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

1 We thank Statistics Austria (ST.AT) for providing the data and especially Paul Haschka for valuable information on the data. This study has been conducted in the context of the 'Eurosystem Inflation Persistence Network (IPN)'. We are indebted to the members in this network, especially to IPN – RG2 members, the participants of the Workshop 'Empirical Industrial Organization', held on June 18, 2004 at the Vienna University of Economics and Business Administration, and an internal seminar at the OeNB, Thomas Url, Michael Pfaffermaier and an anonymous referee for valuable comments and Ingrid Haussteiner for the translation of the goods basket. The views expressed in this paper are those of the authors and do not necessarily reflect those of the Oesterreichische Nationalbank (OeNB) or the Eurosystem. All remaining errors and shortcomings are our responsibility alone.
2 Austrian Institute of Economic Research (WIFO), P.O. Box 91, 1103 Vienna, Austria; tel. +43 1 798 26 01, ext. 230, e-mail: Josef.Baumgartner@wifo.ac.at.
3 Oesterreichische Nationalbank, Otto Wagner Platz 3, P.O. Box 61, 1010 Vienna, Austria; tel. +43 1 40420, ext. 7423, e-mail: Ernst.Glatzer@oenb.at.
4 Oesterreichische Nationalbank, Otto Wagner Platz 3, P.O. Box 61, 1010 Vienna, Austria; tel. +43 1 40420, ext. 7444, e-mail: Fabio.Rumler@oenb.at.
5 Oesterreichische Nationalbank, Otto Wagner Platz 3, P.O. Box 61, 1010 Vienna, Austria; tel. +43 1 40420, ext. 7435, e-mail: Alfred.Stiglbauer@oenb.at.

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

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CONTENTS

Abstract

Based on individual price records collected for the computation of the Austrian CPI, average frequencies of price changes and durations of price spells are estimated to characterize price setting in Austria. Depending on the estimation method, prices are unchanged for 10 to 14 months on average. We find strong heterogeneity across sectors and products. Price increases occur only slightly more often than price decreases. The typical size of a price increase (decrease) is 11 (15) percent. The aggregate hazard function of prices is decreasing with time. Besides heterogeneity across products and price setters, this is due to oversampling of products with a high frequency of price changes. Accounting for unobserved heterogeneity in estimating the probability of a price change with a fixed-effects logit model, we find a positive effect of the duration of a price spell. During the Euro cash changeover the probability of price changes was higher.

JEL classification: C41, D21, E31, L11

Keywords: Consumer prices, sticky prices, frequency and synchronization of price changes, duration of price spells.

Non-technical summary

In this paper we analyze the patterns and determinants of price rigidity present in the individual price quotes collected to compute the Austrian consumer price index. We calculated direct and implied estimates for the average frequency of price changes and the duration of price spells for 639 product categories.

We found that consumer prices are quite sticky in Austria. Depending on the method, our estimates for the weighted average duration of price spells for all products range from 10 to 14 months. The sectoral heterogeneity is quite pronounced: prices for services, health care and education change rarely, typically approximately once a year or even less frequently. For the product types food, energy (transport) and communication, prices are adjusted on average every 6 to 8 months. Temporal promotions and sales have a considerable impact on the frequency of price adjustments for food, clothing and footwear, where promotions and end-of-season sales are a common commercial practice.

With respect to the synchronization of price changes a similar sectoral pattern occurs as for the durations of price spells: The prices of products with a longer duration are also adjusted in a more synchronous way. This reflects the fact that prices of some of these products are either directly administered or strongly influenced by public authorities (including social security agencies).

Price increases occur slightly more often than price decreases, except for communication items, where due to the technical progress and high competition in the electronic equipment sector price decreases appear much more frequent than increases. Price increases and decreases are quite sizeable when they occur: on average, prices increase by 11 percent whereas prices are reduced on average by 15 percent. Especially for clothing and footwear (due to seasonal sales) and again for communication and electronic items (personal computers) price decreases are very pronounced (34 and 26 percent, respectively).

Like in similar studies conducted within the Eurosystem inflation persistence network, we find that the aggregate hazard function for all price spells (i.e. the duration of a price spell) is decreasing with time which is somewhat at odds with common price-setting theories. However, this is to a large extent a consequence of aggregating over product types with different spell durations. A re-weighted version of the hazard function which ensures that each product category basically has the same weight (and which is also adjusted for CPI weights) is not monotonously decreasing but has its most marked spike at a duration of 1 year which again indicates that for a substantial proportion of all goods infrequent price adjustment occurs. Using Kaplan-Meier estimates of survivor and hazard functions, we show that there is substantial heterogeneity across goods and product types: Energy and unprocessed food show high hazards during the first months. For services and non-energy industrial goods, on the other hand, the hazards are highest after one year.

In contrast to other European studies, we find a small, but highly significant positive effect of the duration of a price spell on the probability of a price change if we account for unobserved heterogeneity in a conditional logit model with fixed elementary product effects. We observe also a positive link between the probability of a price change and the accumulated value of inflation at the product level. Additionally, we find a seasonal (January) effect and a negative impact on the probability to change a price if it is currently set as an attractive price. The euro cash changeover seems to have temporarily increased the probability to change prices.

Although some time dependent aspects have a significant impact on the probability of a price change, our evidence does not support pure time dependent representations of the price setting process at the outlet level, as some of the state dependent variables also show a significant influence on the probability to observe a price change.

1. Introduction

The frequency of price changes or its counterpart the duration for which prices remain unchanged play a major role in the assessment of the impact of various shocks on the economy. Most macroeconomic models assume sluggish price and/or wage adjustment to generate real effects of monetary policy at least in the short run. The literature on the microeconomic foundation of price stickiness is vast (see Ball and Mankiw 1995, Taylor 1999 for an overview). However, due to the lack of individual price data and/or a restrictive practice of Statistical offices with respect to the use of the data for academic research, the empirical evidence on the relevance and patterns of price stickiness is sparse. In this paper we present empirical results on the characteristics of price stickiness for the Austrian economy.

Several papers have shown that for some products or product groups prices remain unchanged for many months. Cecchetti (1986), who looked at 38 U.S. news-stand magazine prices from 1953 to 1979, reported 1.8 to 14 years (!) since the last price change. Kashyap (1995), who studied the price changes of 12 mail order catalogue goods, found that on average prices were unchanged for 14.7 months. A series of papers by Lach and Tsiddon (1992, 1996) analyzes the price-setting behavior of firms by looking at the prices of 26 food products at grocery stores. However, all these studies faced the problem of small samples including only a (very) limited number of products and one has to make extremely strong assumptions on the sectoral (or product group) homogeneity for economy wide generalizations of their results.

Bils and Klenow (2004) used a much broader set of unpublished individual price data collected by the Bureau of Labor Statistics (BLS) for the calculation of the U.S. consumer price index (CPI). They found much more frequent price changes of consumer prices in the US than the studies mentioned above. For about half of the consumption goods, prices remain constant for less than 4.3 months. They also found that the frequency of price changes differs dramatically across goods.

For Euro Area countries until recently very limited evidence on this issue was available. Notable exceptions being Campiglio (2002) on Italy, Suvanto and Hukkinen (2002) on Finland, and Aucremanne et al. (2002) on Belgium. Thanks to the initiative of the Eurosystem Inflation Persistence Network (IPN) for 10 of the 12 Euro Area countries micro data evidence on frequencies of price changes and the duration of prices based on CPI data is now available. Dhyne et al. (2005) provide a summary of the research efforts in the analysis of individual consumer price data within the Inflation Persistence Network.[1]

In this paper we examine the frequency of consumer price changes and its counterpart the duration of price spells (i. e. the time span a price is unchanged) in Austria, using a unique data set of individual price quotes collected for the calculation of the Austrian consumer price index. The major aim is to analyze the degree and characteristics of the nominal rigidity present in Austrian consumer prices and trying to explain some factors influencing this rigidity.

We find that (depending on the estimation method), the average duration of price spells is 10 to 14 months, but that the duration varies considerably across sectors and products. For various fuel types and seasonal food products the average (median) duration – assuming price adjustment in continuous time – is less than one month, whereas for several services and administered prices, as e. g. banking, parking or postal fees, it is 50 (34) months and more. Like in similar studies, we find that the aggregate hazard function for all price spells is decreasing with time which is somewhat at odds with all relevant price-setting theories. However, apart from heterogeneity across products and price setters, one important reason for this is aggregating over product types with different spell duration because products with a large frequency of price changes contribute a disproportionately high number of spells. We show that - using an appropriate weighting scheme which attaches to each product category its CPI weight - the aggregate hazard function has its most marked spike at a duration of one year. Using Kaplan-Meier estimates of survivor and hazard functions, we show that there is substantial heterogeneity across goods and product types. Taking into account the unobserved heterogeneity in estimating the probability of a price change with a conditional logit model with fixed elementary product effects, we find a small positive

[1] For detailed results on each country see Aucremanne and Dhyne, (2004A) for Belgium, Dias et al. (2004) for Portugal, Baudry et al. (2004) for France, Álvarez and Hernando (2004) for Spain, Fabiani et al. (2004) for Italy, Jonker et al. (2004) for The Netherlands, Vilmunen and Paloviita (2004) for Finland and Hoffmann and Kurz-Kim (2005) for Germany.

but highly significant effect of the duration of a price spell on the probability of a price change. We also find that in the months before and after the Euro cash changeover the probability of price changes is higher than in other periods.

The paper is organized as follows. In Section 2 the data base and data manipulations like imputation of missing observations, outlier detection and correction, and other data issues (such as the problem of censored price spells) are discussed. The methodology of our analysis and the empirical results which summarize the vast information in the data are presented in Section 3. We compute direct and indirect estimates for the frequency of price changes and the duration of price spells. We also address the issue of the synchronization of price changes within product categories. In a first step to explain the stylized facts of price setting in Austria we adopt methods of survival analysis and run panel logit regressions in Section 4. The paper concludes with a summary of the main results.

2. Data

2.1 Data set and definitions

For investigating individual price dynamics in Austria, we use a longitudinal micro data set of monthly price quotes collected by Statistics Austria (ST.AT) in order to compute the national index of consumer prices, CPI.[2] The sample spans over the time period from January 1996 to December 2003 (96 months) and contains between 33,800 (1996) and 40,700 (2003) elementary price records per month. As can be seen from Table 1, overall around 3.6 million price quotes which cover roughly 90% of the total CPI are included in the *raw data set*. Only the COICOP groups 2 and 7 (alcoholic beverages and tobacco and transport) are underrepresented in our sample.[3] Table 1 also reveals that

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[2] See Statistics Austria (2001A). We distinguish between the 'raw data set' (the data set we received from ST.AT) and 'our data set' (the data set after some manipulations as exclusions, imputations etc. – see below for details) in the text.

[3] COICOP stands for 'Classification Of Individual COnsumption by Purpose' (see Statistics Austria 2001B). Tobacco products, cars, daily newspapers and mobile phone fees were not included in the raw data set for confidentiality reasons by Statistics Austria.

most observations in our sample can be attributed to the food and clothing sectors (COICOP groups 1 and 3). About 40 percent of the product categories are collected centrally (e.g. regulated services, energy, housing), which account for only 7 percent of the observations in our data set. The main portion of price quotes is collected in a decentralized way in 20 major Austrian cities.

Each individual price quote consists of information on the product category, the date, the outlet and the packaging (quantity) of the item (see Table 2). As the product category we define the products at the elementary level which are contained in the CPI basket (e.g. milk). The raw dataset contains a total of 668 products categories. For each product category the product variety denotes the specific variety and brand of the product. For confidentiality reasons the raw dataset has been anonymized with respect to the variety and brand of the product, i.e. we do not have any information on the brand.

With the information on the date (t), the outlet (k) and the product category (j) we can construct a *price trajectory* $P_{jk,t}$, that is a sequence of price quotes for a specific product belonging to a product category in a specific outlet over time. To be more specific, two consecutive price quotes belong to the same *elementary product* if the following conditions are fulfilled:

(i) the time difference between the records is one month,

(ii) the outlet[4] and the product codes are equal for t and t+1 and,

(iii) no product replacement and no outlet replacement occurred in t or t+1.

A *price spell* is defined as the sequence of price quotes (for a specific product in a specific outlet) with the same price.

[4] In the data set we received from Statistics Austria only an outlet code but no store code was included. The latter is usually defined as a location where different kinds of products are sold. In our data set stores selling goods that belong to different sectors (a total of 48 sectors according to the kind of product have been defined by ST.AT) cannot be identified as the same store across goods because stores are coded as individual outlets according to the sector of the good reported. E.g. if bread and stationery are sold in the same store, both products get a different outlet code as these products belong to different sectors.

2.2 Data manipulations: imputations, outliers and sales

For the calculation of the descriptive statistics the raw data set has been modified and data manipulations have been carried out with the aim of bringing them in an appropriate form for our numerical analysis. In the case of temporal unavailability of a price quote the price has been *imputed* with the previous price quote for at most one month. Filling the (one-month) gaps of missing observations mitigates the problem induced by censored price spells (see next sub-section). In case the price quote was unavailable for more than one month it has not been imputed, because the chance of missing an unobserved price change becomes more and more likely with the duration of missing observations. On the other hand, individual prices quotes which were imputed by the statistical office due to temporal and seasonal unavailability of an item (codes F, G and V in Table 3) were excluded from our data set, however with the disadvantage of creating additional censored spells. We do not regard them as true price observations but as "pseudo observations", which unintentionally would introduce an upward bias in the estimation of the duration of price spells. Additionally, all price quotes are converted into prices per unit in order to account for package changes and temporary quantity promotions. The prices around the cash changeover to the Euro have been converted into common currency to make them comparable over the cash changeover.

Some products which display systematically unrealistic price movements were removed as outliers from the data set mainly on a judgmental basis. The nature of these products as outliers was reflected by the fact that they all displayed *average* price increases or decreases of more than 60 percent, some of them considerably more (according to the log price difference, $\ln(P_{jk,t}) - \ln(P_{jk,t-1})$). On this basis, 14 products (e.g. kindergarten fees, public swimming pool, refuse collection, public transport day ticket) have been excluded representing a weight of 1.4 percent in the total CPI. In addition, very large individual price changes exceeding a pre-defined threshold value have been identified as outliers and disregarded in the analysis. We applied a combined rule specifying an absolute value for the log price change and a distribution dependent upper and lower bound as the threshold for outliers. Specifically, all individual price changes with $\left|\ln(P_{jk,t}) - \ln(P_{jk,t-1})\right| \geq 1$ as well as those exceeding the upper and lower quartile of the distribution of price changes plus 3 times the interquartile range have been defined as outliers. This rule turned out to be a rather conservative way of outlier detection such that only a few observations had to be excluded.

In addition, based on information from Statistics Austria, 14 products whose price quotes already contain aggregated information have been removed for the purpose of our analysis as they do not represent price quotes on the micro level (e.g. rents and operating costs for houses are derived from the micro-census of Austrian households, and a few medical services are obtained from the social insurance institution). After the exclusion of these products together with the outlier products, individual price quotes for 639 product categories are included in our data, covering 80 percent of the CPI. Figure 1 shows the price change distribution where observations of zero price changes (which would produce very large spikes of 70 percent or more) were dropped. The five histograms differ considerably: Goods in the unprocessed food and processed food categories have a comparably large dispersion of price changes with especially unprocessed food items being characterized by many large price changes. A similar observation can be made for non-energy industrial goods. Services and energy goods have a much smaller variance of price changes. The distribution for energy is almost symmetric, whereas for services it is markedly skewed towards positive price changes.

With the introduction of a *revised goods basket* for the CPI data collection in January 2000 (see Statistics Austria, 2000A, 2000B, 2001B), definitions and reporting practices were changed for many products. This makes a comparison of prices reported in December 1999 and January 2000 unfeasible for many products. As a consequence, all price changes from December 1999 to January 2000 have been disregarded in the computation of the descriptive statistics, given the large number of products affected by the revision of the Austrian CPI basket.[5] In the econometric analysis in Section 4.3

[5] Specifically, all observations in January 2000 have been disregarded in the nominator as well as in the denominator to compute the frequency of price changes (see Section 3.1), i.e. they have been disregarded regardless if a price change occurred or not (and therefore are neither treated as non-price changes nor as missing values). This effectively reduces the number of months which are included in our descriptive analysis in Section 3.1 by one. It turns out that including January 2000 as a regular month would not affect the frequency of price changes much, as the frequency in January 2000 is about in the range of the frequencies in the first month of other years, but rather the size of price changes: The average price increase (decrease) of 18.7 (22.6) percent in January 2000 is about 7 percentage points higher than the average over all other January figures.

these price changes have not been excluded but a dummy variable has been introduced to account for this effect.

Concerning the price changes associated to promotions or (seasonal) *sales* we decided to follow a dual approach: In the baseline version of the results we treat promotions and sales as regular price changes which terminate a price spell. However, it can be argued that these price changes merely reflect noise in the price setting process and are not due to changes in fundamental price determining factors (as e.g. monetary policy and business cycle developments) and therefore they should be ignored from the viewpoint of monetary policy analysis. Therefore, we also provide an alternative set of results in Section 3 without taking into account the price changes induced by temporary promotions and sales. The information in our data set allows us to identify observations that are flagged as sales (code A in Table 3). In order to exclude price changes induced by flagged sales from our analysis, we replaced all flagged sales prices with the last regular price, i.e. the price before the sale or promotion started. As the reporting of sales and promotions is generally up to the interviewer and therefore cannot be expected to be complete and consistent across all products, we additionally tried to identify also those temporary price promotions which have not been coded as sales. We define "unflagged" temporary promotions and sales as a price sequence $P_{jk,t-1}$, $P_{jk,t}$, $P_{jk,t+1}$, where $P_{jk,t-1} = P_{jk,t+1}$, and $P_{jk,t-1} \neq P_{jk,t}$, i.e. price changes that are reversed in the following period. As in the case of flagged sales, the price changes induced by unflagged promotions and sales have been excluded from the analysis by replacing all identified prices ($P_{jk,t}$) with the preceding regular prices ($P_{jk,t-1}$).[6]

2.3 Censoring, product replacement and weighting

At the beginning and at the end of the sample period all price trajectories are *censored*, as we do not know the true starting date of the first price spell and the ending date of the

[6] Flagged and unflagged sales and promotions are a quite common feature in the data, in particular in the food and cloth sectors. Overall, about 4 percent of all prices in our data set are flagged as sales prices while the share of prices identified as unflagged sales and promotions amounts to about 1.5 percent of the total number of observations. The effect of excluding all price changes that are due to (flagged and unflagged) sales and promotions can be assessed by comparing the results in Tables 4 and 5 (see Section 3.1).

last price spell. A price spell is left (right)-censored if the date of the beginning (end) of the spell is not observed, and double-censored if both the start and the end date of the spell are unknown. Censoring entails a downward bias in the estimation of the duration of price spells, as longer spells are more likely to be censored.

The products underlying the price observations are sometimes *replaced* in the database by others for two reasons: When a product is no longer available in a particular outlet (attrition), it is usually replaced by another product of the same product category which terminates the price spell (and the trajectory). However, products are sometimes also replaced due to the sampling strategy, e.g. when Statistics Austria defines another elementary product to be more representative for the product category. Unfortunately, we have no information on the nature of the product replacements, in particular not if they are forced or voluntary. As according to Statistics Austria, the major part of product replacements in our database are forced replacements due to attrition, we count the end of each price spell associated with a product replacement as a price change.

In order to compute aggregate measures of the statistics described in Section 3, we applied the same *weighting scheme* that is used to calculate the CPI. As these weights are not defined at the individual store level, we use an unweighted average over price records within a product category. All statistics at the elementary products level are then aggregated to 12 COICOP groups and 5 product types based on the CPI weights. As our data set spans over two goods baskets (1996, 2000) and the products included do not completely coincide, the average weights of the two weighting schemes are used, with a weight of zero at times when an elementary product was not included in the respective CPI basket. The individual weights which initially do not sum to one as not 100% of the CPI is covered in our sample, are then rescaled such that the sum of the weights equals 1 and the relative weights among the goods are preserved.

