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Optimal Price Setting During a Currency Changeover: Theory and Evidence from French Restaurants

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Résumé

Ce papier étudie le comportement de fixation des prix des entreprises lors d'un changement de monnaie. Les difficultés des acheteurs face au nouveau prix nominal créent à court terme des incitations pour les vendeurs à augmenter leurs prix, mais induisent aussi un risque sur la réputation des vendeurs à long terme. Nous modélisons cet arbitrage et étudions les conditions pour lesquelles il est optimal d'augmenter ou de diminuer un prix au moment de la conversion. Une variable essentielle dans cette décision est l'information disponible aux acheteurs sur la conversion opérée par le vendeur; (i) la taille de l'entreprise, (ii) la proportion d'acheteurs réguliers, et (iii) la visibilité du prix d'un produit modifient cette information et affectent la décision de conversion des prix. À partir de prix individuels de produits vendus dans les restaurants en France, nous estimons l'effet du passage à l'euro par une méthode de différence-de-différences et montrons que les résultats sont cohérents avec les prédictions du modèle théorique. En effet, au moment du passage à l'euro dans l'Union Monétaire Européenne, les prix ont eu tendance à moins augmenter dans les restaurants de grande taille, dans les restaurants où les touristes sont moins nombreux, surtout quand les prix sont particulièrement visibles.

Codes classification JEL: E31, F33, M39.
Mots-clés: fixation des prix, passage à l'euro, inflation.

Abstract

This paper studies firms' price-setting decision during a currency changeover. Buyers' difficulties with the new nominal price level create incentives to raise real prices temporarily but doing so comes at the risk of damaging a seller's reputation in the long run. We model firms' trade-off and study under which conditions increasing or decreasing prices is optimal. A key variable in the decision is buyers' information about a firm's conversion, which in turn is affected by (*i*) a firm's size, (*ii*) the proportion of regular buyers, and (*iii*) the visibility of a good's price. Difference-in-differences analyses based on micro-data of French restaurants strongly support the model's predictions empirically. Indeed, prices around the 2002 changeover in the European Monetary Union are less likely to rise in larger and non-tourist restaurants, especially when prices are advertised.

JEL classification: E31, F33, M39.
Keywords: price setting, changeover, euro, inflation.

1 Introduction

This paper studies firms' optimal price setting during a currency changeover and tests the model's predictions using firm-level data of French restaurants during the currency changeover in the European Monetary Union in January 2002. The changeover had a noticeable impact on firm's price setting behavior. Baudry et al. [2007], for example, estimate that firms changed around 35 percent of their prices between December 2001 and January 2002 while the average share during other months is 15 percent. Most price changes were upward, reflecting to some extent the low but positive inflation during this period, but more than a third of the price changes were downward.[1] This paper shows that the heterogeneous response is not random, but follows a clear pattern explainable by optimizing behavior.

A good way to understand why a currency changeover may temporarily affect relative prices is to consider the situation of a consumer in the days of a changeover who, facing new nominal prices, fears being 'cheated' by sellers, just like a tourist in a foreign country with an unfamiliar currency. Like a tourist, the buyer may have to pay higher prices until some familiarity toward the new nominal price level sets in. Unlike a tourist, however, the typical consumer is likely to have a regular relationship with sellers. This makes 'cheating' risky, as the seller may damage his reputation as a 'fair' trader. A firm, thus, faces a trade-off between short-run profits from higher prices and the risk of damaging its reputation in the long-run.[2]

We model this trade-off by following the advertising literature and assuming that demand is a function of both price and goodwill where goodwill comprises anything that affects demand other than the good's price. This may be the seller's friendliness, the services offered in addition to the actual good on sale and, central for our case, the seller's reputation as a fair trader. Consumers deem price increases during the changeover unfair and reduce demand when they find out. This way of modeling the trade-off is similar in spirit to Rotemberg's [2005] model of customer anger. Our model is symmetric in that lowering a price at the changeover increases goodwill. In this sense, lowering prices may be viewed as a substitute to an advertising campaign.

Buyers' difficulties with the new nominal price level are modeled by assuming that the price elasticity of demand is temporarily reduced. This is a simple way of modeling

[1]Baudry et al. [2007] use firm-level data of French consumer prices. The estimates for other countries are similar, see for example Hoffmann and Kurz-Kim [2006].

[2]If consumers perceive a price increase as an attempt of the firm to take advantage of the changeover, it is not necessary that the firm actually had the intention to 'cheat'. Any rounding up, thus, bears this risk to some extent.

a possibly higher price mark-up during the changeover. In Gaiotti and Lippi [2004], the higher mark-up is generated by the assumption that a fraction of consumers is unable to observe the new prices allowing firms to raise prices without affecting demand.[3] A different mechanism is proposed by Dziuda and Mastrobuoni [2009] who set up a search model in which search itself is costless, but buyers receive signals about whether an observed price is high or low. Assuming that the changeover increases the variance of this signal generates a higher price mark-up.[4]

This paper contributes to the theoretical literature by showing that both increasing and decreasing prices may be optimal during a changeover and that this decision depends on (observable) characteristics of the firm. The model also predicts that a firm's market power (as measured by the equilibrium price elasticity of demand) does not help to predict the direction of the impact, only whether or not we can expect to observe one. A price increase should, therefore, not be viewed as a sign of lack of competition (as argued in Gaiotti and Lippi [2004]) because a firm charging a high markup may just as well reduce prices at the changeover.

In the model, a key variable influencing a firm's decision is customers' information about the firm's conversion. Lowering prices is only optimal if sufficiently many customers are aware of it. On the other hand, when a firm is sure that its customers are unable to realize price increases, it is more likely to raise its price. We cannot directly observe customers' information about a firm's conversion but the model suggests that we can deduce it indirectly from (*i*) a firm's size, (*ii*) the composition of its clientele, and (*iii*) from the visibility of a good's price. These comparative static results are driven by the way customers acquire information about a firm's conversion and how this information proliferates. Price increases are less likely when the firm is large and when many of its customers are regulars. The visibility of a good's price has a similar effect.

We test these predictions on the restaurant sector using a large data set of individual price quotes (micro-data) collected by the French statistical office to compute the Consumer Price Index. The data set contains more than 600 000 individual monthly price quotes and covers a period of 8 years (1996-2003). In this paper we focus on the restaurant sector because the impact of the 2002 changeover was larger in services and especially in restaurants than in other industries (see for instance Attal-Toubert et al. [2002], Gallot [2002] and Fougere et al. [2007]). An advantage of our focus is that the results are easily compared with the literature on restaurant prices.

[3]In this set-up, the price needs to be restricted from above to avoid firms choosing an infinite price.

[4]The literature has suggested two other explanations for price increases during a changeover: menu costs (Hobijn et al. [2006]) and multiple equilibria (Adriani et al. [2009] and Eife [2012]). Observing price decreases in an environment of a positive trend inflation is difficult to reconcile with menu costs.

