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Regular Adjustment: Theory and Evidence[1]

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Abstract

We ask why, in many circumstances and many environments, decision-makers choose to act on a *time-regular* basis (e.g. adjust every six weeks, etc.) or on a *state-regular* basis (e.g. change an interest rate by 0.25%, etc.), even though such an approach appears suboptimal. The paper attributes regular behaviour to adjustment cost heterogeneity. The reasons for this heterogeneity are discussed. We show that, given the cost heterogeneity, the likelihood of adopting regular policies depends on the shape of the benefit function: the flatter it is, the more likely, *ceteris paribus*, is regular adjustment. In general, however, there is no clear relationship between the degree of cost and benefit function heterogeneity and the incidence of regular adjustment. We provide sufficient conditions under which the less frequent are adjustments, the greater is the incidence of regular policies.

To test the model we use a large Austrian data set, which consists of the direct price information collected by the statistical office and covers 80% of the Consumer Price Index (CPI) over eight years. We run cross-sectional tests, regressing the proportion of attractive prices and, separately, the excess proportion of price changes at the beginning of a year and at the beginning of a quarter, on various conditional frequencies of adjustment, inflation and its variability, dummies for good types, and other relevant variables. The results provide strong support for the model: the lower is the conditional frequency of price changes in a given market, the higher is the incidence of time- and state-regular adjustment.

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1. Introduction

"The Federal Reserve on Tuesday raised U.S. interest rates [by 0.25%] a 14th straight time." Reuters, January 31, 2006.

In many circumstances and in many environments, decision-makers choose to act on a regular basis and, in particular, on a calendar-regular basis (e.g., once a week, on the first day of each quarter, etc.) even though such an approach appears suboptimal. Similarly, some decision-makers appear to prefer some values of the variables under their control (e.g. prices ending with a 9, interest rates which are multiples of 0.25% etc.). The above quote provides a good example. Over the last year and a half the Federal Reserve raised the federal funds rate from 1% to 4.5%. It did so by raising the rate by 1/4% at every, regularly scheduled, policy meeting. The focus of this paper is to analyze a simple explanation of such behaviour.

A common feature of the environments in question is their dynamic structure. The policymaker(s) maximizes a stream of benefits, which depends on the values of some state variables. Over time these values change, or *deteriorate*.[2] The policymaker can reset the state variables but doing so involves a cost. Therefore adjustment is infrequent.

The motivation and focus of the paper is nominal price adjustment at the firm level. In this application, a firm posts the nominal price for the product(s) it sells. Due to general inflation the real price falls over time. The real price can be reset by choosing and posting a new value of the nominal price. But similar problems arise in many other environments. Therefore we begin by describing issues related to regular adjustment using examples from various potential applications. In section 2 we set up the theoretical model in general terms. We then refocus our analysis on the issue of optimal pricing policies, testing the model using a large set of pricing data.

Several further examples of this environment follow:

1. *Wage adjustment*. Under general inflation, the purchasing power of contractually-set wages declines over time and can be increased in a new contract.
2. *Machinery refurbishing.* The capital stock deteriorates over time due to physical use or obsolescence. It is improved by refurbishing or replacing the machinery.
3. *Inventory reordering*. A firm holds an inventory of the product(s) it sells. The level of the inventory falls over time. It is replenished by a new delivery.

[2] Alternatively, the current values of the state variables are constant while the optimal values drift over time. These problems are similar and so we will focus mostly on environments with constant optimal values.

4. *Monetary policy.* The Central Bank sets the interest rate appropriate for the current conditions. Over time the match between the current and the optimal value deteriorates. The interest rate can be readjusted through a decision of the Bank's policy-making body.
5. *Fiscal policy.* The fiscal authority sets spending and taxation priorities in the budget. Over time the desired fiscal structure changes. It is reset in a new budget.
6. *Financial reporting and shareholders meetings.* Financial reports allow investors to evaluate firm's prospects. Over time the quality of the information held by investors declines; also, management may follow strategies that benefit them rather than shareholders. The information is refreshed in financial reports; shareholder meetings help realigning the interests of management and shareholders.
7. *Information*. Newspapers and magazines allow the public to update their information. New events lead to its deterioration. A new issue brings the information up to date.
8. *Elections.* An election aligns the preferences of the government and the population. Over time the preferences diverge. A new election allows realigning the preferences.
9. *Monitoring patients.* A patient's visit allows the physician to undertake a proper course of action. Over time the health of the patient or the effectiveness of the treatment may decline. A repeat visit allows the doctor to review and adjust the treatment.

These problems are fairly common. As discussed below, they often lead to *state-contingent* adjustment policies. The decision maker monitors the state variable and applies the control whenever it has deteriorated to the threshold point. Hence the timing of adjustment does not depend solely on time and, in general, adjustments are not regular.

In practice, however, we observe many cases where controls are applied at specific moments of time. U.S. grocery stores adjust prices on Wednesdays (Levy et al., 1997); drugstores adjust prices on Fridays (Dutta et al., 1999). Seasonal sales are held every January and July. Many firms get regular deliveries; machinery is often refurbished on a regular basis. Labour contracts are signed for a fixed number of years. Financial reporting is quarterly and shareholders meet yearly. Magazines and newspapers appear with fixed frequency. In most political systems elections are held regularly. Medical associations provide guidelines on the frequency of checkups and so on.

In many cases some decision-makers follow regular policies while others do not. While some firms change prices at predetermined dates, others follow state-contingent optimal pricing policies (Cecchetti, 1986). Car firms offer incentives on a state-contingent basis (depending on inventory levels). Machinery is often refurbished when predetermined technical requirements are met. Many firms follow just-in-time delivery schedules, leading to inventory deliveries that are not regular. Outsiders receive financial information at equally spaced intervals but internal information flow is often organized on just-in-time basis. In political systems based on the British parliamentary tradition the timing of elections can be chosen by the government.

Even when the policy is formally regular, it sometimes contains specific provisions for deviating from the schedule if needed. Firms may hold extraordinary shareholder meetings, the interest rate may be changed between the regular meetings of the policy makers, the government may introduce a mini-budget and so on. Recent examples of such special arrangements are Proposition 8 in California, and the Constitution of Venezuela, which allow an early election.

Furthermore, policymakers sometimes switch between regular and irregular policies. Several years ago the Bank of Canada moved from weekly to less frequent meetings. Car producers switched to just-in-time delivery policies. List prices for cars are no longer set for a year; most airlines nowadays use sophisticated pricing schedules etc.

Finally, some policymakers follow different policies for different activities. Paper versions of newspapers are published regularly, but electronic versions are not.[3] Some supplies may be obtained regularly while others are procured on just-in-time basis. Doctors set regular, routine visits for some patients but not for others, etc.

