EUROPEAN CENTRAL BANK

WORKING PAPER SERIES
NO. 413 / NOVEMBER 2004

EUROSYSTEM INFLATION PERSISTENCE NETWORK

AN EMPIRICAL ANALYSIS OF PRICE SETTING BEHAVIOUR IN THE NETHERLANDS IN THE PERIOD 1998-2003 USING MICRO DATA

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AN EMPIRICAL ANALYSIS OF PRICE SETTING BEHAVIOUR IN THE NETHERLANDS IN THE PERIOD 1998-2003 USING MICRO DATA [1]

by Nicole Jonker [2], Carsten Folkertsma [3] and Harry Blijenberg [4]

1 This research is part of a joint effort of euro area central banks and has been initiated as part of the Eurosystem Inflation Persistence Network (IPN). We thank Rob Vet for his outstanding research assistance. We would like to thank an anonymous referee, Bouke Buitenkamp, Peter van Els, Peter Vlaar, Marco Hoeberichts, Ad Stokman, participants of the Inflation Persistence network and the members of the IPN research groups 2 and 3 in particular for their helpful comments on earlier versions of this paper. We would like to express our gratitude towards Statistics Netherlands and in particular to Cécile Schut and Jan Walschots for giving us the opportunity to use the database on Dutch consumer prices. Statistics Netherlands cannot be held responsible for neither the research methods nor the research outcomes. All remaining errors are exclusively the authors' responsibility. Views expressed are those of the individual authors and do not necessarily reflect official positions of neither De Nederlandsche Bank nor Statistics Netherlands.

2 Corresponding author: De Nederlandsche Bank, Payments Policy Department , P.O. Box 98, 1000 AB Amsterdam, The Netherlands, e-mail: n.jonker@dnb.nl

3 De Nederlandsche Bank, Research Department, P.O. Box 98, 1000 AB Amsterdam, The Netherlands.

4 Statistics Netherlands, P.O. Box 4000, 2270 JM Voorburg, The Netherlands.

The Eurosystem Inflation Persistence Network

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

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

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

Address
Kaiserstrasse 29
60311 Frankfurt am Main, Germany

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

Telephone
+49 69 1344 0

Internet
http://www.ecb.int

Fax
+49 69 1344 6000

Telex
411 144 ecb d

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

CONTENTS

Abstract

In this paper we examine pricing behaviour of retail firms in the Netherlands during 1998-2003 using a large database with monthly price quotes of 49 articles, representing different product types. We have conducted this study in order to gain insight in the degree of nominal rigidity of consumer prices in the Netherlands. We find that prices of energy and unprocessed food are most flexible, whereas prices of services are stickiest. A multivariate analysis shows that firm size matters with prices being stickiest in small firms and most flexible in large firms and in retail firms consisting of the owners only. Furthermore, we investigate pass-through effects of VAT changes in prices. We find that VAT increases are almost completely passed on to consumers. Finally, there is some evidence indicating that pricing behaviour of retail firms was different during the introduction of the euro than in the period directly preceding it.

Keywords: nominal rigidity of prices, frequency of price change, Cox regression

JEL classification: E31, D49, C41

NON-TECHNICAL SUMMARY

This paper presents the empirical results of a study on pricing behaviour in the Netherlands in the period 1998-2003. It has been conducted as part of the Eurosystem Inflation Persistence Network. The study is, as far as we know, the first to map pricing behaviour of retail firms in the Netherlands using a unique large micro dataset with monthly product prices. By means of duration analysis we also assess the effects of outlet and product group characteristics on the duration of price quotes. Furthermore, we pay attention to the occurrence of state dependent pricing strategies by assessing the effects of the euro cash changeover and changes in VAT on prices.

We focus on prices of 49 products, representing 9 COICOP categories. Excluded from the analysis are products related to health, telecommunication and education. The 49 products have a total weight of almost 8% in the Dutch CPI.

The average price duration in the Netherlands is almost 10 months. However, there is much variation in price duration across sectors. The frequency of price changes is highest in energy (every month) and in the unprocessed food sector (every three months), whereas prices of non energy industrial goods and services change about once a year. These sector effects are significant according to the estimation results of the Cox proportional hazard model. Price increases occur more often than price decreases, but the difference in occurrence is rather small, indicating that nominal prices are not downward rigid. On average, the magnitude of price decreases is somewhat higher. This picture also emerges in other European countries.

Cox regression results also show that there are differences in the duration of price spells across outlets of different sizes. Price adjustment is fastest in large firms (100 employees or more) and slowest in small firms (1-9 employees). A similar result has been found by France. An explanation of the size effect may lie in menu costs with menu costs declining by firm size. Remarkable is that price adjustment in one-man businesses takes place almost as often as in the large firms.

We also pay attention to price effects due to changes in VAT rates. Here, it seems there may be some asymmetry in price adjustments. According to Cox regression results, changes in VAT rates shorten the duration of price spells. This holds both for increases and decreases in VAT. Yet, an increase in VAT seems to be completely passed on to consumers, but a decrease in VAT only partially. However, evidence for this latter finding is rather limited

Another interesting finding regarding state dependent price effects, is that during the euro cash changeover the frequency of price changes, both increases and decreases in price, was higher than in the period before the cash changeover. This holds especially for non energy industrial goods and services and to a lesser extent for unprocessed food (probably also partly due to cattle diseases and poor harvests). Cox regression shows an

increased probability of price change in November and especially for December 2001 as well as in March (month double pricing ended) and April 2002. Generally, the magnitude of the price increases was somewhat smaller during the euro cash changeover than before this period. The magnitude of the price decreases differed less. However, comparing price statistics before and during the euro cash changeover for processed goods, non energy industrial goods and services shows that inflation for these product groups was relatively high during the introduction of the euro. This is also supported by a comparison of December 2001 prices with January 2002 prices. Usually, prices are lower in January than in December in the previous year because of winter sales, but this was not the case in 2002.

The finding that Dutch price setters follow both time- and state-dependent pricing strategies suggests that macroeconomic models for monetary policy should combine both price adjustment mechanisms. Developing these hybrid models may lead to significantly better models of the monetary transmission mechanisms.

1 INTRODUCTION

This paper analyses price setting behaviour of retail firms in the Netherlands in the period 1998-2003 using a unique data-set. It is the first empirical study on this topic using Dutch data and it provides unique new insights about the price adjustments in the Netherlands at the micro level. The main purpose of this study is to map out the degree of nominal rigidity of consumer prices in the Dutch economy at the sector level. We use monthly price data of 49 products included in the Dutch Consumer Price Index (CPI), representing 9 out of 12 COICOP[3] categories (excluding health, education and telecommunication[4]). Apart from the consequences of monetary policy on prices we also pay attention to asymmetric price effects of changes in indirect taxes, distinguishing between VAT increases and VAT decreases. Another special feature of this paper is that we analyse the effect of the euro cash changeover on prices.

Prices of most articles and services do not change continuously but are usually kept constant by firms for a certain period of time. One of the reasons for this is that changing prices in response to changes in supply or demand factors do not always immediately outweigh the costs involved with changing prices, the so called menu costs. If price rigidities are present, then monetary policy may affect real variables in the short term. In this sense it is important to understand to what extent price rigidities are present in the CPI. Therefore, describing and explaining nominal rigidity is essential for understanding the implications of monetary policy on short term economic developments.

Several macroeconomic models for monetary policy have been developed incorporating alternative price adjustment processes allowing for nominal price rigidities. The Taylor model (prices are set for a fixed number of months, Taylor, 1999) and the Calvo model (each period a fixed proportion of firms may adjust its prices with the distribution of opportunities to adjust prices following a Poisson process, Calvo, 1983) are the best known time dependent pricing models. According to these models, monetary shocks have not immediately their full impact on inflation, because of price stickiness. Instead, a gradual and prolonged effect is predicted by these models. The truncated Calvo model is a combination of the Taylor and the Calvo model and assumes that each period a fixed proportion of the firms sets its prices during the lifetime of a contract. If a contract expires each firm will always set a new price, i.e. the duration of a price quote can't exceed the duration of the contract. According to this model the probability of a price change is constant during the duration of the contract, but is equal to 1 when the contract expires (see e.g. Wolman, 1999). In state dependent pricing models, like in Caplin and Spulber (1987), the probability that a firm changes a product price depends on the difference between the actual price and the firm's target price. Firms do not continuously adjust their prices because of menu costs. When the difference between target price and actual

[3] COICOP is an abbreviation of Classification Of Individual Consumption by Purpose. This product classification is maintained of the European Union (Eurostat).
[4] Health and education are excluded from the analysis because prices of products in these categories are mainly set by the government. Telecommunication is excluded because the product (telefax machine) representing this category is not included in

price is large enough to make a price adjustment profitable for the firm, the price is adjusted. Menu costs are positively related with the general level of inflation. Dotsey et al. (1999) present a model combining the Calvo approach with state dependent pricing features. In their model firms face random menu costs. Firms with relatively low menu costs choose to adjust prices frequently whereas firms with higher menu costs wait longer before adjusting their prices. An increase in general inflation speeds up the price adjustment process.

The effects of monetary policy on inflation and the real economy if price setting is described by one of these theories have been extensively discussed in the macroeconomic literature. However, these discussions mainly focussed on theoretical issues and not so much on microeconomic evidence. Some papers have been devoted to analyse price stickiness empirically, like Cecchetti (1986), Estrada and Hernando (1999), Chevalier et al. (2000), Hall et al. (2000), Bils and Klenow (2002) and Fougère et al. (2004). Bils and Klenow use the BLS consumer price data and study retail price stickiness using monthly price data for 1995-1997 on 350 categories of goods and services. Fougère et al. have conducted a very interesting study in which various theoretical pricing models are tested using advanced duration models using French CPI data.

However, most empirical work focuses on price stickiness in the US and the UK and on small numbers of products. Little is known yet about price stickiness in the Euro Area. This study has been conducted as part of the Eurosystem Inflation Persistence Network (IPN). We use data from the period November 1998 until April 2003. This enables us to study pricing behaviour during the introduction of the Euro in the Netherlands. Other countries represented within this network, for which similar studies have been conducted, are Austria, Belgium, France, Finland, France, Italy, Luxembourg, Portugal and Spain. Results for Belgium (Aucremanne and Dhyne, 2004), France (Baudry et al., 2004), Italy (Fabiani et al., 2004) and Portugal (Dias et al., 2004) have recently been published.

The remainder of this paper is organised as follows: Section 2 gives a description of the data. Section 3 consists of two subsections. The first subsection introduces the pricing statistics for the Netherlands which have been calculated by all countries participating in the IPN. The second subsection gives a brief introduction into duration analysis and more specifically on the Cox regression. Section 4 presents and discusses the pricing statistics for the individual products as well as the aggregated results. Section 5 does the same for the results from the Cox regressions. Results of the Cox regressions are given for the whole sample and by product category. This section also pays attention to state dependency of price changes by comparing pricing behaviour in the Netherlands during the introduction of the euro with the period just before as well as by analysing price effects in case of changes in VAT. Finally, section 6 summarises the paper and concludes.

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the Dutch CPI basket. Unfortunately, we did not have a close substitute in the sample at our disposal. However, we think that the absence of the fax machine in our data, will not alter the main results substantially.

2 DATA

The price data we use in this paper are from Statistics Netherlands (Centraal Bureau voor de Statistiek, CBS). The data-set includes monthly information for the period November 1998- April 2003 on prices of 49 individual products collected at different outlets in the Netherlands. These 49 products have a total weight of almost 8% in the Dutch CPI. Table 1 shows which information is available for each product in the micro dataset and table 2 lists the 49 products of the common sample. Table 2 shows the classification of the products by COICOP[5] group and product type[6]. The CPI weights of base year 2000 are also reported in table 2[7]. Not all countries participating in the consumer price study of IPN had access to the price data of all products in their national CPI. Therefore, it has been agreed that all participating countries analyse the prices of a well-defined subset of products in the national CPI baskets[8]. This approach ensures the comparability of the research results across participating countries. By focusing on a subset of the CPI basket, we were able to tackle data problems in a comprehensive way. The subset of goods and services has been selected to represent a wide spectrum of goods and services, including processed food, unprocessed food, energy, transport, non energy industrial goods, various kinds of services, seasonal products, etc. So although the total weight of the goods and services in the sub sample is only 8% of the Dutch CPI, the prices analysed in our study still provide valuable information on the price setting behaviour of firms in the market. Goods and services related to health care and education are excluded from the common sample. Prices of products falling in these categories are often administered or regulated prices. Explaining the behaviour of these prices is beyond the scope of this study.