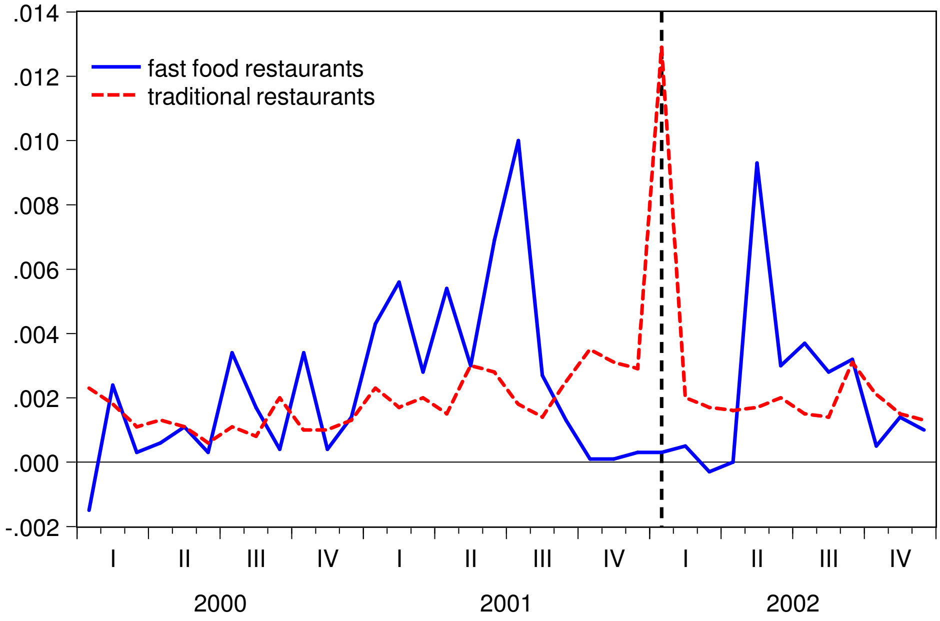

Figure 1: Monthly inflation of French restaurants around the changeover (denoted by the vertical line).

We contribute to the empirical literature on the subject by analyzing a longer and more detailed data set than previous studies. More importantly, the data set is constructed to be representative of the entire industry. In two interesting studies, Gaiotti and Lippi [2004] and Adriani et al. [2009] use data from the Michelin Red Guide, a restaurant guide that is published on a yearly basis covering restaurants of relatively high quality. Both studies report a positive effect of the changeover on restaurant prices. Focusing on data provided by The Economist Big Mac Index (published biannually), Parsley and Wei [2008] do not find a significant effect of the changeover. Attal-Toubert et al. [2002] and Gallot [2002] examine the changeover's impact on the entire CPI basket and report that particularly many price movements were observed in the services sector and especially in the restaurant sector. This is in line with the studies that use aggregate HICP data from Eurostat (Ehrmann [2011], Dziuda and Mastrobuoni [2009] and Ercolani and Dutta [2007]).

To assess the effect of the changeover on prices, we rely on a difference-in-differences strategy and use the labels provided by the statistical office to identify certain characteristics of a restaurant or of a product. One label, for example, indicates the type of product (e.g. coffee or main course). Other labels indicate the type of restaurant (fast food or traditional) or whether a restaurant was closed when the data were collected. Figure 1 plots monthly inflation rates of fast food restaurants and traditional restaurants for the three years around the changeover. In the case of traditional restaurants, the changeover coincides with a sizeable price increase while no such behavior is observed for fast food restaurants.[5] In fact, fast food restaurants are more likely to lower than to raise prices in January 2002. It is this heterogeneity at the level of the individual product that helps to understand firms' motivation to raise or lower real prices at a currency changeover - an event that in principle could be purely nominal. Overall and in line with the literature, we estimate an impact between 0.5 and 2.5 percent on the average price in the sector.

Using the information provided by the labels and following the predictions of the theoretical model, we study how a firm's decision is affected by its size, by the visibility of a product's price and by the composition of a restaurant's clientele (i.e. whether the restaurant's clients tend to be tourists and are thus unlikely to return or whether clients tend to be regulars). For all three samples we find a significant effect of the product's or the restaurant's characteristics on the firm's decision.

The paper is organized as follows. Section 2 discusses our assumptions about how a changeover affects firms' optimization problem and presents the household problem, the firm problem and the equilibrium. The section then studies how the equilibrium is affected by the changeover. Section 3 presents the data and analyzes the effect of the 2002 changeover using French restaurant prices. A discussion in section 4 concludes the paper.

2 The Model

In our economy, the changeover affects two consecutive periods, called period (1) and period (2). The changeover takes place at the beginning of period (1) and consumers have difficulties with the new nominal price level during this period. Firms may take advantage of this and raise prices. In period (2), some consumers realize how a firm converted its price in the previous period and adjust their demand accordingly. Increasing prices in period (1), thus, comes at the risk of lower demand in period (2). In the third period, the economy will return to its pre-changeover equilibrium. There is no entry or exit of firms caused by the changeover. This implies that firms may make profits or losses during period (1) and (2).

We assume in particular that the changeover affects the economy in three ways. First,

[5]Two spikes in the fast food series (six months before and six months after the changeover) are visible in figure 1. A possible explanation is that the reluctance of the average fast food restaurant to raise prices during the months around the changeover drove these restaurants to anticipate and postpone price changes.

buyers' difficulties with the new nominal prices are modeled by assuming that in the period of the changeover buyers' demand elasticity is reduced. Let $\bar{\varepsilon}$ be the normal price elasticity of demand and let $I^{(1)}$ be an indicator that equals one in period (1) and zero otherwise. The demand elasticity is given by

$$\varepsilon \equiv \bar{\varepsilon} - I^{(1)}\kappa > 1,$$

where $\kappa \geq 0$ measures buyers' difficulties.

Second, a firm may damage its reputation once buyers find out that the firm raised its price at the changeover. We assume that in period (1) consumers do not have information about a firm's conversion but in period (2) some (to be specified below) consumers have this information. As discussed in the introduction, we follow the advertising literature and assume that demand is a function of both price and goodwill. Firms can invest in goodwill and increasing prices at the changeover may damage goodwill while lowering prices may have a positive effect on goodwill. This second assumption introduces a trade-off between short-run gains from raising prices and future (*i.e.*, period (2)) losses of customers' goodwill.

Let firm i's goodwill be denoted by G_i and let $\eta_i = p_i^{(1)}/\bar{p}_i$ be a firm's conversion where $p_i^{(1)}$ is firm i's price in period (1) and $\bar{p}_i$ is a reference price.[6] The elasticity of goodwill with respect to η_i is given by $q_i I^{(2)} \geq 0$ where $I^{(2)}$ is an indicator that equals 1 in period (2) and zero otherwise.[7] That is, we assume that during the changeover - in addition to the effect measured by ε - a price change affects demand indirectly through goodwill. Converting correctly ($\eta_i = 1$) has no effect on goodwill. We interpret the elasticity q_i as the probability that a second-period buyer has information. The more consumers know a firm's conversion, the stronger the effect on goodwill and thus on demand.

Third, while we abstract from information asymmetries in the pre-changeover equilibrium, we assume that buyers' information during the changeover is not perfect. As mentioned above, no buyer has information about η_i in period (1) and in period (2), a customer has information about η_i

- if she was a customer of the firm in period (1), or

[6] Below we take the equilibrium price as reference price. A different assumption would be to take $p_i^{(0)}$, the price immediately before the changeover, as reference. This has the interesting consequence that it creates incentives to raise $p_i^{(0)}$.

[7] It is plausible that customers may judge a price increase as unfair all the time as in Rotemberg [2005] - not only during the changeover. This more general set-up can be modeled by assuming that the elasticity is given by $q_i I^{(2)} + \bar{q}_i$, with $\bar{q}_i > 0$. Since our focus is on the changeover and to keep the equilibrium simple, we assume that $\bar{q}_i = 0$.

- if she meets someone who was a customer in period (1), or
- if she happens to observe the price without being a customer.