Understanding of regular policies is important as they reduce flexibility by limiting the ability of the policymaker to react to past, current and future events. It is important to note that the distinction between expected and unexpected events is not crucial here. Once the system is set up to adjust on a regular basis, the policymaker may not be able to alter the course of action for a range of both expected and unexpected changes. For example, a political system that uses regular elections may not be able to react to predictable changes in the environment if the politicians are unable to master enough votes to change the constitution.

The explanation of these phenomena we propose here is simple. Adjustment of the state variable is costly, but the adjustment costs are not constant over time (or over values of the state variable). They are lower at some points of time (or at, or to, some values of the state variable). When the lower values of the costs occur regularly, for some policymakers regular adjustment dominates the state-contingent policy that would have been optimal if costs were homogeneous.

The proposed explanation may, at first thought, appear trivial. But it is no different than the explanations, popular in economics, of infrequent changes based on the presence of adjustment (or *menu*) costs. The logic of the menu cost approach is as follows. If price adjustment were costless, nominal prices would have been changing continuously. Since they do not, adjustment must be costly. Moreover,

[3] We are grateful to Magdalena Konieczna for suggesting this example.

such costs are easily identified. The theoretical task then becomes to explain the observed pattern of behaviour under the assumption that price adjustment is costly.

We follow the same logic here. If adjustment costs were always the same, observed behaviour would not be regular. Elections would be held when the difference between the government's and the population's preferences crosses certain thresholds; a firm would order new delivery when its inventory falls below a certain level; newspapers would be published after a sufficient amount of events worth writing about has taken place etc. But since adjustments are regular, their cost cannot be constant. Moreover, the benefits of regular behaviour are easily identified. It allows planning and organization of activities, enhances the reputation of the decision-maker and so is, in general, less costly than irregular behaviour.

We start the paper by showing an existence result: when the costs of adjustment are lower at regular moments of time, an optimizing policymaker will (except in unlikely circumstances), sooner or later, take advantage of the lower costs. In general, however, there is no clear relationship between the degree of cost heterogeneity and the incidence of regular adjustment. We then show that, given the cost heterogeneity, the likelihood of adopting regular policies depends on the shape of the benefit function: the flatter it is, the more likely, *ceteris paribus*, is regular adjustment. We provide sufficient conditions under which the less frequent are adjustments, the greater is the incidence of regular policies.

The model is applied to nominal price adjustment. The distinction between the time contingent, regular nominal price adjustment policies as in Fischer (1977) and in Taylor (1980), and state-contingent policies as in Sheshinski and Weiss (1977) is crucial, given their different implications for effectiveness of monetary policy (Caplin and Spulber, (1989), Caplin and Leahy, (1992)).

There are two aspects of regular nominal price adjustment we are interested in: time-regularity and state-regularity. A disproportionate proportion of price changes takes place at the beginning of periods, rather than within periods. Several studies in the Inflation Persistence Network (IPN) report a high proportion of prices are held constant for a year (see Álvarez et al. (2005) for Spain, Aucremanne and Dhyne (2005) for Belgium, Baudry et al. (2004) for France, Baumgartner et al. (2005) for Austria, Dias et al. (2005) for Portugal, Veronese et al. (2005) for Italy, Lünnemann and Mathä (2005) for Luxembourg, Hoffmann and Kurz-Kim (2005) for Germany). Konieczny and Skrzypacz (2002) report that, in price data collected three times a month, over a half of all changes take place in the first 10 days of a month. Similarly, several IPN studies, as well as Levy et al. (2006) find a large proportion of prices charged are *attractive* prices.[4]

[4] Attractive prices – which sometimes are also called threshold prices or pricing points – include psychological prices (prices ending in 9), fractional prices (prices which are convenient to pay, such as 1.50) and round prices (defined as whole number amounts, such as 10.00).

To test the model we use a very large Austrian data set, which consists of the direct price information collected by the statistical office and covers about 80% of the CPI over eight years. We run cross-sectional tests, regressing the proportion of attractive prices and, separately, the excess proportion of price changes at the beginning of a year and at the beginning of a quarter on various conditional frequencies of adjustment, inflation and its variability, dummies for good types, and other relevant variables. The results are consistent with model's implications: the lower is, in a given market, the conditional frequency of price changes, the higher is the incidence of time- and state- regular adjustment.

The paper is organized as follows. The model is analyzed, and empirical predictions described, in the next section. In section 3 we discuss the empirical evidence. The last section concludes.

2. The Model

We consider a class of optimization problems where the value of instantaneous benefits depends on state variables that change over time. At any moment the policymaker can adjust the values of the state variables by incurring discrete costs. More formally, the instantaneous value of the benefits is $\tilde{B}[\vec{x}(t),\vec{y}(t),\vec{a}]$, where $\vec{x}(t)$ is a vector of state variables, $\vec{y}(t)$ is a vector of exogenous variables and $\vec{a}$ is a vector of parameters.. This formulation implies that the benefit function depends on time only indirectly.

We assume that $\tilde{B}[\vec{x}(t),\vec{y}(t),\vec{a}]$ is twice continuously differentiable and has a unique global maximum:

Assumption 1:

For every t , $\vec{y}(t),\vec{a}$ there exists $\vec{x}^*(\vec{y}(t),\vec{a})$ such that, for every $\vec{x}(t)\neq\vec{x}^*$:
$\tilde{B}[\vec{x}(t),\vec{y}(t),\vec{a}]<\tilde{B}[\vec{x}^*,\vec{y}(t),\vec{a}]$

Assumption A1 implies that, as long as $\vec{y}$ and $\vec{a}$ do not change, the optimal instantaneous values of the state variables are constant. The policymaker would like to maintain the state variables continuously at the level $\vec{x}^*$ or, if that is not possible, to keep them close to $\vec{x}^*$. Changes in $\vec{x}(t)$ over time will be called the *deterioration* of the state variables. The policy maker can adjust $\vec{x}(t)$ at any time to any desired level (perhaps within some bounds), but doing so involves a discrete cost.[5]

The cost of adjusting the state variable, suggested by the examples above, includes the time, or the opportunity cost of the time needed to set up the decision-making process (e.g. organizing an election and counting votes, the doctor's and the patient's time etc.), the time needed to make and implement the decision (e.g. the time needed to set up and implement a new budget, union/employer bargaining time etc.), physical resources (e.g. new machinery, printing a new price list etc.) and non-time opportunity costs (e.g. potentially beneficial decisions forgone due to election campaign duties, foregone output whenever production is affected by the refurbishing process etc.).

To simplify the analysis, and in line with earlier literature (Scarf, (1960), Sheshinski and Weiss, (1977)), we assume that the cost is lump-sum: independent of the size or of the frequency of adjustment. This is a reasonable assumption in some cases (elections, shareholder meetings, monetary policy decisions, printing a new price list etc.).[6]

In general, the optimal solution to the optimization problems described above is *state-contingent*. The policymaker observes the values of the state variables and, when they reach certain thresholds, incurs the discrete cost and adjusts them to new, optimally chosen levels. State-contingent policies imply, generally, adjustment at intervals of differing length. Thresholds, as well as the new values of state variables are computed optimally and can take on any values (from an admissible range).