Statistics Netherlands collects data of product prices as follows. Each month interviewers visit specific outlets and register prices and package sizes of the articles included in the CPI shopping basket. If different varieties of a product fit the description of the article to be sampled interviewers are instructed to register the price of the best selling brand of the outlet. If the exact item sampled last period is not available any more the

[5] There are 12 COICOP codes, i.e. 1=food and non-alcoholic beverages, 2=alcoholic beverages and tobacco, 3=clothing and footwear, 4=housing, water, electricity and gas, 5=furnishings, household equipment and routine maintenance of the house, 6=health, 7=transport and fuels, 8=Communication, 9=recreation and culture, 10=education, 11=Restaurants and hotels and 12 miscellaneous goods and services. Products of the COICOP codes 1, 2, 3, 4, 5, 7, 9, 11 and 12 are included in the sample. The product fax-machine with COICOP code 8 (Communication) has not been included in our sample because Statistics Netherlands does not collect price data of fax machines. We could not find a suitable substitute for the fax machine in the data-set at our disposal. A second Eurosystem classification is also used in this paper. It distinguishes five subcategories, i.e. unprocessed food (UPF), processed food (PF), energy (E), Non energy industrial goods (NEI) and services (S).

[6] A second Eurosystem classification is also used in this paper. It distinguishes five product types, i.e. unprocessed food (UPF), processed food (PF), energy (E), Non energy industrial goods (NEI) and services (S).

[7] The CPI weights in table 2 refer to the weights of the lowest level COICOP of the individual articles. They don't refer to the weights of the individual articles. If two articles in the sample are in the same lowest level COICOP group we have divided the corresponding weight over these two articles.

[8] Some countries have access to the product prices of the entire CPI. However, we only had access to a subsample of the Dutch CPI and focussed our analyses on the common sample.

collectionner is supposed to substitute the 'old' best selling item with the 'new' best selling item fitting the description of the product[9].

Table 2 presents an overview of these articles together with their COICOP code, the number of price trajectories, the number of price spells, the number of left censored price spells, the number of right censored price spells and the number of observations. There are 204,404 observations in our data-set. Each combination of a price of a specific article in a specific outlet at a given date is an **observation** (hypothetical example: a 1,5 litre bottle mineral water of brand X bought in supermarket Y in April 2002, costing € 0,99). A **price trajectory** refers to a series of price quotes for a specific article of a specific brand observed in a specific outlet. A price trajectory can be divided into different **price spells, i.e.** the time periods in which the price of a product of a specific brand at a specific outlet does not change. A price spell is treated as being **left-hand censored** at the beginning of a price trajectory; the start date of the price observed at the beginning of a trajectory is not known to Statistics Netherlands. Analogously, price spells ending at the end of the observation period April 2003 are **right-hand censored.** Censoring may lead to a downward bias in the estimation of the duration of an event, since there may be relatively many 'long duration' spells among the censored ones. Just omitting censored spells from the analysis would lead to a data set with relatively many spells of short duration.

Our data set contains 7,214 price trajectories and 45,697 price spells. On average, there are 6.33 price spells within a price trajectory. Regarding censored price spells; first price spells of all price trajectories are considered to be left-hand censored. Furthermore, there are 3,301 price spells which end on April 2003 and we consider as right-hand censored. The number of price observations and price trajectories differ between products. Men's shirts range with more than 10,000 price quotes among the most frequently sampled goods, followed by socks and lettuce which both have over 8,000 price quotes in our data set. For fuel we have less than 1,000 price quotes, while for heating gas not more than 21 observations are available. Collection of price quotes for the latter two articles is somewhat different than data collection for the other articles in our sample. From October 2000 onwards, fuel prices are collected by Statistics Netherlands via internet and not by interviewers visiting petrol stations. The price of heating gas was – until recently - regulated in the Netherlands and changed at most twice a year by at most 3 guildercents per cubic meter of gas (excluding changes in tax-rates), depending on the development of the oil price.

[9] We don't think that this may lead to an upward bias in price change frequencies, since a specific item is only replaced by a new item of the best-selling brand when the old item isn't available anymore (involuntary replacement). The price trajectory of the old item ends and a new price trajectory for the new item starts. Involuntary replacement and price trajectories are going to be defined in this section.

3 DEFINITIONS

3.1 Pricing statistics

P_{ijt} represents the price of one particular article i (i = 1 to n_j where n_j represents the total number of individual articles in the product classification j) of the product classification j (j = 1 to 50) at time t (t falling in the period November 1998-April 2003). An individual article is defined by its characteristics (individual article code) and its selling point (location and outlet). The monthly frequency of a price change, increase or decrease, of product j is denoted by F_j. This frequency statistic can also be refined by distinguishing between the frequencies in price increases and price decreases, F_j^+ respectively F_j^-. On top of that we also include variables measuring the average magnitude of the change in price, also broken down into separate variables for price increases and price decreases of product j, $\overline{\Delta}_j^+$ and $\overline{\Delta}_j^-$. The precise formulas of these pricing statistics can be found in the appendix of this paper.

The monthly frequency of price changes of product j can be used to derive the median duration of a price for product j as well as the average duration. The definitions given below are valid under the assumption that the durations of prices follow an exponential distribution. An advantage of constructing duration measures by using frequencies is that the statistics are not biased by censored observations. All observations, both censored and uncensored, can be used to estimate the monthly price change frequencies.

Median price duration: $$T_{50,j} = \frac{\ln(0.5)}{\ln(1-F_j)} \qquad (1)$$

Average price duration: $$\overline{T}_j = -\frac{1}{\ln(1-F_j)} \qquad (2)$$

There are a few things worth mentioning with respect to the collection of price data and the construction of the variables related to price changes:

- In our sample all guilder prices until December 2001 are converted to euro prices. Small changes (at 2nd decimal level) in prices due to the guilder euro conversion are not regarded as price changes in the analysis.

- Sometimes, within a price spell a price is not recorded in month t, but in both months t-1 and t+1 the same price was recorded. Instead of creating two time spells we thought it more reasonable to impute the price of month t-1 and t+1 in month t.

- Statistics Netherlands' definition of articles may be narrow (e.g. a specific chocolate bar) but it may also be rather broad, especially for articles within the categories clothing or furnishing. This may cause spurious price changes because articles may be replaced with other articles that also fall within the definition. E.g. the definition of a men's shirt is: white, cotton, long sleeves. It may happen that in month t a (slightly) different men's shirt (with a different price) is bought in a particular outlet although the original shirt is still available at this outlet with unchanged price. This may lead to an upward bias in the frequency of price changes and an unknown bias in the magnitude of the price changes of articles in categories with relatively broad product definitions.

- Regarding **attrition** due to **product replacement**, the Dutch database doesn't provide unambiguous information on the nature of product replacement (**voluntary** or **involuntary**[10]). In the Dutch data set, when a voluntary article replacement occurs a new price trajectory starts for the replacement. In case of voluntary article replacement the last price spell of an article is actually right-hand censored, whereas the first price spell of its replacement may be left-hand censored. Regarding the calculation of pricing statistics (see sections 4 and 5) we consider all replacements to be forced and we assume the last price spell of an article in case of replacement to have ended, also when they are actually right-hand censored. This may lead to an upward bias in the reported frequencies on price changes and a downward bias in the estimated lengths of price spells (section 4, tables 3, 4 and 5). First price spells are always considered to be left-hand censored and are removed from the duration analyses (section 5, table 6). Because of the large number of observations in the data set we do not believe this will alter the main research results substantially, although long duration spells may be underrepresented.

3.2 Basic concepts duration analysis

In this section we introduce some basic concepts often used in duration analysis (for a more extensive exposition see Greene, 1997, or Lancaster, 1990). Duration analysis has its roots in biomedical research where it is also known as survival analysis. There it is, for example, used in the analysis of survival times after the diagnosis of a disease or after a medical treatment. At the end of the seventies Lancaster (1979) and Nickell (1979) introduced duration models in empirical labour economics, for analysing the time the unemployed needed to find themselves a new job. From then on, the use of duration analysis became more and more important in economics.

In duration analysis the variable of interest is the length of time that elapses from the beginning of the event under investigation until either its end or until the end of the observation period. The durations in the sample

[10] Voluntary product replacement refers to a particular product or selling point not being considered anymore by the statistical agency to be representative of the consumption habits of the population. The article and selling point still exist but are replaced by another article or selling point that is believed to be more representative. Involuntary or forced replacement occurs when a product can't be bought anymore at a particular selling point or when a particular selling point stops to exist.

may have started at the same point in calendar time but they may also have started at different points in calendar time. The duration of events which are not completed at the end of the observation period are said to be right-hand censored. The precise duration of right-hand censored durations is not known. However, what is known is the minimal duration. If an observation i is censored after t_i periods of time, the duration amounts at least to the observed t_i periods of time. In the estimations right-hand censoring will be taken into account. We consider all 3,025 price spells, having an April 2003 price quote to be right-hand censored. Regarding left-hand censoring, in the micro data-set there are also about 7,000 first price spells, most of them starting at November 1998. For most of these price spells the month of first observation may not be the actual starting date of the first price quote observed. Since our data-set is quite large we have decided to exclude these probably left-hand censored price spells from the duration analysis. This leaves us with 151,920 price quotes, 38,483 price spells of which 35,458 are completed before April 2003.

Suppose that the random variable T, measuring the duration of a certain event, has a density distribution f(t). The corresponding cumulative distribution function gives the probability that the duration of the event is lower or equal to t

$$F(t) = \int_{s=0}^{t} f(s)ds \tag{3}$$

In duration analysis, it is quite common to look at the probability that the length of the spell is at least t periods. This probability is given by the complement of F(t), known as the survival function

$$S(t) = 1 - F(t) \tag{4}$$

Another concept often used in duration analysis is called the hazard rate. The hazard rate reflects the conditional probability that, given the spell has lasted until t, the spell will end in the short interval of time (t, t+Δ). Another interpretation of the hazard rate is the rate at which spells are completed after duration t, given that they lasted at least until t (see e.g. Lancaster, 1990, p. 7-8 for a derivation).

$$l(t) = \lim_{dt \to 0} \frac{P\left(t \le T \le t + dt \mid T \ge t\right)}{dt} = .. = \frac{f(t)}{S(t)} \tag{5}$$

The hazard rate shows the pattern of the distribution of completed spells over time. The exponential (constant hazard rate), the Weibull (hazard rate increases or decreases over time) and the Log-logistic distribution (hazard rate first increases over time and then decreases) are the most simple specifications of the distribution function of the duration under study (again, see Lancaster, 1990 for an overview of more complex specifications). However, it is also possible to leave the distribution of T unspecified and to focus on the effects of explanatory variables on T.

There are two ways to incorporate the effect of explanatory variables into the hazard model. The first is known as the accelerated failure time (AFT) model in which the natural logarithm of the survival time is related linearly with the explanatory variables and an error term. The distribution of the duration under investigation depends on the assumption about the distribution of the error term z. The general idea of AFT models is to change the time scale by a factor $\exp\left(x_j b\right)$. A factor smaller than one decelerates passing of time, whereas a value larger than one accelerates the passing of time.

$$\ln\left(t_j\right) = x_j b + z_j \qquad (6)$$

Another branch of duration models is known as proportional hazards (PH) model. PH models divide the hazard function λ(t) in two parts

$$l\left(t_j\right) = l_o\left(t\right) g\left(x_j\right) \qquad (7)$$

A baseline time pattern for the hazard rate $\lambda_0(t)$ is multiplied by a nonnegative function g of the explanatory variables. A common assumption for g is the exponential distribution: $g\left(x_j\right) = \exp\left(x_j b\right)$. You can proceed by specifying a distribution for the baseline hazard $\lambda_0(t)$, but you may also decide to leave the baseline hazard unspecified. The advantage of leaving $\lambda_0(t)$ unspecified is that the estimation of the effects of the explanatory variables on the event under investigation, does not get clouded due to imperfections in the parameterisation of the baseline hazard. This might play a role here, because it is not unthinkable that baseline hazards corresponding to the duration of price quotes display multiple peaks (seasonal effects, time dependent pricing, sales, etc.).

