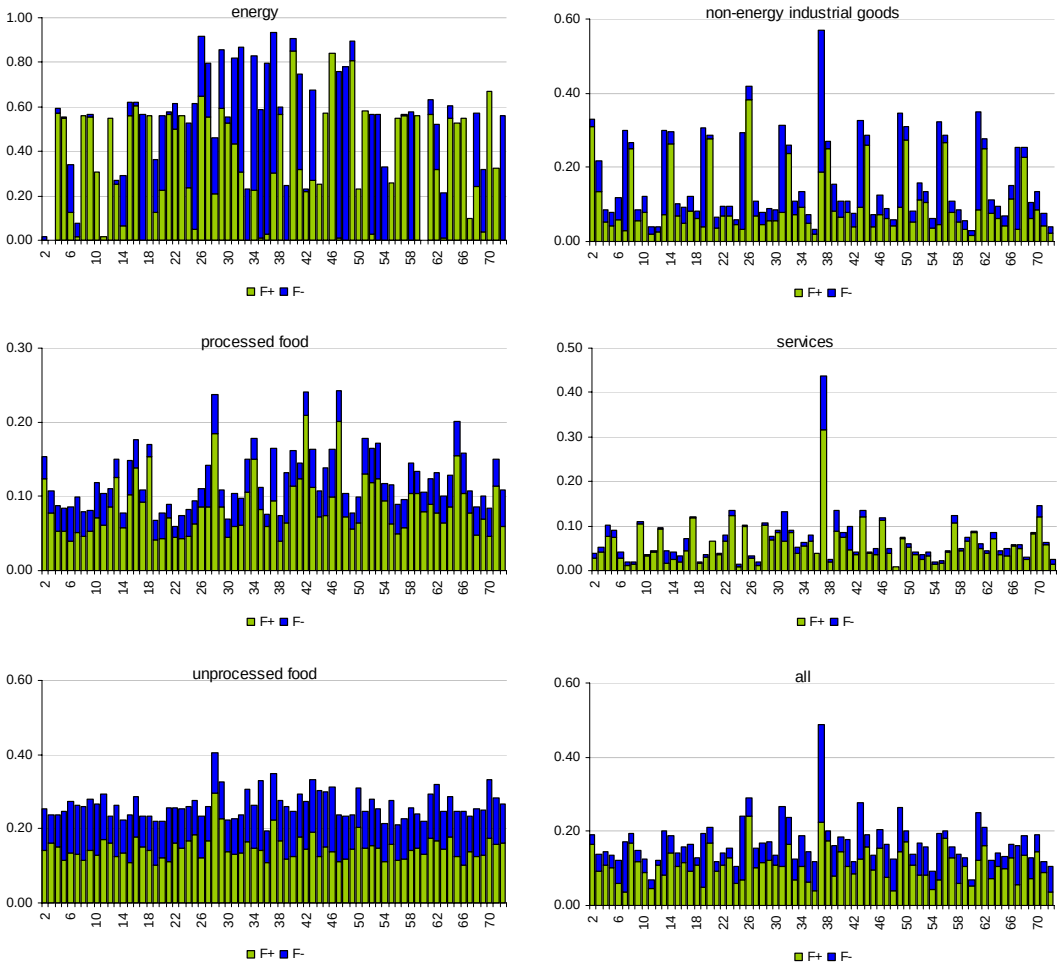


Figure 7: Price characteristics in euro: Last 2 decimal