2.4 Exemplary price trajectories

Figures A1 to A5 in Annex III show exemplary price trajectories for the five product groups unprocessed food, processed food, energy, non-energy industrial goods and services. For each of the groups, the twelve product categories with the highest CPI weights were chosen. Then, from each of the product categories, an elementary product was chosen randomly and its price trajectory is shown where all price observations were converted to Euros (the period of the cash changeover is highlighted). Price observations are marked by symbols and connected by lines except in the case where price observations were temporarily unavailable or missing at all (see, for example product category "Turkey breast" in Figure A1). Note that in quite a number of cases trajectories are shorter than 96 months (e.g. "Rye-wheat bread" in Figure A2, "Gasoline (super)" and "Heating oil extra light" in Figure A3) and/or are characterized by repeated gaps (apparently due to seasonal unavailability such as "Woman's jacket" or "Woman's overcoat" in Figure A4). Pricing behavior is very heterogeneous: Prices are adjusted infrequently (such as "pressed ham" in Figure A1, "Whole milk" in Figure A2). In some cases, these infrequent price changes appear to happen regularly every twelve months (such as "Rye-wheat bread" in Figure A2, "Aerated concrete blocks" in Figure A4 or "Apartment cleaning services" in Figure A5). Especially fuel prices change practically every month (Figure A3). Other goods prices are changed also very often but appear to be oscillating (such as "Salami" in Figure A1 or "Coffee beans" in Figure A2).

3. Methodology and empirical results

The descriptive analysis of the degree of price rigidity at the micro level using individual price quotes is either based on the frequency of price changes or vice versa on the duration of price spells. Prices are considered as rigid if they show a low frequency of price changes and therefore a long duration of price spells. In this Section we briefly describe the applied methodologies for both approaches and discuss the results. In Section 3.1 we describe how the frequency of price changes (with and without sales and promotions) is calculated and the implied duration of price spells is derived from these numbers. In Section 3.2 we present the results for the synchronization of price changes and price increases and decreases separately. Finally, we also present a direct computation of the duration of price spells and derive its counterpart, an implied frequency measure (in Section 3.3).

3.1 The Frequency Approach

According to the frequency approach the *frequency of price changes* (F) is computed directly from the data and the duration of price spells (T) is derived indirectly from the frequency. An advantage of the frequency approach is that it uses the maximum amount of information possible, implying that it can be used even if the observation period is very short and if specific events, such as the revision of the CPI basket or the Euro cash changeover, need to be excluded from the analysis. In addition, it does not require an explicit treatment of the censoring of price spells. The frequency approach and the duration approach are equivalent only if all spells are uncensored.

For each product category *j*, the frequency of price changes (F_j) is computed as the ratio of observed price changes to all valid price records. Thus, the measure F_j is an average incorporating price changes of all firms where the product *j* has been recorded and over all periods of time. The implied duration of price spells could be calculated as the inverse of the frequency of price changes $T = \frac{1}{F}$. However, for this estimator to be consistent homogeneity of observations in the cross-sectional dimension is required. Another issue to be considered for the derivation of the implied duration of price spells is the discrete timing of observations: We observe only one price per month and implicitly assume, if we observe a price change, that the price change occurred at the end of the month and the price remained unchanged for the rest of the month. Relaxing this assumption and allowing for continuous timing and assuming that the durations of price spells follow an exponential distribution, the *implied average duration* of price spells can be estimated as

$$T_j^{F,avg} = \frac{-1}{\ln(1 - F_j)} \tag{1}$$

and the *implied median duration* as

$$T_j^{F,med} = \frac{\ln(0.5)}{\ln(1 - F_j)}. \tag{2}$$

In other words, these expressions are unbiased estimates of the mean and median duration of price spells in continuous time under the assumption of a constant hazard rate, i.e. assuming that the probability of a price change is constant within a month (see Baudry et al., 2004, and Bils and Klenow, 2004).

In Tables 4 to 6 the results aggregated on the COICOP and product type level are presented. If not otherwise stated, all statistics in these tables are computed by accounting for product replacements.

Price rigidity as measured by the implied duration based on the frequency of price changes varies considerably (see Table 4). On average, 15 percent of all prices are changed every month, which implies an average (median) duration of price spells of 14 (11) months. Unprocessed food and energy products display a rather high frequency of price changes (24 and 40 percent) and thus a short implied duration (6.5 and 8.3 months, respectively). Within these categories seasonal food products and fuels of different types show the highest flexibility. Due to the continuous time assumption to derive formula (1) and (2), for these products the implied durations are smaller than one month, although the observation frequency is monthly. However, this is not unrealistic since fuel prices are indeed changed with a very high frequency – sometimes even on a daily basis, so estimates of a duration of less then one month are not unreasonable.

In contrast, some service items as well as products with administered prices display a (very) low frequency of price changes and, on average, a duration which is almost three times as long as for unprocessed food. So, e.g. banking, parking and postal fees show an estimated average duration of 50 months or longer.

If we analyze price increases and decreases separately, we realize that prices increase slightly more often than they decrease: the frequency of price increases is 8.2 percent compared to 6.6 percent for price decreases. Exceptions from this pattern can be found in the category communication (especially personal computers), where price decreases appear much more frequent than price increases. Concerning the size of price changes, price increases and decreases appear to be quite sizeable when they occur. The average price increase is 11 percent whereas prices are reduced on average by 15 percent. Especially for cloth and footwear (due to seasonal sales) and again for communication and electronic items (personal computers) price decreases are very pronounced.

As has been mentioned before, the results on the frequencies of price changes and the implied duration of price spells are also computed *without sales and promotions* (Table 5). For all product groups the frequency of price changes have to be smaller (or equal) compared to the figures in Table 4. It also turns out that also the average size of price changes is smaller without sales and promotions reflecting the fact that price cuts due to seasonal sales especially in the clothing sector are usually quite sizeable. As expected, these effects are most pronounced for food and alcoholic beverages where temporary promotions are a common practice to attract new customers, as well as for cloth and footwear where end of season sales are a common practice to clear inventories. For the latter category the average price decrease (in absolute terms) is almost 15 percentage points lower if sales are disregarded.

In addition to the frequency of price changes calculated as an average over all periods, it can also be calculated for each *t*. When looking at the frequency of price changes over time we can see that there is a clear seasonal pattern visible in Figure 2: The spikes in January 1998, 1999, 2001, 2002 and 2003 indicate that most prices are changed in January.[7] Starting with the year 2000 price changes have been more frequent than before which seems to correspond with higher aggregate inflation in the period 2000-2003 than in the period 1996-1999. Apart from this shift in 2000, there is no trend in the frequency of price changes visible over the period considered. Furthermore, price increases and decreases show a similar seasonal pattern.

Equivalent to the frequency of price changes, also the size of the price changes can be calculated for each period in time. Figure 3 plots the weighted average (over all goods) of the absolute size of all price changes as well as the sizes of price increases and decreases over time. The graph reveals a strong seasonal pattern especially for price decreases: Price decreases appear to be more pronounced in January and February as well as in July and August of each year which clearly reflects end-of-season sales usually taking place in that period of the year. Consequently, also price increases display a seasonal pattern as the price decreases which are due to sales are usually reversed in the following period implying higher price increases in March and September, but this pattern appears to be less clear-cut than the seasonal pattern of price decreases. The most striking observation from the figure is the decrease in the size of price changes in the second half of 2001 reaching a low of less than 10 percent in January 2002 and increasing again thereafter. This development is clearly attributable to the euro cash changeover which obviously induced many small price changes when prices had to be converted from the old to the new currency. In addition, the size of price increases and the size of price decreases turned out to be roughly equal in January

[7] Note that price changes in January 2000 have been excluded from the analysis, see Section 2.2.

2002 which is at odds with the seasonal regularity of larger decreases than increases normally observed in January. Disregarding the smaller than average price changes in 2001 and 2002, there is no upward or downward trend visible in the development of the size of price changes in Figure 3 with the average size of price changes fluctuating around 15 percent most of the time.

Taking together the evidence for the frequency of price changes and the size of price changes in Figures 2 and 3 we find that in the period surrounding the cash changeover (from about mid 2001 to mid 2002) consumer prices were adjusted more frequently but by smaller amounts than in other times. In addition, price adjustment with respect to both the frequency and the size of price changes was quite symmetric during the cash changeover period. This implies that our dataset – to the extent that it is representative for the total CPI – does not suggest a sizeable positive nor negative impact of the cash changeover on aggregate inflation.

3.2 Synchronization of price changes

For each product the *synchronization of price changes* (SYNC_j) is measured by the approach proposed by Fisher and Konieczny (2000) which is given as the ratio of the empirical standard deviation of the frequency of price changes for product category *j* (numerator) to the theoretical maximum standard deviation in the case of perfect synchronization of price changes (denominator)

$$SYNC_j = \frac{\sqrt{\frac{1}{\tau - 1}\sum_{t=2}^{\tau}\left(F_{jt} - F_j\right)^2}}{\sqrt{F_j\left(1 - F_j\right)}} \tag{3}$$

where τ is the total number of periods for which the ratio is calculated. Perfect synchronization of price changes occurs when either all stores change their price at the same time or none of them changes a price. Consequently, synchronization of price changes is high if the synchronization ratio is near 1 and low if it is near 0. Analogous expressions are applied for price increases and decreases, with the only difference that in the calculation of the frequencies of price increases and decrease we did not account for product replacements because price changes cannot seriously be divided into price increases and decreases as the new price corresponds to a different product in the case of a product replacement.

The results in Table 7 show that the average synchronization ratio of price changes for all products amount to 42 percent which constitutes an intermediary degree of price synchronization. However, this number greatly masks the heterogeneity across sectors and products: There is a wide range from 20 percent for alcoholic beverages to 87 and 94 percent for health care and communication items, respectively. Prices in education and health care are regulated to a large extent, and in most cases these changes are price increases.[8] For food items the synchronization ratios are also very low, with an average of 21 percent. Furthermore, we observe that the synchronization ratio is generally higher for price increases than for decreases. This could reflect price changes that are triggered mainly by supply shocks, as the observed asymmetry is especially pronounced for energy products.

With the exception of alcoholic beverages and cloth and footwear the results calculated without the price changes induced by sales and promotions are very similar (see Table 8). As expected, for these products the exclusion of promotions and seasonal sales results in a synchronization ratio for price decreases which is considerably lower (by 4 and 7 percentage points, respectively) compared to the results including sales and promotions.

3.3 The Duration Approach

According to the *duration approach* the duration of price spells is directly computed from the price trajectories in the data and prices are considered rigid if the durations are long and flexible if they are short. The implied frequency of price changes is then derived from the duration of spells using formulas (4) and (5). In contrast to the frequency approach, the duration approach directly deals with the issue of censoring of price spells which has a considerable influence on the results.

From Table 9 it can be seen that a majority of 63 percent of the price spells is not censored, and that left-censored spells (25.3%) are much more frequent than right-

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[8] As already mentioned, the synchronization ratios for price increases and decreases are based on calculations without accounting for product replacements. As a consequence, the value for all changes (with replacements) need not necessarily lie within the range given by ratios for increases and decreases for each product category.

censored spells (7.7%). The asymmetry between left-censored and right-censored spells is a consequence of product replacements: when a (forced) replacement occurs, the ending spell is considered to be non-censored while the new spell is typically left-censored. Among all spells the double-censored spells tend to be the longest (15.7 months), while the non-censored spells are shorter on average (8.5 months), which might be due to the fact that longer price spells are more likely to be censored than short ones.[9] The median spell duration is much shorter than the average duration of spells (only 1 month for non-censored spells and 4 months for left-censored spells) which can be explained by the dominance of spells with a duration of 1 month indicating an extremely skewed distribution of spell durations (see Figure 4).[10]

In the period January 1996 to December 2003 a total of 520,041 price spells (including also censored spells) are observed (see Table 10). Most price spells are observed in the food and clothing sectors. The weighted mean duration of a price spell is about 10 months; the weighted median spell duration, however, is only 2 months, which is determined by the result that 38% of all spells have a duration of 1 month (see Figure 4). The longest price spells can be found in the service sector, whereas durations are relatively short for food and energy items. Compared with the results of Table 4, the direct computation of spell durations yields a similar pattern across COICOP groups and products types, but the durations are shorter, in several cases considerably shorter (services, education, health care), where prices are rather sticky and therefore the downward bias induced by censoring becomes more severe. Apart from that, part of the difference in spell durations between the two approaches can also be explained by the non-linear transformation of frequencies into durations in the frequency approach where

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[9] From these figures it can be seen that our database exhibits a large proportion of in-sample censoring: Besides product replacements also missing observations, which seem to be quite frequent for a number of products, by interrupting the trajectories bias the duration of spells downward and additionally induce censoring of the spells before and after the missing value.

[10] Note that for the duration approach the median duration of price spells is calculated directly from micro observations while under the frequency approach the implied median duration is converted from the frequency of price changes, which is already an aggregate number, according to formula (2). This explains the large difference between the two measures.

low frequency products yield very long implied durations by the transformations shown in equations (1) and (2).

The implied frequencies of price changes (F^{imp}_j) in the last two columns of Table 10 are again calculated with and without accounting for product replacements. According to Aucremanne and Dhyne (2004A), equation (4) shows an unbiased estimator for the frequency of price changes only if the number of left-censored and right-censored spells is equal

$$F_j^{imp} = \frac{n_{j,nc} + n_{j,lc}}{n_{j,nc}T_{j,nc} + n_{j,lc}T_{j,lc} + n_{j,rc}T_{j,rc} + n_{j,dc}T_{j,dc}}, \tag{4}$$

where $n_{j,nc(lc)[rc]\{dc\}}$ is the number of non-censored (left-censored), [right-censored] and {double-censored} spells

and $T_{j,nc(lc)[rc]\{dc\}}$ is the corresponding duration of non-censored (left-censored), [right-censored] and {double-censored} spells for product category j.

In the case of product replacements, i.e. taking into account of replacements induces an asymmetry between left-censored and right-censored spells, an unbiased estimator of the implied frequency of price changes is given by

$$F_j^{imp,rep} = \frac{n_{j,nc} + n_{j,lc}}{n_{j,nc}T_{j,nc} + n_{j,lc}T_{j,lc} + n_{j,lc}T_{j,rc} + n_{j,lc}\left(\frac{n_{j,dc}}{n_{j,rc}}\right)T_{j,dc}}. \tag{5}$$

As can be seen from Table 10, the estimates for the implied frequency accounting for product replacements are consistently lower than without controlling for product replacements which again illustrates the bias induced by asymmetric censoring of price spells. Given that product replacements are an important feature of our database, the implied frequencies in the last row of the table should be the more appropriate ones. Comparing the results with those from the direct estimation of the frequency of price changes (see Table 4), we observe the same pattern across COICOP groups and product types and also similar magnitudes of the frequencies for both approaches: The average implied frequency of all products amounting to about 15 (12) percent compares to the average frequency of 15 percent calculated according to the frequency approach.

The unweighted distribution of all spell durations is shown in Figure 4. As can be seen the distribution is extremely skewed, with a concentration of spells with very short durations: about 38 percent of all spells last only one month and about 86 percent of all spells are terminated within one year. In addition, local modes are visible at durations of 6, 12 and 24 months. This suggests that a substantial portion of firms change their prices at fixed intervals, which is consistent with a Taylor-type price setting behavior.

3.4 Discussion

As a result, depending on the estimation method, price spells last on average 10 to 14 months with a considerable variation among products. Compared with other European countries price adjustment in Austria depicts similar patterns across product groups. Also for the aggregate, the duration of price spells and the frequency of price changes are similar to the other countries as they are not far from the average of all euro area countries considered; see Dhyne et al. (2005). Further estimates of durations by product categories (Table 11), based on the Kaplan-Meier estimator of the survivor function, are presented in the next section.

4. The Probability of Price Changes

As is shown in the previous section price setting is very heterogeneous among products and also within a product group. To gain further insight in the determination of the frequency of price changes we present estimates of hazard functions (Section 4.2) and of a panel logit model for the probability of a price change (Section 4.3). For similar studies for other euro area countries see Álvarez and Hernando (2004) for Spain, Baudry et al. (2004) and Fougère et al. (2004) for France, Aucremanne and Dhyne (2004B) for Belgium, Dias et al. (2005) for Portugal and Jonker et al. (2004) for the Netherlands.

4.1 Data and sample selection

After excluding several product categories (see subsection 2.2), our full sample contains 639 different product categories, 1,888 product varieties and 49,766 combinations of product categories (j) and outlet codes (k), with a total of 3.6 million price observations. As regards the panel structure, the most common case is that the records span the full period from Jan. 1996 to Dec. 2003 (46.1 percent of all combinations of product categories and outlets). Because our data contain two CPI baskets, many such combinations show up only from Jan. 1996 to Dec. 1999 (1996 CPI basket; 10.8 percent of all products-outlet combinations) or from Jan. 2000 to Dec. 2003 (2000 CPI basket, 14.1 percent of all combinations). Other patterns (including price trajectories with gaps) account for the rest.

As already explained in Section 2 censoring constitutes a problem for the estimation of the average length of price spells. While dealing with right-censored spells is rather easy, left-censoring is a more serious problem. For each elementary product, the first price spell is left-censored because we cannot know for how long the price has been unchanged. Furthermore, every spell after a product replacement is also regarded as left-censored.[11] This comes close to "stock sampling" which constitutes a sample selection problem. A way to overcome this bias is to omit all left-censored spells from the analysis. Then only those spells are considered where we know exactly when the spell started. This is also called "flow sampling" and does not constitute a selection problem if at least one price change for every elementary product is observed (see Dias et al. 2005).[12]

After dropping left-censored spells, we are left with a dataset that consists of 42,832 product-outlet combinations, contributing to 366,102 price spells or 1,879,929 monthly price observations.

Product-outlet combinations, however, are not identical to "elementary products" as defined in Section 2 because they do not consider product and store replacements. As can be seen from Table 3, in particular product replacements occur quite often. For the panel logit regressions below we construct a subject variable which should correspond closely to the definition of an elementary product over time: In any case where a product or a store replacement is observed we change the identifier of the product-outlet combination. This results in 72,892 elementary products which is considerably higher than 42,832, the number of different product-outlet combinations.

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[11] Only when a product is replacement by another product newly introduced to the market the starting point of the spell is known. But this information is not available.

[12] Other studies within the IPN also follow this approach (e. g. Aucremanne and Dhyne, 2004A, B, and Fougère et al., 2005).

In Figure 5 the effect of the exclusion of left-censored spells on the average frequency of price changes is shown. As can be seen, for the first two years of the sample the frequency of price changes is much higher and shows a falling trend. This is due to the fact that at the beginning of the sample all longer spells started at an unknown date before January 1996 and are therefore excluded. As during the first months only the very short spells are kept in the data set, the frequency of a price change has to be higher, but falling as time passes and then also longer spells, which started after January 1996, are taken into account. However, the frequency of price changes after excluding left-censored price spells is also higher on average, as left-censored spells occur also within the sample period, e. g. because of a product replacement, a gap in the collection of the price of an elementary product or a temporary unavailability of a price quote, which are also excluded from the dataset.

4.2 Kaplan-Meier estimates of survivor and hazard functions

Unweighted and weighted versions of aggregate survivor and hazard functions

In survival analysis, the time until an event occurs (in our case, a price change) is the variable of particular interest. Its advantage for the topic of price setting behavior is that the duration of price spells and the shape of the hazard function are treated more explicitly. We present Kaplan-Meier estimates of the survivor and hazard functions for all products and separately for product groups. Particular emphasis is given to the question how the weighting of spell observations influences the results.