In many environments, however, we observe behaviour inconsistent with state-contingent policies: adjustment often takes place at regular intervals and some values of the state variables are chosen more often than others. This paper therefore focuses on adjustment policies, which we call regular policies. We distinguish between *time-regular* policies, which involve adjustment on a regular basis (e.g. a firm orders new inventory every 52 days, monetary policy decision making body meets every six weeks, machinery is refurbished once every two years etc.) and *state-regular* policies, in which newly chosen values of the state variables belong to a small subset of all possible values (e.g. inventory is ordered by a truckload, a firm selects new prices ending in a nine: 0.69, 0.79 etc.) or when the thresholds are specific numbers (e.g. a hedge-fund manager's compensation rule changes if the

[5] In an equivalent problem, the optimal values change over time and the goal of the policymaker is to maintain the state variable as close as possible to the drifting optimal value, given the adjustment costs.

[6] Adjustment costs often include, in addition, a component which depends on the size of adjustment (refurbishing machinery, delivering a mini-budget etc.). We do not consider such cases here.

return exceeds 20% per year).[7] An important subset of time-regular policies are *calendar time-regular* policies, which involve adjustment at calendar-related intervals (e.g. an election is held every four years on the Tuesday next after the first Monday in November, a new price list is issued once a year etc.) or where the time of applying the control is related to the calendar (e.g. sales are held at the beginning of each January and each July)

To make the analysis tractable we make several simplifying assumptions:

Assumption 2:
Over the relevant range, and for any values of $\vec{y}(t), \vec{a}$, the effect of the vector $\vec{x}(t)$ on the benefit function $\tilde{B}[\vec{x}(t), \vec{y}(t), \vec{a}]$ can be completely summarized by a single state variable $x(t)$.[8] i.e. there exists $B[.]$ such that

$B[x(t), \vec{y}(t), \vec{a}] \equiv \tilde{B}[\vec{x}(t), \vec{y}(t), \vec{a}]$ and,

for every t, $\vec{y}(t), \vec{a}$ there exists $x^*(\vec{y}(t), \vec{a})$ such that, for every $x(t) \neq x^*$:

$B'[x(t), \vec{y}(t), \vec{a}] \cdot [x^* - x(t)] < 0$.

where $B'[.]$ denotes the derivative of the benefit function with respect to its first argument. Assumption A2 means that the problem is equivalent to one in which the benefit function is a smooth, quasiconcave function of a single state variable.

The crucial assumption, which differentiates the model from earlier literature, is that the cost of adjusting $x(.)$ may depend on time or/and on the level of x. We now consider the former case; the latter is similar and is discussed below.

To make matters as simple as possible, we divide time into periods and assume that the cost of adjustment can take on only two values: high, c_h, and low, c_l The cost is equal to the high value whenever adjustment takes place within a period; it is equal to the lower value at the end of each period. Some notation will be helpful. Let $\Im \equiv \{\tau_0, \tau_1, ...\}$ consist of the ends of each period. The interval $\{\tau_{i-1}, \tau_i\}$, i=1, 2… will be called *period i*. Whenever the adjustment takes place at $t \in \Im$, its cost is c_l. Such adjustment will be called regular adjustment and the *incidence of regular adjustments* will be the proportion of all adjustments which are regular.

Assumption 3:
The cost of adjustment is:

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[7] What we call time-regular policy is usually called a time-contingent policy. For clarity we avoid the latter term; this allows us to distinguish between state-regular and state-contingent policies.

[8] A somewhat stronger restriction is that all but one (say, the first) of the elements of the vector of state variables $\vec{x}(t)$ are fixed, i.e. $\vec{x}(t) \equiv (x(t), x_2^0, x_3^0, ..., x_k^0)$.

$$c(t) = c_h + I(t) \cdot (c_l - c_h), \quad c_h \geq c_l \tag{1}$$

where $I(t)$ is an indicator function, given by:

$$I(t) = \begin{cases} 1 & \text{for } t \in \Im \\ 0 & \text{for } t \notin \Im \end{cases} \tag{2}$$

As the focus of the paper is regular behaviour, we further assume that periods are of the same length, i.e. τ_i's are evenly spaced over time:

$$\tau_i = \tau_0 + n \cdot \tau, \quad n = 1,2,.... \tag{3}$$

Obviously, the larger is the difference between the high and low values of costs, the more tempting is regular adjustment and so a large value of c_h - c_l makes the problem trivial. Therefore we are careful not to make any assumptions about the size of the difference. All results hold even if the c_h - c_l is arbitrarily small.

In this paper we concentrate on the simple nonstochastic case. In particular:

Assumption 4:
The state variable $x(t)$ is assumed to change over time at a constant rate:[9]

$$x(t) = x(t_0) \cdot e^{-\alpha(t-t_0)} \tag{4}$$

Without loss of generality, we assume α>0.

At the time of the first adjustment the policymaker's goal is to pick the sequences of times of adjustment and the new values of the state variable, $W \equiv \{x_0, (t_1, x_1), (t_2, x_2), ...\}$ so as to maximize the present value of the benefits:

$$\underset{\{t_i\}_{i=0}^{\infty}, \{x_i\}_{i=0}^{\infty}}{\text{maximize}} \quad PV(W) = \sum_{i=0}^{\infty} \left\{ \int_{t_i}^{t_{i+1}} B[x_i e^{-\alpha(t-t_i)}, \vec{y}(t), \vec{a}] e^{-\rho t} dt \quad - \quad c(t) e^{-\rho t_{i+1}} \right\} \tag{5}$$

where $PV(W)$ denotes the present value of policy W, t_0 is the time of the first adjustment, ρ is the discount factor, and the first adjustment is assumed to be costless.[10]

[9] As already mentioned, an equivalent problem is when the optimal value of the state variable changes over time and adjustments are needed to keep the actual value close to the optimal value. This problem can be converted into the time–dependent problem by normalizing the drifting optimal value by its trend.

[10] As we consider the nonstochastic case here, we omitted expectations from equation (5).

We start the analysis by summarizing the well-known case when $\vec{y},\vec{a}$ do not change over time and the cost of adjustment is constant and equal to its higher value, i.e. $c_l=c_h$. To differentiate this case from the main one, the choices under this assumption are denoted with a " ˆ ".