Customers may be regulars or randomly assigned whereas meeting someone or observing the price without being customer is purely random. We expect that, for example, the composition of a firm's clientele affects a firm's decision. With the third assumption, we want to take into account that some prices are advertised and easier to observe than others. Loss-leader products are an example of such a pricing strategy. An example of this strategy in our data is the French *menu*, a fixed-price meal that includes two or more courses whose price is regularly displayed on chalkboard outside the restaurant well visible to passers-by.

A firm is assumed to know its size (the number of its customers) and how many of its customers are regular. However, since some customers are assigned randomly, the firm does not know who exactly of the economy's customers will enter its shop in a given period. This assumption will play an important role during the changeover.

2.1 The Economy

We modify a standard Dixit-Stiglitz-economy with a large number of monopolistically competitive firms in two ways. First, as mentioned above, we follow the advertising literature and assume that households have preferences over goods as well as over sellers. Sellers can invest in goodwill and sellers with high goodwill are able to sell more for a given price.

Second, we want to allow for the possibility that the firm is uncertain about its customers' identities. In the standard Dixit-Stiglitz model, all consumers purchase all products so that this uncertainty is absent. In our model, households consume only a fraction of the existing range of products and firms have regular as well as non-regular customers. Regular customers return in each period whereas non-regular customers are assigned randomly. This set-up also allows us to have firms differ in size without disturbing the symmetry of the standard Dixit-Stiglitz model.

2.1.1 Households

Consider an economy that admits a representative household with preferences given by

$$U\left(c_1, c_2, ..., c_N, c_0\right) = U\left(C, c_0\right) \tag{1}$$

where

$$C=\left(\sum_{i=1}^{N}I_iG_i^{\frac{1}{\varepsilon}}c_i^{\frac{\varepsilon-1}{\varepsilon}}\right)^{\frac{\varepsilon}{\varepsilon-1}}$$

is a consumption index of N differentiated varieties $c_1,...,c_N$ of a particular good and c_0 a generic good that embodies the rest of the economy. Let c_0 be the numeraire good. G_i is goodwill of variety i (with $i=1,2,...,N$) and $I_i=\{0,1\}$ is an indicator function.[8] The function $u(\cdot,\cdot)$ is strictly increasing, differentiable in both of its arguments, and jointly strictly concave which requires in particular that the elasticity of substitution between the different varieties is greater than one ($\varepsilon>1$). The household problem is static; we therefore suppress all time indices.

In order to avoid that the supply of goods produced by the monopolistically competitive firms automatically generates its own demand, households have the choice between the produced good (C) and the numeraire good (c_0). Households consume only a fraction n of all available products but all purchase the same amount $\left(\sum_{i=1}^{N}I_i\equiv n\leq N\quad\forall j\right)$. Thus, the households' consumption baskets differ in their composition but not in size. We assume that $M\geq N$, where M is the economy's number of households; that is, no consumer purchases the same item more than once in one period. Let the price of C be defined as

$$P\equiv\left[\sum_{i=1}^{N}I_iG_ip_i^{1-\varepsilon}\right]^{1/(1-\varepsilon)}.$$

Then, the choice between C and c_0 can be found by maximizing (1) subject to $PC+c_0\leq Y$, where Y is the household's income (which includes potential profits generated by the monopolistically competitive sector). This maximization yields the first order condition $\frac{\partial u(C,c_0)/\partial c_0}{\partial u(C,c_0)/\partial C}=\frac{1}{P}$, which assumes that the solution is interior, an assumption we maintain throughout. The strict joint concavity of u, combined with the budget constraint, implies that this first order condition can be expressed as

$$\begin{aligned}c_0&=h(P,Y)\\PC&=Y-h(P,Y)\equiv y\end{aligned}$$

for some function $h(\cdot)$. In words, the fraction y of income is devoted to the differentiated goods and the rest to the numeraire good. To derive the demand for individual varieties, denote the price of variety i by p_i. Then the budget constraint of the individual takes

[8] We abuse notation slightly by using the same symbol to denote the set and its cardinality.

the form

$$\sum_{i=1}^{N} I_i p_i c_i \leq y. \tag{2}$$

If $I_i = 0$ for some i, the good does not enter preferences and demand will be zero. If $I_i = 1$, the good provides utility and demand is positive. Maximizing equation (1) subject to equation (2) implies the following demand function for good i.

$$c_i = G_i \times Q(p_i)$$

where $Q(p_i) = \left[\frac{p_i}{P}\right]^{-\varepsilon} C$.

Two comments are in order. First, since the household's expenditure devoted to the differentiated good sector (PC) is independent of goodwill, a firm can only raise demand (by investing in goodwill) at the expense of its competitors. Aggregate goodwill is normalized to one, $\sum_{i=1}^{N} I_i G_i = 1$. Second, note that demand is separable in goodwill (G_i) and the 'primitive' demand function $Q(p_i)$. This separability combined with the assumption of constant marginal costs implies that investing in goodwill increases the quantity sold but not the good's price. The model, thus, represents what the literature calls the 'informative' view of advertising (see Bagwell [2007] for a discussion).

2.1.2 Firms

Suppose that each variety i can only be produced by a single firm, which is thus an effective monopolist for this particular commodity. Also assume that all monopolists maximize profits and are owned by the representative household. Let superscripts denote time periods. Firms maximize profits by choosing the price and by investing in goodwill. Investment in goodwill in period t is denoted by $A_i^{(t)}$ and per unit costs of investment in goodwill is normalized to one. The elasticity of goodwill with respect to A_i is $\varepsilon_A \in (0,1)$. We assume that the number of firms is sufficiently large in order to ignore the effect of p_i on P and C. Let $m_i \leq M$ be the firm's number of buyers. By assumption, the firm knows m_i. The problem of the firm is to maximize a stream of profits choosing price and investment in goodwill optimally:

$$\max_{p_i^{(t)}, A_i^{(t)}} = m_i \sum_{t=1}^{\infty} \beta^t \left(\left(p_i^{(t)} - \phi \right) \times c_i \left(p_i^{(t)}, A_i^{(t)} \right) - A_i^{(t)} - f \right) \tag{3}$$

where marginal costs (ϕ) are assumed constant, $\beta \in (0,1)$ is a discount factor and f are per-buyer fixed costs.[9] The fixed costs are expressed in terms of the firm's own output. Rearranging the first order conditions, we find that

$$p_i^{(t)} = \frac{\varepsilon}{\varepsilon - 1}\phi \quad (4)$$

$$\frac{\varepsilon_A}{\varepsilon} = \frac{A_i^{(t)}}{p_i^{(t)} c_i^{(t)}}. \quad (5)$$

Equation (4) is the familiar mark-up pricing equation and equation (5) is the Dorfman and Steiner [1954] condition that states that the proportion of sales revenue that a profit-maximizing monopolist spends on investment in goodwill is determined by a simple elasticity ratio.

2.1.3 Equilibrium

Equilibrium requires that the number of transactions in the differentiated good sector equals the total number of produced goods, that is,

$$(M \times n) = \sum_{i=1}^{N} m_i. \quad (6)$$

Equation (6) is a market clearing condition. As long as this condition is satisfied, the model offers some freedom regarding the size-distribution of firms without disturbing the symmetry of the standard Dixit-Stiglitz model. In equilibrium, all firms set the same price (eq. 4) and invest the same amount in goodwill (eq. 5). Profits, if expressed in per-buyer terms, are the same across firms. Households purchase the same quantity (c_i) of the same number of goods (n) and thus expenditure is the same across households.[10] Since there are no information asymmetries in equilibrium, we do not have to distinguish between regular or non-regular customers. In equilibrium, per-customer profits are given by

$$\frac{\pi_i}{m_i} = \frac{1 - \varepsilon_A}{\varepsilon} \frac{PC}{n} - f. \quad (7)$$

Imposing zero profits, equation (7) determines the size of the consumption basket n and through the market clearing condition (eq. 6) the total number of firms (N).