Figure 6 shows the Kaplan-Meier estimate of the survivor function for all price spells of all elementary products in our data. Panel (a) is the "unweighted" version whereas panel (b) is "weighted" in a sense which will be explained in a moment. The Kaplan-Meier estimator is a non-parametric estimate of the survivor function *S(t)*, the probability of "survival" of a price spell until time *t*. For a dataset with observed spell lengths $t_1,\ldots,t_k$ where *k* is the number of distinct failure times (time until a price change) observed in the data, the Kaplan-Meier estimate at any time *t* is given by

$$\hat{S}(t) = \prod_{j|t_j \le t} \left(\frac{n_j - d_j}{n_j} \right) \tag{6}$$

where n_j is the number of price spells "at risk" of exhibiting a price change at time t_j and d_j is the number of price changes at time t_j. The product is calculated over all observed spell durations less than or equal to t (see, for example, Cleves et al., 2002). The interpretation of the survivor function is as follows: For each analysis time t, the step function gives the fraction of price spells which have a duration of t months or more. Note that the function in Figure 6 (a) decreases quickly during the first months which means that most price spells have a low duration.

Figure 6 (a) gives equal weight to each price spell. This implies that its shape is dominated by elementary products which exhibit a high number of spells, i. e. which have short durations. Table 11 gives values for quantiles (25th percentile, the median, and 75th percentile, respectively) of Kaplan-Meier duration estimates for both the "unweighted" and the "weighted" version. The table also shows the number of spells per COICOP group and product type and compares the share of spells to the weights in the CPI baskets. Both classifications indicate clearly that COICOP food items have a much higher share of spells (59%) than indicated by their CPI weight (17%). Non-energy industrial goods and services, on the other hand, contribute a comparably small share of spells but much higher CPI weights.

Dias et al. (2005) show formally how the relatively higher share of spells of product categories with higher frequencies of price changes creates a bias when estimating the duration of price spells. They suggest, as one way to solve this problem, to use only a fixed number of spells per product category.[13] As the authors note themselves, such a sampling scheme does not use all the available information and will hence not be efficient. As an alternative, we apply a weighting scheme where (1) each product category is weighted with the inverse of the total number of price spells for that product category which ensures that each product category has the same weight in the results; (2) in addition, we attach to each product category its CPI weight. This is the basis for our "weighted" Kaplan-Meier estimates of survivor and hazard functions. Adjustment (1) makes a big difference because the enormous weight of food products is reduced whereas adjustment (2) changes the picture not very much.

[13] For example, when calculating the Kaplan-Meier estimate of the aggregate hazard function Dias et al. (2005) work with a *single* spell for each product randomly drawn from the full set of spells (after excluding the left-censored ones). See also Fougère et al. (2005) applying a similar approach.

Panel (b) of Figure 6 shows the survivor function where each spell was reweighted as described. Compared to panel (a) this new survivor function is shifted upwards. Moreover, it has a marked drop at a duration of twelve months which indicates that price changes every year are an important phenomenon. Table 11 indicates that the unweighted median duration over all spells is merely 2 months which is mainly due to the low duration of food item price spells. The weighted median over all products categories is 11 months which is approximately the same result as obtained by the frequency approach in Section 3. According to the survivor function in (b), for almost half of all products (adjusted for different CPI weights), prices are adjusted at a frequency of less than once a year.

The hazard rate based on the Kaplan-Meier estimator is displayed in Figure 7.[14] Again, panel (a) is an unweighted version, and as expected, its overall shape is decreasing with time. But it also displays peaks, for example at durations of 12, 24, and 36 months, respectively. Unconditional, aggregate hazards which are decreasing with analysis time are a typical result of duration studies on micro CPI data (see Fougère et al., 2005, Álvarez et al., 2004 and Dhyne et al., 2005). At first sight, this result is puzzling in the light of price-setting theories, as it could be interpreted that a firm will have a lower probability to change its price the longer it has been kept unchanged.

However, there are several reasons for a decreasing hazard function, none of which is at odds with prevailing theories of price setting behavior. All explanations focus on the heterogeneity of price setters or products. Apparently, a major reason for the decreasing hazard function is the oversampling problem described above, namely that product categories with a high frequency of price changes and thus a higher number of spells wrongly suggest that the probability of a price change is highest after 1, 2, or 3 months (such as in panel (a) of the figure). Panel (b), however, shows that after re-weighting the likelihood of a price change is highest 12 months after the last price change.

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[14] According to the Kaplan-Meier estimator, the hazard rate is estimated as d_j / n_j, i. e. the rate at which spells are completed after duration *t*. Note that if for a duration t_j no price changes are observed the estimator is not defined.

An additional reason for downward sloping hazard functions comes from aggregating firms with different (time-dependent) price setting behavior. As Álvarez et al. (2004) point out, the aggregation of different types of time dependent price setters almost always leads to a decreasing aggregate hazard function. Another related rationale for falling hazards is that the CPI is the result of the aggregation of heterogeneous products: For some products, prices are adjusted infrequently (e. g. services) whereas for others many price changes are observed (e. g. energy). Even if there is no oversampling of products categories, the hazard function may still be decreasing. For example, Dias et al. (2005) and Fougère et al. (2005) only use one spell per product in the estimation of hazard functions, but in both cases the hazard functions are still declining.

Breakdown by product groups

Panels (a) and (b) of Figure 8 show separate survivor functions for the five different product types.[15] Again, the quartiles of the estimated durations are contained in Table 11. The heterogeneity of products is substantial: The solid line (unprocessed food) shows that price spell durations for this product group are rather short. Even shorter spells are indicated by the dotted line (energy): The corresponding survivor function decreases very quickly towards zero which means that almost all price spells have a duration of just a few months at most. The short-dashed line (processed food) implies price spells of intermediate length for that product category. At the other end of the spectrum, price spells of non-energy industrial goods (dashed-dotted line) and especially the services (long-dashed line) are relatively long-lived. Although the survivor functions are shifted upward for each product group, a comparison of panels (a) and (b) indicates that re-weighting within product categories does not change the results as much as for all product categories.

This is also the impression of Figure 9 which displays separate hazard functions for each of the five product groups where the weighted and unweighted versions do not differ much. Hazard rate estimates for different product groups (Figure 9) show some

[15] A picture which is remarkably similar to that in Panel A of Figure 7 can be found in Fabiani et al. (2004) for Italy.

interesting patterns: For example, for services the hazard is highest when the duration is approximately 1 year. The corresponding hazard function also displays noticeable spikes at 24, 36, and 48, months, respectively. Energy items, on the other hand, have a very high hazard when the spell duration is low. Non-energy industrial goods have high probabilities of price changes both at short durations and after one year whereas, for both unprocessed and processed food, the hazard rates are highest at short durations.

Hazard functions for single product categories

Additionally, Figures A6 to A10 in Annex III show Kaplan-Meier estimates of hazard functions for single product categories. For these graphs, of each product group, the 24 product categories with the highest CPI weights were selected and the hazards computed.[16] For each analysis time t_j the hazard rate is marked by a symbol. If the hazard rate is defined at two subsequent failure times, then the symbols were connected by lines. The results confirm the impression from the more aggregate analysis: For many food items, hazards appear to be decreasing with analysis time. A possible reason for this is heterogeneity which could still be present at this rather disaggregrate level (e. g. differences across store types etc.). Within energy, the picture is heterogeneous: For fuels, hazard rates are very high for short durations (and longer durations cannot be observed at all). Motor oil or electricity and heating fees, on the other hand, have lower hazards. Hazards are mostly low for non-energy industrial goods and rather constant over time. The same is true for a number of services where also very high hazards at durations of 12, 24, or 36 months can be observed.

4.3 Logit estimates

In order to control for unobserved characteristics of individual units (i), which is a major reason for downward sloping (unweighted) hazard functions, we estimate a panel conditional logit model with fixed elementary product effects, where an elementary product is the combination of the product category (j) and the outlet code (k), taking

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[16] With the exception of energy items where only 18 product categories (of which one did not exhibit a price change) are in the data.

into account product and store replacements (72,892 elementary products). The cross-section dimension (j*k) is indexed by (i). This allows us to control for the fact that within the same product category firm A can adjust its price more or less frequently than firm B. For this exercise also all left-censored price spells are excluded, as some explanatory variables like the duration of a price spell and the accumulated inflation for a product category since the last price change are not defined when the starting date of the spell is unknown.

The Model

Inspired by Cecchetti (1986) and Aucremanne and Dhyne (2004B), we specify the following fixed effects conditional logit model.[17] The dependent variable is binary indicating the occurrence of a price change at the beginning of next month (or at the end of the current period t, $Y_{it}=1$),

$$Pr(Y_{it}=1 \mid \mathbf{x}_{it}) = F(\boldsymbol{\beta}'\mathbf{x}_{it} + \alpha_i) \qquad (7)$$

with

$$\begin{aligned}\boldsymbol{\beta}'\mathbf{x}_{it} = {} & \beta_0 + \beta_1 * TAU_{it} + \beta_2 * INF\_ACC\_J_{it} + \beta_3 * ATTR_{it} + \beta_4 * LDLNP\_UP_{it} \\ & + \beta_5 * LDLNP\_DW_{it} + \beta_6 * LDLNPDW_{it} + \textstyle\sum_{h=1}^{5} \beta_{6+h} * DURh_{it} + \textstyle\sum_{h=1}^{2} \beta_{11+h} * EUROh_{it} \\ & + \textstyle\sum_{h=1}^{11} \beta_{13+h} * MONTH\_h_{it} + \textstyle\sum_{h=2}^{8} \beta_{23+h} * YEAR\_h_{it} + \beta_{32} * DUM\_00\_03_{it} \\ & + \beta_{33} * LS96\_97_{it}\end{aligned} \qquad (8)$$

and $i = 1,\ldots,N$ is the cross-section dimension (the number of elementary products), $t = 1,\ldots,T_i$ is the time-series dimension, α_i are the fixed effects and F stands for the cumulative logistic distribution function

$$F(z) = \frac{\exp(z)}{1-\exp(z)}. \qquad (9)$$

As explanatory variables we included several state and time dependent variables described below.[18] One explanatory variable, namely the duration of price spells gained

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[17] See Baltagi (2001, Chapter. 11) for a discussion of the properties of panel logit models.

[18] See Annex I for more details on the definitions of all variables included.

a lot of attention in related studies, as the sign of its coefficient reflects the panel data estimate of the direct time effect which was described by the hazard functions in Section 4.2. We argued there, that the downward sloping hazard functions are a consequence of aggregating over (very) heterogeneous products. After controlling for unobserved heterogeneity with a fixed effects model, we therefore expect a positive sign for the coefficient of the duration of price spells (TAU).

As another state dependent explanatory variable we included the absolute value of the accumulated sectoral rate of inflation for the product category (j) to which the elementary product (i) belongs (INF_ACC_J). For each elementary product at time (t) the sectoral inflation rate is accumulated over the period since its last price change. This variable is a proxy for the relative price position of outlet (k) selling product (j) to the average of all other outlets selling a product of the same category. Therefore, if the accumulated inflation increases, on average the competitors already increased their prices, whereas outlet (k) held its price unchanged, which puts it in a better position to increase its price without losing customers. Consequently, for this variable our expectation is also a positive coefficient.

We consider the impact of common commercial practices (as psychological pricing, sales and promotions) on the price setting behavior by including a dummy variable reflecting the fact that a price was set in attractive terms (ATTR) as well as by measuring the impact of the magnitude (LDLNP) and the direction of the last price change (LDLNPDW=1 if the last price change was a price decrease). The variable LDLNP_UP (defined as LDLNP*(1-LDLNPDW)) captures the magnitude of price increases, whereas LDLNP_DW (defined as LDLNP*LDLNPDW) contains the size of price reductions. For attractive prices we expect a dampening effect on the probability of a price change, i. e. a negative sign of the coefficient.

The hazard functions (Figures 7 and 9) and the distribution of the duration of price spells shown in Figure 11 highlights the fact that there are local modes at specific durations, noteworthy 1, 6, 12, 24 and 36 months. We interpret this fact as some kind of truncated Calvo or Taylor pricing behavior and try to capture this with a set of dummy variables (DUR_*h*, *h* = 1, 6, 12, 24, 36).

Two variables capture the effects of the Euro cash changeover: one dummy for the direct effect in January 2002 (EURO1), and a second defined over the period 6 months before and 5 months after the month of the changeover in January 2002 (EURO2).[19]

In addition, several dummy variables indicating time dependent aspects as the seasonal pattern (MONTH_h, $h = 1,...,11$) and yearly dummies to control for structural and/or cyclical economic effects not captured by other variables (YEAR_h, $h = 2,...,8$) are included.

As can be seen from Figure 5 the average frequency (probability) of price changes shows a downward trend, which is due to the exclusion of left-censored spells at the beginning of the sample period. We account for this with a trend variable (LS96_97, descending from 24 in January 1996 to 1 in December 1997, and 0 afterwards). To control for effects due to the revision of the CPI basket in January 2000 an additional dummy variable is included (DUM_00_03 = 1 for the period after December 1999).[20]

Results

The unweighted estimation results are reported in Table 12. We present the estimated coefficients (β), the standard errors (S.E.), the significance levels of the estimated coefficients (p-value, β), and the marginal effects (slope) and its significance levels (p-

[19] In another specification we divided the EURO2 period in 2 sub-periods: a dummy for the period of the compulsory dual product pricing (October 2001 to March 2002, excluding January 2002, EURO21) and another dummy for the period of three months before and three months after the period of double product pricing (EURO22). As during the EURO21 period price changes due to the Euro introduction were more visible, highly debated in public and consumers or consumer associations could appeal against unjustified price increases at the 'Euro Price Commission' (appointed by the federal government), we expected that firms anticipated this debate and either changed prices before or postponed them to the period after dual pricing (EURO22). However, the coefficients for EURO21 and EURO22 were statistically not distinguishable. Consequently, we added EURO21 and EURO22 to one dummy variable (EURO2).

[20] In addition to the variables discussed above, we also experimented with several other state dependent variables as the industrial production index and the aggregate consumer price index (both variables were included either as month-on-month or year-on-year rates of change) and a tax variable. But none of these variables showed any significant effect.

value, slope). The marginal effects are defined as the first derivatives of the probability function with respect to the explanatory variables, evaluated at the mean of the variables. The reference probability is a price change in January 1996.

As can be seen from Table 12, the probability of a price change slightly increases the longer a price quote has been unchanged.[21] An increase in the duration of a price spell (TAU) by one month increases the probability of a price change by roughly 0.6 percentage points. We interpret this result as evidence that, after controlling for unobserved heterogeneity at the elementary product level, (slightly) increasing hazard rates are obtained through a direct duration impact. In addition, there is an indirect duration effect working through the role of the accumulated inflation variable as the sign of the coefficient for the accumulated inflation (INF_ACC) is positive as one would expect, i. e. the probability of a price change increases as inflation in the same product category rises. An increase in the accumulated monthly inflation rate by 1 percentage point increases the probability for a price change by 18 percentage points. This effect seems high at the first sight, but one has to take into consideration, that over the estimation period the average yearly inflation was about 2 percent (see Figure 2). Thus, the average month-on-month rate of inflation was 0.16 percent, indicating that the monthly increase of the accumulated rate of inflation on average (over all products and months) was also 0.16 percentage points. Therefore, an average increase in the accumulated rate of inflation leads to an increase in the probability of a price change by 2.9 percentage points.

The occurrence of attractive prices (ATTR) reduces the probability to change prices, as expected, and the opposite is true for the dummy indicating that the last price change was a price reduction (LDLNPDW). Both results are in line with commercial practices, especially with promotions and seasonal sales. The effect of the size of the last price decrease (LDLNP_DW) on the probability of a price change is much stronger than the effect of the size of a price increase (LDLNP_UP). This finding is also consistent with the practice of temporal promotions as large price reductions due to promotions are usually quickly reversed by (large) price increases.

[21] In the following subsection we discuss the estimates of the marginal effects. We therefore refer to the columns 'slope' and 'p-value slope' in Table 12.

Concerning the time-dependent and Taylor-type phenomena mentioned in Section 4.2, our logit estimates reinforce this evidence: especially for the duration of 12 months and to a lesser extent for durations of 1 month, 2 and 3 years we find a higher probability of a price change.

For the Euro cash changeover the time dummies are indicating a higher probability of a price change in January 2002 (EURO1), and to a lesser extent in the 6 months before and 5 months after the month of the Euro introduction as a physical mean of payment (EURO2).

According to our results, there is a strong seasonal pattern in the price setting process. The probability to change prices in January (MONTH_12, the reference month) is larger, as the coefficients for all other seasonal dummies are negative and highly significant.[22] Aucremanne and Dyhne (2004A,B), Baudry et al. (2004), Jonker et al. (2004), Dias et al. (2004) report similar results for other euro area countries. Furthermore, the seasonal dummies are jointly highly significant, indicating some time-dependent elements in the price setting process.

The establishment of a new CPI basket in January 2000 and the thereby introduced new definitions and reporting practices, measured by the variable D_00_03 did have a significant impact on the probability of a price change: For the period starting with January 2000 this probability was 4 percentage points higher.

From the results of the logit regression (Tables 12) we conclude that, although some time dependent aspects can be observed in the data, our evidence does not support pure time dependent representations of the price setting process (as Calvo, truncated Calvo or Taylor contracts) at the micro CPI level, as some of the state dependent variables have a significant effect on the probability of a price change.

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[22] In the interpretation of the seasonal dummies, one has to consider the definition of the price change variable Y: As Y defines a price change in the next period, the seasonal December-dummy (MONTH_12) accounts for price changes in January (of the next year).

5. Summary

In this paper we analyze the patterns and determinants of price rigidity present in the individual price quotes collected to compute the Austrian CPI. We calculated direct and implied estimates for the average frequency of price changes and the duration of price spells for 639 product categories.

We found that consumer prices are quite sticky in Austria. The weighted average (implied) duration of price spells for all products is 10 (14) months. The sectoral heterogeneity is quite pronounced: prices for services, health care and education change rarely, typically approximately once per year (or even less according to the implied duration measure). For the product types food, energy (transport) and communication prices are adjusted on average every 6 to 8 months. Promotions and sales have a considerable impact on the frequency of price adjustments for food, clothing and footwear, where temporal promotions and end-of-season sales are a common practice.

With respect to the synchronization of price changes a similar sectoral pattern occurs as for the durations of price spells: The prices of products with a longer duration are also adjusted in a more synchronous way. This reflects the fact that prices of some of these products are either directly administered or strongly influenced by public authorities (including social security agencies).

Price increases occur slightly more often than price decreases, except for communication items, where due to the technical progress and high competition in the electronic equipment sector price decreases appear much more frequent than increases. Price increases and decreases are quite sizeable when they occur: on average, prices increase by 11 percent whereas prices are reduced on average by 15 percent. Especially for clothing and footwear (due to seasonal sales) and again for communication and electronic items (personal computers) price decreases are very pronounced (34 and 26 percent, respectively).

Like in similar studies conducted within the European inflation persistence network, we find that the aggregate hazard function for all price spells is decreasing with time (i.e. the duration of a price spell) which is somewhat at odds with most price-setting theories. However, this is to a large extent a consequence of aggregating over product types with different spell durations. A re-weighted version of the hazard function which ensures that each product category basically has the same weight (and adjusted for CPI weights) is not monotonously decreasing, but has its most marked spike at a duration of 1 year which indicates that for a substantial proportion of all goods infrequent price adjustment occurs. Using Kaplan-Meier estimates of survivor and hazard functions, we show that there is substantial heterogeneity across goods and product types: Energy and unprocessed food show high hazards during the first months. For services, on the other hand, the hazard is highest after one year.

In contrast to other European studies, we find a positive and significant effect of the duration of a price spell on the probability of a price change if we account for unobserved heterogeneity in a panel logit model with fixed elementary product effects. We observe also a positive link between the probability of a price change and the accumulated inflation at the product level. Additionally, we find a seasonal (January) effect and a negative impact on the probability to change a price if it is currently set as an attractive price. The euro cash changeover seems to have increased the probability to change prices.

Although some time dependent aspects have a significant impact on the probability of a price change, our evidence does not support pure time dependent representations of the price setting process at the outlet level, as some of the state dependent variables always show a significant influence on the probability to observe a price change.

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Annex I: List of variables

Identifier of an observation:

PCODE	Product identifier (*j*)
LCODE	Outlet code (*k*)
MONTH	Time identifier (*t*)
ID	Elementary product code (i = j*k), accounting for product and store replacements.
SPELLID	Identifier of a price spell for an elementary product (*s*).
P_{it}	Price quote for product (*j*) in outlet (*k*) at time (*t*).