points, per product type

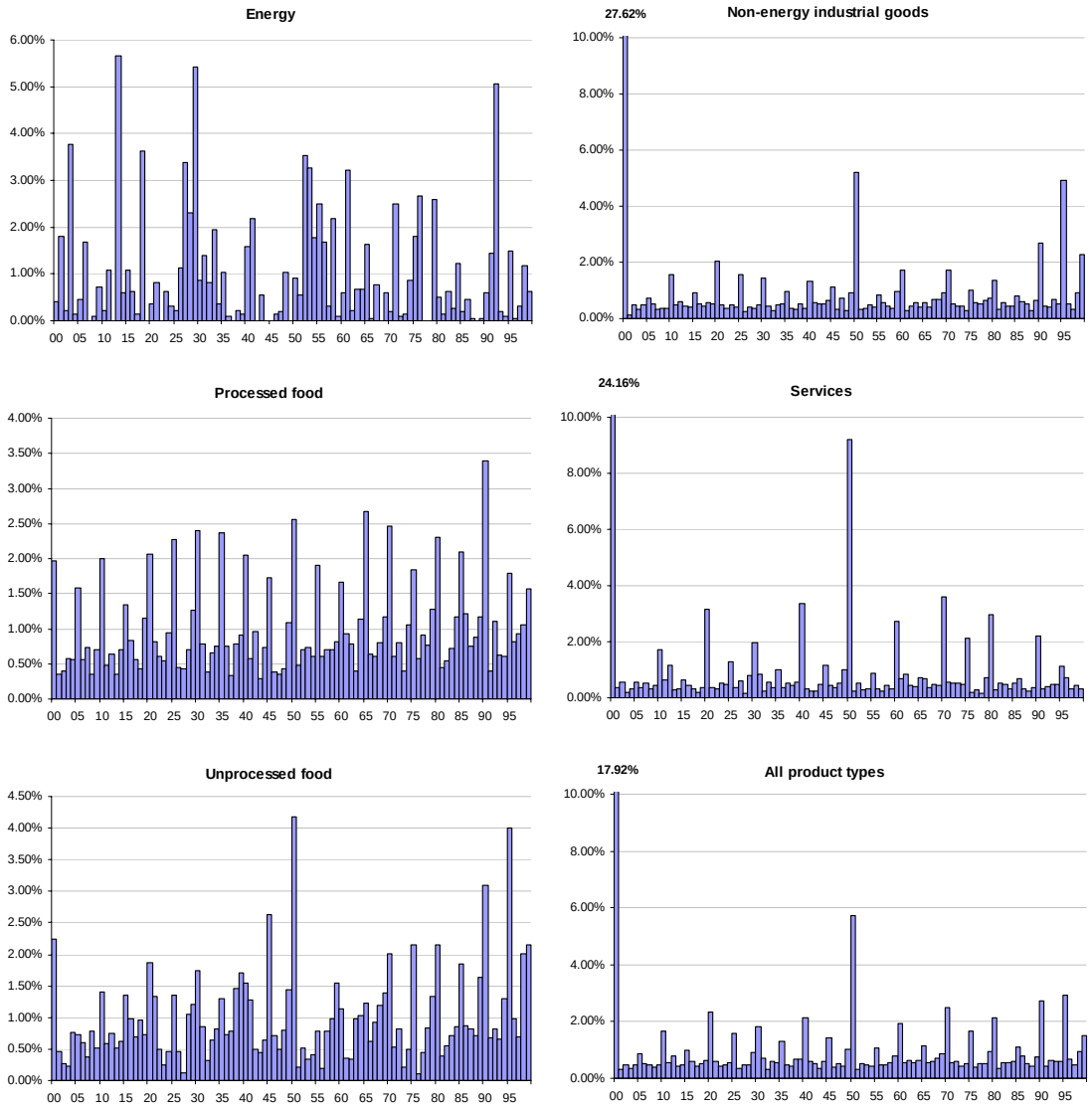
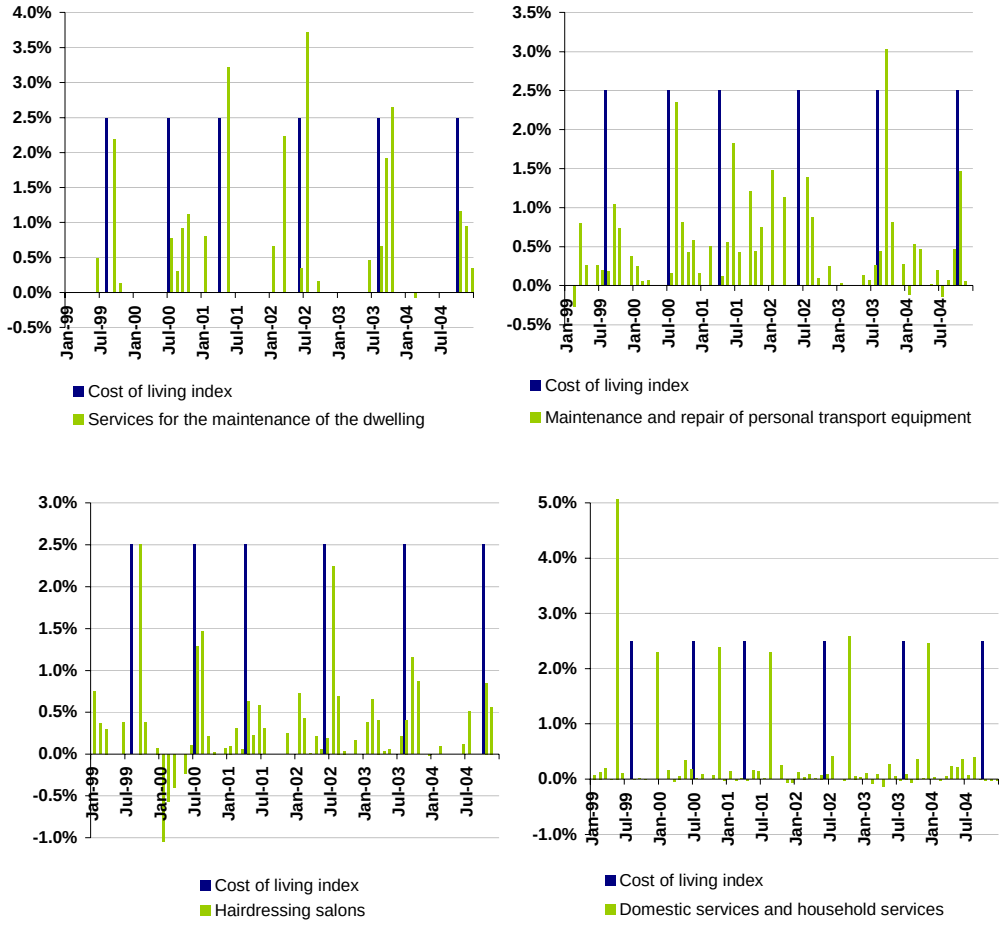