[9] The fixed costs may be interpreted as the costs of setting up branches. A firm that serves more customers needs to open more branches.

[10] If all firms are of equal size ($m_i = m$) and if households consume all existing products ($n = N$), we are back in the standard Dixit Stiglitz set-up where each firm serves the entire pool of customers ($m = M$).

2.2 The Changeover

As outlined above, the changeover affects the economy by (i) making information imperfect and by (ii) altering two elasticities. Setting $I^{(1)} = I^{(2)} = 1$ in the firm's problem, the firm now takes into account that its period (1) decision affects period (2) profits. In addition to the assumptions made in section 2.1.1, concavity requires $\varepsilon - q_i > 1$ and $\varepsilon_A < 1 - \frac{q_i}{\varepsilon - 1}$, which we assume to hold. We can now state the following proposition.

Proposition 1. *When the firm's decision in period* (1) *affects both current and future (second-period) demand, the actual demand elasticity is given by*

$$\varepsilon^{(1)} \equiv \bar{\varepsilon} - \left(\kappa - q_i \beta \frac{A_i^{(2)}}{A_i^{(1)}} \right).$$

A proof can be found in the appendix. Depending on the parameters, the actual elasticity may be larger or smaller than the pre-changeover elasticity $\left(\varepsilon^{(1)} \gtrless \bar{\varepsilon}\right)$. The term in brackets illustrates a firm's trade-off. As before, price is a mark-up over marginal costs and the Dorfman and Steiner [1954] condition takes the same form as above where in both equations the elasticity is replaced by $\varepsilon^{(1)}$. In period (2), the price returns to its pre-changeover level (since $\varepsilon^{(2)} = \varepsilon$) but second-period profits may still be affected by the changeover's effect on second-period goodwill. In the third period, the economy returns to its pre-changeover equilibrium.

A firm's decision to raise, lower or to keep prices constant depends on several parameters. For example, the larger households' confusion with the new price level (the larger κ), the more likely is a price increase. Since firms' action at the changeover affects future demand, firms' time preference (β) enters the elasticity. Of particular interest is q_i, the parameter measuring the probability that a second-period customer has information about a firm's conversion. As q_i decreases, $p_i^{(1)}$ increases reflecting the fact that lowering prices is only profitable if many consumers are aware of it. Let $\rho_i \leq m_i$ be firm i's number of regular customers, $E \leq M$ be the number of customers that exchange information and $v_i \leq M$ be the number of households that observe firm i's price independently of purchase. Given the assumptions about how information proliferates and given that firms may differ in size and in the composition of their clientele, q_i is both firm-specific and endogenous. In particular, q_i is the probability of the union of three events, $q_i = \Pr\left(A_1 \cup A_2 \cup A_3\right)$, with

- $A_1 = j \in m_i$,
- $A_2 = \cup_{k=1}^{M \setminus \{j,k\}} \left((k \in m_i) \cap (k \in E)\right)$ for $k \in M$,

- $A_3 = j \in v_i$.

A_1 is the event that household j was customer in the first period. This may happen either if j is a regular customer ($j \in \rho_i$) or if she happens to be customer in period (1) without being regular ($j \in m_i \backslash \rho_i$). A_2 is the event that customer j meets some other customer k and k happens to have been a customer of the shop. With the third event, we allow for the possibility that a customer observes a firm's price by chance. The events are not mutually exclusive. The following proposition summarizes the main comparative statics results of the model.

Proposition 2. *Given the assumption about how the changeover affects demand and given the assumptions about how information proliferates, the incentives to raise (lower) prices decrease (increase)* (*i*) *with firm i's size,* m_i*,* (*ii*) *with the visibility of its price,* v_i*, and* (*iii*) *with the firm's proportion of regular customers,* $\frac{\rho_i}{m_i}$.

$$\frac{dp_i^{(1)}}{dm_i} < 0, \quad \frac{dp_i^{(1)}}{dv_i} < 0, \quad \frac{dp_i^{(1)}}{d\left(\rho_i/m_i\right)} < 0.$$

A proof can be found in the appendix. Proposition 2 suggests the direction of the empirical analysis. The effect of ρ_i on $p_i^{(1)}$ is what we called repeated purchases. Firms with only few regular customers are more likely to raise prices. The firm's size enters because of the way information proliferates. A random exchange of information generates a bias toward larger firms. The larger a firm, the less likely are price increases. In this model, firms' market power (as measured by ε) does not affect the direction of the changeover's impact, only whether or not we can expect one. As the market becomes more competitive, the changeover effect disappears, that is, $\lim_{\varepsilon\to\infty} \eta_i = 1$.

3 Empirical Analysis

Using micro data on prices, this section explores whether the predictions of proposition 2 help to explain the heterogeneity in pricing strategies during the euro changeover. One challenge in this exercise is that the variables of the theoretical model (such as a firm's size) are not directly observable but need to inferred indirectly from the information provided by the statistical office. The underlying assumptions are discussed in detail below. The section starts with a description of our data. Section 3.2 provides information about our estimation strategy, and section 3.3 presents the results.

3.1 Data

To test the predictions of the theoretical model, we use a large longitudinal data set of restaurant prices. The sample is extracted from the database of monthly price quotes collected by the French National Statistical Institute (INSEE) to compute the French Consumer Price Index (CPI). Each observation is the price of a specific item sold in a given outlet in a given month. Prices are always inclusive of service and value-added tax (VAT).[11] Prices prior to January 2002 are converted from French franc to euro and rounded to the second decimal.

Our data set consists of more than 600 000 monthly price quotes for more than 20 000 items sold in about 4 500 different restaurants. Besides the large number of restaurants surveyed, the INSEE sample has the advantage of being designed to be representative of the restaurant sector.

The time dimension of the data set is also quite long, our sample covers the period from January 1996 to February 2003.[12] The monthly frequency gives us some flexibility about the point in time we assume firms to convert prices into the new currency. This point is discussed in detail in the next section. Moreover, an individual code is associated with a specific product in a given outlet. This allows us to follow a price trajectory over time and might reduce measurement biases.

Finally, the number of product-types available in our data set is rather large: more than 10 different products are reported by INSEE for the restaurant sector. Examples are meals ('*menus*' in French, which is a fixed-price bundle of a starter plus a main course, or a main course plus a dessert), starters, main courses, desserts, drinks sold in traditional restaurants and meals sold in fast food restaurants (which typically consist of a hamburger, French fries, and a soft drink) or other self service restaurants.

3.2 Empirical model

The aim of the empirical model is to assess the effect of the changeover on price dynamics. Using a difference-in-differences strategy, we test the impact of firm and product characteristics by comparing differences in inflation rates before and after the changeover (the changeover is defined as the treatment). In our three empirical analyses, the treated group consists of restaurants or products that, following proposition 2, we expect to be more affected by the changeover: (*i*) small restaurants, (*ii*) less visible prices, and (*iii*)

[11] See Baudry et al. [2007] for more details on the database.