Dependent variables:

Y_{it}	Binary variable for a price change: if a price change occurs in t+1, i.e. $[\ln(P_{it+1}) - \ln(P_{it}) \neq 0]$ then $Y_{it} = 1$, otherwise = 0

Explanatory variables:

TAU_{it} — Duration of a price spell: number of months since the price P_{it} is unchanged.

$INF\_ACC\_J_{it}$ — Absolute value of the accumulated inflation for product category (*j*) since the price for product (*j*) was changed the last time in outlet (*k*).

$ATTR_{it}$ — Dummy to indicate an attractive or psychological price. Attractive prices are defined for ranges of the mean absolute price in order to take account of different attractive prices at different price levels. For products with a mean of valid prices

- from 0 to 100 ATS all prices ending at 0.00, 0.50 and 0.90 ATS
- from 100 to 1,000 ATS prices ending at 00.00, 5.00 and 9.00 ATS have been defined as attractive (where prices have been rounded up to the *x.00 digit)
- exceeding 1,000 ATS prices ending at 000.00, 50.00 and 90.00 (where prices have been rounded up to the *x0.00 digit)

have been defined as attractive. If this definition is met, ATTR = 1, otherwise ATTR = 0. An equivalent rule has been defined to identify attractive prices in Euro after the cash changeover.

$LDLNP_{it}$ — Absolute value (magnitude) of the last price change of product (*j*) in outlet (*k*).

$LDLNPDW_{it}$ — Dummy variable indicating that the last price change was a price decrease

$LDLNP\_UP_{it}$ — Magnitude of the last price increase of product (*j*) in outlet (*k*): $LDLNP_{it}*(1-LDLNPDW_{it})$.

$LDLNP\_DW_{it}$ — Magnitude of the last price decrease of product (*j*) in outlet (*k*): $LDLNP_{it}*LDLNPDW_{it}$.

$DUR1_{it}$ — $DUR\_1_{it} = 1$ if $TAU_{it} = 1$, otherwise $DUR\_1_{it} = 0$.

$DUR6_{it}$ — $DUR\_6_{it} = 1$ if $TAU_{it} = 6$, otherwise $DUR\_6_{it} = 0$.

$DUR12_{it}$ — $DUR\_12_{it} = 1$ if $TAU_{it} = 12$, otherwise $DUR\_12_{it} = 0$.

$DUR24_{it}$ — $DUR\_24_{it} = 1$ if $TAU_{it} = 24$, otherwise $DUR\_24_{it} = 0$.

$DUR36_{it}$	DUR_36_{it} = 1 if TAU_{it} = 36, otherwise DUR_36_{it} = 0.
$EURO1_{it}$	Dummy variable for the Euro cash changeover: if t = December 2001 EURO1 = 1, otherwise EURO1 = 0 (Y indicates a price change in the next month. Thus, EURO1 indicates that a price change occurred in January 2002).
$EURO21_{it}$	Dummy variable for the period of compulsory dual product pricing in Austria: for t = September 2001 to February 2002 (except Dec. 2001) EURO21 =1, otherwise EURO21 = 0 (the defined period therefore represents the time interval October 2001 till March 2002, the period for which dual product pricing was compulsory).
$EURO22_{it}$	Dummy variable for a period of 3 months before and 3 months after dual product pricing in Austria: for t = June to August 2001 and March to May 2002 EURO22 =1, otherwise EURO22 = 0 (this period represents the time intervals July to September 2001 and April to June 2002).
$EURO2_{it}$	Dummy variable for the period of 6 months before and 5 months after the month of the euro introduction: for t = June 2001 to May 2002 (except Dec. 2001) EURO2 =1, otherwise EURO2 = 0 (i. e. $EURO2_{it} = EURO21_{it} + EURO22_{it}$).
$MONTH\_h_{it}$	h=1 to 12, seasonal dummy variables; if h is a specific month MONTH_h =1, otherwise MONTH_h = 0.
$YEAR\_h_{it}$	h= 1 to 8, yearly dummies for 1996 to 2003. if h is a specific year YEAR_h =1, otherwise YEAR_h =0.
$DUM\_00\_03_{it}$	Dummy variable to control for the revision of the CPI basket in January 2000: $DUM\_00\_03_{it}$ = 1 for the period after December 1999; before that period (i. e. the period of the 1996 CPI basket) DUM_00_03 = 0.
$LS96\_97_{it}$	Trend variable, descending from 24 in January 1996 to 1 in December 1997, and LS96_97 = 0 afterwards.

Annex II: Tables and Figures

Table 1: Classification of elementary price quotes by COICOP groups and type of product

	Price records		Weights of products included in our sample		Weights in CPI basket
	# of observations	in %	actual weights	weights rescaled	
By COICOP					
COICOP 01: Food and non-alcoholic beverages	1,343,690	37.4%	13.4%	15.0%	13.8%
COICOP 02: Alcoholic beverages and tobacco	93,101	2.6%	1.4%	1.5%	3.3%
COICOP 03: Clothing and footwear	547,013	15.2%	7.3%	8.1%	7.3%
COICOP 04: Housing, water, gas and electricity	92,769	2.6%	19.3%	21.4%	19.3%
COICOP 05: Furnishing & maintenance of housing	348,748	9.7%	9.1%	10.1%	9.1%
COICOP 06: Health care expenses	31,998	0.9%	2.9%	3.2%	2.9%
COICOP 07: Transport	255,449	7.1%	8.0%	8.8%	13.6%
COICOP 08: Communications	4,504	0.1%	2.8%	3.1%	2.6%
COICOP 09: Leisure and culture	304,548	8.5%	10.6%	11.8%	11.6%
COICOP 10: Education	12,307	0.3%	0.7%	0.8%	0.7%
COICOP 11: Hotels, cafés and restaurants	294,974	8.2%	7.2%	8.0%	6.9%
COICOP 12: Miscellaneous goods and services	265,202	7.4%	7.2%	8.0%	8.9%
By Product type					
Unprocessed food	627,972	17.5%	5.6%	6.3%	5.7%
Processed food	808,819	22.5%	9.2%	10.2%	11.5%
Energy	73,177	2.0%	7.7%	8.6%	7.7%
Non energy industrial goods	1,314,899	36.6%	29.6%	32.9%	35.3%
Services	769,436	21.4%	37.8%	42.1%	39.8%
Total	3,594,351	100.0%	89.9%	100.0%	100.0%

Table 2: Information available for each elementary price quote

Date of the price quote	Month and year of the quote
Outlet code	Each outlet can be identified by a specific code (anonymized)
Product category code	639 product categories, incl. a sub-code for the variety (no direct information on the brand of the product)
Packaging of the item	Weight of or number of items in the package
Price of the item	in Austrian Schilling (till Dec. 2001) and Euro (from Jan. 2002)
Sales prices	Code indicating a sales price (Codes A, N)
Temporal unavailability	Code indicating seasonal and other temporal unavailability of the item (Codes F, G, V, X,)
Product replacements	Code indicating that a product variety has been replaced by another (no information on the nature of the replacement, i.e. whether it is forced or voluntary; Code S)
Quality adjustment	Code for quality adjustment (for hedonic prices; Code Q)
Store replacements	Code indicating the replacement of a specific store by another (Code W)

Table 3: Repartition of the records: occurrence of informational codes by year

Year	1996	1997	1998	1999	2000	2001	2002	2003
Total observations	405,427	412,771	416,465	424,690	480,452	480,379	486,102	488,065
Code								
A	11,761	14,473	14,967	16,081	19,583	20,318	23,299	26,167
N	2,562	2,203	4,007	4,804	4,858	4,465	4,793	5,508
F	4	28	1,667	4,084	4,489	2,537	2,615	2,626
G	3,993	5,316	5,493	5,629	12,823	20,428	22,218	22,276
V	0	0	0	0	6	3,140	4,087	3,260
X	5,422	11,065	13,906	26,410	16,619	11,248	13,511	14,464
S	0	0	0	0	6,005	6,067	7,677	7,413
QZ	0	0	0	0	732	1,224	1,169	1,243
Q0	0	0	0	0	104	296	476	804
Q1	0	0	0	0	54	172	253	189
Q2	0	0	0	0	97	504	767	681
Q3	0	0	0	0	25	159	286	189
Q4	12,324	4,860	4,219	5,540	3,652	2,780	1,675	572
W0	0	0	0	0	30	485	776	106
W1	0	0	0	0	3	1	5	0
W2	0	0	0	0	6	4	68	5
W3	0	0	0	0	0	11	0	1
W4	0	0	0	0	2,145	1,154	916	8

Codes:		
	A	sales price
	N	normal price - end of sales price
	F, G	seasonal unavailability
	V	temporal unavailability
	X	longer unavailability
	S	replacement of product
	QZ,Q0-Q4	quality adjustment - equivalent to product replacement
	W0-W4	store replacement

Table 4: Frequency of price changes by COICOP classification and product type
Weighted average of the entire basket

	Frequency of price changes	Average duration of price spells	Median duration of price spells	Frequency of price increases	Frequency of price decreases	Average price increase	Average price decrease
By COICOP							
COICOP 01: Food and non-alcoholic beverages	17.3%	7.9	7.9	9.1%	7.9%	16.9%	18.7%
COICOP 02: Alcoholic beverages and tobacco	14.6%	6.5	5.9	7.4%	7.0%	14.6%	14.9%
COICOP 03: Clothing and footwear	12.0%	9.4	7.9	6.4%	5.0%	23.1%	33.7%
COICOP 04: Housing, water, gas and electricity	11.2%	14.7	11.3	6.9%	4.0%	6.6%	8.7%
COICOP 05: Furnishing & maintenance of housing	6.9%	17.8	16.0	4.1%	2.5%	9.3%	13.6%
COICOP 06: Health care expenses	5.6%	18.8	19.7	4.4%	1.1%	4.0%	6.7%
COICOP 07: Transport	36.5%	11.2	9.6	18.8%	17.7%	8.3%	8.8%
COICOP 08: Communications	8.9%	16.0	10.5	1.8%	6.9%	15.5%	26.0%
COICOP 09: Leisure and culture	24.2%	15.8	11.2	12.3%	11.2%	11.1%	12.3%
COICOP 10: Education	4.5%	23.2	20.2	4.1%	0.4%	4.9%	0.5%
COICOP 11: Hotels, cafés and restaurants	8.3%	19.3	21.3	5.4%	2.6%	7.3%	8.4%
COICOP 12: Miscellaneous goods and services	7.1%	18.7	15.2	4.9%	2.0%	7.6%	11.4%
By Product type							
Unprocessed food	24.0%	6.5	7.5	12.6%	11.1%	19.6%	22.0%
Processed food	12.8%	8.5	7.9	6.8%	5.8%	14.8%	16.1%
Energy	40.1%	8.3	4.8	20.7%	19.3%	5.1%	4.4%
Non energy industrial goods	10.2%	13.7	11.5	5.4%	4.3%	13.2%	18.6%
Services	12.6%	19.4	18.5	7.4%	5.0%	8.1%	10.9%
Total	15.1%	14.1	11.1	8.2%	6.6%	11.4%	14.7%

Frequency: average proportion of prices changes per month, in percent - Duration: in months
Sample period: January 1996 - December 2003

Table 5: Frequency of price changes by COICOP classification and product type
Weighted average of the entire basket - without sales and promotions

	Frequency of price changes	Average duration of price spells	Median duration of price spells	Frequency of price increases	Frequency of price decreases	Average price increase	Average price decrease
By COICOP							
COICOP 01: Food and non-alcoholic beverages	11.3%	13.0	13.9	6.1%	4.8%	12.3%	13.3%
COICOP 02: Alcoholic beverages and tobacco	8.0%	12.1	11.9	4.1%	3.6%	10.3%	10.0%
COICOP 03: Clothing and footwear	8.5%	12.5	11.1	4.7%	2.9%	16.3%	19.2%
COICOP 04: Housing, water, gas and electricity	10.5%	15.2	11.3	6.6%	3.7%	6.5%	8.5%
COICOP 05: Furnishing & maintenance of housing	5.9%	19.5	17.1	3.5%	2.0%	7.8%	11.1%
COICOP 06: Health care expenses	5.5%	19.1	19.7	4.4%	1.0%	4.0%	6.9%
COICOP 07: Transport	34.4%	11.5	10.4	17.7%	16.6%	8.2%	8.5%
COICOP 08: Communications	8.1%	16.4	10.5	1.5%	6.5%	21.4%	26.2%
COICOP 09: Leisure and culture	21.3%	17.2	11.7	10.8%	9.7%	10.7%	11.6%
COICOP 10: Education	4.5%	23.2	20.4	4.0%	0.4%	4.9%	0.5%
COICOP 11: Hotels, cafés and restaurants	7.7%	19.7	21.9	5.1%	2.3%	7.3%	7.8%
COICOP 12: Miscellaneous goods and services	6.4%	21.1	15.5	4.6%	1.7%	6.5%	8.7%
By Product type							
Unprocessed food	17.4%	10.6	12.5	9.2%	7.7%	15.0%	16.4%
Processed food	7.1%	14.3	14.0	4.0%	2.8%	10.4%	10.8%
Energy	37.9%	8.6	5.5	19.6%	18.2%	5.1%	4.5%
Non energy industrial goods	8.4%	15.7	13.7	4.5%	3.2%	10.6%	13.5%
Services	11.6%	20.2	19.3	6.9%	4.5%	8.2%	10.5%
Total	12.8%	16.1	14.0	7.0%	5.3%	9.6%	11.5%

Frequency: average proportion of prices changes per month, in percent - Duration: in months
Sample period: January 1996 - December 2003

Table 6: Distribution of price increases and decreases by COICOP classification and product type
Median, 1. Quartile and 3. Quartile

	1. Quartile Increase	Median Increase	3. Quartile Increase	1. Quartile Decrease	Median Decrease	3. Quartile Decrease
By COICOP						
COICOP 01: Food and non-alcoholic beverages	12.4%	16.9%	20.4%	14.2%	18.4%	22.4%
COICOP 02: Alcoholic beverages and tobacco	12.4%	14.0%	17.1%	13.3%	14.2%	16.6%
COICOP 03: Clothing and footwear	19.5%	22.6%	26.7%	28.1%	34.8%	40.0%
COICOP 04: Housing, water, gas and electricity	4.8%	6.1%	7.9%	3.9%	6.7%	11.4%
COICOP 05: Furnishing & maintenance of housing	7.2%	8.5%	10.3%	9.7%	13.0%	17.6%
COICOP 06: Health care expenses	0.9%	2.5%	7.4%	0.8%	5.4%	12.1%
COICOP 07: Transport	3.1%	6.8%	9.1%	2.7%	8.0%	16.2%
COICOP 08: Communications	0.4%	21.1%	24.9%	3.2%	32.8%	34.0%
COICOP 09: Leisure and culture	7.2%	11.0%	12.5%	7.0%	10.7%	15.5%
COICOP 10: Education	4.3%	4.9%	5.2%	0.3%	0.3%	0.8%
COICOP 11: Hotels, cafés and restaurants	5.4%	6.6%	7.6%	5.6%	6.7%	9.4%
COICOP 12: Miscellaneous goods and services	3.6%	6.4%	10.2%	5.6%	10.5%	14.7%
By Product type						
Unprocessed food	16.1%	17.6%	22.5%	17.5%	22.0%	24.8%
Processed food	11.2%	14.3%	17.9%	13.3%	15.9%	19.0%
Energy	3.1%	3.8%	6.4%	2.7%	3.4%	5.1%
Non energy industrial goods	7.4%	10.2%	19.0%	10.3%	14.8%	24.8%
Services	4.7%	6.6%	10.2%	5.1%	7.8%	12.8%
Total	6.2%	8.8%	15.6%	6.5%	12.6%	19.5%

Sample period: January 1996 - December 2003

Table 7: Synchronisation ratios by COICOP classification and product type
Weighted average of the entire CPI basket

	Synchronisation ratio of price changes	Synchronisation ratio of price increases	Synchronisation ratio of price decreases
By COICOP			
COICOP 01: Food and non-alcoholic beverages	21.1%	21.2%	19.9%
COICOP 02: Alcoholic beverages and tobacco	20.3%	16.6%	23.5%
COICOP 03: Clothing and footwear	26.0%	18.9%	22.8%
COICOP 04: Housing, water, gas and electricity	53.6%	58.3%	39.9%
COICOP 05: Furnishing & maintenance of housing	27.7%	25.5%	21.2%
COICOP 06: Health care expenses	86.7%	84.0%	54.6%
COICOP 07: Transport	51.4%	54.9%	56.5%
COICOP 08: Communications	93.8%	85.3%	91.7%
COICOP 09: Leisure and culture	51.1%	51.3%	42.9%
COICOP 10: Education	80.8%	81.7%	44.7%
COICOP 11: Hotels, cafés and restaurants	30.9%	29.0%	25.4%
COICOP 12: Miscellaneous goods and services	50.1%	47.6%	36.5%
By Product type			
Unprocessed food	20.5%	22.7%	22.4%
Processed food	21.3%	19.6%	18.9%
Energy	51.6%	62.7%	49.1%
Non energy industrial goods	34.2%	30.8%	28.3%
Services	58.6%	55.8%	45.4%
Total	41.9%	40.4%	34.2%

Sample period: January 1996 - December 2003

Table 8: Synchronisation ratios by COICOP classification and product type
Weighted average of the entire CPI basket - without sales and promotions

	Synchronisation ratio of price changes	Synchronisation ratio of price increases	Synchronisation ratio of price decreases
By COICOP			
COICOP 01: Food and non-alcoholic beverages	21.8%	22.6%	19.5%
COICOP 02: Alcoholic beverages and tobacco	22.2%	18.1%	27.3%
COICOP 03: Clothing and footwear	26.5%	17.4%	15.9%
COICOP 04: Housing, water, gas and electricity	53.7%	58.6%	39.1%
COICOP 05: Furnishing & maintenance of housing	28.3%	25.8%	21.4%
COICOP 06: Health care expenses	86.4%	83.6%	55.2%
COICOP 07: Transport	52.0%	54.5%	56.8%
COICOP 08: Communications	93.9%	79.8%	91.8%
COICOP 09: Leisure and culture	49.3%	49.9%	41.4%
COICOP 10: Education	80.7%	81.6%	44.7%
COICOP 11: Hotels, cafés and restaurants	31.0%	28.9%	25.2%
COICOP 12: Miscellaneous goods and services	50.3%	47.8%	36.2%
By Product type			
Unprocessed food	20.9%	23.8%	21.0%
Processed food	22.4%	21.2%	19.8%
Energy	51.9%	62.7%	47.9%
Non energy industrial goods	34.6%	30.4%	26.6%
Services	58.1%	54.6%	44.9%
Total	42.0%	40.0%	33.2%

Sample period: January 1996 - December 2003

Table 9: Number of spells and duration by type of censoring (taking product replacements into account)

Left censored	**Right censored**	**Number of spells**		**Duration of spells**	
		Observations	in %	Weighted mean	Weighted median
0	**0**	326,002	62.7%	8.5	1
0	**1**	40,100	7.7%	11.0	7
1	**0**	131,661	25.3%	10.7	4
1	**1**	22,278	4.3%	15.7	6
Total		520,041	100.0%	10.0	2

Sample period: January 1996 - December 2003; duration: in months

Table 10: Statistics on price spells – duration and implied frequency of price changes