Lemma 1

Let $\hat{W}^* \equiv \left\{\hat{x}_0^*,(\hat{t}_1^*,\hat{x}_1^*),(\hat{t}_2^*,\hat{x}_2^*),...\right\}$ denote the optimal policy, and $\hat{T}^* = \left\{\hat{t}_1^*,\hat{t}_2^*,...\right\}$ denote the set of the optimal adjustment times, when $c_h=c_l$. Then:

$\hat{W}^*$ is recursive: $\forall i : \hat{x}_i^* = x^*$ and , for all i $\hat{t}_{i+1}^* = \hat{t}_i^* + \Delta\hat{t}^*$. Also, $\hat{W}^*$ is unique. [11]
Finally, $\hat{B}(\hat{x},\vec{y},\vec{a}) - \hat{B}(\hat{x}e^{-\alpha\Delta\hat{t}},\vec{y},\vec{a}) = \rho c_h$
The proof is essentially the same as in Sheshinski and Weiss (1977).

Next, we consider the case when $c_l < c_h$. Proposition 1 below shows sufficient conditions under which, when $c_l < c_h$, it is optimal for the policymaker to take advantage of the lower adjustment costs. The proof is based on the following approximation of real numbers with rational numbers:

Lemma 2

For every
$x, K > 0$ there exist integers N_1, N_2 such that $N_2 \le K$ and $\left|N_2 x - N_1\right| < 1/K$

Proof: see Niven (1961).

The lemma can be applied to the problem considered here by setting $x = \Delta\hat{t}^* / \tau$. It implies that, if the policymaker follows a policy of adjusting once every $\Delta\hat{t}^*$ (which is optimal when costs of adjustment are constant), eventually an adjustment will take place arbitrarily close to the end of a period. Given the notation, N_2th adjustment will be within $1/K$ of the end of period N_1.

Since N_2th adjustment is close to the end of a period, the firm needs to alter its timing just a little to take advantage of the lower end-of-period adjustment costs. It will do so as long as the reduction in adjustment costs exceeds the loss in benefits. Obviously, as already mentioned, we do not want the result to depend on the difference c_h - c_l . A sufficient condition for the results to hold regardless of the

[11] Note that, since the optimized present value of benefits may be negative, no additional restrictions are placed on the values of the parameters and the momentary benefit function B.

size of c_h - c_l is that the slope of the benefit function be bounded; this is the motivation for assumption (b) below:

Proposition 1

Let $W^* \equiv \left\{x^*_0,(t^*_1,x^*_1),(t^*_2,x^*_2),...\right\}$ denote the optimal policy, and $T^* = \left\{t^*_0,t^*_1,t^*_2,...\right\}$ denote the set of the optimal adjustment times, when $c_l<c_h$.

Assume:

(a) $c(t)$ meets (1)–(3);

(b) for every $\vec{y},\vec{a}$ there exists $A<\infty$ such that, for every t: $\left|B'(x(t))\right|<A$;

(c) the time of the first adjustment $t_0 \in \Im$.

Then $\left\{T^* \setminus \{t_0\}\right\} \cap \Im \neq \varnothing$.

Proof

Without loss of generality let the time of the first adjustment be $t_0 = \tau_0$. The proof is by contradiction. Assume that $T^* \cap \Im = \{t_0\}$. Therefore, by Lemma 1, the set of optimal adjustment times is $\hat{T}^*$, with $\hat{t}^*_0 = \tau_0$. By Lemma 2, setting A=K, there exist two positive integers N_1 and N_2 such that:

$$N_2\Delta\hat{t}^* - N_1\tau < (1/\rho)\ln(c_h/c_l) \tag{6a}$$

$$N_2\Delta\hat{t}^* - N_1\tau < \rho\left(c_h - c_\ell\right)/(2A) \tag{6b}$$

When (6) are met we have:

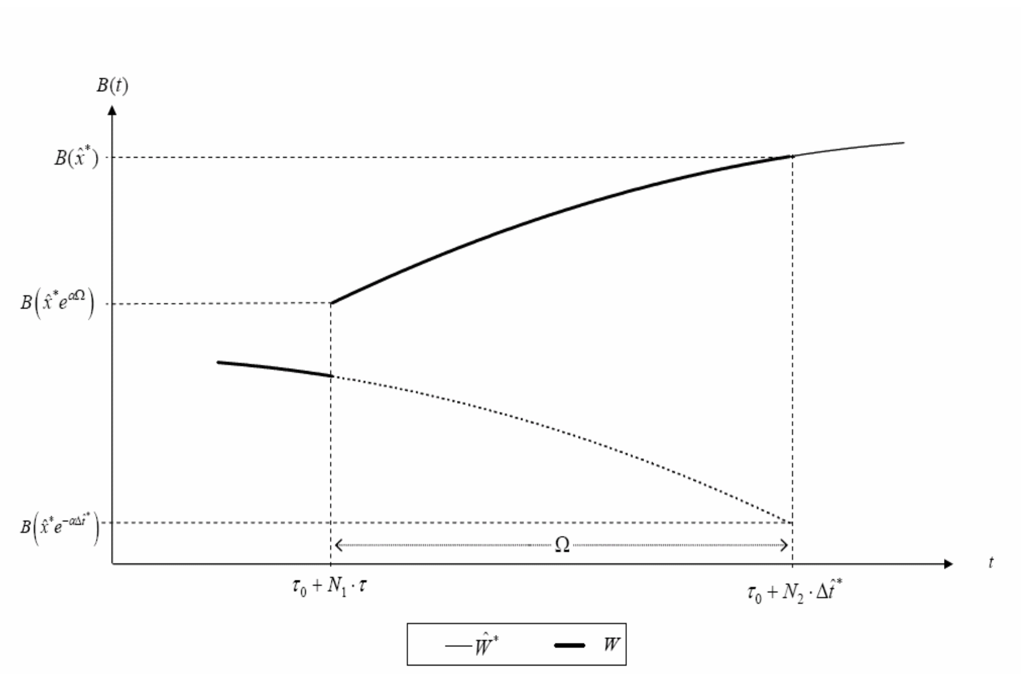
$$PV(\hat{W}^*) < PV\Big((\tau_0,\hat{x}^*),...,(\tau_0+(N_2-1)\Delta\hat{t}^*,\hat{x}^*),(\tau_0+N_1\tau,\hat{x}^*e^{-\alpha\Omega}),\tau_0+(N_2+1)\Delta\hat{t}^*,\hat{x}^*),...\Big) \leq PV(W^*)$$

where $\Omega = N_2\Delta\hat{t}^* - N_1\tau$. The second inequality follows from the fact that the middle policy need not be optimal for $c_l < c_h$. ■

Proposition 1 is illustrated in chart 1. It describes the situation in which the N_2th adjustment falls $\Omega = N_2\Delta\hat{t}^* - N_1\tau$ after the end of period N_1. To take advantage of the lower adjustment costs, the policymaker accelerates the N_2th adjustment to

$\tau_0 + N_1 \cdot \tau$ from $\tau_0 + N_2 \cdot \Delta\hat{t}^*$. Under policy *W*, she follows $\hat{W}^*$ until $\tau_0 + N_1 \cdot \tau$, when she adjusts *x* to such a value that, from $\tau_0 + N_2 \cdot \Delta\hat{t}^*$ on, *W*= $\hat{W}^*$. Inequalities (6) provide sufficient conditions[12] for the present value of *W* (the middle term in the above inequality) exceeds the present value of $\hat{W}^*$.