[12] The original database begins in July 1994, but in August 1995 a VAT increase occurred concerning only traditional restaurants, so we restrict our sample to January 1996 - February 2003.

tourist restaurants. The control group consists of the counterparts: (*i*) large restaurants, (*ii*) advertised prices, and (*iii*) non-tourist restaurants. In all three empirical analyses, the variable of interest is the inflation rate.

Our baseline empirical model is the following difference-in-differences specification:

$$\Delta p_{i,t} = \alpha + \beta\ EURO + \gamma\ TREAT + \delta\ EURO * TREAT + \tau\ Z_{i,t} + u_i + \epsilon_{i,t} \quad (8)$$

where $\Delta p_{i,t}$ is the log difference of the price of a product sold in a given restaurant between two dates (see below for details), α is a constant term, *EURO* is a dummy variable equal to one around the changeover (this variable captures temporal effects common to the treated and the control group), *TREAT* is a dummy variable equal to one if i belongs to the treated group, 0 if it belongs to the control group (this variable captures systematic differences between the two groups), and the dummy variable *EURO*TREAT* is the interaction between these two variables. We also add some control variables $Z_{i,t}$, like regional dummies and year dummies. u_i is an individual random effect and $\epsilon_{i,t}$ is an idiosyncratic shock which is normally distributed with zero mean and standard deviation σ_ϵ. The parameter δ captures the effect of the changeover on the treated group.

For each of the three analyses, we run three different exercises. First, the linear specification above; second, an estimate of the dynamic impact which is also based on the linear specification; and third, a Tobit regression. The Tobit model allows us to assess the probability of a correct conversion and in case of a price change, we can measure to which extent the price change is larger for the treated group.

A key issue here is the moment a firm converts prices from the old to the new currency. The changeover took place on January 1st 2002 but firms may have converted prices several weeks or months before or after that date (denoting prices in two currencies was possible).[13] The actual moment of a firm's conversion is unobservable so that in order to compute inflation at the changeover ($\Delta p_{i,t}$), we have to specify time windows in which we assume firms converted prices.

In our baseline window, the variable of interest (the inflation rate) is calculated over a one-year period: we compute the log-difference between the price observed six months before the changeover and the price observed 6 months after the changeover (June 2001 - June 2002), and compare these price changes with price changes that occurred between June of the year $t-1$ and June of year t over the 1996-2001 period. The 12 month window is our baseline because the typical duration of prices in restaurants is one year

[13]The inflationary trend in the restaurant sector starting in January 2001 and strengthening toward the end of that year is consistent with the possibility that price adjustment started well before January 2002. See also the discussion in Eife [2012].

(Fougere et al. [2010]). Overall we consider five different windows. All windows are symmetric around the changeover. The first four windows have a length of 2, 6, 12, and 18.[14] months, respectively.

The fifth window has a length of 2 months but unlike in the case of the other four windows, we compare price dynamics at the changeover with price dynamics *after* the changeover (December 2001/January 2002 with December 2002/January 2003). The idea is to check whether firms' pricing behavior during the changeover was different not only from previous price evolutions, but from the subsequent one as well. This comparison is only possible for the shortest (2 month) window because our data ends in February 2003.[15]

3.3 Results

In this section, we present the results of our empirical analyses, looking respectively at the effect of a firm's size, the visibility of a product's price, and the proportion of regular customers on a restaurant's pricing behavior at the changeover.

3.3.1 Firm size

The first empirical analysis examines the role of a firm's size on inflation at the changeover. The labels provided by the statistical office do not give direct information about a restaurant's size. But we argue that the label indicating the type of restaurant (fast food versus traditional) provides indirect information.

While both types or restaurants share many characteristics they often differ in size. Many fast food restaurant belong to larger chains whereas the typical traditional restaurant is an independent business. Using sectoral national accounts and the information given in a report by Parniere and Pollet [2003], we find that sales of chains represented 85 percent of total sales in the fast-food restaurant sector (including cafeterias) in the 2001-2002 period. In the traditional restaurant sector this share is only 16 percent. In addition, the fast food market is relatively concentrated, with the first five largest chains having a market share of 75 percent. We use meals sold in traditional restaurants as the treated group and meals sold in fast food restaurants as the control group.[16]

[14] In order to avoid overlapping in the 18 month window, we exclude from the sample price changes calculated between March 2000 and September 2001, between March 1998 and September 1999 and between March 1996 and September 1997.

[15] In principle, such a comparison could shed light on the hypothesis of Adriani et al. [2009] and Eife [2012] of a persistent and structural impact of the changeover but because of the short sample our results remain inconclusive.

[16] We also consider cafeterias as part of the fast food category. Cafeterias often belong to larger chains

Table 1: Price changes of meals in traditional restaurants versus price changes of meals in fast food restaurants around the changeover (size effect).

	Windows: $Dec_{t-1} : Jan_t$ (2 months)	$Sep_{t-1} : Mar_t$ (6 months)	$Jun_{t-1} : Jun_t$ (12 months)	$Mar_{t-1} : Sep_t$ (18 months)	$Dec_{t-1} : Jan_t$ (2002-03 control)
Euro	-0.007	-0.699***	-0.678***	3.219***	0.118
	(0.102)	(0.167)	(0.227)	(0.294)	(0.156)
Trad.Rest	-0.042	-0.370***	-0.279***	0.264	0.236*
	(0.048)	(0.078)	(0.101)	(0.196)	(0.124)
Euro*Trad.Rest	0.977***	2.408***	2.170***	0.194	0.655***
	(0.109)	(0.179)	(0.243)	(0.325)	(0.176)
Intercept	0.358***	1.370***	2.891***	1.738***	0.257
	(0.110)	(0.173)	(0.219)	(0.379)	(0.207)
N	12,693	10,906	9,768	4,247	3,945
R^2	0.031	0.055	0.063	0.119	0.041

Note: Estimates are obtained using OLS technique. Year and regional dummies are included.
Standard Errors in brackets. Significance levels: *: 10% **: 5% ***: 1%.

Table 1 reports our estimation results obtained in the simple linear model. Each column corresponds to one of our five windows. In our baseline window, price changes are calculated over a period of one year. The estimates associated with the interaction term *EURO*TRAD.REST* provide the effect of the changeover on traditional restaurants. This effect is significant and rather large. Prices in traditional restaurants increased by 2.17 percent more than in fast food restaurants around the changeover.

For the 2 month window in the second column, we still obtain a positive effect close to 1 percent. This effect is however smaller than the one found for the one-year window, which is consistent with the observation that some firms may have converted their prices before December 2001 or after January 2002. The effect for the 6 month window is larger (2.4 percent) and close to the effect obtained with our baseline window. For the 18 month window, the effect is still positive but not significant. This result is consistent with the possibility that fast food restaurants may have anticipated their price increases before the changeover in order to keep them stable or even decrease them at the date of the changeover.

As discussed above, the fifth window (2002-03 control) compares inflation between December 2001 and January 2002 with inflation between December 2002 and January 2003. For this window, we obtain a positive and significant effect. This result suggests that the inflation effect singled out in January 2002 is not observed in January 2003.

and, like fast food restaurants, typically have no waiting staff table service.