	Price spells				Implied frequency of price changes	
	Number of spells	in %	duration of spells - weighted mean	duration of spells - weighted median	F_j^{imp}	$F_j^{imp,rep}$
By COICOP						
COICOP 01: Food and non-alcoholic beverages	270,072	51.9%	7.2	2	17.3%	14.7%
COICOP 02: Alcoholic beverages and tobacco	14,863	2.9%	6.3	3	14.6%	12.4%
COICOP 03: Clothing and footwear	64,067	12.3%	7.6	4	13.3%	4.4%
COICOP 04: Housing, water, gas and electricity	12,057	2.3%	11.1	2	11.1%	9.9%
COICOP 05: Furnishing & maintenance of housing	32,044	6.2%	12.0	6	7.2%	4.3%
COICOP 06: Health care expenses	2,467	0.5%	12.1	12	5.6%	4.8%
COICOP 07: Transport	41,841	8.0%	7.7	1	36.1%	35.2%
COICOP 08: Communications	876	0.2%	9.0	3	8.2%	5.5%
COICOP 09: Leisure and culture	33,431	6.4%	9.4	1	24.5%	18.6%
COICOP 10: Education	668	0.1%	9.9	10	8.3%	2.9%
COICOP 11: Hotels, cafés and restaurants	24,591	4.7%	14.3	8	8.8%	6.3%
COICOP 12: Miscellaneous goods and services	23,063	4.4%	13.0	10	7.1%	5.9%
By Product type						
Unprocessed food	175,009	33.7%	6.2	1	23.7%	20.4%
Processed food	109,926	21.1%	7.6	3	12.9%	10.9%
Energy	30,743	5.9%	6.7	1	39.7%	38.9%
Non energy industrial goods	141,727	27.3%	9.9	4	10.6%	5.6%
Services	62,635	12.0%	12.6	3	12.8%	10.8%
Total	520,041	100.0%	10.0	2	15.3%	12.3%

Duration: in months; implied frequency, in percent
Sample period: January 1996 - December 2003

Table 11: Spell durations – Kaplan-Meier estimates (weighted and unweigthed)

	no. of spells (completed or right-censored)	share of spells	share of product categories	Average CPI weight	unweighted			weighted		
					25p	median	75p	25p	median	75p
By COICOP										
COICOP 01: Food and non-alcoholic beverages	214,650	58.6%	20.7%	16.9%	1	2	5	1	3	10
COICOP 02: Alcoholic beverages and tobacco	12,000	3.3%	1.7%	1.7%	1	3	9	1	3	9
COICOP 03: Clothing and footwear	26,049	7.1%	8.9%	9.2%	2	7	24	4	10	29
COICOP 04: Housing, water, gas and electricity	10,005	2.7%	6.4%	12.6%	1	2	7	4	11	20
COICOP 05: Furnishing & maintenance of housing	17,019	4.6%	11.7%	11.4%	3	9	22	6	12	32
COICOP 06: Health care expenses	1,478	0.4%	4.1%	3.3%	9	12	24	12	12	26
COICOP 07: Transport	34,283	9.4%	9.7%	9.9%	1	1	10	1	5	13
COICOP 08: Communications	462	0.1%	2.5%	3.5%	2	4	11	10	11	n. def.
COICOP 09: Leisure and culture	18,234	5.0%	15.2%	13.0%	1	5	18	2	10	28
COICOP 10: Education	299	0.1%	1.7%	0.7%	12	12	24	12	12	24
COICOP 11: Hotels, cafés and restaurants	16,819	4.6%	6.1%	8.7%	2	8	22	7	15	29
COICOP 12: Miscellaneous goods and services	15,043	4.1%	11.3%	9.1%	3	11	19	11	12	24
By Product type										
Unprocessed food	140,953	38.5%	9.4%	7.1%	1	2	3	1	2	7
Processed food	85,697	23.4%	13.0%	11.6%	1	3	10	1	4	12
Energy	29,348	8.0%	2.8%	9.4%	1	1	2	1	2	9
Non energy industrial goods	68,768	18.8%	41.5%	37.0%	2	8	20	4	12	25
Services	41,575	11.4%	33.3%	34.9%	3	11	22	8	12	28
Total	366,102	100.0%	100.0%	100.0%	1	2	10	3	11	22

Note: Left-censored spells and spells with gaps have been dropped.

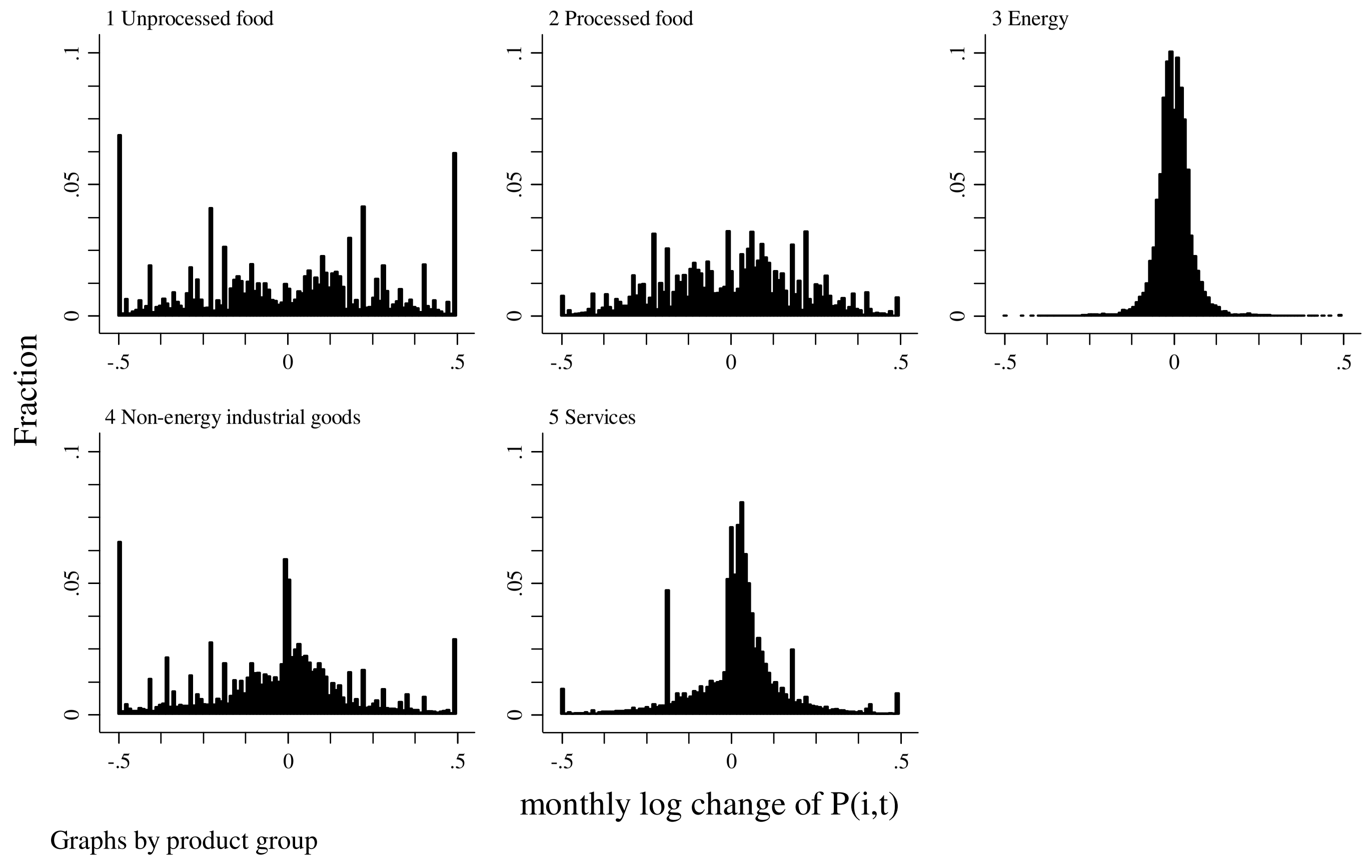
Table 12: Probability of a price change - conditional fixed effects logit model

	β	S .E.	p-value (β)	Slope	p-value (slope)
TAU	0.02	0.001	0.00	0.006	0.00
INF_ACC_J	0.73	0.049	0.00	0.179	0.00
ATTR	-0.20	0.010	0.00	-0.049	0.00
LDLNP_UP	0.09	0.027	0.00	0.022	0.00
LDLNP_DW	1.50	0.030	0.00	0.367	0.00
LDLNPDW	0.25	0.009	0.00	0.061	0.00
DUR1	0.45	0.006	0.00	0.107	0.00
DUR6	0.08	0.012	0.00	0.019	0.00
DUR12	0.92	0.014	0.00	0.199	0.00
DUR24	0.54	0.035	0.00	0.125	0.00
DUR36	0.37	0.066	0.00	0.088	0.00
EURO1	0.81	0.022	0.00	0.179	0.00
EURO2	0.17	0.009	0.00	0.040	0.00
MONTH_01	-0.29	0.014	0.00	-0.071	0.00
MONTH_02	-0.52	0.014	0.00	-0.128	0.00
MONTH_03	-0.39	0.014	0.00	-0.097	0.00
MONTH_04	-0.53	0.014	0.00	-0.131	0.00
MONTH_05	-0.53	0.014	0.00	-0.132	0.00
MONTH_06	-0.52	0.014	0.00	-0.128	0.00
MONTH_07	-0.66	0.014	0.00	-0.163	0.00
MONTH_08	-0.40	0.013	0.00	-0.099	0.00
MONTH_09	-0.48	0.013	0.00	-0.118	0.00
MONTH_10	-0.47	0.013	0.00	-0.116	0.00
MONTH_11	-0.68	0.014	0.00	-0.168	0.00
YEAR_2	0.22	0.025	0.00	0.053	0.00
YEAR_3	0.36	0.037	0.00	0.086	0.00
YEAR_4	0.25	0.037	0.00	0.060	0.00
YEAR_5	0.09	0.048	0.06	0.022	0.05
YEAR_6	0.19	0.049	0.00	0.047	0.00
YEAR_7	0.13	0.049	0.01	0.031	0.01
YEAR_8	0.06	0.050	0.26	0.014	0.26
D_00_03	0.17	0.026	0.00	0.042	0.00
LS96_97	0.04	0.002	0.00	0.009	0.00

No. of observations: 1,579,553 LR (β=0, p-value)= 0.000, log likelihood = -470,733.01
No. of groups: 44,192 elementary products, LR (pooling, p-value) = 0.000
Slope: dy/dx at the mean of the explanatory variable
Reference: January 1996 (MONTH_12, YEAR_1)
Dependent variable: Y = 1 if a price change occurs in the next month

Figure 1

Price change distribution within product groups



Note: Price changes with absolute values of more than 0.5 were replaced by -0.5 and +0.5, respectively. Bin width 0.01. 374,143 total observations (zero values are excluded).

Figure 2: Frequency of price changes over time, weighted average (in p.c.), and aggregate inflation (right axis)

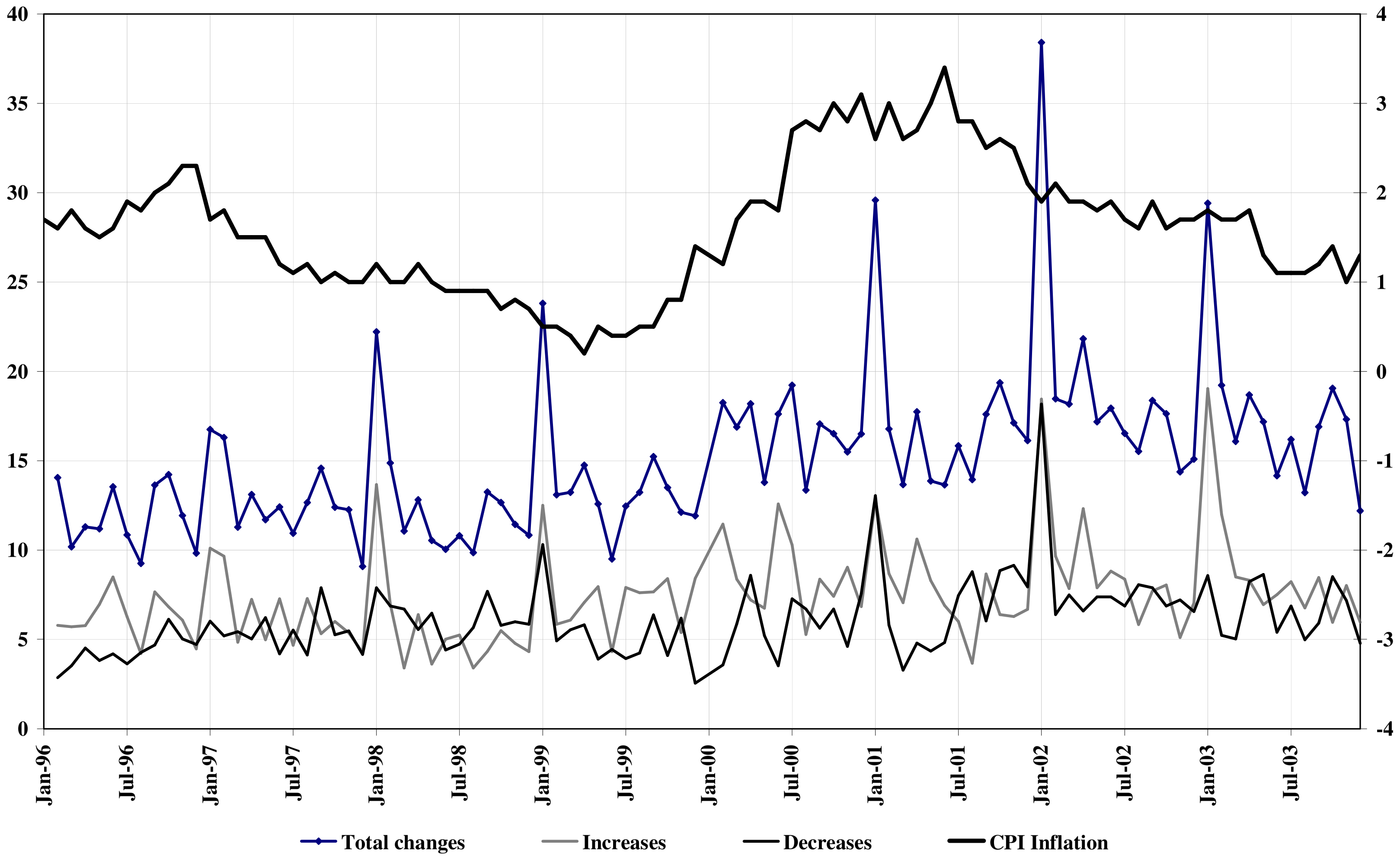


Figure 3: Size of price changes over time, weighted average (in percent)

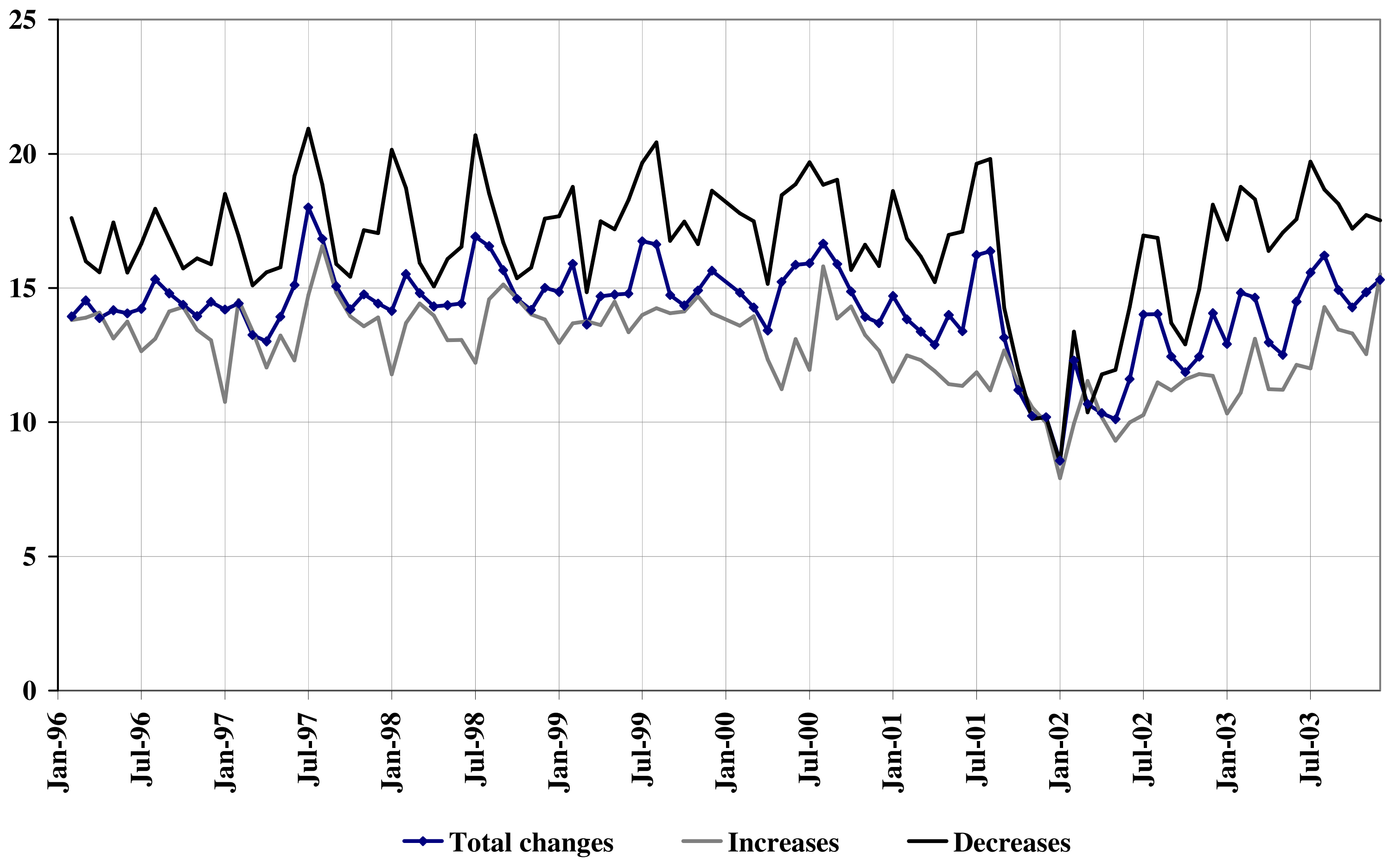


Figure 4: Distribution of the duration of price spells (frequencies in percent)

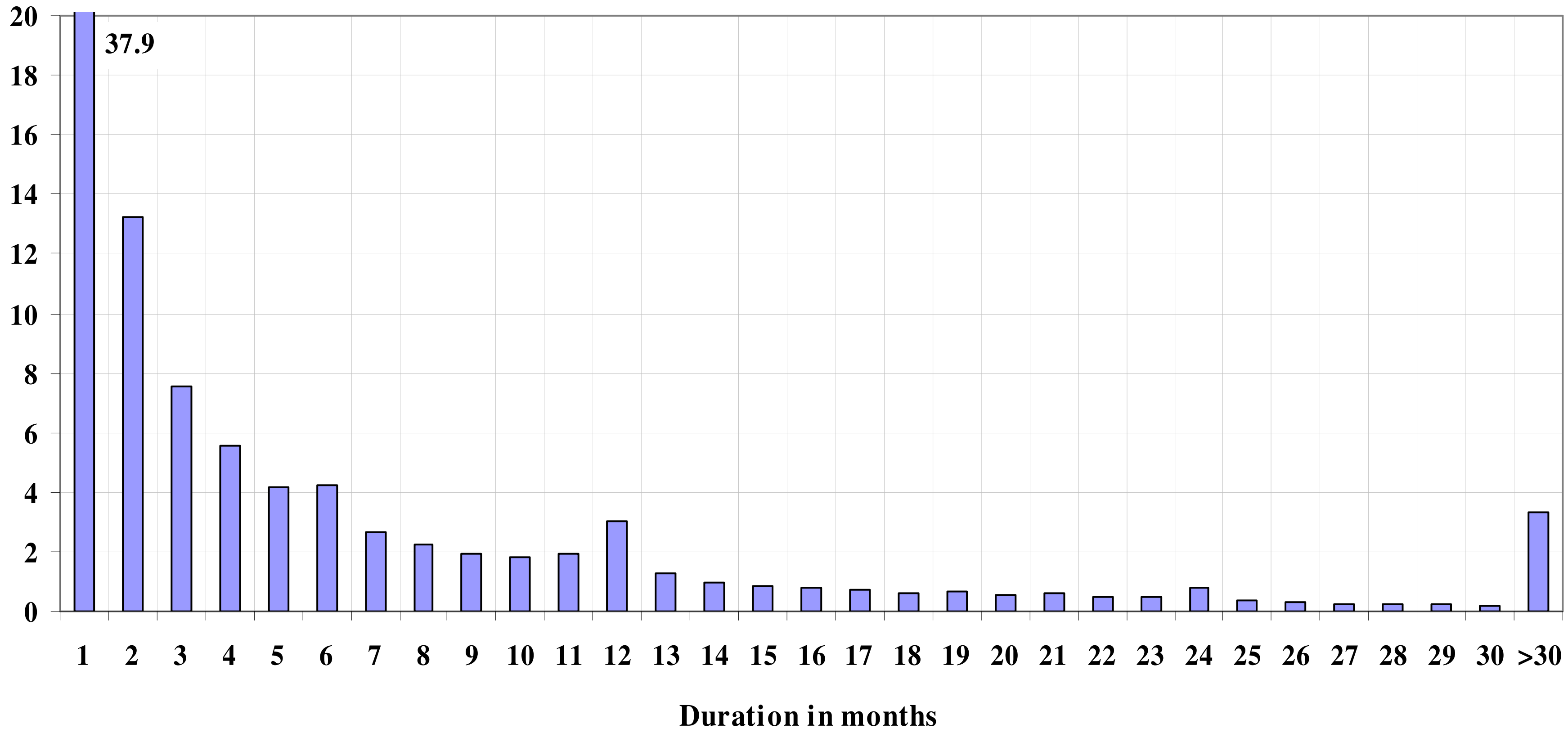


Figure 5: Unweighted frequency of price changes (in percent)