Chart 1: Profits as a Function of Time



Proposition 1 shows that, when the adjustment costs vary over time as postulated in *Assumption 3* and the first adjustment is at the beginning of period 0 ($t_0 \in \Im$.), under general conditions the policymaker would, sooner or later, take advantage of the lower costs of adjustment. Assumption (c) requires a discussion. If the time of the first adjustment $t_0 \notin \Im$, it is possible that the policymaker will never take advantage of lower adjustment costs. This would be the case if, for example,

[12] Inequalities (6) provide sufficient conditions also for the case when adjustment is delayed.

$\Delta\hat{t} = i \cdot \tau$ (i.e. when the optimal time between adjustments under constant costs is an integer number of periods) and the difference between c_l and c_h is small.

In many environments, however, $t_0 \notin \Im$ is an unlikely outcome. This is because the timing of the whole sequence of subsequent adjustment times, $T^* \backslash \{t_0\}$, often depends on the time of the first adjustment. For example, the timing of subsequent visits to a doctor is set relative to the initial visit, the timing of elections is set relative to the first election, dates of subsequent delivery depend on initial delivery and so on.[13] From now on we will assume that $t_0 \in \Im$.

We now characterize the optimal policy W^*. By Proposition 1, at least one time of adjustment under W^* coincides with the end of a period. To set notation, assume that the first such adjustment is the *N*th adjustment, and it takes place at the end of period *k*. Denote such a policy as $W^*_{N,k}$. This means that, under $W^*_{N,k}$, $t^*_N = \inf\left\{\{T^* \setminus \{t_0\} \cap \Im\right\} = \tau_k$,

It is easy to see that, for a given benefit function and adjustment costs, the optimal policy need not be unique. It is possible that $\tau_k < \hat{t}^*_N < \tau_{k+1}$ and that $PV(W_{N,k}) = PV(W_{N,k+1})$, i.e. the policymaker is indifferent between accelerating or delaying the *N*th adjustment.

The analysis of multiple equilibria in the current framework is complex. We therefore assume that, if $PV(W_{N,k}) = PV(W_{N,k+1})$ then $W^* = W_{N,k}$, i.e. whenever two policies yield the same present value of benefits, the policymaker chooses the policy with earlier adjustments.

Proposition 2

(a) W^* is recursive:

$$W^* = \left\{\left[(\tau_0, x^*_0), (t^*_1, x^*_1), \ldots, (t^*_{N-1}, x^*_{N-1})\right], \left[(\tau_k, x^*_N), (t^*_{N+1}, x^*_{N+1}) \ldots, (t^*_{2N-1}, x^*_{2N-1})\right], \ldots\right\}$$

(b) for every i: $t^*_{i+1} - t^*_i = t^*_1 - \tau_0$

[13] In environments in which the timing of adjustment is dictated by custom this need not be the case. For example a clothing store which opens in June may not be willing to have a sale shortly after the opening.

Proof

W^* can be written as: $W^* = \left\{ \left[(\tau_0, x_0^*), (t_1^*, x_1^*), ..., (t_{N-1}^*, x_{N-1}^*) \right], W^*(\tau_k^+) \right\}$, where $W(\tau_k^+)$ is the remainder of the optimal policy from period τ_k forward. Since W^* is optimal and unique, by the principle of optimality $W(\tau_k^+)$ is the solution to the problem of maximizing the present value of the benefits, starting in period τ_k. But this problem is identical to the original problem, as can be checked by substituting, $t'_i = t_{i-N}^*$. Therefore, $t_{2N}^* = \tau_{2k}$ and for every i such that $N < i < 2N : t_i^* \notin \Im$.
The proposition follows by induction.
The proof of part (b) is straightforward. ■

Proposition 1 shows an existence result: as long as the benefit function is not too steep, and subject to the discussion above, the timing of at least some of the adjustments will be dictated by the heterogeneous adjustment costs.

While the result in Proposition 1 is interesting, it has little empirical content, especially given the fact that the starting point of the analysis is the observation that many policies are, indeed, regular: some prices are changed at the beginning of the year, firms sometimes order a delivery of multiple truckloads etc..

The crucial question arising in this framework is the empirical incidence of adjustment at times in $\Im$, i.e. the proportion of all adjustments that are done at the beginning of a period (say, in January). By proposition 2, every 1/*N*th adjustment is in $\Im$ since the first adjustment in $\Im$ is the *N*th adjustment and the optimal policy W^* is recursive. Of particular interest is the special case $T^* \subseteq \Im$, i.e. when *N*=1 and the firm never pays c_h.[14] This incidence depends on two types of factors. The first is the empirical distribution of the exogenous variables $\vec{y}$ and of the parameters $\vec{a}$ across observations; the second is related to the shape of the benefit function *B*[.] and the difference $c_h - c_l$. The first type of factors determines the empirical distribution of the optimal length of time between adjustment under constant adjustment costs, $\Delta\hat{t}_i^*$; the second type determines the willingness of a policymaker to shift adjustment time to the end of a period to take advantage of the lower cost.

The existence result in Proposition 1 provides little information on the second question. Furthermore, whatever information it provides may be quite misleading. Consider a given problem in which $t_0 = \tau_0$ and $\Delta\hat{t}^*$ is a well-defined, continuous function of the exogenous variables $\vec{y}$ and the parameter vector $\vec{a}$. Assume further that, for some specific values of the exogenous variables and parameters, $\vec{y}_0$

[14] Of course, T^* may be a proper subset of $\Im$ (i.e. $T^* \subset \Im$) when N=1, for example if the optimal adjustment frequency is once every two periods.

and $\vec{a}_0$, we have $\Delta\hat{t}^* = \tau$, i.e. under constant adjustment costs it is optimal for the policymaker to adjust at the end of the period. In this case the policy is completely regular (*N*=1) in a neighbourhood of $(\vec{y}_0, \vec{a}_0)$ but *N*>1 outside this neighbourhood. Since there is, in general, nothing special about $(\vec{y}_0, \vec{a}_0)$, the resulting policy is regular just by coincidence.