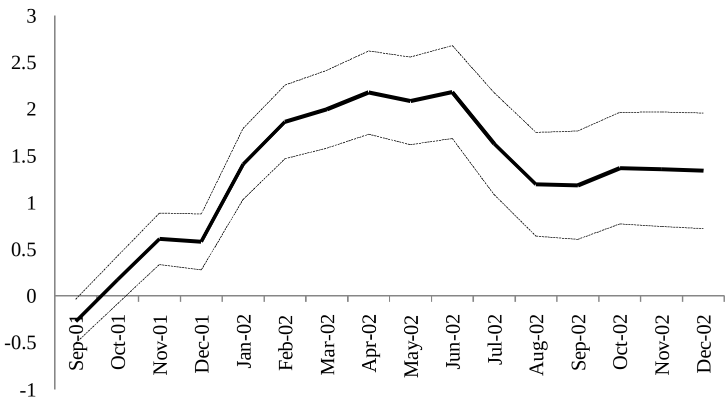

Figure 2: Cumulative effect over time of the changeover on price changes of meals in traditional restaurants versus price changes of meals in fast food restaurants (size effect).
Note: Over the period from September 2001 to December 2002, we estimate for every month a difference-in-differences exercise and we plot the point estimates obtained for the interaction term (which capture the euro effect) and their 95 percent confidence interval (dashed lines)

Moreover, it shows that our results are robust to using the pre-changeover years as the control group.

To examine more precisely the dynamic effect of the changeover, we run the following exercise: we calculate price changes using an incremental window (from September 2001 to December 2002) and measure the cumulative effect of the changeover on prices in each month. In particular, we construct different samples, always considering June 2001 (for price changes at the changeover) or June year $t-1$ (for price changes before the changeover) as dates of the initial price observation.[17] Then, we consider September 2001 (respectively September year $t-1$), October 2001 (respectively October year $t-1$), ..., December 2002 (respectively December year $t-1$) as dates for the final price observations. As before, we compute price changes as the log-difference between the initial and final price observations and estimate the difference-in-differences specification (8) for each sample. Again, the estimates associated with the interaction term corresponds to the changeover effect on inflation in traditional restaurants. Using the series of point estimates, we can observe the evolution of the changeover effect over time.

In figure 2 the solid line indicates the estimates for the interaction term while the dashed lines shows the 95 percent confidence interval. The horizontal axis indicates the endings of the windows. The effect shown in the figure can be interpreted as a cumulative effect of the changeover on inflation. As expected from the previous exercise, the effect is hump shaped, thus supporting our choice of 12 months for the baseline window. The effect reaches a maximum between March and June 2002 after which it markedly declines

[17] As robustness checks, we also consider March year $t-1$ and September year $t-1$ as initial dates for the calculations of price changes and we find that estimation results remain very similar.

but it is still significant at the end of 2002. Thus, the estimates suggest that the price increases in traditional restaurants during the changeover were not completely offset by price decreases during the year 2002.

Table 2: Probability and size of price changes of meals in traditional restaurants versus price changes of meals in fast food restaurants around the changeover (size effect).

	Estimates	Marginal effect on price change	
		probability	size
Euro	-0.247	-0.015	-0.100
	(-0.422)	(-0.026)	(0.168)
Trad.Rest	-1.742***	-0.110***	-0.765***
	(-0.203)	(-0.013)	(0.095)
Euro*Trad.Rest	4.528***	0.282***	2.349***
	(-0.451)	(-0.027)	(0.283)
Intercept	1.282***		
	(-0.439)		
N	9768		
Log likelihood	-16386		

Note: Standard Errors in brackets. Significance levels: *: 10% **: 5% ***: 1%.
Estimates are obtained using a Tobit model estimated by maximum likelihood.
Year and regional dummies are included.
Price changes are defined on the period from June of year $t-1$ to year t.

As a third exercise, we estimate a Tobit regression. This allows us to decompose the effect of the changeover into (1) the effect on the probability of a price change and (2) the effect on the size of a price change. We report results obtained with our benchmark sample (12 month window) in table 2 (estimates and marginal effects). In the Tobit model, $\Delta p^*_{i,t}$ is unobserved and, $\Delta p_{i,t}$, the actual price change is zero if $\Delta p^*_{i,t} < 0$ and equal to $\Delta p_{i,t}$ if $\Delta p^*_{i,t} > 0$. The coefficients estimated are equal to $\frac{dE(\Delta p^*_{i,t})|x_{i,t}}{dx_{i,t}}$ and capture the effect of the covariates on the unobserved price change $\Delta p^*_{i,t}$ and can, therefore, not be interpreted as the effect on the observed value $\Delta p_{i,t}$. Following the decomposition suggested by McDonald and Moffitt [1980], we calculate the marginal effect of $x_{i,t}$ on the probability of $\Delta p^*_{i,t} > 0$ and the marginal effect of $x_{i,t}$ on price changes when prices are increased, *i.e.* $E(\Delta p^*_{i,t}|\Delta p^*_{i,t} > 0, x_{i,t})$. We find that the changeover has a positive and significant effect on the probability of a price change. The effect on the probability is very large and close to 30 percentage points.[18] The estimates imply that restaurants did not convert prices from francs to euros correctly but decided to raise them. Moreover,

[18]See Puhani [2012] and Ai and Norton [2003] for a discussion on difference-in-differences estimates in non-linear models.

Table 3: Price changes of products with less versus more advertised prices in traditional restaurants around the changeover (visibility effect).

	Windows: $Dec_{t-1} : Jan_t$ (2 months)	$Sep_{t-1} : Mar_t$ (6 months)	$Jun_{t-1} : Jun_t$ (12 months)	$Mar_{t-1} : Sep_t$ (18 months)	$Dec_{t-1} : Jan_t$ (2002-03 control)
Euro	0.980***	1.735***	1.857***	3.448***	0.775***
	(0.059)	(0.105)	(0.149)	(0.185)	(0.095)
Other Prod.	0.001	0.000	0.064	0.215	-0.083
	(0.031)	(0.052)	(0.074)	(0.157)	(0.098)
Euro*Other Prod.	0.311***	0.433***	0.638***	-0.028	0.413***
	(0.074)	(0.126)	(0.179)	(0.257)	(0.139)
Intercept	0.283***	0.976***	2.224***	2.187***	0.483***
	(0.082)	(0.130)	(0.183)	(0.339)	(0.182)
N	18519	18003	16084	5971	5796
R^2	0.051	0.069	0.067	0.097	0.042

Note: Standard Errors in brackets. Significance levels: *: 10% **: 5% ***: 1%.

Estimates are obtained using OLS technique. Year and regional dummies are included.

restaurants that increased their prices at changeover, increased them by more (around 2.5 percent) than what one would expect had the changeover not taken place.

3.3.2 Visibility of Product Price

The second analysis examines the effect of the visibility of a product's price on inflation around the changeover. Again, we do not have direct information about whether a price is advertised or otherwise easy to observe for customers. But we take advantage of the custom to display the price of a *menu* outside restaurants well visible to passers-by. A French *menu* is a fixed-price meal that includes two or more courses. In the marketing literature, these fixed-price meals are an example of product bundling and they are cheaper than the sum of their individual components.

In this section, we restrict our analysis to traditional restaurants only and compare the price of *menus* with the price of other products whose prices only appear inside the restaurant written on the menu (in the English sense). This restriction allows us to control even more precisely for possible differences in other factors that can have played a role in the observed price dynamics. Summing up, our main assumption here is that the price of meals (*menus*) is easier to observe for customers than the price of the other products sold by the restaurant (e.g. desserts, starters, coffee, and bottles of wine). *Menus* are the control group and the other products the treatment group.