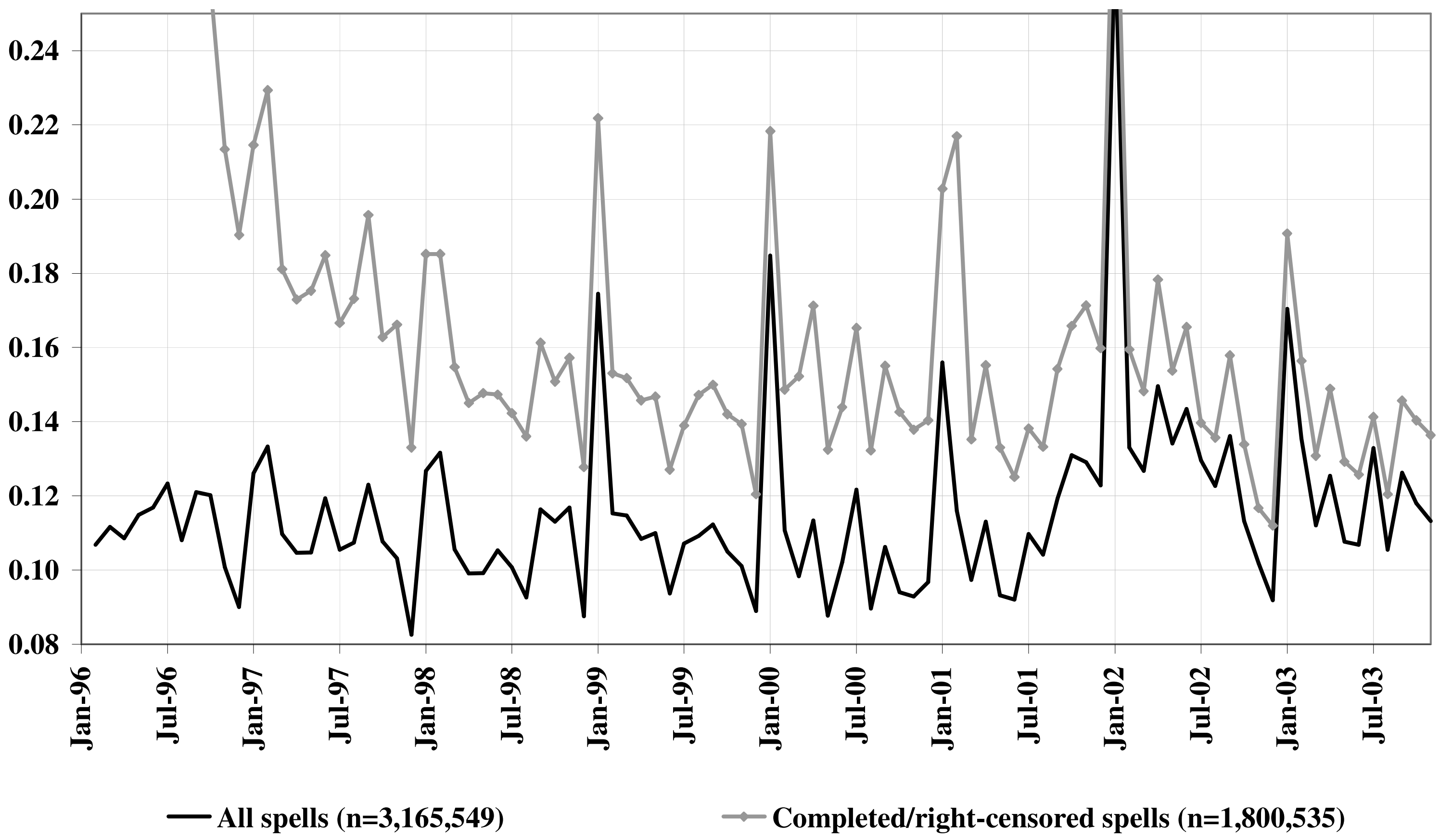


Figure 6

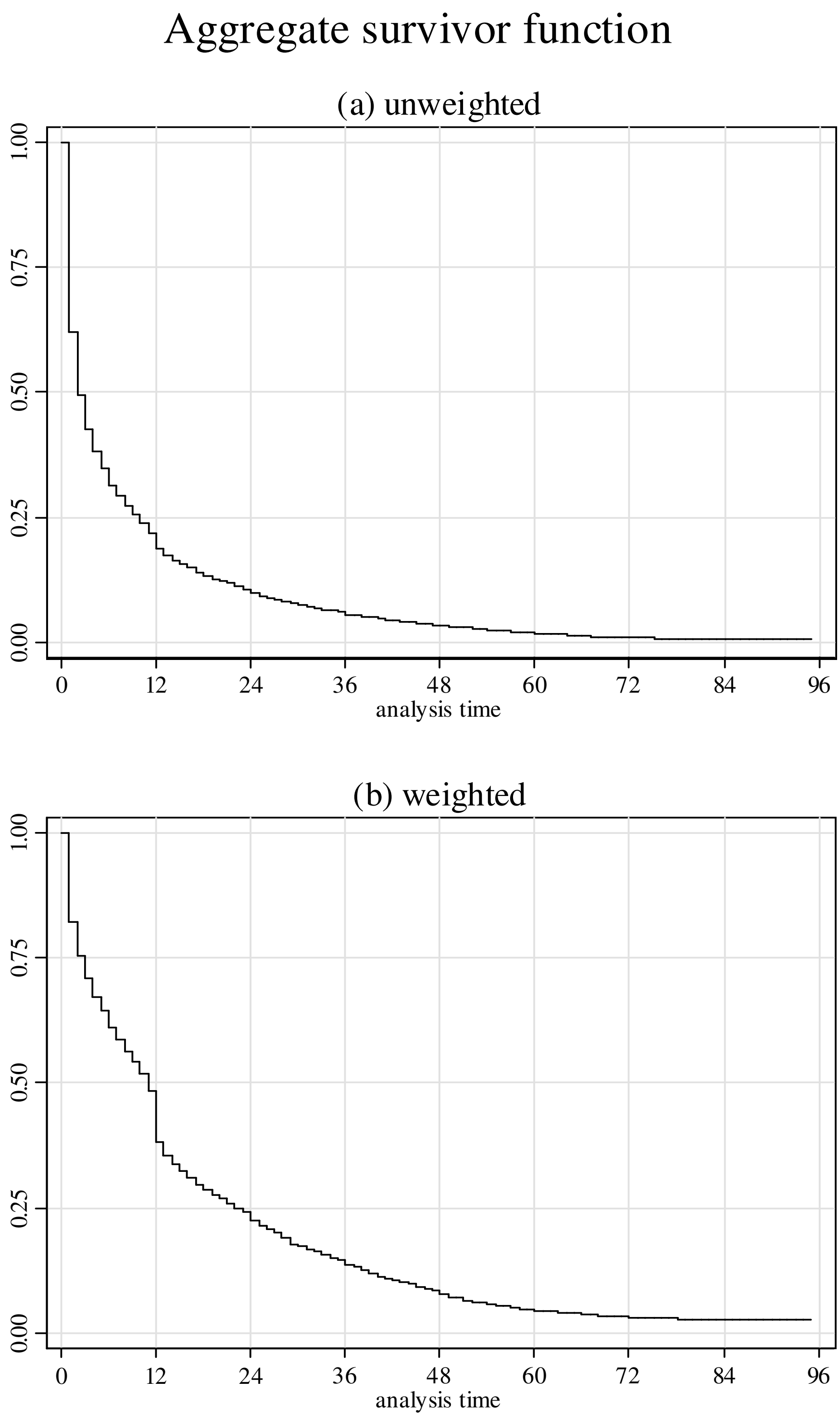
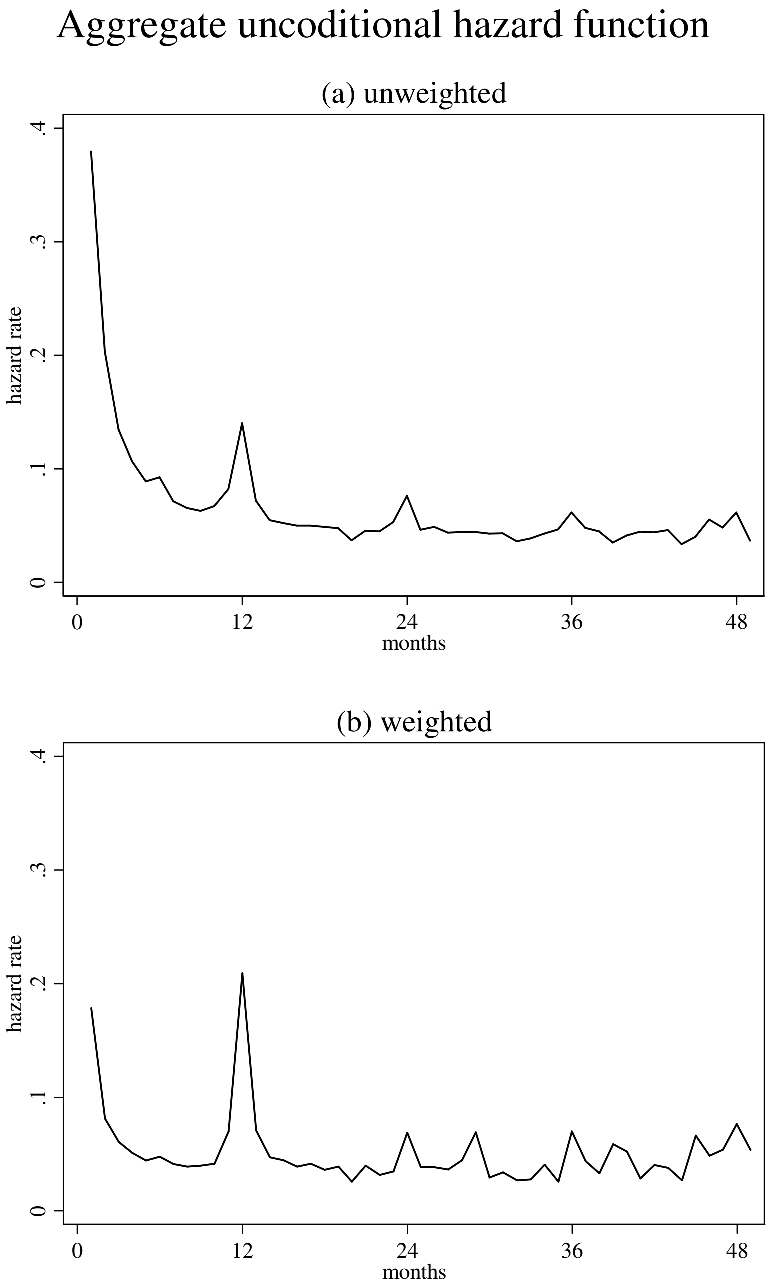

Aggregate survivor function
(a) unweighted
1.00
0.75
0.50
0.25
0.00
0
12
24
36
48
60
72
84
96
analysis time
(b) weighted
1.00
0.75
0.50
0.25
0.00
0
12
24
36
48
60
72
84
96
analysis time

Figure 7



Aggregate uncoditional hazard function
(a) unweighted
hazard rate
0
.1
.2
.3
.4
0
12
24
36
48
months
(b) weighted
hazard rate
0
.1
.2
.3
.4
0
12
24
36
48
months

Figure 8

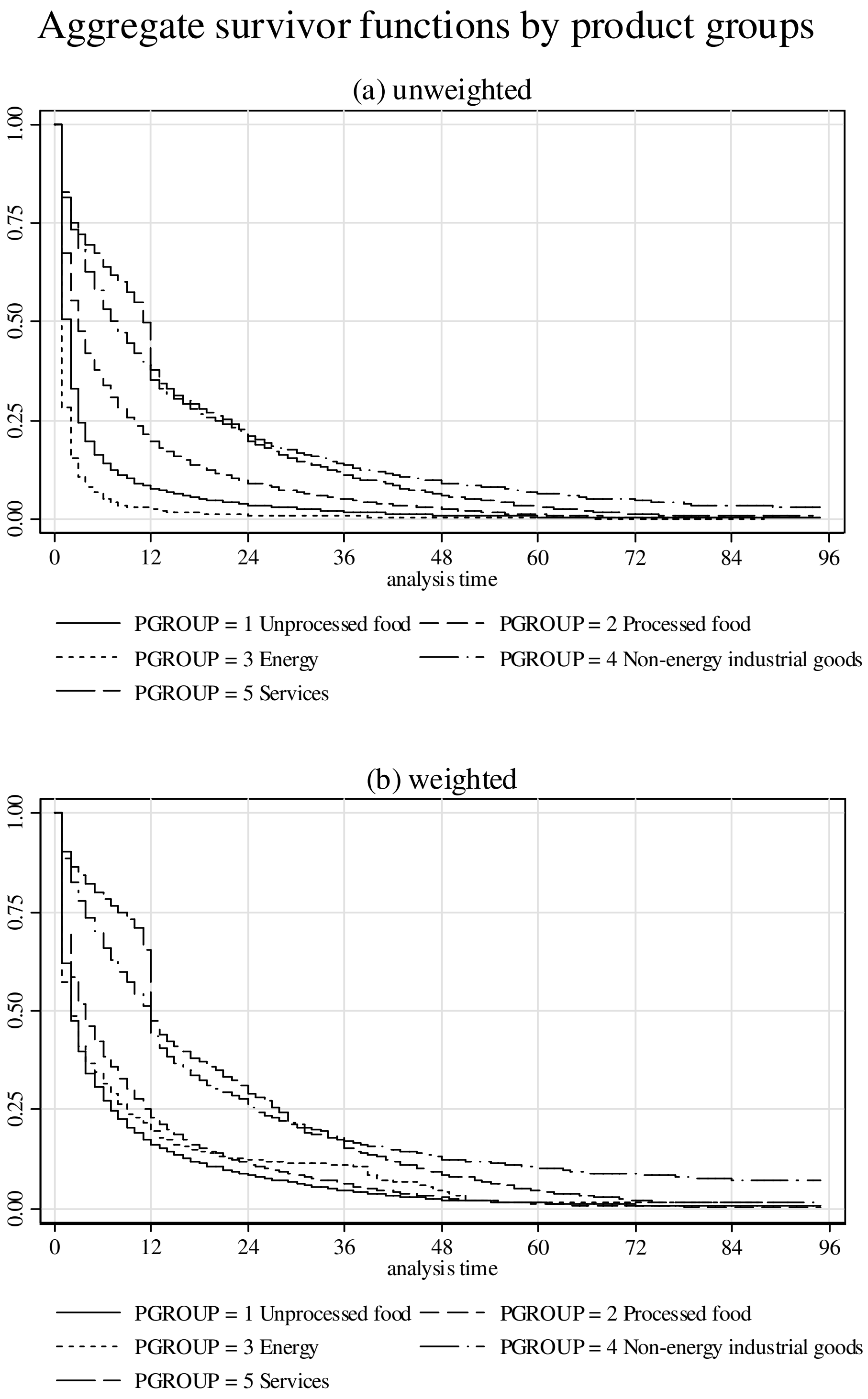
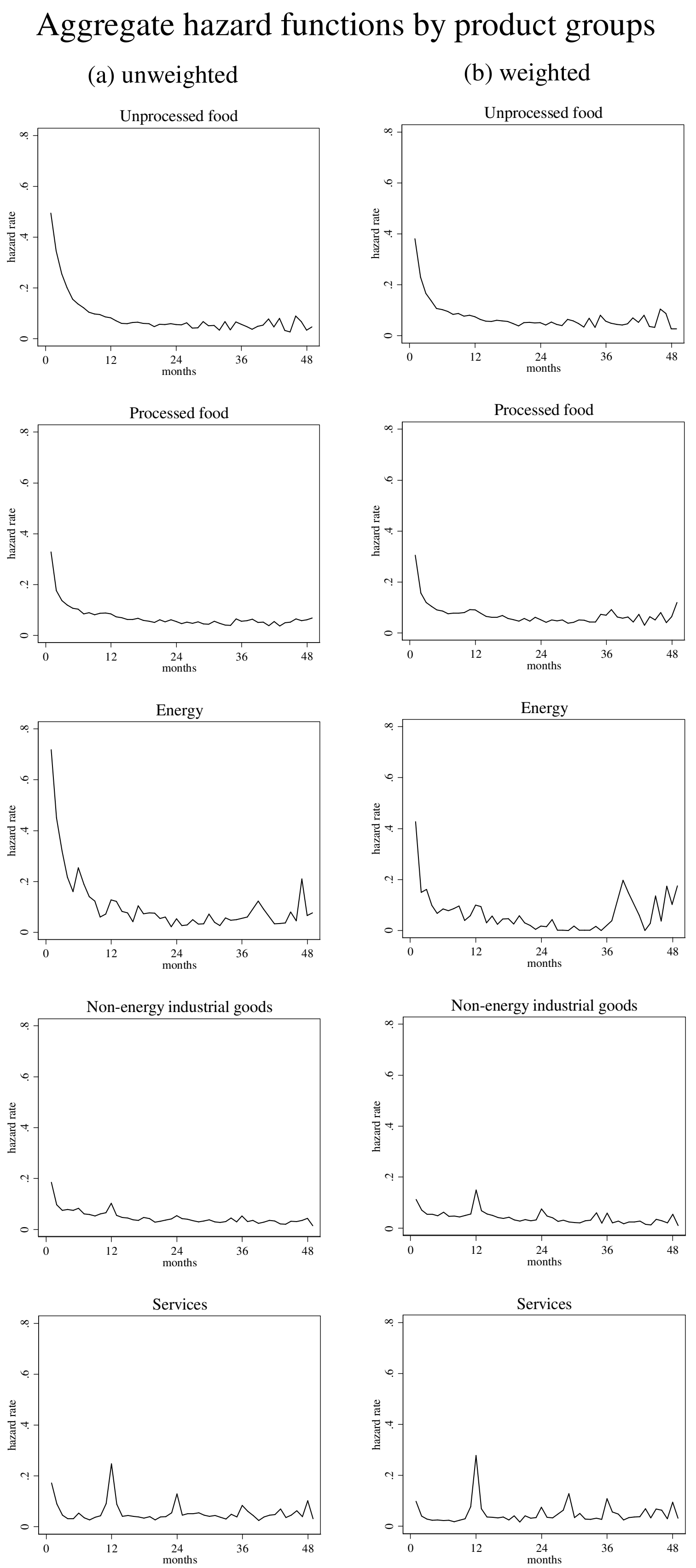