The basic log likelihood function for analysing duration data which takes right-hand censoring reads as follows (eq. 8). In this specification time varying covariates are not taken into account yet:

$$\ln L = \sum_{uncensored} \ln\left(f\left(t \mid b\right)\right) + \sum_{censored} \ln\left(S\left(t \mid b\right)\right) = \qquad (8)$$

In duration analysis it is convenient to reformulate this log likelihood function and use the hazard function instead of the density function, using the relation f(t)=λ(t)S(t). The log likelihood function can be rewritten into two parts. The first part consists of the contributions to the likelihood function of the uncensored price spells, i.e. the observations of which the price spell is completed. The second part consists of the contributions of all observations in the sample.

$$\ln L = \sum_{uncensored} \ln\left(l\left(t \mid b\right)\right) + \sum_{all} \ln\left(S\left(t \mid b\right)\right) \tag{9}$$

Plugging the expression for the hazard function in eq.7 in eq. 9 while leaving out the unspecified baseline hazard λ_0 gives us the partial log likelihood of the Cox proportional hazards model

$$\ln L = \sum_{uncensored} x_i' b - \sum_{all} \exp(x_i' b) \tag{10}$$

In the equation above it is implicitly assumed that only one observation exits at each distinct exit time. If we extend eq. 10 by allowing for multiple exits and for time-varying covariates we get:

$$\ln L = \sum_{t=1}^{D} \left[\sum_{k \in D_t} \ln(x_{kt} b) - d_t \ln \left\{ \sum_{i \in R_t} \exp\left(x_{it} b\right) \right\} \right] \tag{11}$$

The following new symbols are introduced in eq. 11:

D : D denotes the month. It ranges from November 1998 until April 2003.

D_t : D_t denotes the set of price spells k that are completed in month t. It may be empty in case no spells are completed in month t

d_t : number of price spells that are completed at t

R_t : the set of price spells i at risk in month t

x_{it} : vector of covariates of price spell i at month t

We have used the statistical package Stata 7 for optimising the likelihood function. We have estimated the robust variance-covariance matrix for the parameter vector β by the method devised by Lin and Wei. We have taken right-hand censoring into account in the estimations of the Cox duration model. Spells ending in April 2003 are all considered to be right-hand censored. The efficient score residuals have been summed within price spell cluster before using the robust variance estimator.

4 EMPIRICAL RESULTS: FREQUENCY AND MAGNITUDE OF PRICE CHANGES 1998- 2003

In this section the pricing statistics are shown and discussed. Table 3 presents the statistics for the individual products and table 4 focus on aggregated statistics[11],[12]. These aggregated statistics are weighted once to get an estimator for the COICOP or product type aggregates and are weighted twice for estimating the CPI statistic (bottom row table 4). Table 5 provides the values, both for the single weighted sample and for its CPI basket representative (article statistics, double weighted), of the 5^{th}, 25^{th}, the median, the 75^{th} and 95^{th} percentiles of the frequency in price changes and the duration of prices.

Table 3 shows that the monthly frequencies of price changes and the magnitudes of these monthly price changes. Frequency and the size of price changes vary widely across the 49 products. The frequency of changes in fuel prices is close to one, meaning that fuel prices change almost on a monthly basis. However, with 3% the average change in fuel price is relatively moderate. Other articles that change price relatively often are the unprocessed fresh food articles lettuce (frequency 0.72) and bananas (frequency 0.46). These prices are the most volatile in the common sample; they change relatively often and they change a lot. Their prices change by about 30% each month (both up- and downwards price adjustments). All products with at most one price adjustment per year belong to services or non energy industrial (NEI) goods, namely domestic services, a car wash, hiring a video tape, a football, drinks and food in restaurants/cafes and a suitcase.

We have estimated the correlation between the frequency of price changes and their magnitude using the figures in table 3 to shed some light on the role of menu costs in price setting behaviour of firms. We distinguish between price increases and price decreases. We assume that the magnitude of the price change is a proxy of the magnitude of the menu costs. Generally, the more frequently prices change, the lower menu costs are likely to be. Hence, we regard a negative correlation between the frequency of price changes and the magnitude of these price changes as an indication for the importance of menu costs in the price adjustment process. The estimated correlation between the frequency of price increases and their magnitude

[11] The formula used to construct CPI representative statistics S from COICOP/product type statistics S_k is: $S=\sum_{k=1}^{n_k} w_k S_k \,/ \sum_{k=1}^{n_k} w_k$ with w_k the weight of COICOP category k or product type category k in the Dutch CPI. A similar formula is used to estimate S_k using the product level statistics S_j: $S_k=\sum_{k=1}^{n_k}\sum_{j=1}^{n_j} I_{jk} w_j S_j \,/ \sum_{j=1}^{n_j} I_{jk} w_j$, with S_j the estimated statistic for product j from the micro data and w_j is the weight of product j in the Dutch CPI. The aggregated statistics by COICOP or product type are weighted *once* (CPI weights of products within COICOP or product type group). The statistics representing the CPI are weighted *twice*: within COICOP group or product type using CPI weights for individual products *and* by the CPI weights for the COICOP groups or product types in the Dutch CPI. In the text we refer to the common sample as the *single* weighted sample and to the CPI representative as the *double* weighted sample.

[12] In RG2 we have agreed that each country uses CPI weights. In an earlier stage of this research we used HICP weights instead of CPI weights. Changing the weights hardly altered the aggregated statistics. The main differences between the Dutch CPI and HICP are the exclusions of prices related to the costs of home ownership, private health insurances, consumption related taxes (e.g. VAT) and services offered by the public sector in the HICP, whereas they are included in the Dutch CPI.

is +0.1. The estimated correlation between the frequency of price decreases and their magnitude is even lower, namely +0.02. These figures suggest there is no clear relation between the frequency of price changes and the magnitude of these price changes. However, if we focus on non-food products the correlation between the frequency of price increases and their magnitude becomes –0.3 and the correlation between the frequency of price decreases and their magnitude becomes –0.2. These latter figures suggest that there is a negative relation between the frequency of price changes and the magnitude of price changes for non-food products. This indicates that menu costs are likely to be a factor in the price setting behaviour of firms selling non food products and provides some empirical support for state dependent pricing models.

Table 4 shows that the average duration of a price spell is about 9.7 months for the double weighted sample, representing the average duration of a price spell of products included in the Dutch CPI. There is a lot of variation in average duration of price quotes across COICOP categories/product types, ranging from an average duration of 1.5 months in energy (due to the fuel prices) to almost a year in NEI goods and services. Looking at the pricing statistics in table 3 and 4 which distinguish between price increases and price decreases, we see that although prices are usually changed upwards, downward price adjustments are by no means an exception. Price cuts are least likely in services and most likely in energy due to frequent changes in the oil price. The average frequency of price increases is with 10% almost twice as high as the average frequency of price decreases. Both tables also reveal that the magnitude of the average price decreases is larger than the magnitude of the price increases. This, together with the frequent occurrence of price decreases, indicates that there is no clear evidence of downward price stickiness in the Netherlands.

There are some product categories in which price decreases occur almost as often as price increases, namely COICOP 4 (Housing, water, electricity and fuels), COICOP 7 (Transport) and COICOP 9 (Recreation and culture). At the product type level this is the case for unprocessed food, energy and non-energy industrial goods.

The average magnitude of price increases for products in the Dutch CPI is estimated at 11.6% and for price decreases at 15.1%. Large differences between the magnitude of price decreases and price increases can be found in the following COICOP categories: 2 Alcoholic beverages, 3 Clothing and footwear, 4, Housing, water, electricity etc and 9, recreation and culture. Of these categories only COICOP 4 has relatively large price increases whereas the other COICOP categories have relatively large price decreases.

In table 5 the estimated distribution of price changes for the CPI (distribution of price changes for double reweighed products in the sample) is presented. Looking at the frequency distribution it becomes clear that there is quite some variation in the duration of price quotes, ranging from less than a month at the 5th percentile to almost 16 months for the 95th percentile. The CPI reweighed median duration is estimated at 8.7 months. Other Euro area countries findings of CPI representative median duration of price quotes are 6 months for Finland, almost 6 months for France, 6 months for Italy and 8 months for Spain. Belgium has a relatively high median price duration of 13 months. It seems that the speed of price adjustment in the Netherlands is relatively moderate compared to other Euro area countries.

Regarding the magnitude of price changes, results similar to ours have been found for Belgium (Aucremanne et al., 2004), Italy (Fabiani et al., 2004), Portugal (Dias et al., 2004) and Spain (Àlvarez et al., 2004). There, price increases also occur more often than price decreases but the magnitude of price decreases is relatively large. The French results (Baudry et al., 2004) also show a higher frequency of price increases than price decreases, but the magnitude of the price decreases turns out to be relatively small. These results suggest that nominal downward price adjustments are somewhat less common in Europe than nominal upward price adjustments but the magnitude of price decreases might be relatively large, indicating that retailers may have a tendency to postpone price decreases more than price increases.

5 DURATION ANALYSIS

In this section we present the results of a multivariate analysis on product price changes. We use a duration analysis framework in which we focus on the time a product has a particular price. We have adopted the Cox proportional hazard approach to focus on the effects of variables/events on price changes and leave the baseline hazard λ_0 unspecified. We believe that this baseline hazard may behave in a non-monotonous way (having several spikes) and can't be captured by a standard parametric specification, like the exponential or the Weibull. We have estimated a Cox model using the complete sample, excluding first price spells of price trajectories and we have estimated separate Cox models by COICOP group and by product type. The results of the latter regressions enable us to see whether there are differences in the baseline hazards and in the way covariates affect price setting behaviour between COICOP groups/product types.

We have included outlet size and product group dummies, but also time dummies indicating months of the year and time dummies indicating when products faced a change in VAT-rate. We have distinguished between VAT increases and VAT decreases. We have also included month dummies indicating the euro introduction period, July 2001-June 2002 in order to highlight firms' pricing strategy during this period. On top of that we have also added the macro economic variables inflation and wage, (both y-o-y change) to the list of explanatory variables so that macro economic influences do not interfere with the estimated effects for the other variables.

5.1 Exploratory graphs of product type and size effects

Figure 1 shows the distribution of the duration of price quotes for the whole sample of goods and services. This graph suggests that firms are heterogeneous in their price setting behaviour and use a mixture of pricing strategies, both time and state dependent. In order to gain more insight in the existence of multiple pricing strategies we focus in figures 2 and 3 on the distribution of price quote durations of different product types and different outlet sizes. We distinguish the five already mentioned product categories unprocessed food, processed food, non-energy industrials, energy and services and four size categories. Size 0 denotes one man businesses (no employees), size 1 denotes outlets with 1-9 employees, size 2 denotes outlets with 10-99 employees and size 3 denotes outlets with 100 or more employees.

Figure 1 shows that the distribution of price durations in the sample is highly skewed to the left. This is generally in line with the predictions from most theoretical pricing models. Some features in the graph suggest support for specific pricing models. Almost ½ of the prices in the sample lasted only one month and ¾ of the prices in the sample changed within three months time (both findings indicating low menu costs). The fraction of price durations of 7-11 months is quite stable (Calvo) at about 2-3%, followed by a peak for price durations lasting 1 year (truncated Calvo or Taylor)).