Table 3 reports our estimation results obtained for the OLS model. As in the previous

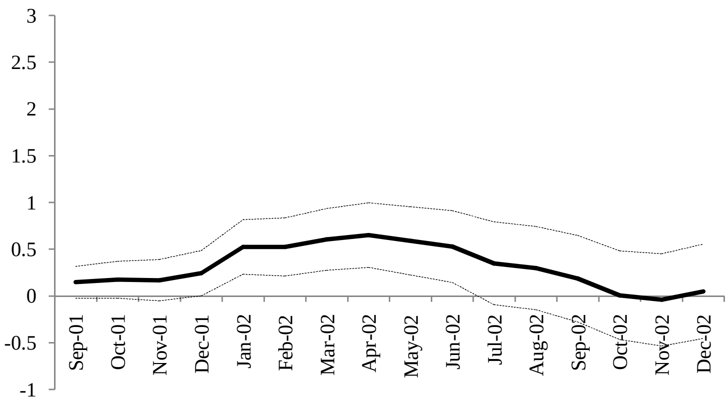

Figure 3: Cumulative effect over time of the changeover on price changes of products with less versus more advertised prices in traditional restaurants (visibility effect).
Note: Over the period from September 2001 to December 2002, we estimate for every month a difference-in-differences exercise and plot the point estimates obtained for the interaction term (which capture the euro effect) and their 95 percent confidence interval (dashed lines).

section, each column refers to a different window. For our baseline window (12 months), we find that the coefficient on the interaction term *EURO*OTHERPROD* is 0.64 percent. The effect is smaller than in the previous section but still significant. The incentives to increase prices at the changeover are higher for less visible products. For the shorter 2 and 6 month windows, the interaction term is positive and significant but for the longest window of 18 months, the effect is negative and non-significant. This implies that the euro effect tends to disappear and prices go back to their pre-changeover path.

In order to control for product specific rounding effects related to different price levels, we also restrict our sample to products (meals and other items) whose price levels are similar in size and run regressions where we add some controls for price levels.[19] Differences in price levels, however, do not appear to affect our results.

The second exercise is, as before, an estimation of the dynamic effect of the changeover. Figure 3 shows the point estimates of the interaction term with a 95 percent confidence interval. Again, the effect is hump shaped but significantly smaller in magnitude than in the previous section. Moreover, between July 2002 and December 2002, the euro effect is no longer significant suggesting that the effect is temporary with firms reverting to their normal price-setting after a couple of months.

Finally, table 4 reports estimates of the euro effect obtained with a Tobit model. We find that the changeover had a positive and significant effect on the probability of a price increase. This effect is smaller (6 percentage points) than the one obtained with fast food and traditional restaurants. The estimate of the interaction term suggests that less

[19]We introduce dummy variables for the position of prices in the price distribution (in terms of quartiles).

visible prices are more likely to increase at the changeover. Moreover, when prices are increased, they are likely increase more when they are less visible. The marginal effect obtained on the size of a price change is positive and significant (+0.5 percentage points).

Table 4: Probability and size of price changes of products with less versus more advertised prices in traditional restaurants around the changeover (visibility effect).

	Estimates	Marginal effect on price change	
		probability	size
Euro	6.082***	0.292***	2.723***
	(0.321)	(0.015)	(0.176)
Other Prod.	-0.783***	-0.036***	-0.271***
	(0.184)	(0.008)	(0.064)
Euro*Other Prod.	1.330***	0.062***	0.493***
	(0.372)	(0.018)	(0.148)
Intercept	-2.544***		
	(0.445)		
N	16084		
Log likelihood	-25315		

Note: Standard Errors in brackets. Significance levels: *: 10% **: 5% ***: 1%.

Estimates are obtained using a Tobit model estimated by maximum likelihood.

Year and regional dummies are included.

Price changes are defined on the period from June of year $t-1$ to year t.

3.3.3 Regular customers

In our last analysis, we study how a restaurant's clientele affects its pricing decision at the changeover. Restaurants with many regular customers may behave differently during a changeover than restaurants visited only by tourists. Unlike in Adriani et al. [2009], our data set does not indicate whether a given restaurant is in a tourist location. In order to identify tourist restaurants, we proceed in two steps. First, we restrict attention to traditional restaurants in seacoast *départements*.[20] The idea is to focus on areas where summer (May until October) tourism is an important economic factor.

In the second step, we take advantage of the custom of many restaurant owners to go on extended summer vacation. Restaurants that expect many tourists in summer are unlikely to close during this period. On the other hand, restaurants that do not expect many tourists during summer are more likely to close during that period. The French

[20] A *département* is an administrative zone of which there are 96 in France. Each has approximately the same geographical size (6000 km^2). Around a quarter of the French *départements* are on the seacoast.

Table 5: Price changes of meals in tourist restaurants versus price changes of meals in non-tourist restaurants around the changeover (regular customers effect).

	Windows: $Dec_{t-1} : Jan_t$ (2 months)	$Sep_{t-1} : Mar_t$ (6 months)	$Jun_{t-1} : Jun_t$ (12 months)	$Mar_{t-1} : Sep_t$ (18 months)	$Dec_{t-1} : Jan_t$ (2002-03 control)
Euro	1.155***	1.306***	1.192***	2.022***	1.225***
	(0.148)	(0.274)	(0.378)	(0.460)	(0.332)
Tourist	0.022	0.057	0.004	-0.034	0.433
	(0.073)	(0.136)	(0.172)	(0.332)	(0.304)
Euro*Tourist	0.011	0.483	0.668	1.746***	-0.357
	(0.164)	(0.302)	(0.411)	(0.538)	(0.394)
Intercept	0.271**	1.123***	2.455***	2.046***	0.231
	(0.121)	(0.222)	(0.283)	(0.454)	(0.352)
N	3364	2998	2685	1195	844
R^2	0.068	0.073	0.070	0.107	0.044

Note: Standard Errors in brackets. Significance levels: *: 10% **: 5% ***: 1%.

Estimates are obtained using OLS technique. Year and regional dummies are included.

statistical office records on one of the labels whether a restaurant is open or closed in a given month.

Summing up, restaurants on the seacoast that are closed during summer are assumed to have more regular customers (less tourists) than restaurants on the seacoast that are open during summer. Following the convention from before, tourist restaurants are the control group and non-tourist restaurants are the treatment group.

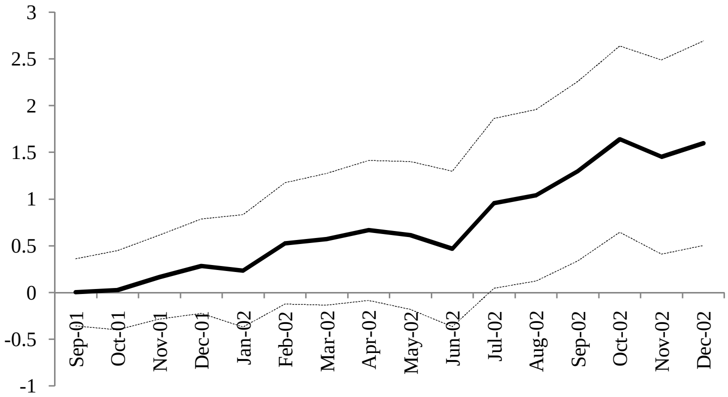

Figure 4: Cumulative effect over time of the changeover on price changes of meals in tourist restaurants versus price changes of meals in non-tourist restaurants (regular customers effect).

Note: Over the period from September 2001 to December 2002, we estimate for every month a difference-in-differences exercise and plot the point estimates obtained for the interaction term (which capture the euro effect) and their 95 percent confidence interval (dashed lines).