Figure 8: Automatic wage indexation and m-o-m price changes, 4 selected COICOP-10 indices



Appendix

Table A1: Product (-indices) supplied by quasi-publicly institutions or subject to price regulation

COICOP10	Description
04.04.01.01.01	Water supply
04.04.02.01.01	Refuse collection
04.04.03.01.01	Sewerage collection
04.04.04.01.01	Other services related to the dwelling n.e.c.
04.05.01.01.01	Electricity
04.05.02.01.01	Town gas and natural gas
04.05.02.02.01	Liquefied hydrocarbons
04.05.03.01.01	Liquid fuels
04.05.04.01.01	Solid fuels
04.05.05.01.01	Heat energy
06.01.01.01.01	Pharmaceutical products
06.01.02.01.01	Other medical products
06.01.03.01.01	Therapeutic appliances and equipment
06.02.01.01.01	Medical services
06.02.02.01.01	Dental services
06.02.03.01.01	Paramedical services
06.03.00.01.01	Hospital services
07.02.04.01.01	Toll facilities and parking meters
07.02.04.01.02	Roadwothiness tests, etc.
07.03.01.01.01	Passenger transport by railway
07.03.02.01.01	Passenger transport by road (bus, etc.)
07.03.02.01.02	Passenger transport by road (taxi, etc.)
08.01.00.01.01	Postal services
08.03.00.01.01	Telephone and fax services
09.04.02.01.02	Theaters, opera houses, musea, libraries, etc.
10.09.00.01.01	Education
12.04.00.01.01	Retirement houses, residences for disabl. persons, rehab. centres, etc.
12.04.00.01.02	Wet nurses, crèches, play schools & other child minding facilities