Figure 2 shows that prices of energy products change very quickly; after just one month most energy prices have changed. This holds especially for fuel prices. Prices of heating gas change every 6 months (Taylor), but the number of observations of gas prices is rather low. Therefore, their price change peak after 6 months isn't visible in the graph. Prices of unprocessed food change somewhat less often. After one month about 60% of the prices have changed and after three months this figure has increased to 90% of the prices. Unprocessed food includes many seasonal and/or non-storable food products that are sold at auctions to firms. The frequent changes in cost prices are translated into frequent changes in consumer prices. Pricing of these goods are influenced by short-term purchase contracts and low menu costs. This explains the variability of the prices. Prices of unprocessed food and non-energy industrial goods change at an almost similar rate. After one month about 30% of the prices have changed, cumulating to 50-60% of the prices after three months. The distribution of price quote durations for unprocessed food price shows a small peak after 6 and 12 months (truncated Calvo or Taylor). Prices of services change slowest, although 20% of the prices did change already after one month. Here, the second highest peak of price changes occurs after one year! These peaks are examples of time dependency in pricing with prices are maintained for a fixed or a maximum number of months.

Similar patterns of heterogeneity in the distribution of price durations between different product types have also been found in Belgium (Aucremanne et al., 2004), France (Baudry et al., 2004, Fourgère et al., 2004) and Portugal (Dias et al., 2004). This may suggest that the co-existence of firms in a country, which use different pricing strategies, is not just a Dutch or a European phenomena but might also exist in other countries.

From figure 3 it becomes clear that in large outlets prices change more quickly than in smaller firms. This relation has also been found by Portugal (Dias et al., 2004). This may be explained by menu costs decreasing through economies of scale. After one month already 70% of the prices have changed. The distribution of price quote durations for medium sized firms resembles the distribution for large firms quite well, except that 50% of the prices have changed after one month instead of 70%. Prices in small outlets seem to change at a somewhat slower rate than prices in one-man businesses. The distributions of price duration in the three smallest size classes all show a peak after 12 months indicating that in these outlets some of the prices are adjusted only or at most once a year (Taylor or truncated Calvo time dependency).

5.2 Results Cox proportional hazard model

Table 6a displays regression results explaining the duration of price quotes of the whole sample. Tables 6b and 6c show the results of the regressions by COICOP group and product type. Figures 4a-4c show the corresponding estimated baseline survival functions. Variables not included in the regression since they serve as reference variables are the non-energy industrial goods (table 6a) and large outlets. The presented figures under the column headed "hazard ratio" are exponentiated β's. They reflect the proportional changes in the baseline hazard (=conditional probability of not surviving given survival until time t) as a result of the effects of the explanatory variables on the event of interest, i.e duration of a price quote. If a variable does not affect duration, β equals 0 and its exponent equals 1. If a variable increases (decreases) the duration of a price quote, it decreases (increases) the probability of a change in price, resulting in a negative (positive) value of β and a value between 0 and 1 (larger than 1) for its exponent.

Figure 4a shows that the survival function of the whole sample declines sharply during the first months of a price quote. After 1 month, 20% of the price quotes in the sample have changed. After 1 year 80% of the price quotes has changed declining further to over 90% after duration of 2 years. However, the right wing tail of the baseline hazard function seems to be rather thick.

Figure 4b shows the different estimated baseline survival functions by COICOP group. There are clear differences between these baseline survival functions. Food prices (COICOP 1), clothing and footwear prices (COICOP 3) and transport prices including fuel prices (COICOP 7) change very quickly. After half a year less than 20% of the prices hasn't changed yet. Prices of alcoholic drinks (COICOP 2) change much more gradually; after two years about 15% of the prices hasn't changed, just like the product prices in COICOP group 9 (recreation and culture) and 12 (miscellaneous goods and services). Prices of products in the COICOP categories 4 (housing, water, heating gas, etc.), 5 (Furnishings, housing equipment and maintenance), and 11 (Restaurants and hotels) change at an even more modest pace; after three years 20-50% of the prices still hasn't been adjusted. The same picture emerges from graph 6c, showing the estimated baseline survival functions by product type. Prices of unprocessed food and energy change very rapidly whereas prices of services change very gradually. The speed of the price adjustment process of processed food prices and prices of non energy industrial goods lie between these extremes.

Insert figures 4a-4c Estimated baseline survival functions for the whole sample and by COICOP category or product type

Size and product type effects

First we discuss the results in table 6a, after which we turn to regression results by product type/COICOP group. Product type effects are quite pronounced and emphasise what we already saw in figure 2. The hazard ratio of a fuel price change is 2.5 times higher than the hazard ratio corresponding to changes in NEI goods prices and prices of unprocessed food have a 1.7 higher hazard than NEI good prices. At the other end of the price adjustment spectre we find services. Hazard ratios of prices of services are 40% lower than the hazard ratio of NEI goods prices. Prices of unprocessed food and energy excluding fuel have almost equal estimated hazard ratios that lie somewhat below 1, indicating that they have an almost equal hazard for changing price as NEI goods.

Size effects are much smaller than product type effects. The picture emerging from figure 3 is also present here. All three size variables are significantly different from zero, although the accompanying hazard ratios do not differ that much from the benchmark. Small outlets have the smallest and most significant estimated hazard ratio of 0.8, indicating that the conditional probability of a change in price quote in a small outlet is 20% lower than in a large outlet. Medium sized outlets adjust prices with a 10% lower hazard than large outlets and in one man businesses prices are adjusted with only a marginally (4.5%) lower hazard than in large outlets. It seems plausible that larger firms change prices more often than smaller firms, because of menu costs. Menu costs may decline by firm size because of economies of scale. The smallest firms may be very flexible in price setting their products because they are so small and may offer custom made goods and services. The owner is then free to set a new price for each good or service sold.

Table 6b shows that the size effects estimated for each COICOP group separately may differ a lot from the size effects shown in table 6a. For the COICOP groups 2 (alcoholic drinks), 4 (housing water, heating and gas), 7 (transport), 9 (recreation and culture), 11 (restaurants and hotels) and 12 (miscellaneous goods and services) and the product types unprocessed food, processed food, energy and services, large firms have higher hazards than the other firms and there is a clear positive relation between the outlet size and the conditional probability to change prices. For COICOP groups 1 (food and non-alcoholic drinks), 3 (clothing and footwear) and especially 5 (furnishings, household equipment and maintenance of the house) and for NEI goods in table 6c, the one man businesses have a higher hazard than the large outlets. This latter result supports the idea that the smallest firms selling COICOP 1, 3 or 5 products, set prices very frequently. Some of them may even set prices individually for each product sold.

Changes in prices and wages

Between 1998 and 2003 the Dutch economy experienced both a peak and a trough in economic growth. Contractual wages increased by 2.5 to 4.6% between 1998 and 2003 and experienced the highest growth in 2001. The Dutch CPI was below 2% in 1998 and 1999 and peaked in 2001 when it was about 4.5%. The most important contributor to inflation in 2001 was unit labour costs with a contribution to CPI inflation amounting to 2.6 percentage points.

We have included both contractual wage growth and inflation on the list of explanatory variables. Significant results for these variables indicate that time-dependent pricing rules cannot account for all observed pricing behaviour and state-dependent pricing rules should also be considered. Portugal (Dias et al., 2004) and Spain (Álvarez et al., 2004) found that periods of high inflation in their countries were also characterised as periods with frequent price changes, which indicates that price setting by firms in these countries is affected by general inflation.

In the regressions wages turn out to be significant. A one percentage point higher yearly wage rise increases the conditional probability to change prices by 12%. On the whole, the duration of price quotes doesn't seem to be affected by the general inflation in consumer prices. This may be due to the still relatively low and stable inflation rate during the sample period. We have explored the possibility of multicollinearity between wage growth and price inflation. In estimations in which we didn't include wage growth as a covariate the magnitude of the effect of CPI inflation on the duration of price quotes hardly altered and only became mildly significant. So, it doesn't seem that wages caught up the effect of inflation on the duration of price quotes

The effect of wages and inflation on price duration differs with COICOP group and product type. Prices of processed food, energy and notably services react relatively strong on wage changes This holds especially for COICOP group 4 (housing, water, heating gas etc.), 5 (furnishings, household equipment and maintenance of the house), 11 (restaurants and hotels) and 12 (miscellaneous goods and services). The general inflation level has a positive significant effect on the hazard of transport related goods and services. However, also for these products the effect of wages on the hazard is stronger.

Summarising, based on the Cox regression using information of all products in the common sample we can't confirm that the probability of a price change increases with general inflation. However, we have found evidence for the Netherlands that the probability of price changes increases with wage growth, which was one of the main contributors to Dutch inflation in the period 1998-2002. Furtermore, we have also found that in certain sectors general inflation is a significant factor. In our view, these findings support the importance of menu costs in price setting.

Changes in VAT

In the period 1998-2003 some changes in consumer price increasing taxes (like VAT) occurred. They are listed below. The increase in VAT rates from 17.5% to 19% in 2001 contributed a full percentage point to the CPI inflation that year.

January 1999:	Increase tax on energy
January 2000:	Increase tax on energy
	Change from high (17.5%) to low (6%) VAT tariff for labour intensive services, like hairdressing
January 2001:	Increase high VAT tariff from 17.5% to 19%
April 2002:	Increase tariff excise duty for alcoholic drinks

Since changes in VAT rate are announced in advance firms may adjust prices gradually, or not all at the same time, stretching the effect of a change in VAT rate on prices over time. Therefore, we have included six dummies reflecting changes in VAT, three for increases in VAT and three for decreases in VAT. One of each threesome equals 1 a month before the change in VAT takes place, another one equals 1 in the month of the change in VAT rate and the last one equals 1 one month after the change in VAT. A significant coefficient for (at least one of) these dummies shows that some firms used a state dependent pricing strategy.
Note that only two products in our sample, haircuts for men and haircuts for women, got a decrease in VAT during the observation period. It is unknown to what extent the price setting behaviour of hairdressers reflects the price setting behaviour of other firms in case of a VAT decrease.

The estimated effects in table 6a show that a change in VAT results in an increased hazard ratio in the month the VAT change takes place. These results may be interpreted as follows: a VAT increase/decrease leads to an increased probability to change a price. On the whole, firms do not seem to spread their price changes due to changes in VAT rates over time. In case of an increase in VAT rate, the hazards of the preceding and the following month are not affected. Looking at the results of the individual COICOP groups and product types we see that in services, and more specifically in transport and in recreation and culture, the passing through of the increase in VAT rate in consumer prices is spread over two months.

When a VAT decrease takes place the hazards of the two surrounding months are much lower than usual (table 6a and table 6b, COICOP 12). This indicates that the hairdressers adjusted their prices when the decrease in VAT rate became effective. The results in table 6b show that the month in which the VAT rate changed (January) is a month in which haircut prices are usually adjusted.

Table 7 presents some price statistics which give an indication of the magnitude of the price changes. It shows the average price of hairdressing from January 1999 until January 2003 and the change in price compared to the December price of the previous year. By comparing these statistics for 1999-2003 we may get an idea on how changes in VAT are passed through via product prices.

A VAT decrease only occured in hairdressing (so the results on the effects of a VAT decrease should be treated cautiously since it concerns just two products in our sample). In general, prices increase 4-5% each year. However, in 2000 prices decreased by 2% compared to the December 1999 prices and, roughly speaking, they increased 7% less than in the other years of the observed period; this suggests that the hairdressers shared the decrease in VAT approximately at a 50%-50% base with their customers.

The effect of a VAT increase on prices is not easily to deduce from table 7. We focus on the general VAT increase from 17.5% to 19% on January 2001. A problem is that in January prices usually fall due to the winter sales. However, in 1999 and 2000 we observe an about 1% lower price in January than in December, whereas in 2001 prices went up with 0.3%, suggesting a 1.3% higher price increase than in the two previous years. This indicates that the 1.5%-point increase in VAT was almost completely passed through in the consumer prices, whereas we see a sharing of the benefits of 11.5% arising from the VAT decrease.

Spain (Álvarez and Hernando, 2004) and Belgium (Aucremanne and Dhyne, 2004) also examined the effects of changes in VAT rate on pricing. Just like in the Netherlands, changes in VAT-rate or excise duties had a clear upward effect on the frequency of price increases in Belgium and Spain. Álvarez and Hernando also examined the impact of changes in VAT-rate and excise duties on the size of price changes and found that these changes didn't affect the size of price changes very much. It would be quite interesting to learn more about the effects of changes in VAT rates on prices in the other countries.