Table 6: Probability and size of price changes of meals in tourist restaurants versus price changes of meals in non-tourist restaurants around the changeover (regular customers effect).

	Estimates	Marginal effect on price change probability	size
Euro	4.748***	0.277***	2.261***
	(0.755)	(0.042)	(0.430)
Tourist	0.080	0.005	0.030
	(0.395)	(0.022)	(0.150)
Euro*Tourist	0.664	0.038	0.262
	(0.798)	(0.046)	(0.327)
Intercept	-1.317**		
	(0.640)		
N	2685		
Log likelihood	-4345		

Note: Standard Errors in brackets. Significance levels: *: 10% **: 5% ***: 1%.
Estimates are obtained using a Tobit model estimated by maximum likelihood.
Year and regional dummies are included.
Price changes are defined on the period from June of year $t-1$ to year t.

Table 5 reports estimation results obtained with an OLS specification where the interaction term *EURO*TOURIST* provides the estimates of the effect of the changeover on inflation in tourist restaurants. For our baseline window of 12 months, the effect is positive and similar in magnitude to the one obtained in the section on visibility but it is not significant. The same result is obtained for the two shorter windows. However, a large positive and significant effect is found for the longest window of 18 months where the effect is estimated to be around 2 percent.

We find a similar picture when we estimate the dynamic effect of the changeover (Figure 4), where the estimated effect is small and non-significant at first but increases over time to reach a maximum towards the end of the sample. One possible explanation for this delayed effect may be different pricing seasons. In order to compensate for the lack of an extended summer vacation, many tourist restaurants on the seacoast close in winter. In our sample, 82 percent of the tourist restaurants are closed from November to March. During that period, no price observations are recorded and the opening of a restaurant after the winter holidays often coincides with a price change.

Note that in this exercise, we find that the dynamic effect is still significant in December 2002. The effect appears to be rather persistent and price increases are not compensated by price decreases. We should also note that the standard errors are quite large and even larger at the end of our sample period. This uncertainty is due to the small size of our sample and may also be due to measurement issues associated with how we identify tourist restaurants.

The final exercise is again a Tobit regression (see table 6). As in the previous sections, we find a positive effect of the changeover on the probability of a price increase (+4 percentage points) and also on the size of a price change when prices are increased. However, the effect on the size of a price change is relatively small (0.25 percent) and not significant.

Here too, focusing on the longer 18 month window changes the results. For the 18 month window, the marginal effect of the euro on the probability of a price increase is large (+13 percentage points) and significant, and the effect on the size of a price increase when prices are increased is also large (equal to +1.3 percent) and significant. All in all, tourist restaurants on the seacoast are more likely to raise prices at the changeover than restaurants on the seacoast with more regular customers.

4 Conclusion

This paper studies firms' price setting during a currency changeover and shows that the heterogeneous response - with some firms increasing and others decreasing prices - is not random but follows a clear pattern explainable by optimizing behavior. In the theoretical model, firms face a trade-off between short-run profits from raising prices and trying to take advantage of customers' difficulties with the new prices and the risk of damaging their reputation in the long-run. A key variable in a firm's decision is customers' information about the firm's conversion. Lowering prices is only optimal if customers realize the price decrease. On the other hand, a firm that is sure that its customers do not realize when it increases prices is more likely to do so. We cannot observe customers' information directly, but the model suggests that we can deduce it indirectly from a firm's size, the composition of its clientele, and from the visibility of a good's price.

We test the model's predictions using a large data set of individual price quotes from French restaurants. Using a difference-in-differences strategy, we find that smaller firms are more likely to raise prices than larger firms and that when a price is advertised and easy to observe for customers, it is less likely to rise. Moreover, restaurants with many regular customers are less likely to raise prices than tourist restaurants. Overall, the magnitude of the changeover effect is between 0.5 and 2.5 percent. This is in line with the literature on the subject. There is some evidence that the changeover effect was persistent but our data set is not long enough to make inferences about a long-run effect. Using Tobit regressions, we find that the changeover not only affected the size of price adjustments but also the probability to observe a price increase.

In the empirical part, we focused on the restaurant sector, but the model's predictions are more general. For example, the prediction that a firm's size affects the firm's conversion seems plausible in general. For instance, Fougere et al. [2007] find that similar to the fast food chains, large retailers avoided price increases in the six months around the changeover. The empirical test of the model's theoretical predictions in other contexts is left to further research.

Appendix

Proof of Proposition 1 The system to solve is

$$
\begin{aligned}
c_i^{(1)} &= \frac{A_i^{(1)}}{\varepsilon_A\left(p_i^{(2)}-\phi\right)} \\
p_i^{(1)} &= \frac{\left(\varepsilon+q_i\beta\frac{A_i^{(2)}}{A_i^{(1)}}\right)}{\left(\varepsilon+q_i\beta\frac{A_i^{(2)}}{A_i^{(1)}}\right)-1}\phi \\
c_i^{(2)} &= \frac{(\varepsilon-1)\,A_i^{(2)}}{\beta\varepsilon_A\phi},
\end{aligned}
$$

where $c_i^{(t)}$ is given by (2.1.1) and $d\ln G_i/d\ln\eta_i = I^{(2)}q_i$. We solve the system by Cramer's rule after linearizing it by total differentiation. Let the vector of endogenous variables be $b=\left(dp_i^{(1)}, dA_i^{(1)}, dA_i^{(2)}\right)$. The system to solve is then given by $\Omega b' = \chi$, where Ω is a 3×3 matrix of first partials and χ is a vector of exogenous variables. Showing that the determinant Ω is positive, reduces to showing that

$$
\left(\varepsilon^{(1)}-\varepsilon-q_i\right)\frac{\left(\varepsilon^{(1)}-\varepsilon\right)}{\varepsilon^{(1)}} > -\left(\varepsilon^{(1)}-1\right)(1-\varepsilon_A).
$$

Recall that concavity requires $q_i < \left(\varepsilon^{(1)}-1\right)(1-\varepsilon_A)$. Combining both equations and rearranging gives

$$
-\varepsilon q_i < \left(\varepsilon^{(1)}-\varepsilon\right)^2
$$

which holds since all parameters are assumed positive.

Proof of Proposition 2 Given our assumptions, the probabilities that one of the three events occurs are Laplace probabilities: $\Pr(j\in v_i)=\frac{v_i}{M}$, $\Pr(j\in m_i)=\frac{m_i}{M}$ and if we allow for regular customers

$$
\Pr(j\in m_i) = \frac{\rho_i}{M}+\left(1-\frac{\rho_i}{M}\right)\left(\frac{m_i-\rho_i}{M-\rho_i}\right).
$$

The probability that consumer j receives information from consumer k is given by

$$
\Pr\left(\cup_{k=1}^{M\setminus\{j,k\}}\left((k\in m_i)\cap(k\in E)\right)\right) = \frac{m_i}{M}\left(1-\left(1-\frac{E}{M}\right)^{M-2}\right)\equiv t_i.
$$

Combining, we find that the probability that a second period costumer has information is given by

$$q_i = \frac{m_i}{M} + \left(1 - \frac{m_i}{M}\right)\left(\frac{v_i}{M} + \left(1 - \frac{v_i}{M}\right) \times t_i\right).$$

It is straightforward to show that $\frac{dp_i^{(1)}}{dm_i} < 0$, $\frac{dp_i^{(1)}}{d(\rho_i/m_i)} < 0$, $\frac{dp_i^{(1)}}{dv_i} < 0$, as stated in the main text.

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