As a more specific example, assume that *B*=*B*(*x*,*a*), i.e. the benefit function depends on the state variable and one parameter. Assume that the parameter is observable and its value is positively related to $\Delta\hat{t}^*$. This is the setup considered by Sheshinski and Weiss (1977), where *B*[.] is the real profit function of a monopolist, *x* is the real price and *a* is the inflation rate. Adjustment costs vary as postulated here. A researcher studies six policymakers and the observable parameter *a* is distributed across policymakers in such way that their (unobservable) optimal period of adjustment under constant cost, $\Delta\hat{t}_i^*$, are equal 10+*i*/32 months, *i*=15,…,20. Assume also that the difference between the high and low level of adjustment costs is small so that they never depart from *W**. Then, as the researches studies policymakers ordered by *a*, she observes the following incidence of regular policies: 1/8, 1/32, 1/16, 1/32, 1/2, 1/32. While the monthly frequency of price changes varies between 9.41% and 9.55%, the proportion of regular prices varies between 3.13% and 50%. Seemingly small changes in the parameter *a* have dramatic, nonmonotonic effects on the incidence of regular policies. This issue is, essentially, a number problem that is irrelevant to the questions considered here. We return to it below.

As we are interested in the reasons for different incidence of regular adjustment between policymakers, the question is what property of the benefit function determines the willingness of the policymaker to take advantage of the lower adjustment costs. The idea is straightforward. The policymaker faces a trade-off between the saving on the adjustment cost, and the profits foregone by not following *W**. The loss depends on how far profits decline as the time of adjustment varies. This, in turn, depends on the slope of the benefit function. A benefit function that is, at a given distance from its maximum, flat, makes the loss small and so the policymaker is willing to vary adjustment time to lower the adjustment cost. Definition 1 describes this intuition more precisely:

Definition 1

For any two twice continuously differentiable concave functions *f, g* such that there exists $x_0 : f'(x_0) = g'(x_0) = 0$, *f* is *more strongly concave* than *g* if and only if, for all *x*: $f''(x) - g''(x) < 0$.

Before we proceed it is useful to define precisely when a policymaker will deviate from the optimal policy *W** (i.e. the policy that she would have followed if adjustment costs were constant) to take advantage of the lower costs. We call it the *shift range.*

Definition 2

The shift range, S_i , is the interval $\left\{\tau_i - a_i, \tau_i - b_i\right\}$ such that, if and only if $\tau_i - a_i < \hat{t}_j^* < \tau_i + b_i$, the policymaker moves the jth adjustment at τ_i to save on adjustment costs.

In other words the policymaker moves the timing of adjustment to the end of period i if and only if the optimal timing under constant adjustment costs falls in the shift range. The size of the shift range S_i determines the willingness of the policymaker to take advantage of the lower adjustment costs at τ_i .

Proposition 3

Let B^1 and B^2 be two benefit functions, $\Delta\hat{t}*^1$, $\Delta\hat{t}*^2$ be the respective optimal times of adjustment when the costs of adjustment are constant and S_i^1, S_i^2 be there respective shift ranges. If B^1 is more strongly concave than B^2 then:

(a) $\Delta\hat{t}*^1 < \Delta\hat{t}*^2$
(b) $S_i^1 \subset S_i^2$

Proof

The benefit from extending $\Delta\hat{t}*$ is to reduce the expenditure on adjustment costs; the loss is due to the fact that it increases the range of $x(t)$ between adjustments. The optimality of $\Delta\hat{t}*^1$ under B^1 means that, under constant adjustment costs, the loss and benefit are equal. Under B^2 the benefit is the same but, since B^1 is more strongly concave than B^2, the cost is smaller. Hence $\Delta\hat{t}*^1$ is not optimal under B^2 and, for $\Delta\hat{t}* = \Delta\hat{t}*^1$, $PV(B^2)$ is increasing. This proves (a).

To prove (b) assume $\hat{t}_j^{*1} \in S_i^1$ so that $t_j^{*1} = \tau_i$, i.e. under B^1 the optimal policy involves shifting jth adjustment to the end of period i. Assume now that $\hat{t}_k^{*2} = \hat{t}_j^{*1}$, i.e. under constant costs the kth adjustment under B^2 coincides with the jth adjustment under B^1, $k<j$. Since B^1 is more strongly concave than B^2, the benefit of

shifting adjustment time from $\hat{t}_k^{*2} = (\hat{t}_j^{*1})$ to τ_i under B^2 exceeds the benefit under B^1. Therefore $t_k^{*2} = \tau_i$ which implies $S_i^1 \subset S_i^2$. ■

Propositions 1 and 3 summarize what can be said unequivocally about the incidence of regular policies. As long as $t_0 \in \Im$, regular behaviour is observed even under arbitrarily small difference between c_h and c_l. The smaller is the curvature of the benefit function, the less frequent are adjustments and the wider are shift ranges. This means that if, under constant adjustment costs, two policymakers would make adjustment at the same time, the policymaker who adjusts less frequently is more likely to move the adjustment to the end of the period.

It would be incorrect to conclude that Proposition 3 implies that the less frequent is adjustment, the greater is the incidence of regular policies. This is because there is, in general, no reason for adjustments to occur simultaneously. For example, assume that $\Delta\hat{t}^{*1} = \tau$ and $\Delta\hat{t}^{*2} = 10.5\,\tau$. Then, even though B^2 is much flatter than B^1, the adjustment policy under B^1 is completely regular, while, as long as $c_h - c_l$ is not too large, only every second adjustment under B^2 is at the end of a period.

There is no easy way around this number problem. One solution is to assume that there are many policymakers who differ with respect to the (unobserved) parameter a, which is distributed across policymakers in such a way that the following (sufficient) conditions are met:

Assumption 5:

(a) the empirical distribution of $\Delta\hat{t}^*$ on $\{\tau_{i-1}, \tau_i\}$ is independent of *i*;

(b) $\max_i(\Delta t_i) - \min_i(\Delta t_i) >> \tau$

Under the first assumption, the probability of finding a policymaker for whom the timing of the *k*th adjustment, $k\Delta\hat{t}^*$, is within a given distance from the end of the period is the same for all periods. This means that the proportion of adjustments at the end of a period would be larger the further away is the period from τ_0. The second assumption is needed so the effect of truncation of the range of $k\Delta\hat{t}^*$ "averages out".

Condition A5 (a) is not met in practice due to truncation of the range of $k\Delta\hat{t}^*$ both from below and above. The truncation from below is due to the fact that, first, $\Delta\hat{t}^*$ is bounded away from zero under lump-sum costs but $\Delta\hat{t}^*$ is not bounded away from above from τ, 2 τ, …The truncation from above is due to the fact that the limited length of the sample makes it impossible to observe policies $W^*_{N,k}$ for which $k\tau$ exceeds the length of the sample. Therefore it is possible for results of empirical tests of the model to be dominated by the number problem. This makes it difficult to interpret rejections of the model since the empirical tests of the model is a joint test of the relationship between benefit function shape and the incidence of regular policies as well as the fact that the number problem is "averaged out" in the data set.

Of course the number problem becomes irrelevant if the results of empirical tests are consistent with the model.