Table A2: Variables used in econometric investigation

Identifiers	
Product identifier	I
COICOP10 identifier	J
Period identifier	T
Spell identifier	S
Product price i at time t	p_{it}
Product type identifier	K
T	Starting date of price spell
Dependent variables	
y_{it}	Binary variable indicating a price change in $t+1$ (1, if $d\ln(P_{it+1}) \neq 0$, 0 otherwise)
y_{it}^{+}	Binary variable indicating a price increase in $t+1$ (1, if $d\ln(P_{it+1}) > 0$, 0 otherwise)
y_{it}^{-}	Binary variable indicating a price decrease in $t+1$ (1, if $d\ln(P_{it+1}) < 0$, 0 otherwise)
Explanatory variables	
$\Sigma p\pi_{i,t-T}$	Cumulated disaggregate price inflation since last price change (COICOP10)
$\Sigma w\pi_{t-T}$	Cumulated aggregate wage inflation since last price change
Δ_{it}^{+}	Size of last price increase
Δ_{it}^{-}	Size of last price decrease
$lpcdw_{it}$	1 if last price change was negative, 0 otherwise
$cco1_t$	Dummy variable for the euro cash changeover, 1 if t = December 2001, 0 otherwise
$cco2_t$	Dummy variable around the euro cash changeover period, 1 if $t \in$ {September 2001, …, March 2002} and $cco1_t = 0$, 0 otherwise
$cco3_t$	Dummy variable around the euro cash changeover period, 1 if $t \in$ {June 2001, … , June 2002} and $cco1_t = 0$ and $cco2_t = 0$, 0 otherwise
$itr1_t$	1 in t if wages are indexed in t, 0 otherwise
$itr2_t$	1 in t if wages are indexed in t-1, 0 otherwise
$itr3_t$	1 in t if wages are indexed in t-2, 0 otherwise
$itr4_t$	1 in t if wages are indexed in t-3, 0 otherwise
$itr1\_2_t$	1 in t if wages become indexed in t-1 or t, 0 otherwise
$attr_{it}^{ALL}$	1, if p_{it}, expressed in EUR or LUF, is a either round, fractional or psychological price
$attr_{it}^{EUR}$	1, if p_{it}, expressed in EUR, is a either round, fractional or psychological price
$attr_{it}^{LUF}$	1, if p_{it}, expressed in LUF, is a either round, fractional or psychological price
$round_{it}^{EUR}$	1, if p_{it}, expressed in EUR, is a round price
$round_{it}^{LUF}$	1, if p_{it}, expressed in LUF, is a round price
$fract_{it}^{EUR}$	1, if p_{it}, expressed in EUR, is a fractional price
$fract_{it}^{LUF}$	1, if p_{it}, expressed in LUF, is a fractional price
$psycho_{it}^{EUR}$	1, if p_{it}, expressed in EUR, is a psychological price
$psycho_{it}^{LUF}$	1, if p_{it}, expressed in LUF, is a psychological price
$regulated_{it}$	1, if product is supplied by (quasi-)public institutions or subject to price regulation
$leng_{i,t-T}$	Duration in months of a price spell in t
$dur_{i,t-T}^{d}$	1 if $leng_{i,t-T} = $ d, 0 otherwise {$d \in$ 1, 5, 6, 12, 18, 24}
$month_t^m$	1 if month(t) $= m$, 0 otherwise {$m \in$ 1, …, 12}
$year_t^a$	1 if year(t) = y, 0 otherwise {$a \in$ 1999, …, 2004}
$prodtype_k$	1 if product type = k, 0 otherwise {$k \in$ unprocessed food, processed food, non-energy industrial goods, energy, services}
$\Sigma supermkts_j$	Number of suppliers of a particular COICOP10 product category
$\Sigma p_s / \Sigma p$	Number of observations from supermarket s relative to all supermarket observations

European Central Bank working paper series

For a complete list of Working Papers published by the ECB, please visit the ECB's website (http://www.ecb.int)

509 "Productivity shocks, budget deficits and the current account" by M. Bussière, M. Fratzscher and G. J. Müller, August 2005.

510 "Factor analysis in a New-Keynesian model" by A. Beyer, R. E. A. Farmer, J. Henry and M. Marcellino, August 2005.

511 "Time or state dependent price setting rules? Evidence from Portuguese micro data" by D. A. Dias, C. R. Marques and J. M. C. Santos Silva, August 2005.