Euro-conversion

Half a year before the euro conversion the majority of Dutch citizens had good faith in the euro according to a study by Van Renselaar and Stokman (2001). They felt well informed about the cash changeover and the euro itself. However, most people also expected that some retailers might take advantage of the fact that Dutch consumers were not used to euro prices and would raise their prices. The guilder/euro exchange rate was set at 2.20371 and consequently the new euro prices 'looked' very low in comparison to the old guilder prices (money illusion). One of the measures agreed by consumer and retail organisations to give people time to get used to the euro and to unmask price increases was double product pricing, with both guilder and euro prices. This period of double pricing was from July 2001 until February 2002.

Some of the time dummies in the Cox regression shown in table 6a-6c clearly affect price duration. The parameter estimate of the December 2001 dummy shows a doubled hazard ratio[13], one month before the euro replaced the guilder. For the COICOP categories 4 (Housing, water, gas, etc.), 5 (household equipment and maintenance) and 9 (recreation and culture) the increase in the hazard ratio was even higher. In March 2002 the hazard was almost 20% higher than normal, with peaks again for the COICOP categories 5 and 9. Apparently, prices may have increased after the dual display was removed, not allowing consumers anymore to compare the 'old' and the 'new' currency. In the months just after December 2001 and March 2002 we observe less price changes than expected. There are also some categories in which the euro conversion period seemed to had less or no impact on the pricing of products, namely COICOP categories 2 (alcoholic drinks) and 7 (transport) and the product type energy. Prices of food, NEI goods and services also changed relatively often at the end of 2001 and at the end of the first quarter of 2002.

Another striking result is shown in table 7. The January 2002 prices of high VAT products were 0.6% higher than the December 2001 prices, whereas in 1999, 2000 and 2003 January prices of high VAT products were on average lower than their December prices in the year before due to winter sales! This indicates that the guilder-euro conversion may have triggered upward effects on prices.

These results suggest that the pricing strategy of retailers was different during the introduction of the euro than before, suggesting state dependency in pricing behaviour of some of the retailers. In order to shed some more light into the pricing behaviour strategy of retailers during the introduction of the euro we also compare pricing statistics during the introduction of the euro (July 2001-June 2002) with pricing statistics just before the introduction (January 2000- June 2001). The results of this 'back of the envelope' exercise are shown in table 8.

The statistics $\Delta P_j^+ F_j^+$ and $\Delta P_j^- F_j^-$ in table 8 reflect average monthly price increases and decreases. The net monthly price change equals $\Delta P_j^+ F_j^+ - \Delta P_j^- F_j^-$. The bottom part of the table shows the ratio euro introduction statistics over the pre euro introduction statistics. We give these statistics for the five main product categories in order to focus on the general picture and not too much on details at the product level. A ratio larger than one for frequency (magnitude) of price change indicates that the monthly frequency of price changes (change in prices) was larger during the introduction of the euro than in the period just before. The ratios for processed food, non energy industrial goods and services are most important for our analysis, since they give an indication of the effect of the euro introduction on core inflation. The ratios for energy and unprocessed food may be a bit clouded due to developments on the oil market, euro/dollar exchange rate, crop failures and cattle diseases.

We see that the frequencies of price changes were higher during the introduction period than before. This holds especially for NEI goods and services. However, not only the frequency of price increases was higher, also the frequency of price decreases increased, except for processed food. Fabiani et al. (2004) have similar findings for Italy. They also examined pricing behaviour of Italian firms during the euro cash changeover. They found that during the first quarter of 2002 the share of prices that changed was around 20% higher than in previous years.

The higher frequency of price increases is partly compensated by smaller price increases. The combination of more but smaller price changes may be explained partly by rounding the new euro prices to the nearest psychologically attractive euro price. Price increases of NEI goods and services were about 20% smaller than before. Processed food is an exception with 25% higher price increases during the euro introduction period. The net monthly price increase ratio was positive for all sectors except energy with non energy industrial goods taking the biscuit with a ratio of 1.6. However, in this sector and in services the frequency of price decreases also increased considerably. Only the processed food sector faced less monthly price decreases. The ratio for average monthly price decreases was positive for all sectors, with again NEI goods having the largest ratio (1.65) and processed food the smallest (1.06). Overall, the ratio for net monthly price changes was larger than 1 in the processed food sector (2.6), NEI sector (1.2) and in services (1.3), indicating that during the introduction of the euro monthly price change of products included in core inflation were higher than during the pre euro introduction period. This finding is supported by a study of the Nederlandsche Bank, conducted by Folkertsma, Van Renselaar and Stokman and reported in DNB's quarterly bulletin of March 2002. The study shows that in the Netherlands, on average, retail prices went up by 0.5-0.9 percentage point as a result of the changeover (passing on euro conversion costs for retailers to consumers and rounding prices up to attractive psychological prices) and the Dutch CPI by 0.2-0.4 percentage point.

[13] In a regression without the inclusion of wage growth as an explanatory variable the parameter estimate of the December 2001dummy indicated a tripled hazard to change a price, because it also picked up the effect of increasing labour costs on pricing.

6 SUMMARY AND CONCLUSIONS

This paper presents the results of a study on pricing behaviour of retail firms in the Netherlands in the period 1998-2003 using a large micro dataset with monthly price quotes of 49 products, having a total weight of 8% in the Dutch CPI. It has been conducted as part of the Eurosystem Inflation Persistence Network (IPN). We also assess the effects of outlet and product group characteristics on the duration of price quotes. Furthermore, next to time dependent pricing strategies we pay attention to the occurrence of state dependent pricing strategies by assessing the effects of wage growth, the introduction of the euro cash and changes in VAT on prices. Most of the Dutch results are consistent with results found for other euro area countries participating at the IPN.

The average price duration in the Netherlands is 9.7 months and the median duration is 8.7 months, which is somewhat longer than in other Euro area countries. However, there is much variation in price duration across product types and across outlet sizes. Price increases occur more often than price decreases, but the difference in occurrence is rather small, indicating that nominal prices are not downward rigid. On average, the magnitude of price decreases is somewhat higher. This picture also emerges in other European countries.

Product prices change most frequently in the energy (fuel prices change every month) and in the unprocessed food sector (every three months), whereas prices of non energy industrial goods and services change about once a year. These sector effects are significant. The result for services is a clear example of firms using time dependent pricing strategies. There are also significant differences in the duration of price spells across outlets of different sizes. Price adjustment is fastest in large firms and slowest in small firms. Remarkable is that price adjustments in one-man businesses take place almost as often as in the large firms.

Regarding changes in VAT rates, it seems there is some asymmetry in price adjustments. Changes in VAT shorten the duration of price spells. This holds both for increases and decreases in VAT and reveals that with respect to VAT some firms use a state dependent pricing strategy. Yet, an increase in VAT seems to be completely passed on to consumers, but a decrease in VAT only partially. However, evidence for the latter claim is somewhat limited.

Another interesting finding regarding state dependent price effects is that during the euro cash changeover the frequency of price changes, both increases and decreases in price, was higher than in the period before the introduction. This holds especially for NEI goods and services. Generally, the magnitude of the price increases was somewhat smaller during the changeover period than before this period. The magnitude of the price decreases differed less. There are some indications that for certain product groups inflation may have been relatively high during the introduction of the euro.

The finding that Dutch price setters follow both time- and state-dependent pricing strategies suggests that macroeconomic models for monetary policy should combine both price adjustment mechanisms. Developing these hybrid models may lead to significantly better models of the monetary transmission mechanisms.

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APPENDIX

A. Definition pricing statistics

The following definitions used for constructing the pricing statistics are from Dhyne (2003). We define the following binary variables:

Price available at t:

$$DEN_{ijt} = 1 \quad \text{if } P_{ijt} \text{ and } P_{ij,t-1} \text{ are observed or if a forced product replacement occurs in t} \qquad \text{(a.1)}$$
$$0 \text{ otherwise}$$

Price change at t:

$$NUM_{ijt} = 1 \quad \text{if } P_{ijt} \neq \mathrm{P}_{ij,t-1} \text{ or if a forced product replacement occurs in t} \qquad \text{(a.2)}$$
$$0 \text{ otherwise}$$

And more specifically we distinguish between price increases and price decreases

Price increase at t:

$$NUMUP_{ijt} = 1 \quad \text{if } P_{ijt} > P_{ij,t-1} \text{ or if a forced product replacement occurs in t} \qquad \text{(a.3)}$$
$$0 \text{ otherwise}$$

Price decrease at t:

$$NUMDW_{ijt} = 1 \quad \text{if } P_{ijt} < P_{ij,t-1} \text{ or if a forced product replacement occurs in t} \qquad \text{(a.4)}$$
$$0 \text{ otherwise}$$

Based on these binary variables the following pricing statistics can be constructed

Frequency of price changes: $$F_j = \frac{\sum_{i=1}^{n_j} \sum_{t=2}^{t} NUM_{ijt}}{\sum_{i=1}^{n_j} \sum_{t=2}^{t} DEN_{ijt}} \qquad \text{(a.5)}$$

Frequency of price increases: $$F_j^+ = \frac{\sum_{i=1}^{n_j} \sum_{t=2}^{t} NUMUP_{ijt}}{\sum_{i=1}^{n_j} \sum_{t=2}^{t} DEN_{ijt}} \qquad \text{(a.6)}$$

Average price increase in p.c. $$\overline{\Delta}_j^+ = \frac{\sum_{i=1}^{n_j} \sum_{t=2}^{t} NUMUP_{ijt} \left(\ln P_{ijt} - \ln P_{ij,t-1} \right)}{\sum_{i=1}^{n_j} \sum_{t=2}^{t} NUMUP_{ijt}} \qquad \text{(a.7)}$$

Frequency of price decreases: $$F_j^- = \frac{\sum_{i=1}^{n_j}\sum_{t=2}^{t} NUMDW_{ijt}}{\sum_{i=1}^{n_j}\sum_{t=2}^{t} DEN_{ijt}} \quad \text{(a.8)}$$

Average price decrease in p.c.: $$\overline{\Delta}_j^- = \frac{\sum_{i=1}^{n_j}\sum_{t=2}^{t} NUMDW_{ijt}\left(\ln P_{ij,t-1} - \ln P_{ijt}\right)}{\sum_{i=1}^{n_j}\sum_{t=2}^{t} NUMDW_{ijt}} \quad \text{(a.9)}$$

Frequency of price changes at t: $$F_{jt} = \frac{\sum_{i=1}^{n_j} NUM_{ijt}}{\sum_{i=1}^{n_j} DEN_{ijt}} \quad \text{(a.10)}$$

Similar expressions can be derived for the frequency of price increases (F_{jt}^+) or decreases (F_{jt}^-) at time t.

B. Sensitivity analysis Cox regression model

In this appendix we present some robustness checks on the regression results discussed in section 5. We compare the results of alternative specifications with the regression results shown in table 6a. We have adopted the following three specifications. Log likelihood ratio outcomes on homogeneity tests are reported in table 9.

Alternative specifications:

1) Exclusion of the month dummies (January,…,December)
2) Inclusion of 48 product dummies
3) Exclusion of four out of the six VAT dummies (i.e. the dummies equal to one a month before and a month after the change in VAT-rate

The month dummies are statistically significant in the model, which indicates that they should be included in the set of covariates. They reveal the seasonal patterns in price setting behaviour of firms (sales, introduction new collection of goods and services, etc.). They have been included in the analysis to facilitate the examination of the price setting behaviour of firms during the introduction of the euro, by removing any seasonal effects from the euro conversion parameters. In the regression without month dummies the estimated parameter for December 2001 equals 3.3 (t-value 57.0) instead of 2.3 (t-value 27.5) and for April equals 0.9 (t-value 4.1) instead of 1.3 (t-value 7.7). Without the month dummies the December 2001 effect is overestimated whereas the April 2003 (dual pricing just ended) is highly underestimated, due to the usual quietness in price setting in April. Their inclusion also affects the estimates of the VAT-dummies. Other parameters hardly changed.