Aggregate survivor functions by product groups
(a) unweighted
0.00
0.25
0.50
0.75
1.00
0
12
24
36
48
60
72
84
96
analysis time
PGROUP = 1 Unprocessed food
PGROUP = 2 Processed food
PGROUP = 3 Energy
PGROUP = 4 Non-energy industrial goods
PGROUP = 5 Services
(b) weighted
0.00
0.25
0.50
0.75
1.00
0
12
24
36
48
60
72
84
96
analysis time
PGROUP = 1 Unprocessed food
PGROUP = 2 Processed food
PGROUP = 3 Energy
PGROUP = 4 Non-energy industrial goods
PGROUP = 5 Services

Figure 9


Aggregate hazard functions by product groups
(a) unweighted
(b) weighted
Unprocessed food
Processed food
Energy
Non-energy industrial goods
Services
hazard rate
0
.2
.4
.6
.8
0
12
24
36
48
months

Annex III

Figure A1

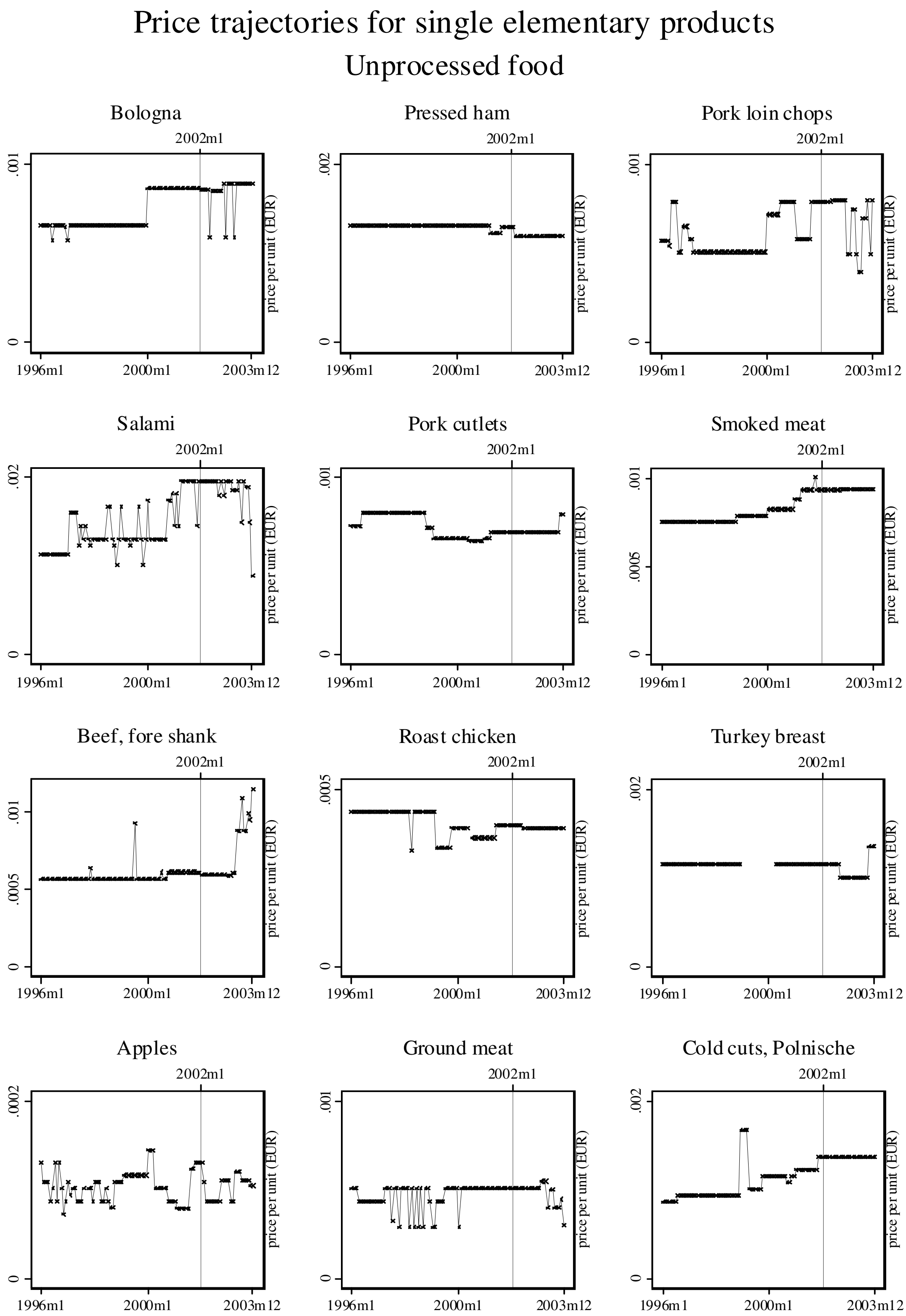


Figure A2

Price trajectories for single elementary products

Processed food

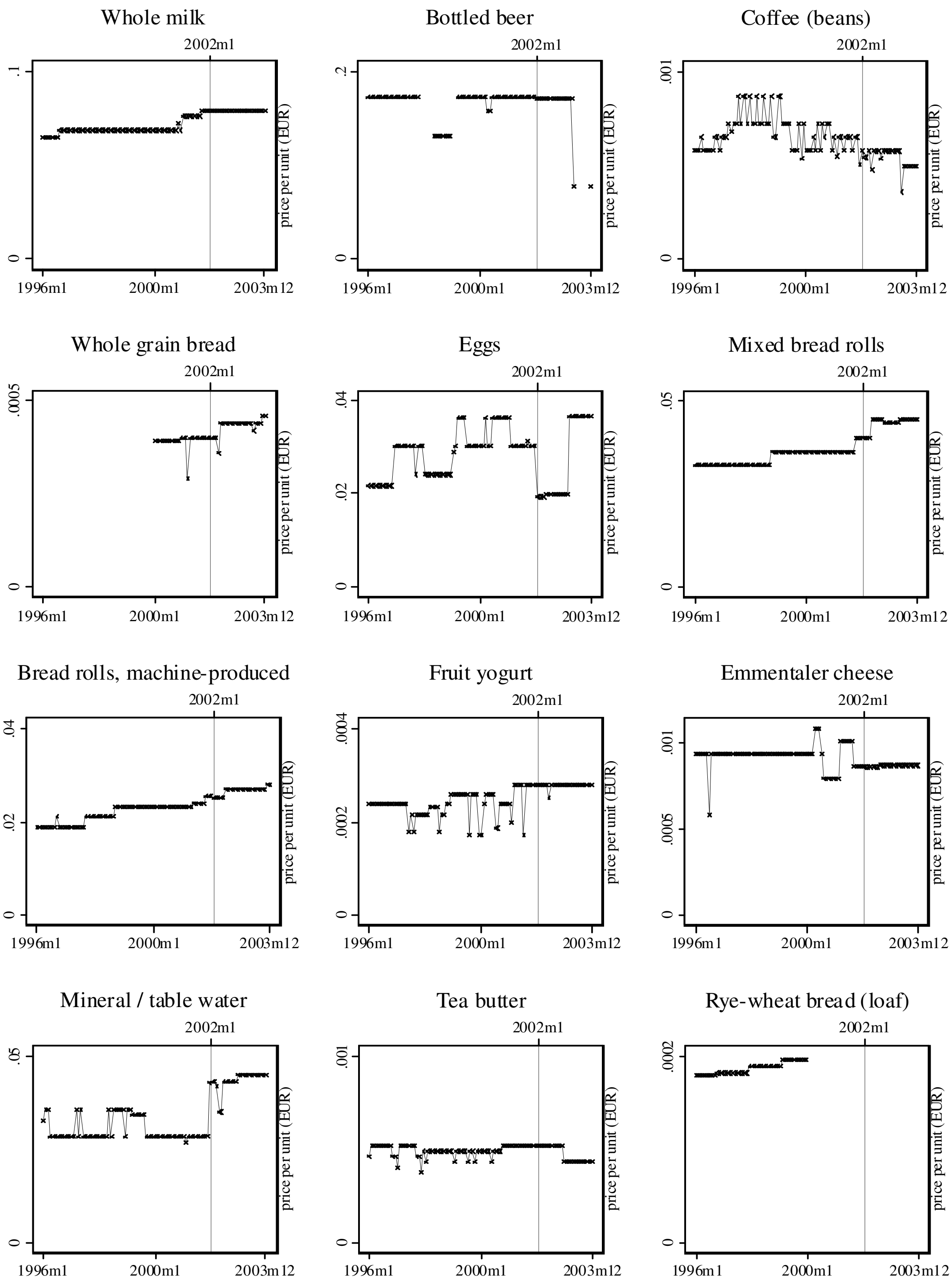
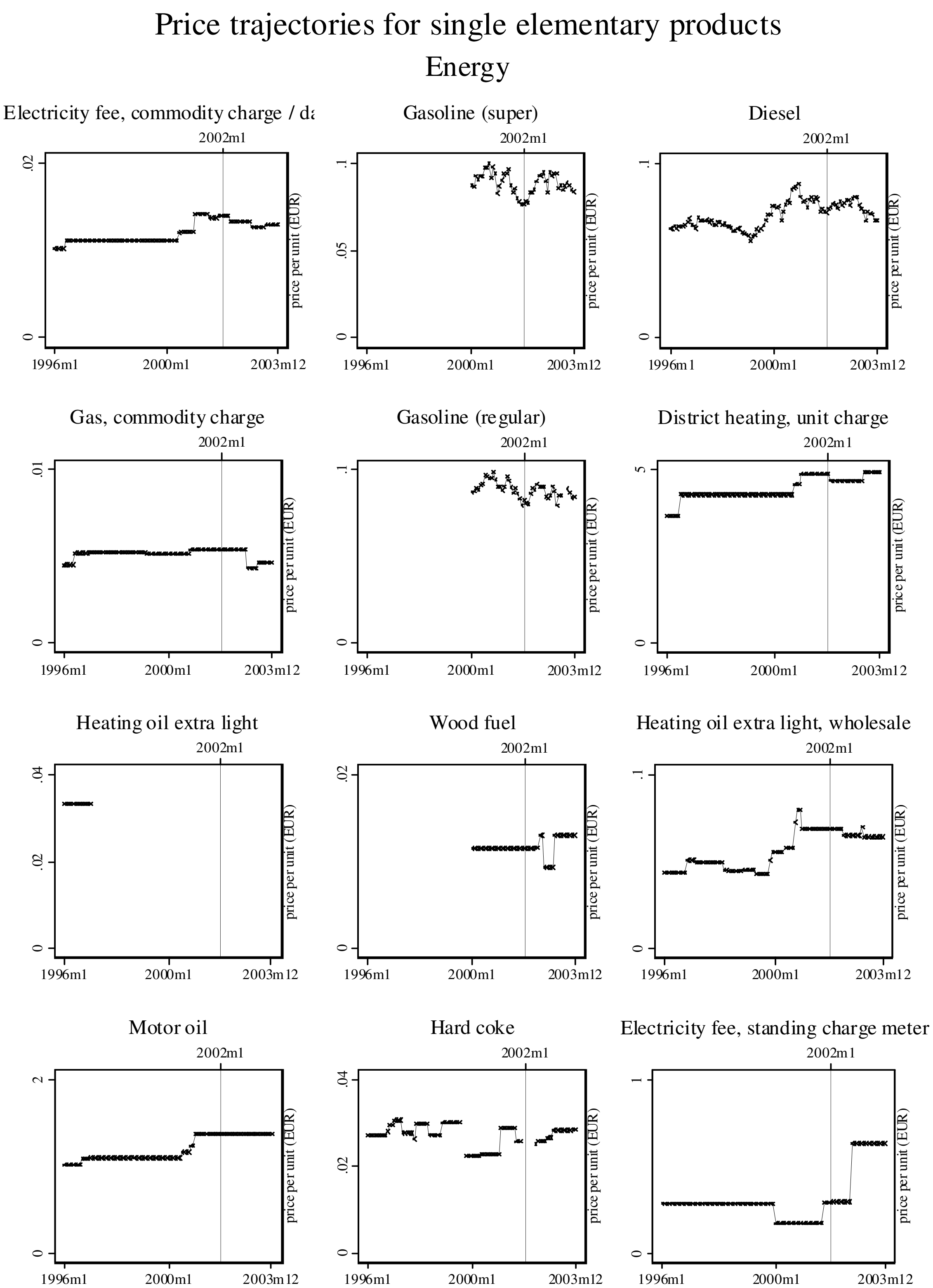


Figure A3



Price trajectories for single elementary products
Energy
Electricity fee, commodity charge / da
Gasoline (super)
Diesel
Gas, commodity charge
Gasoline (regular)
District heating, unit charge
Heating oil extra light
Wood fuel
Heating oil extra light, wholesale
Motor oil
Hard coke
Electricity fee, standing charge meter
2002m1
1996m1
2000m1
2003m12
price per unit (EUR)
0
.02
.05
.1
.01
.5
.04
1
2

Figure A4

Price trajectories for single elementary products

Non-energy industrial goods

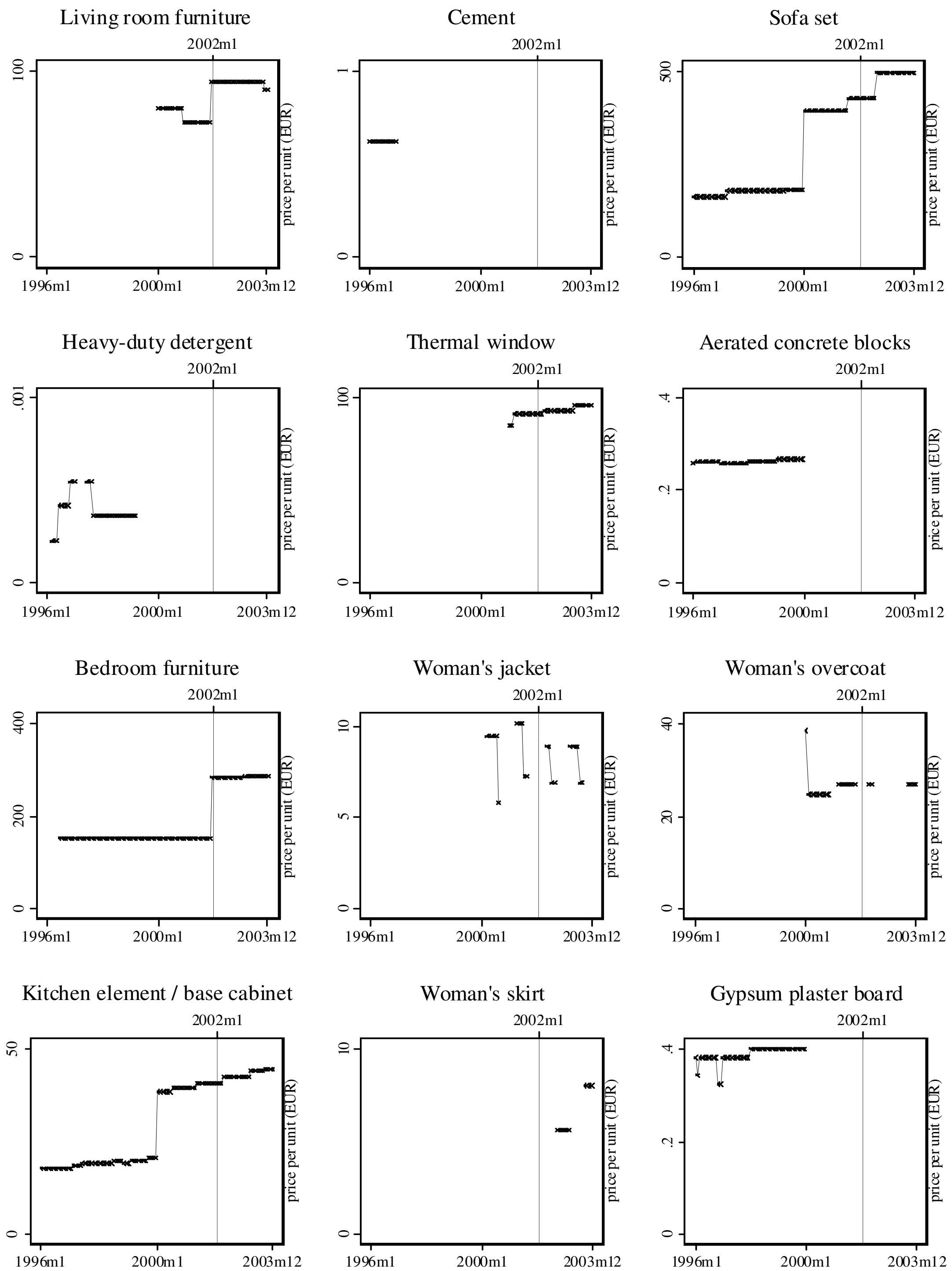


Figure A5

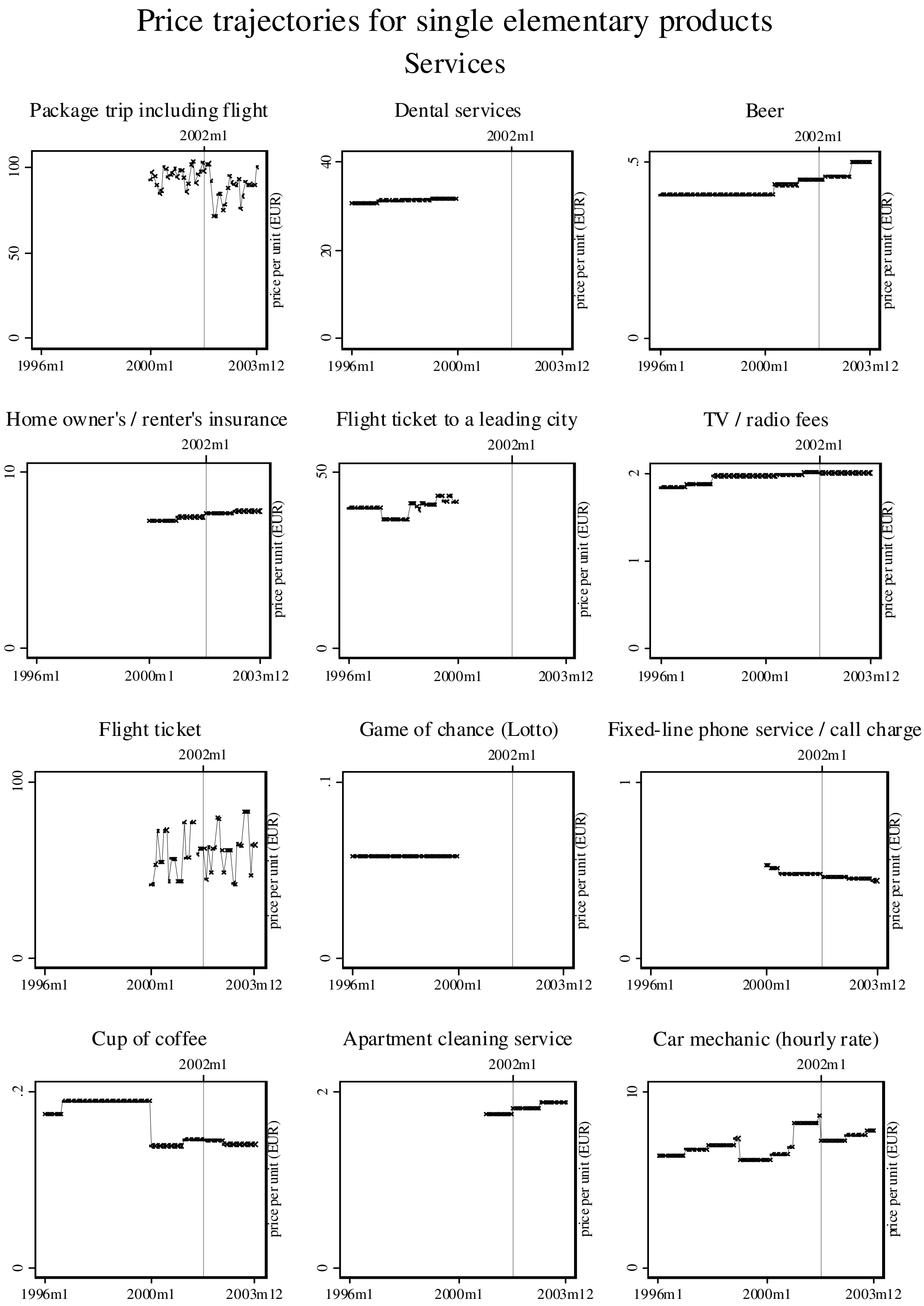

Price trajectories for single elementary products
Services
Package trip including flight
Dental services
Beer
Home owner's / renter's insurance
Flight ticket to a leading city
TV / radio fees
Flight ticket
Game of chance (Lotto)
Fixed-line phone service / call charge
Cup of coffee
Apartment cleaning service
Car mechanic (hourly rate)
2002m1
price per unit (EUR)
1996m1
2000m1
2003m12