512 "Counterfeiting and inflation" by C. Monnet, August 2005.

513 "Does government spending crowd in private consumption? Theory and empirical evidence for the euro area" by G. Coenen and R. Straub, August 2005.

514 "Gains from international monetary policy coordination: does it pay to be different?" by Z. Liu and E. Pappa, August 2005.

515 "An international analysis of earnings, stock prices and bond yields" by A. Durré and P. Giot, August 2005.

516 "The European Monetary Union as a commitment device for new EU Member States" by F. Ravenna, August 2005.

517 "Credit ratings and the standardised approach to credit risk in Basel II" by P. Van Roy, August 2005.

518 "Term structure and the sluggishness of retail bank interest rates in euro area countries" by G. de Bondt, B. Mojon and N. Valla, September 2005.

519 "Non-Keynesian effects of fiscal contraction in new Member States" by A. Rzońca and P. Ciżkowicz, September 2005.

520 "Delegated portfolio management: a survey of the theoretical literature" by L. Stracca, September 2005.

521 "Inflation persistence in structural macroeconomic models (RG10)" by R.-P. Berben, R. Mestre, T. Mitrakos, J. Morgan and N. G. Zonzilos, September 2005.

522 "Price setting behaviour in Spain: evidence from micro PPI data" by L. J. Álvarez, P. Burriel and I. Hernando, September 2005.

523 "How frequently do consumer prices change in Austria? Evidence from micro CPI data" by J. Baumgartner, E. Glatzer, F. Rumler and A. Stiglbauer, September 2005.

524 "Price setting in the euro area: some stylized facts from individual consumer price data" by E. Dhyne, L. J. Álvarez, H. Le Bihan, G. Veronese, D. Dias, J. Hoffmann, N. Jonker, P. Lünnemann, F. Rumler and J. Vilmunen, September 2005.

525 "Distilling co-movements from persistent macro and financial series" by K. Abadir and G. Talmain, September 2005.

526 "On some fiscal effects on mortgage debt growth in the EU" by G. Wolswijk, September 2005.

527 "Banking system stability: a cross-Atlantic perspective" by P. Hartmann, S. Straetmans and C. de Vries, September 2005.

528 "How successful are exchange rate communication and interventions? Evidence from time-series and event-study approaches" by M. Fratzscher, September 2005.

529 "Explaining exchange rate dynamics: the uncovered equity return parity condition" by L. Cappiello and R. A. De Santis, September 2005.

530 "Cross-dynamics of volatility term structures implied by foreign exchange options" by E. Krylova, J. Nikkinen and S. Vähämaa, September 2005.

531 "Market power, innovative activity and exchange rate pass-through in the euro area" by S. N. Brissimis and T. S. Kosma, October 2005.

532 "Intra- and extra-euro area import demand for manufactures" by R. Anderton, B. H. Baltagi, F. Skudelny and N. Sousa, October 2005.

533 "Discretionary policy, multiple equilibria, and monetary instruments" by A. Schabert, October 2005.

534 "Time-dependent or state-dependent price setting? Micro-evidence from German metal-working industries" by H. Stahl, October 2005.

535 "The pricing behaviour of firms in the euro area: new survey evidence" by S. Fabiani, M. Druant, I. Hernando, C. Kwapil, B. Landau, C. Loupias, F. Martins, T. Y. Mathä, R. Sabbatini, H. Stahl and A. C. J. Stokman, October 2005.

536 "Heterogeneity in consumer price stickiness: a microeconometric investigation" by D. Fougère, H. Le Bihan and P. Sevestre, October 2005.

537 "Global inflation" by M. Ciccarelli and B. Mojon, October 2005.

538 "The price setting behaviour of Spanish firms: evidence from survey data" by L. J. Álvarez and I. Hernando, October 2005.

539 "Inflation persistence and monetary policy design: an overview" by A. T. Levin, and R. Moessner, November 2005.

540 "Optimal discretionary policy and uncertainty about inflation persistence" by R. Moessner, November 2005.

541 "Consumer price behaviour in Luxembourg: evidence from micro CPI data" by P. Lünnemann and T. Y. Mathä, November 2005.