The parameters reflecting product specific effects are also jointly significant. Parameters for fresh fish, beer in a shop, domestic services and fuel could not be estimated due to collinearity. The reference product is socks. The products with the highest estimated hazards are 'replacement of brake blocks' and 'car service labour charge', both with hazards twice as high as the reference group. The two products with the lowest hazards are sugar and cement, both having an estimated hazard twice as low as the reference group. The parameters of the other covariates are hardly affected by the inclusion of the product specific parameters. The estimated effect of services is with 0.4 30% lower than in the specification without the product dummies. We decided not to include all the product specific dummies in the set of covariates. Instead, we thought it more informative to present regression results by COICOP group and product type (table 6b and 6c).

The four additional 'change in VAT-rate dummies', which show the effects the month before and the month after a change in VAT- rate on the duration of price quotes, are also jointly significant. As a result of this test we have included them in the set of covariates. Their inclusion hardly altered the estimated parameters of the other covariates.

Figure 1 Duration until price change (in months)

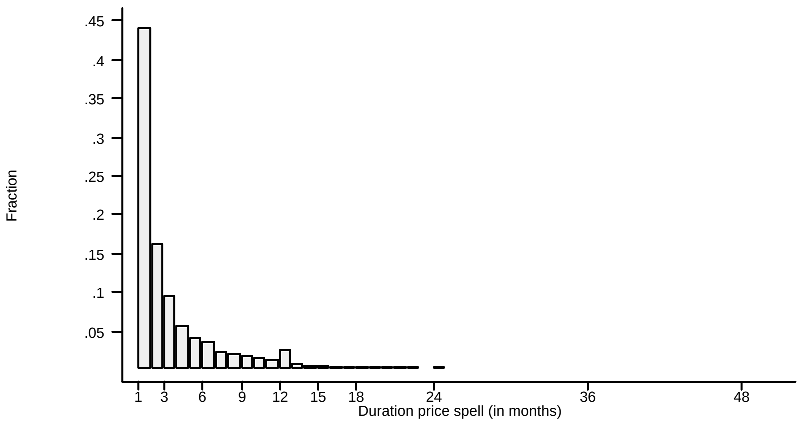

Figure 2 Duration until price change by product type (in months)

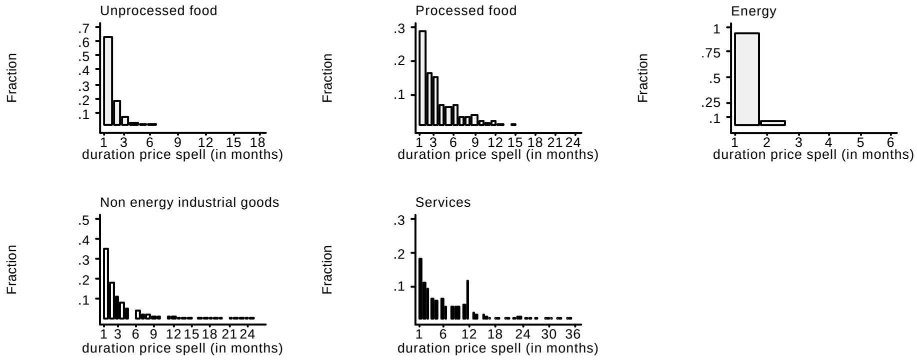

Figure 3 Duration until price change by outlet size (in months)

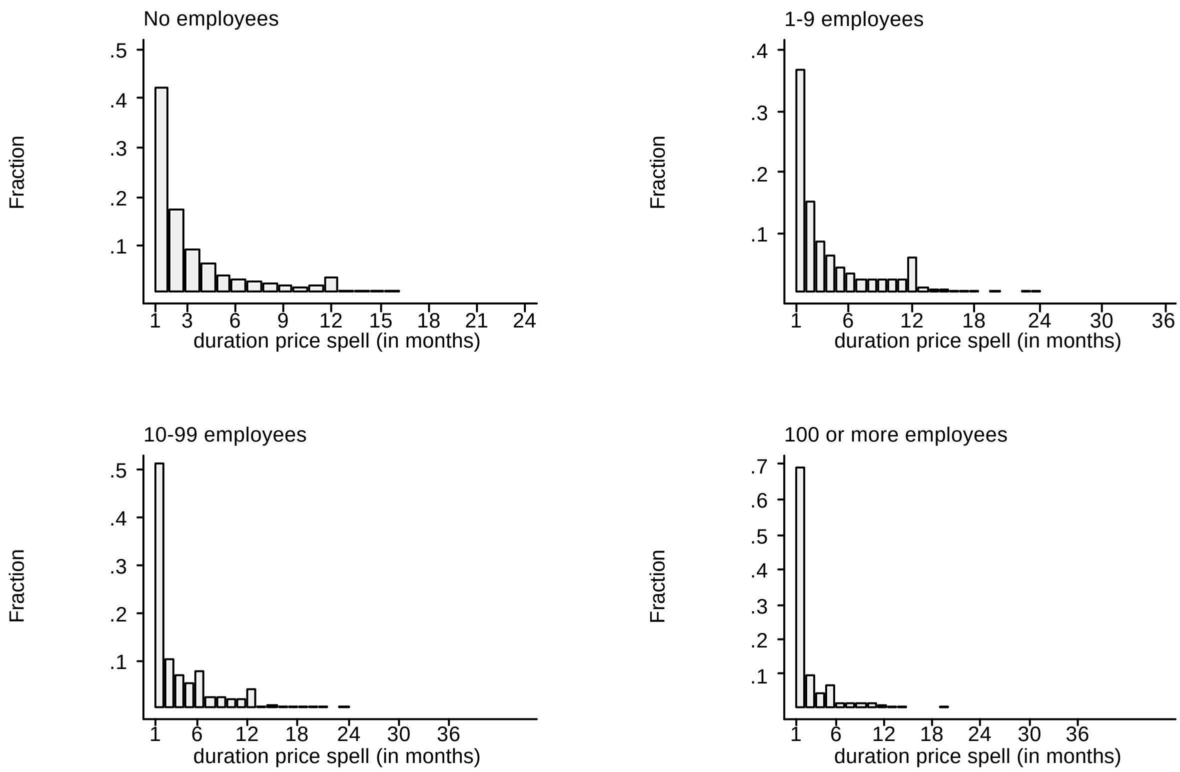

Figure 4a Estimated survival function, whole sample

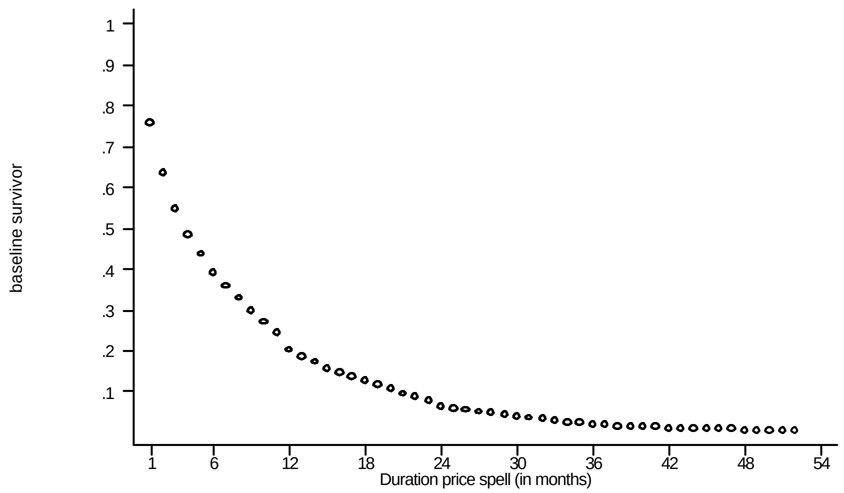

Figure 4b Estimated survival function by COICOP group

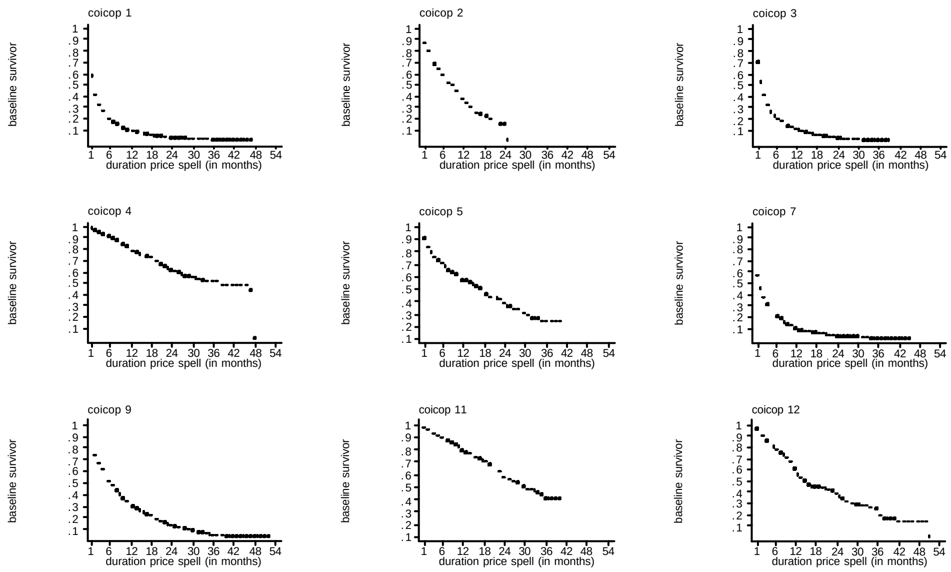

Figure 4c Estimated survival function by product type

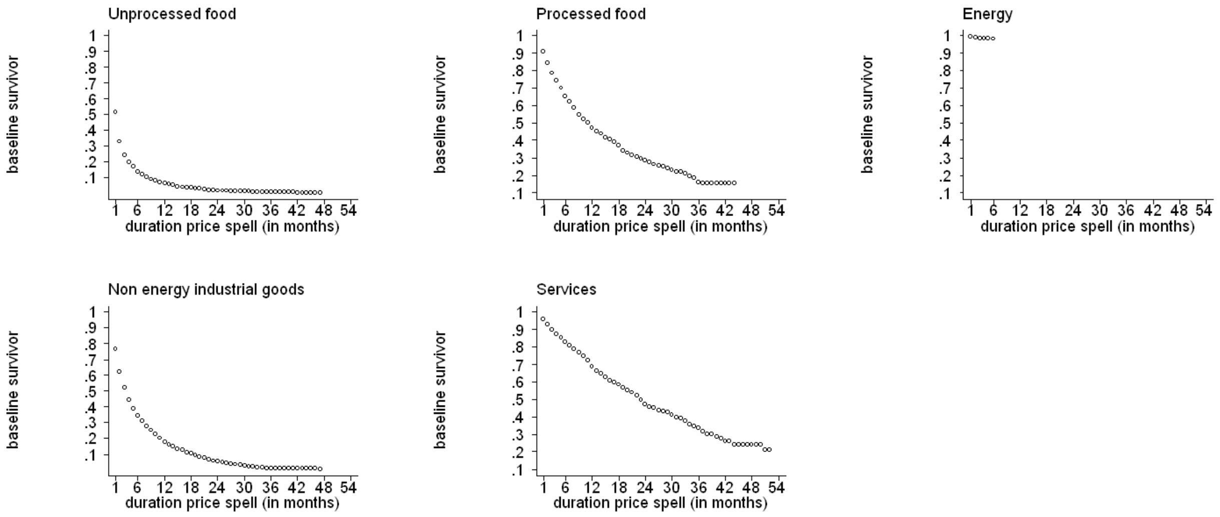

Table 1: Information available in the data base (metadata)

article code	Each product has a 5 digit code
article name	Name of the product
outlet number	Each outlet has a numeric code
date (year+month)	Date of the observation
quantity	
price	Product price
Correction quality code	Dummy indicating a change in product quality
CPI weight	Weight of product in CPI basket (5 digit), base year 2000
product code (COICOP)	Product code according to the COICOP classification
outlet code classification	SBI classification outlet according to Statistics Netherlands
outlet size	Code indicating size class of the outlet according to Statistics Netherlands
interviewer	numerical code i
Price not observed	Dummy variable indicating that price of article was not observed
Non-durables	Dummy variable
Semi-durables	Dummy variable
Durables	Dummy variable
Non-energy industrial	Dummy variable
Electricity, Gas	Dummy variable
Liquid fuels	Dummy variable
Energy	Dummy variable
Industrial goods	Dummy variable
Processed Food	Dummy variable
Seasonal food	Dummy variable
Meat	Dummy variable
Unprocessed food	Dummy variable
Food	Dummy variable
Goods	Dummy variable
Services	Dummy variable