3. Empirical Evidence

We now turn to testing the implications of the model. Empirical testing requires cross-sectional (across policymakers) data on the frequency of adjustment and on the incidence of regular adjustment. Furthermore, the range of the adjustment frequencies in the data needs to be large for the pattern implied by the model to dominate the idiosyncratic actions of firms, i.e. to overcome the number problem.

To test the model we use a very large Austrian data set. It is the data set analyzed in Baumgartner et al. (2005) who studied the stylized facts of price setting in Austria.[15] It contains monthly price quotes collected by the Austrian statistical office, which are used in the computation of the Austrian CPI. The sample spans the period from January 1996 to December 2003 (96 months) and contains about 40,000 elementary price records per month. Overall, the data set contains about 3.6 million individual price quotes and covers roughly 80% of the total Austrian CPI. Each record includes, in addition to the nominal price, the information on the product category, date, outlet (shop) and packaging type.

Testing the model involves the comparison of price behaviour across policymakers. Applied to the pricing set-up, the policymaker is a monopolistic (or monopolistically competitive) seller. She chooses the timing of adjustment as well as the nominal prices to maximize real profits, subject to lump-sum (*menu*) costs of changing nominal prices.

We identify a "policymaker" with a product category, i.e. products at the elementary level included in the CPI basket (e.g. milk), rather than an individual store/product pair. Treating individual store/product pair as a policymaker would require calculating the average frequency of price changes from few observations, especially for stores which change prices infrequently. We need a large number of price changes to compute the conditional frequencies used in the empirical testing. Thus, we implicitly assume that firms operating on the same market (selling the

[15] They describe the data and some manipulations which have been carried out prior to the statistical analysis in detail.

same product) share the same profit function and that the heterogeneity in the profit function is across markets. The original data set (used in Baumgartner et al., 2005) contains a total of 668 product categories. We excluded 151 product categories with administered prices, excessive price changes and products for which we had data for several varieties.[16]. This leaves 517 product categories for our analysis.

The average product category frequency of price changes is between 0.8% per month (chipboard screws) and 91% per month (package holidays). The substantial differences in adjustment frequency and the large number of product categories are promising indications that the number problem may, indeed, "average out".

The main element of the model that determines the incidence of regular policies is the heterogeneity in the curvature of the profit function. Since the curvature is not observable in our data, a direct test of the model is not possible in our framework. However, an indirect test of the model can be performed with other variables of the model, which are observable, acting as instruments for the unobservable variable. This is done by regressing the incidence of regular policies on a set of variables for which the curvature of the profit function implies a certain cross-relation as described in the previous section. If the coefficient signs in this regression are in line with the cross-relations implied by the model, we interpret this as an empirical support of the model. In our case the average frequency of price changes serves as the instrument.

The data allow us to analyze the incidence of both time-regular and state-regular policies. We define a time-regular policy as price adjustment at the beginning of the year, and, separately, as price adjustment at the beginning of a quarter. We will refer to such policies as *seasonal* price setting. State-regular policies involve choosing *attractive* prices: prices that end in a nine or round prices. The definition (values) of attractive prices are in the appendix. The testing involves the analysis of the cross-sectional relationship between the frequency of price adjustment and the excess proportion of seasonal price setting or the excess proportion of attractive prices.

The analysis of this relationship raises the issue of causality. Our model implies that infrequent price changes and high incidence of regular policies coincide because of a common causing characteristic (flat profit function). On the other hand, existing studies in the Inflation Persistence Network imply causation from what we call regular policy to the frequency of price changes. In the data set we are using, Baumgartner et al. (2005) find that the probability of price adjustment, conditional on the last price being an attractive price, is lower than the

[16] We eliminated all products with an average size of price changes of more than 50%.We suspect that, in such cases the definition of the product (on which no direct information is available in the data set) has been changed during the sample period. For some product categories the data set contains prices for several varieties (for example car insurance for different types of cars). These prices are usually changed jointly and so, in such cases, we included only the price for the variety with the highest CPI weight.

unconditional probability. Similar results have been documented by Álvarez and Hernando (2004) for Spain, by Aucremanne and Dhyne (2005) for Belgium, by Veronese et al. (2005) for Italy, by Lünnemann and Mathä (2005) for Luxembourg, by Hoffmann and Kurz-Kim (2005) for Germany and by Dhyne et al. (2005) for a panel of euro area countries. This means that, if we simply looked at the relationship between the frequency of price changes and the incidence of attractive prices, we may discover a negative relationship where causality goes from the proportion of attractive prices to low price changing frequency: in markets in which the proportion of attractive prices is high, the average frequency of price changes will be low.

In order to overcome this potential problem of reverse causality in our regression we have to define a measure for the frequency of price changes that is independent of the proportion of attractive prices. This can be done by conditioning the frequency of adjustment on, separately, attractive and non-attractive prices: for product category *i* we calculate the average conditional frequency of a price change given that the last price is an attractive price, denoted F_i^{att}, as well as the conditional frequency of price changes given that the last price is not an attractive price, denoted F_i^{natt}. We then use both conditional frequencies in the regression as explanatory variables. The use of both conditional frequencies avoids the results being dominated by the mixture of attractive and other prices in the given market.

We suppose the same is true for seasonal price setting: the probability of price adjustment conditional on the previous adjustment taking place at the beginning of the year would be lower than the unconditional probability of adjustment. Therefore we adopt the same approach in the regressions explaining the incidence of seasonal price setting using, as explanatory variables, both the conditional frequency of price change if the last price change was at the beginning of the year/quarter, denoted F_i^{seas}, and the conditional frequency if it was not at the beginning of the year/quarter, denoted F_i^{nseas}.

The estimated regression equations are:

$$Attr_i = f\left(F_i^{att}, F_i^{natt}, \vec{x}_i\right) \tag{7a}$$

$$Seas_i = f\left(F_i^{seas}, F_i^{nseas}, \vec{x}_i\right) \tag{7b}$$

where $Attr_i$ is the proportion of prices in market *i* that are attractive, $Seas_i$ is the proportion of price changes that take place at the beginning of a year (quarter) and $\vec{x}_i$ is the vector of other explanatory variables which are explained below.

We first discuss the results for state-regular policies, i.e. policies under which the price charged is an attractive price. The empirical implementation of the testing requires a definition of attractive prices. There is no universal approach to defining attractive prices. Since results are sensitive to the definition of the phenomenon to be explained, it is important to find a sensible definition of attractive prices, even though it is clear that any definition would be debatable, given its subjective nature. We chose to adopt a broad definition that tries to capture all prices which are used by any firm or retailer as attractive prices. This comes at the risk of classifying too many prices as attractive. We think this is less problematic than missing important attractive prices. We require that the (percentage) differences between attractive prices be not affected by the order of magnitude of the prices (i.e. if 15.90 is an attractive price, so is 159 and 1,590). This is important in our data set as it encompasses the replacement of the Schilling with the euro, which involved the reduction of prices by roughly an order of magnitude (the exchange rate was 13.7603 Schillings/euro). In addition, our definition is specifically tailored to the Austrian retail market as it takes account of the common pricing practices observed there (e.g. prices ending in 75 are not used as attractive prices in Austria). An explanation of the principles of our definition and (an excerpt of) a list of attractive prices are in the appendix. With our definition, the average proportion of attractive prices in the data is 60.7%.