Figure A6

Hazard functions for single product categories

Unprocessed food

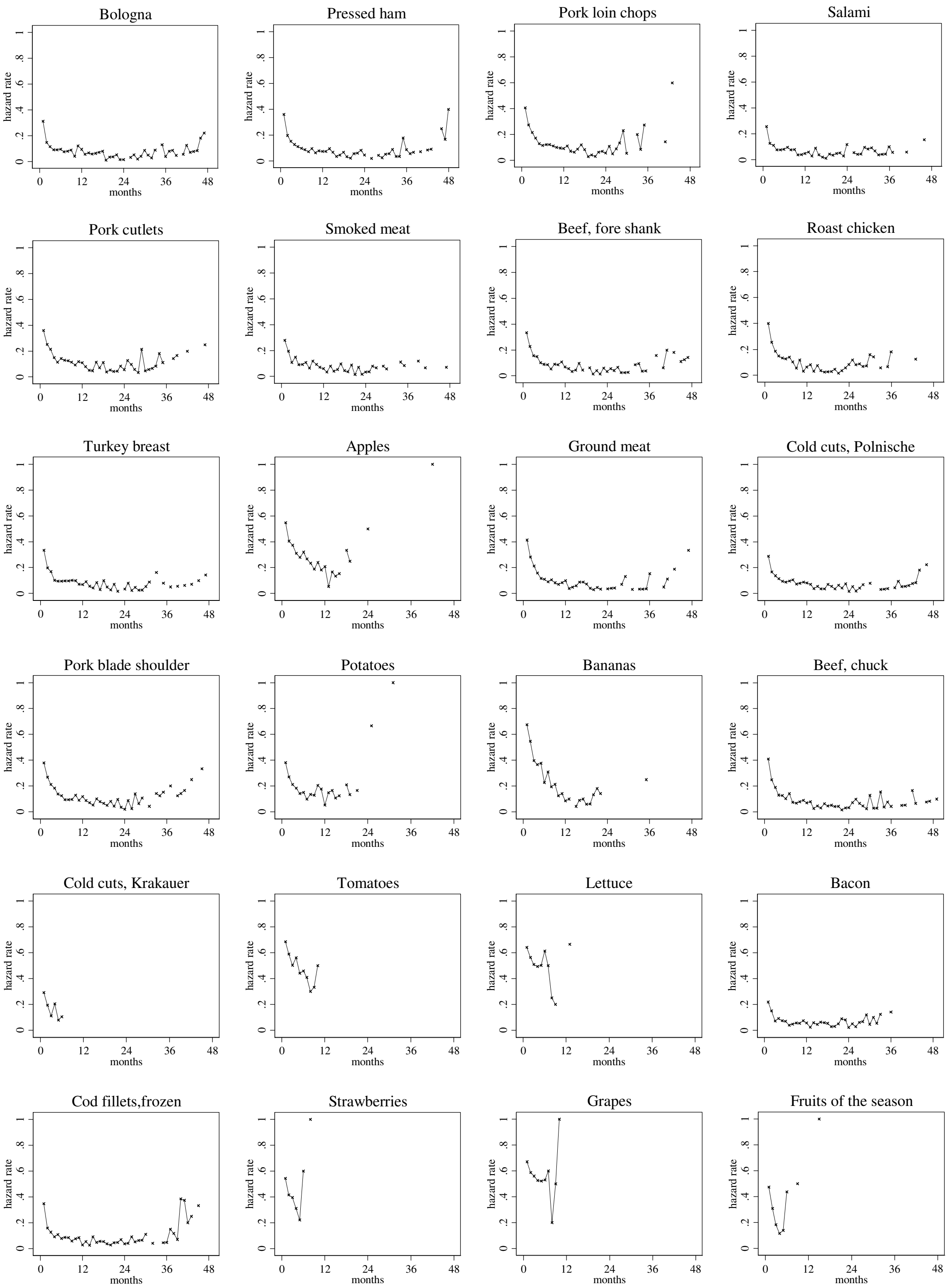


Figure A7

Hazard functions for single product categories

Processed food

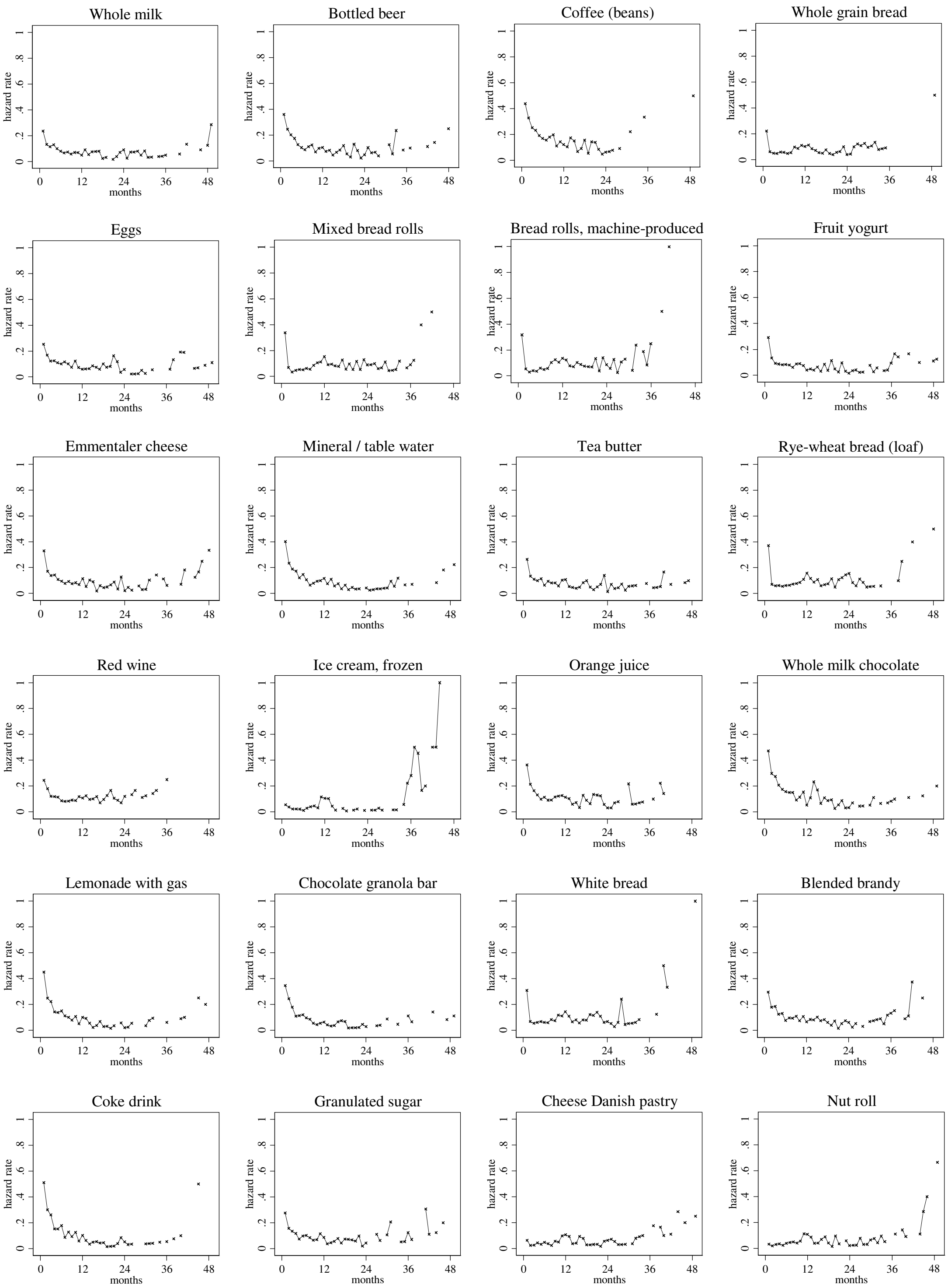


Figure A8

Hazard functions for single product categories

Energy

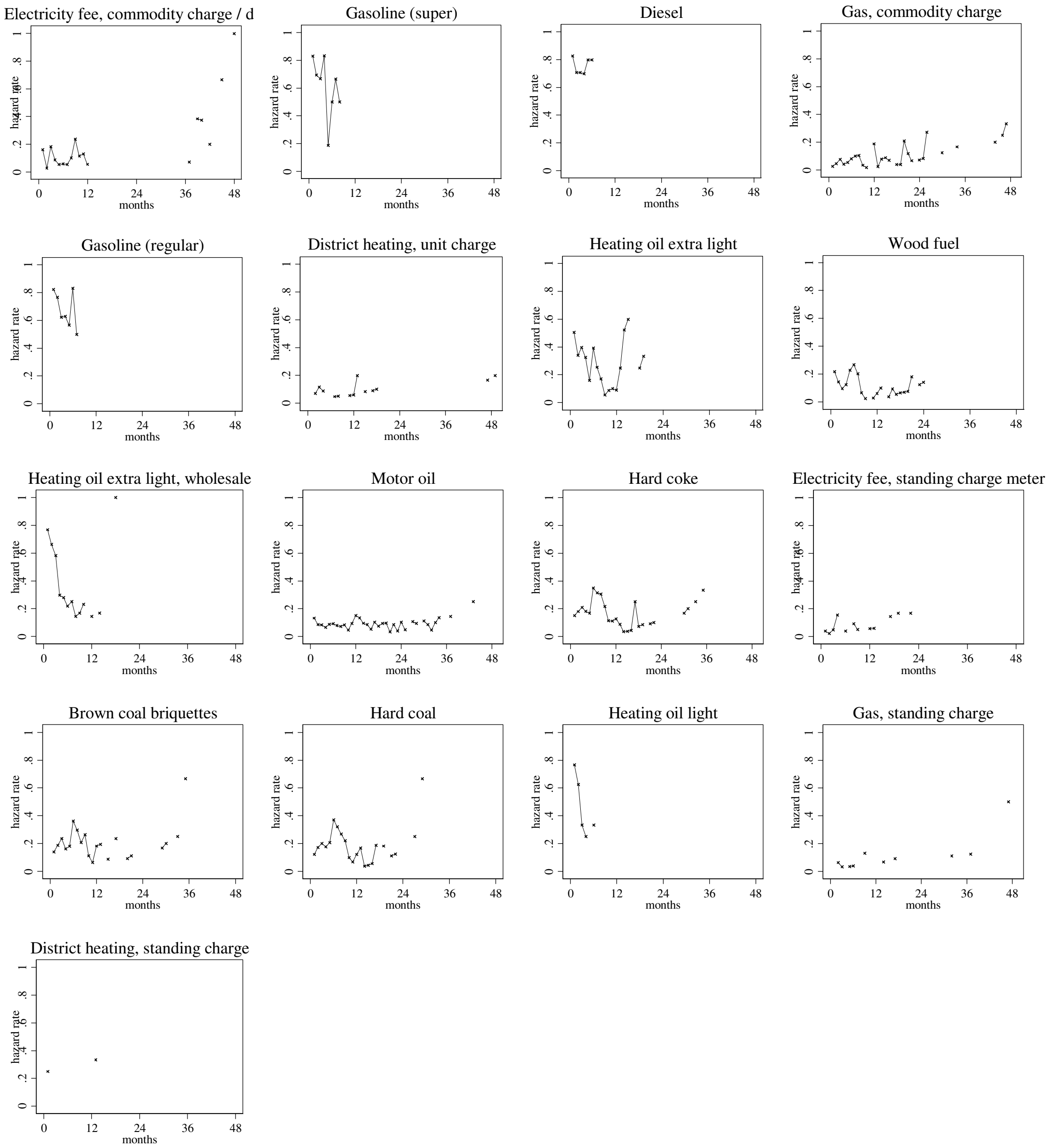


Figure A9

Hazard functions for single product categories

Non-energy industrial goods

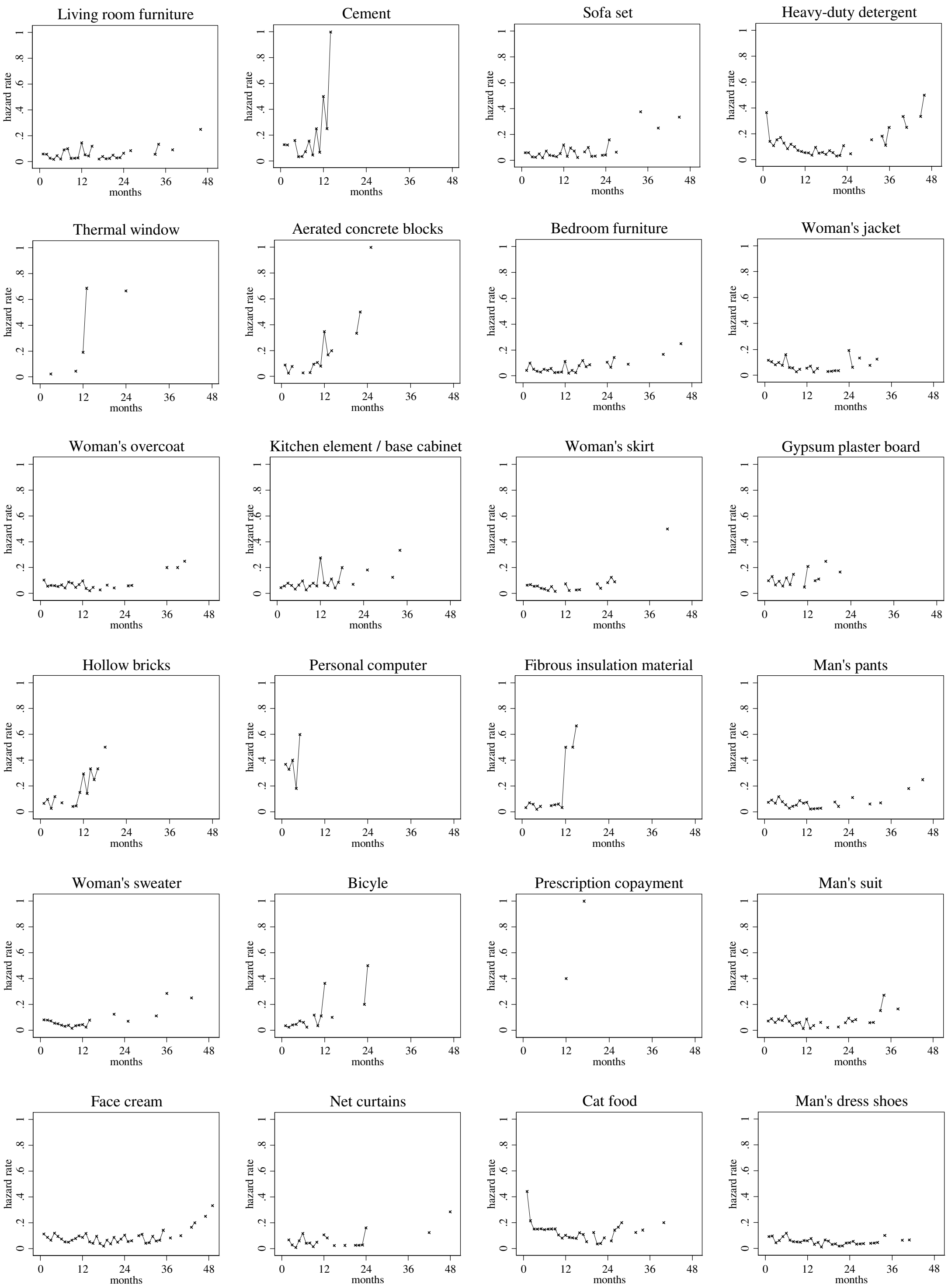
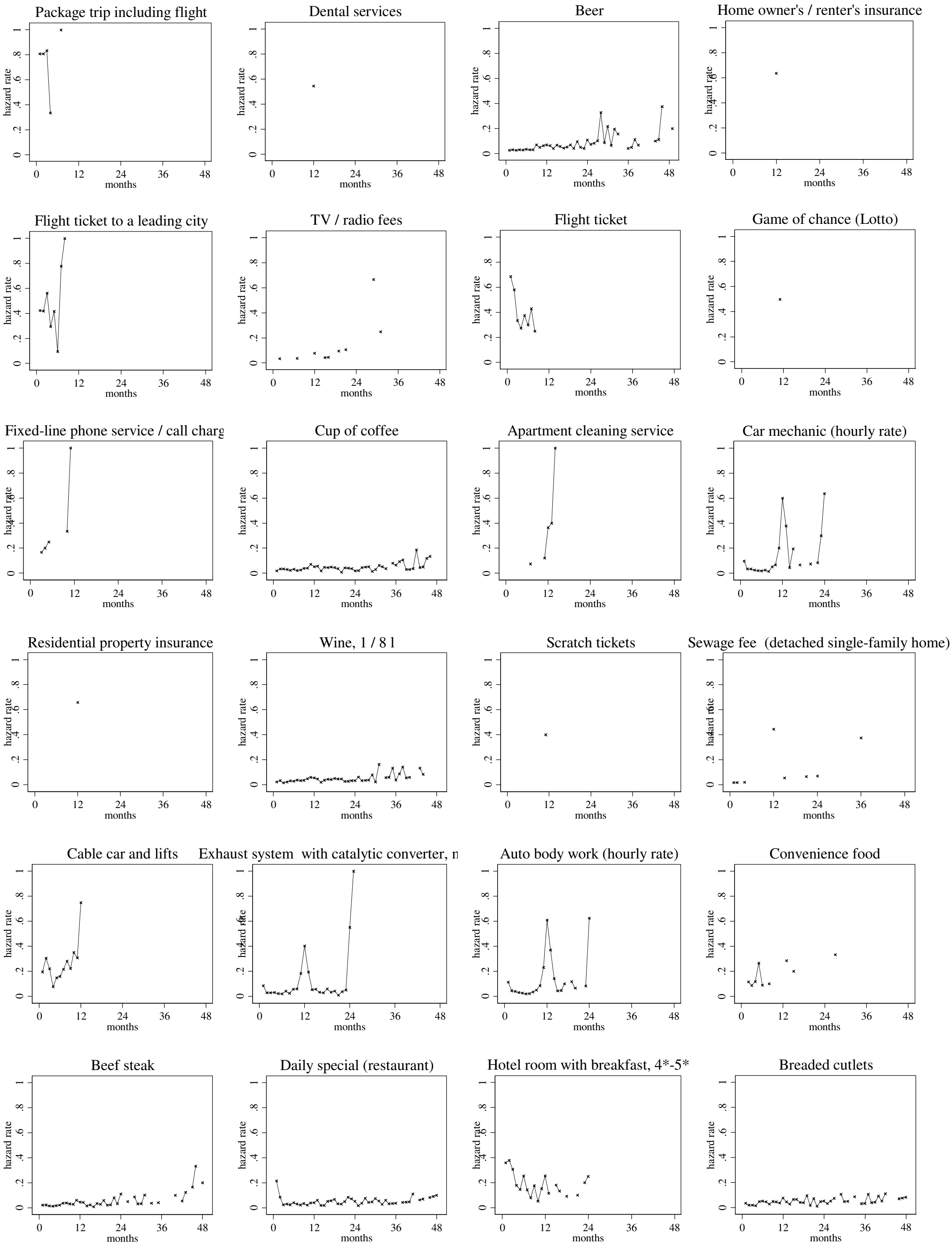


Figure A10

Hazard functions for single product categories
Services



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