Table 2 Summary statistics

COICOP	Product type	Article name	cpi weight year 2000	#trajectories	# price spells	# right cens. spells	# left cens.spells	#obs.
1	UPF	Steak	0.644%	143	1099	74	143	5493
1	UPF	1 fresh fish	0.515%	711	2463	99	711	6524
1	PF	Milk	2.059%	95	692	36	95	3583
1	UPF	Banana	1.068%	217	3756	82	217	7997
1	UPF	Lettuce	0.154%	242	6125	88	242	8456
1	UPF	Frozen spinach	0.425%	178	1061	62	178	6269
1	PF	Sugar	0.798%	95	472	41	95	3710
1	PF	Coffee	1.686%	99	929	44	99	3840
1	PF	Mineral water	0.283%	132	849	55	132	4974
2	PF	Liquor	0.232%	82	1015	65	82	3655
2	PF	Beer in a shop	2.342%	87	758	46	87	3719
3	NEI	Socks	0.386%	209	1550	149	209	8315
3	NEI	Jeans	1.467%	172	1216	132	172	4736
3	NEI	Shirt (men)	0.708%	241	3362	177	241	10276
3	S	Dry cleaning	1.596%	46	275	40	46	2357
3	NEI	Sport shoes	0.605%	182	1545	130	182	7358
4	NEI	Acrylic painting	0.592%	134	705	69	134	4212
4	NEI	Cement	0.296%	146	674	99	146	5893
4	S	Hourly rate of a carpenter	0.605%	57	306	35	57	2090
4	S	Hourly rate of a plumber	1.223%	74	384	49	74	2676
4	E	Gas (heating)	16.190%	1	5	0	1	21
5	NEI	1 type of furniture	0.541%	102	269	22	102	1432
5	NEI	Towel	0.386%	161	1135	93	161	5700
5	NEI	Coffee-maker	0.322%	313	771	56	313	2752
5	NEI	Electric bulb	0.219%	87	373	60	87	3501
5	S	Domestic services	5.586%	82	234	42	82	2586
7	NEI	Car tyre	0.322%	210	1549	123	210	6750
7	E	Fuel type 1	35.650%	31	905	30	31	960
7	E	Fuel type 2	4.801%	31	877	30	31	960
7	S	Car service labour charge	1.274%	53	387	29	53	2008
7	S	Car wash	0.759%	106	396	78	106	3864
7	S	Replacement of brake blocks	0.502%	53	450	29	53	2041
7	S	Taxi	0.322%	49	277	35	49	2251
8	NEI	Fax machine		#N/B	#N/B	#N/B	#N/B	#N/B
9	NEI	Television set	0.502%	469	1011	28	469	2979
9	NEI	Construction game	0.965%	90	429	48	90	2650
9	NEI	Football	0.644%	440	706	75	440	5323
9	S	Dog food	0.528%	164	918	88	164	5955
9	S	Movie	0.425%	68	286	56	68	2375
9	S	Videotape hiring	0.103%	52	171	46	52	2496
9	S	Photo development	0.644%	88	432	70	88	3706

Table 2 continued

COICOP	Product type	Article name	cpi weight year 2000	#trajectories	# price spells	# right cens. spells	# left cens.spells	#obs.
11	S	Glass of beer in a café	1.441%	132	364	88	132	3585
11	S	1 meal in a restaurant	0.772%	133	408	64	133	3665
11	S	Snack	1.030%	73	287	60	73	3274
11	S	Glass of cola in a café	1.107%	126	370	92	126	3871
11	S	Hotel room	0.901%	274	1102	125	274	6976
12	S	Haircut (men)	2.033%	116	630	91	116	5213
12	S	Hairdressing (ladies)	3.256%	114	626	87	114	5065
12	NEI	Toothpaste	0.450%	104	796	51	104	3851
12	NEI	Suitcase	0.644%	150	297	33	150	2461
		sample	**100%**	**7214**	**45697**	**3301**	**7214**	**204404**

Table 3 Monthly frequency of price changes, magnitude of price changes, median and mean duration of prices

COICOP	Article name	Freq.of price changes (per month)	Implied median duration (in months)	Implied average duration (in months)	Freq. price increases r.t. freq price changes	Average price increase (in p.c.)	Freq. price decreases r.t freq. price changes	Average price decrease (in p.c.)
1	Steak	0.179	3.514	5.070	65	10.8	35	17.0
1	1 fresh fish	0.301	1.936	2.792	59	28.9	41	35.5
1	Milk	0.171	3.696	5.332	71	9.6	29	13.7
1	Banana	0.455	1.142	1.648	55	25.7	45	28.8
1	Lettuce	0.716	0.551	0.794	51	36.3	49	36.7
1	Frozen spinach	0.145	4.425	6.384	59	28.7	41	40.6
1	Sugar	0.104	6.312	9.106	68	4.9	32	7.9
1	Coffee	0.222	2.761	3.984	47	6.6	53	6.6
1	Mineral water	0.148	4.328	6.243	60	15.5	40	18.0
2	Liquor	0.261	2.292	3.306	64	8.1	36	10.7
2	Beer in a shop	0.185	3.388	4.888	72	6.1	28	10.9
3	Socks	0.165	3.844	5.546	54	22.3	46	25.7
3	Jeans	0.229	2.665	3.845	58	18.8	42	23.9
3	Shirt (men)	0.311	1.861	2.684	50	33.6	50	33.0
3	Dry cleaning	0.099	6.649	9.592	91	5.5	9	8.1
3	Sport shoes	0.190	3.289	4.746	56	20.8	44	25.5
4	Acrylic painting	0.140	4.596	6.630	71	13.3	29	22.6
4	Cement	0.092	7.182	10.362	90	5.4	10	19.2
4	Hourly rate of a carpenter	0.122	5.327	7.686	84	5.2	16	5.8
4	Hourly rate of a plumber	0.119	5.471	7.893	87	4.8	13	5.9
4	Gas (heating)	0.200	3.106	4.481	50	10.0	50	1.9
5	1 type of furniture	0.126	5.147	7.425	56	15.0	44	23.6
5	Towel	0.176	3.581	5.166	57	24.6	43	28.6
5	Coffee-maker	0.188	3.328	4.802	55	8.3	45	8.1
5	Electric bulb	0.084	7.900	11.397	69	20.8	31	35.6
5	Domestic services	0.061	11.013	15.888	90	8.0	10	7.8
7	Car tyre	0.205	3.021	4.359	63	11.9	37	10.0
7	Fuel type 1	0.941	0.245	0.353	52	2.7	48	3.1
7	Fuel type 2	0.911	0.287	0.413	45	2.9	55	3.4
7	Car service labour charge	0.171	3.696	5.332	79	10.0	21	15.7
7	Car wash	0.077	8.651	12.480	84	20.8	16	12.1
7	Replacement of brake blocks	0.200	3.106	4.481	79	6.7	21	9.2
7	Taxi	0.104	6.312	9.106	80	6.8	20	7.7
8	Fax machine							
9	Television set	0.216	2.848	4.109	42	8.4	0.125	8.6
9	Construction game	0.132	4.896	7.064	63	14.3	0.049	28.6
9	Football	0.054	12.486	18.014	56	17.9	0.024	28.1
9	Dog food	0.130	4.977	7.181	68	18.8	0.041	27.4
9	Movie	0.094	7.022	10.130	76	6.7	0.023	7.2
9	Videotape hiring	0.049	13.796	19.904	77	15.3	0.011	29.0
9	Photo development	0.095	6.944	10.018	57	14.0	0.041	17.8

Table 3 continued

COICOP	Article name	Freq.of price changes (per month)	Implied median duration (in months)	Implied average duration (in months)	Freq. price increases r.t. freq price changes	Average price increase (in p.c.)	Freq. price decreases r.t freq. price changes	Average price decrease (in p.c.)
11	Glass of beer in a café	0.067	9.995	14.420	93	8.4	7	10.5
11	1 meal in a restaurant	0.078	8.535	12.314	79	7.7	21	13.4
11	Snack	0.067	9.995	14.420	87	9.4	13	16.7
11	Cola in a café	0.065	10.313	14.879	97	7.9	3	8.2
11	Hotel room	0.124	5.236	7.553	79	8.4	21	11.3
12	Haircut (men)	0.101	6.510	9.392	84	6.0	16	5.9
12	Hairdressing (ladies)	0.103	6.377	9.200	80	8.4	20	10.5
12	Toothpaste	0.185	3.388	4.888	64	9.9	36	12.8
12	Suitcase	0.064	10.480	15.119	50	9.9	50	8.9

Table 4 Pricing statistics by COICOP classification and product type

	Frequency of price changes (in %)	Average duration of price (months)	Freq. price increases r.t. freq price changes	Average price increase (in p.c.)	Freq. price decreases r.t freq. price changes	Average price decrease (in p.c.)
By COICOP						
1 Food and non- alcoholic beverages	23.23	4.72	58	13.88	42	17.50
2 Alcoholic beverages	19.16	4.75	72	6.29	28	10.87
3 Clothing and footwear	20.52	5.10	58	19.41	42	22.76
4 Housing, water, electricity gas and other fuels	18.87	4.96	53	9.53	47	3.23
5 Furnishings, household equipment and routine maintenance of the house	7.85	14.07	78	9.87	22	11.02
7 Transport	87.98	0.86	51	3.38	49	3.77
8 Communication						
9 Recreation and culture	7.92	14.63	58	16.29	42	25.47
11 Restaurants and hotels	7.79	13.02	86	8.39	14	11.80
12 Miscellaneous goods and services	10.43	9.55	77	7.88	23	9.04
By Product type						
Unprocessed food	32.42	3.39	57	23.34	43	28.99
Processed food	18.17	5.26	64	7.48	36	10.63
Energy	72.65	1.54	51	4.79	49	2.76
Non energy industrial goods	12.35	11.26	57	17.15	43	24.56
Services	9.33	11.43	83	8.55	17	10.17
Total representing CPI (weighted twice)	16.52	9.71	63	11.58	37	15.11

Table 5: Frequency of price changes and price duration

	Sample (weighted once)	Double weighted sample, reflecting CPI basket
Monthly frequency of price changes		
5th percent	0.054	0.054
25th percent	0.126	0.099
Median	0.190	0.179
75th percent	0.911	0.200
95th percent	0.911	0.911
Duration of prices in months		
5th percent	0.120	0.120
25th percent	0.120	3.670
Median	3.670	8.660
75th percent	11.110	12.250
95th percent	15.690	15.690

Table 6a Cox regression results including changes in VAT and month dummies during introduction of the euro, explaining duration until price change (in months, robust standard errors)

No. of subjects = 38483
No. of failures = 35458

Log likelihood = -341995.89

Number of obs = 151920

Wald chi2(35) = 15853.58
Prob > chi2 = 0.0000

	Est. hazard ratio	Standard error	z	p-value
January	1.354	0.046	9.000	0.000
February	1.232	0.041	6.240	0.000
March	1.172	0.040	4.620	0.000
April	0.773	0.028	-7.040	0.000
May	1.094	0.039	2.530	0.012
June	1.236	0.043	6.100	0.000
July	1.207	0.043	5.260	0.000
August	1.229	0.043	5.830	0.000
September	1.339	0.046	8.580	0.000
October	1.167	0.041	4.430	0.000
November	1.112	0.030	3.920	0.000
December	1.726	0.048	19.550	0.000
July 2001	0.852	0.035	-3.900	0.000
August 2001	0.872	0.035	-3.410	0.001
September 2001	0.847	0.031	-4.480	0.000
October 2001	0.945	0.037	-1.460	0.145
November 2001	1.081	0.043	1.960	0.050
December 2001	2.261	0.068	27.310	0.000
January 2002	0.756	0.027	-7.890	0.000
February 2002	0.965	0.031	-1.110	0.269
March 2002	1.174	0.039	4.780	0.000
April 2002	1.348	0.052	7.720	0.000
May 2002	1.006	0.038	0.160	0.874
June 2002	1.559	0.050	13.770	0.000
Vat increase next month=1	0.945	0.043	-1.250	0.211
Vat increase this month=1	1.682	0.056	15.500	0.000
Vat increase previous month=1	1.004	0.042	0.090	0.932
Vat decrease next month=1	0.228	0.112	-3.000	0.003
Vat decrease this month=1	3.056	0.270	12.620	0.000
Vat decrease previous month=1	0.498	0.137	-2.540	0.011
wage_growth	1.120	0.018	7.250	0.000
hicp_growth	0.990	0.006	-1.530	0.126
size0	0.956	0.015	-2.880	0.004
size_small	0.807	0.011	-15.870	0.000
size_med	0.903	0.011	-8.230	0.000
Unprocessed food	1.736	0.022	44.200	0.000
Processed food	0.926	0.014	-4.970	0.000
Services	0.597	0.008	-36.510	0.000
Energy excl. fuel	0.924	0.168	-0.430	0.665
Fuel	2.466	0.041	53.700	0.000

Benchmark: No change in VAT, outlet size is large, product is a NEI good

Table 6b Cox regression results explaining duration until price change by COICOP category (in months, robust standard errors)

	Food and non alc. drinks	Alc. drinks	Clothing and footwear	Housing, water, heating gas,etc.	Furnishings, household equipment and maintenance	Tran-sport	Recrea-tion and culture	Restau-rants and hotels	Misc. goods and services
	Haz. ratio	Haz. ratio	Haz. ratio	Haz. Ratio	Haz. ratio	Haz. ratio	Haz. ratio	Haz. ratio	Haz. ratio
January	1.431*	1.859*	1.204*	3.049*	1.083	0.847	0.985	1.789*	1.874*
February	1.263*	0.857	1.190*	1.825*	0.932	0.993	1.029	1.849*	1.359
March	1.268*	1.042	0.814*	1.794*	1.304	1.088	1.181	2.005*	1.175
April	0.983	0.488*	0.414*	1.190	0.620*	0.584*	0.648*	1.148	0.455*
May	1.198*	0.958	0.770*	1.019	0.880	1.094	1.065	0.960	0.407*
June	1.258*	0.773	1.381*	1.193	1.085	0.768*	1.061	0.817	1.032
July	1.172*	1.448*	1.100	1.630	0.958	0.823*	1.316	1.332	1.589*
August	1.141*	1.502*	1.232*	2.257*	1.126	0.938	0.853	1.700*	1.208
September	1.381*	1.581*	0.960	3.292*	1.361*	1.178*	1.050	1.433	1.260
October	1.328*	1.152	0.743*	1.913*	0.979	1.013	1.023	1.845*	0.867
November	1.010	1.393*	1.083	1.707*	1.101	1.100	0.711*	1.132	1.073
December	1.423*	2.388*	1.408*	2.193*	1.416*	1.706*	1.530*	6.458*	6.266*
July 2001	1.051	0.232*	0.835	0.761	0.812	0.881	0.516*	0.619	0.334*
August 2001	0.969	0.500*	0.998	0.433*	0.616*	0.843	0.952	0.461*	0.476*
September 2001	1.018	0.418*	0.748*	0.543*	0.505*	0.846	0.779	0.479*	0.769
October 2001	1.112*	0.263*	0.783*	0.802	0.576*	0.839	0.779	0.462*	0.965
November 2001	1.230*	0.493*	0.897	0.881	0.670*	0.687*	1.832*	1.401	1.336
December 2001	1.829*	1.347	2.242*	3.839*	2.318*	1.373*	3.264*	2.071*	1.129
January 2002	0.678*	0.299*	0.762*	0.680	0.641*	0.869	0.981	1.029	0.509*
February 2002	0.845*	0.906	0.853*	0.927	1.461*	0.840*	1.296*	0.866	1.065
March 2002	1.320*	1.430	0.890	0.916	0.595*	0.845	0.959	0.935	1.783*
April 2002	1.100	4.159*	1.236	1.097	1.989*	1.330*	1.221	3.081*	1.783*
May 2002	0.996	0.966	0.942	0.891	1.018	0.805*	1.106	1.817*	2.916*
June 2002	1.720*	5.411*	0.995	1.707	1.609*	1.197	1.424*	4.144*	3.282*
Vat increase next month=1		0.593*	0.953	0.846	0.572*	0.827*	0.593*	1.122	3.091*
Vat increase this month=1		1.730*	1.094	2.955*	0.972	1.572*	1.446*	1.599*	0.904
Vat increase previous month=1		0.519*	0.974	0.885	0.592*	1.216*	1.405*	0.994	1.159
Vat decrease next month=1									0.242*
Vat decrease this month=1									0.954
Vat decrease previous month=1									0.475*
Wage growth	1.041	1.395*	1.066	1.447*	1.404*	1.165*	1.193*	1.755*	1.616*
Hicp growth	0.993	0.984	0.998	1.040	1.027	1.066*	0.987	0.931	0.895*
Size 0	1.031	0.657*	1.060*	0.846	1.200*	0.344*	0.687*	0.440*	0.507*
Size small	1.045*	0.690*	0.803*	0.737*	0.656*	0.354*	0.663*	0.339*	0.426*
Size med	0.882*	0.770*	0.878*	0.874	1.009	0.455*	0.805*	0.613	0.591*
no. observations	41188	5818	26579	11440	9058	15982	14736	14004	13115
No. price spells	15534	1504	7098	1552	2037	4308	2582	1793	1865
No. ended price spells	14984	1494	6498	1417	1789	3961	2229	1478	1608
Log likelihood	-133737	-9507	-51996	-8921	-12074	-29741	-15458	-9264	-10189

* indicates significance at the 95% confidence level

Table 6c Cox regression results explaining duration until price change by product type (in months, robust standard errors)

	Unprocessed Food Haz. Ratio	Processed Food Haz. Ratio	Energy Haz. Ratio	NEI Haz. Ratio	Services Haz. Ratio
January	1.328*	2.189*	0.889	1.295*	1.239*
February	1.373*	1.105	0.564*	1.204*	1.137
March	1.341*	1.202	1.443*	0.957	1.356*
April	1.031	0.783	0.205*	0.524*	0.758*
May	1.312*	0.861	0.092*	0.941	0.895
June	1.296*	1.032	0.081*	1.340*	0.792*
July	1.104	1.772*	1.917*	1.167*	1.043
August	1.251*	1.162	1.437*	1.201*	1.213
September	1.530*	1.216	1.147	1.183*	1.348*
October	1.436*	1.036	1.429*	0.936	1.073
November	1.063	1.101	0.805	1.094	0.964
December	1.395*	2.043*	1.805*	1.518*	4.318*
July 2001	1.230*	0.371**	0.075*	0.795*	0.655*
August 2001	0.942	0.694*	0.097*	0.926	0.624*
September 2001	0.900*	1.020	0.180*	0.772*	0.748*
October 2001	0.897*	1.650*	0.125*	0.835*	0.770
November 2001	1.252*	1.052	0.218*	0.888	1.471*
December 2001	1.845*	1.830*	0.088*	2.527*	2.015*
January 2002	0.912	0.233*	0.180*	0.799*	1.048
February 2002	0.908*	0.558*	0.496*	1.116*	0.923
March 2002	1.122*	1.639*	0.126*	1.084	0.913
April 2002	1.182*	1.452*	2.583*	1.470*	1.858*
May 2002	1.045	1.061	11.093*	0.905	1.313*
June 2002	1.451*	4.107*	18.044*	1.140*	2.528*
Vat increase next month=1		0.778*	0.390	0.926	0.924
Vat increase this month=1		2.243*	0.380	1.296*	1.338*
Vat increase previous month=1		0.869	0.514	0.939	1.504*
Vat decrease next month=1					0.229*
Vat decrease this month=1					1.557*
Vat decrease previous month=1					0.677
Wage growth	1.031	1.387*	8.083*	1.092*	1.587*
Hicp growth	0.995	0.950*	0.837	0.998	0.909*
Size 0	0.925*	0.846*		1.106*	0.625*
Size small	0.943*	0.862*		0.782*	0.557*
Size med	0.905*	0.849*		0.874*	0.828*
no. observations	28564	18442	1849	54865	48200
No. price spells	13013	4125	1724	13178	6443
No. ended price spells	12637	3841	1664	11939	5377
log likelihood	-111351	-28149	-12020	-102696	-41101

* indicates significance at the 95% confidence level

Table 7
Pass-through VAT decrease hairdressing (17.5% to 6%) hairdressing, January 2000

	No obs.	Average price	Average change wrt December price (%)
January 1999	194	19.46	5.60
January 2000	196	19.50	-2.18
January 2001	191	20.45	4.36
January 2002	190	21.45	3.44
January 2003	186	22.85	3.81

Pass-through VAT increase 17.5% to 19%, January 2001

	No obs.	Average price	Average change wrt December price (%)
January 1999	1829	39.02	-1.20
January 2000	2152	45.19	-0.80
January 2001	2395	47.66	0.27
January 2002	2364	45.33	0.63
January 2003	2243	42.87	-0.30

Table 8 Price effects introduction of the euro

	F_j	F_j^+	ΔP_j^+	$\Delta P_j^{+*}F_j^+$	F_j^-	ΔP_j^-	$F_j^- \Delta P_j^-$	Net monthly p.c.
January 00-June 01								
UPF	30.36%	17.57%	25.11%	4.41%	12.79%	30.82%	3.94%	0.47%
PF	21.77%	14.76%	6.18%	0.91%	7.01%	10.52%	0.74%	0.17%
Ener	69.51%	40.05%	6.42%	2.57%	29.46%	3.32%	0.98%	1.60%
NEI	9.41%	5.68%	17.25%	0.98%	3.73%	22.99%	0.86%	0.12%
Serv	7.99%	6.88%	9.05%	0.62%	1.12%	11.52%	0.13%	0.49%
July 01-June 02	F_j	F_j^+	ΔP_j^+	$\Delta P_j^{+*}F_j^+$	F_j^-	ΔP_j^-	$F_j^- \Delta P_j^-$	Net monthly p.c.
UPF	37.83%	22.44%	21.77%	4.89%	15.39%	31.05%	4.78%	0.11%
PF	22.64%	16.04%	7.72%	1.24%	6.61%	11.79%	0.78%	0.46%
Ener	94.49%	40.05%	2.71%	1.08%	54.44%	2.66%	1.45%	-0.36%
NEI	18.28%	11.77%	13.29%	1.56%	6.52%	21.71%	1.41%	0.15%
Serv	13.07%	11.34%	6.99%	0.79%	1.73%	10.01%	0.17%	0.62%
Euro/pre-euro	F_j	F_j^+	ΔP_j^+	$\Delta P_j^{+*}F_j^+$	F_j^-	ΔP_j^-	$F_j^- \Delta P_j^-$	Net monthly p.c.
UPF	1.25	1.28	0.87	1.11	1.20	1.01	1.21	0.22
PF	1.04	1.09	1.25	1.36	0.94	1.12	1.06	2.62
Ener	1.36	1.00	0.42	0.42	1.85	0.80	1.48	-0.23
NEI	1.94	2.07	0.77	1.60	1.75	0.94	1.65	1.22
Serv	1.64	1.65	0.77	1.27	1.55	0.87	1.35	1.26

Table 9 Likelihood ratio tests on parameter coefficients (table 6a)

	H_o	H_1	LR-test	p-value
Month dummies	$\vec{b}_{months} = 0$	$\vec{b}_{months} \neq 0$	764.90	0.00
Product dummies	$\vec{b}_{products} = 0$	$\vec{b}_{products} \neq 0$	3000.0	0.00
VAT-rate dummies	$\beta_{vat\_incr\_next\_month} = \beta_{vat\_incr\_prev\_month} = \beta_{vat\_decr\_next\_month} = \beta_{vat\_decr\_prev\_month} = 0$	$\beta_{vat\_incr\_next\_month} = \beta_{vat\_incr\_prev\_month} = \beta_{vat\_decr\_next\_month} = \beta_{vat\_decr\_prev\_month}\ ?\ 0$	23.9	0.00

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