The cross-sectional variations of the share of attractive prices is explained by the variation in the frequency of price changes, conditional on the last price being an attractive price and, separately, on the last price not being an attractive price, the size of price changes and a number of control and dummy variables to account for other factors influencing the incidence of attractive prices. The conditional frequencies of price changes are expected to have a negative effect on the share of attractive prices because, as implied by the model, firms with a relatively flat profit function will change their prices less frequently and will be more likely to choose attractive prices. Similarly, firms with a flat profit function will also change their prices by a larger amount implying that (controlling for inflation) the size of price changes is positively related with the share of attractive prices in the cross section of products.

The control variables include the average price level in the product category, the rate of inflation and its variability (measured by its standard deviation) and the share of sales prices. If attractive prices are more relevant at lower price levels (i.e. for cheaper goods), the average absolute price in a product category should be related negatively to the share of attractive prices. This variable also serves as a check if our definition of attractive prices is reasonable. The coefficient on the average product-specific inflation is expected to be negative since the higher is the average inflation rate in the product category, the more frequent are price changes and the smaller is the share of attractive prices. The model has no implication for the standard deviation of inflation but, in general, we would expect the coefficient to be negative. First, the empirical relationship between inflation and its variability is positive. Second, and perhaps more importantly, in more volatile environment firms can be expected to adopt more flexible policies. Finally, the incidence of attractive prices may be affected by temporary promotions and end-of-season sales; casual observation suggests that these prices are often attractive, and so we include the share of sales prices and promotions in each product category as another control variable in the regressions.

The regression results for the share of attractive prices as the dependent variable are shown in table 1. Note that the share of attractive prices is a fractional response variable (it is bounded between 0 and 1), which implies that estimating a linear model is not appropriate. A common approach in this case, which we follow here, is to transform the dependent variable to the log-odds ratio, $\log\left(\frac{Attr_i}{1-Attr_i}\right)$ which is not bounded, and run an OLS regression on the transformed variable[17]. In order to get the marginal effect of each variable on the dependent variable, the regression coefficients, β_k, have to be converted back by the formula $\frac{dy}{dx}=\beta_k\ \overline{Attr}\left(1-\overline{Attr}\right)$ which usually is evaluated at the sample mean. The results in table 1 are quite consistent with the model. The frequency of price changes (conditional on the last price being an attractive price, F_i^{att}) has a negative impact on the share of attractive prices, as predicted by the theoretical model, and this effect is significant at the 10% level. Specifically, the marginal effect implies that, if the conditional frequency increases by one percentage point, the share of attractive prices is decreased by 0.75 percentage points. The conditional frequency if the last price was not an attractive price (F_i^{natt}), however, has a positive impact on the share of attractive prices. While the model clearly implies a negative sign for the first conditional frequency, F_i^{att}, its implications for F_i^{natt} are less clear. The sign could be negative if the conditioning of the frequency is empirically not relevant. A positive sign is reasonable F_i^{natt} if we assume that firms have a strong incentive to follow an attractive pricing policy, i.e. if they have a very flat profit function, but for some reason sometimes deviate from that policy and choose price that is not attractive. But if they do so, they quickly return to an attractive price afterwards, which increases the conditional probability of a price change when the last price was not attractive.

[17] The log-odds model has been criticized for delivering marginal effects that may be inconsistent. An alternative approach used in Dhyne et al. (2005) is the quasi-maximum likelihood (QML) approach proposed by Papke and Wooldridge (1996). It involves directly estimating a non-linear model of the explanatory variables and maximizing its likelihood function based on a Bernoulli distribution. We also performed estimations according to this approach, but the results (available upon request) are very similar.

Table 1: Explaining the Share of Attractive Prices

Variable	Long Sample (96-03) Marginal Effect		Schilling Sample (96-01) Marginal Effect	
Constant	0.231	***	0.371	***
Frequency cond. on attr (F_i^{att})	-0.745	*	-0.130	
Frequency cond. on not attr (F_i^{natt})	0.649	*	-0.189	
Size of price changes$_i$	0.622	***	0.552	**
Av. Price$_i$ (Schilling)	0.000		0.000	
Av. Price$_i$ (Euro)	0.000			
Av. Inflation$_i$	-0.102	**	-0.132	***
Stdv. Inflation$_i$	0.001		0.006	
Group processed food	0.008		0.023	
Group energy	-0.528	***	-0.611	***
Group industrial goods	-0.284	***	-0.360	***
Group services	-0.315	***	-0.315	***
Share of sales prices$_i$	0.830	**	0.919	**
Number of observations	505		507	
Adjusted R^2	0.417		0.356	

Notes: *Estimation method is OLS on the log-odds ratio of the share of attractive prices; displayed coefficients are marginal effects of each variable on the share of attractive prices evaluated at the sample mean; standard errors are computed using White's correction for heteroskedasticity; inflation is calculated as monthly changes of the corresponding product category's sub-index; the number of products included is lower than the maximum 517 because some variables are not defined for all products; *** denotes significance at the 1%, ** at the 5% and * at the 10% level.*

The average (absolute) size of price changes in a market has a positive impact on the share of attractive prices in this market, as predicted by the model. The average price in the product category, which has been calculated and included in the regression for the Schilling period (1996–2001) and the euro period (2001–2003) separately, does not affect the incidence of attractive prices. This result is reassuring as it indicates that, if attractive prices in the data are equally distributed across the price spectrum, the definition of attractive prices has been chosen appropriately. Furthermore, average (monthly) inflation in a product category has a significant negative impact on the share of attractive prices in that category as predicted by the model, while the volatility of inflation (measured by the standard deviation over the sample period) has no significant impact. Finally, the practice of sales and temporary promotions turns out to be an important additional determinant of attractive prices: the product categories with a higher share of sales and promotions are characterized by a higher share of attractive prices.

The dummy variables for product groups are included to account for product group fixed effects. The group dummies which are included in the regression are defined according to the five product groups used by the ECB to analyze inflation dynamics in the euro area: unprocessed food, processed food, energy, non-energy industrial goods and services. Unprocessed food is used as the reference group and is therefore not included in the regression. It is important to account for these fixed effects as there is extensive evidence that the frequency of price changes varies greatly across product groups (Baumgartner et al., 2005 provide the evidence for the data set we use; Dhyne et al., 2005 summarize these differences for ten euro area countries). The results indicate that the share of attractive prices is significantly lower for non-food items.

To check whether attractive price setting was not systematically different for Schilling and for euro prices, in column 2 we show the regression results obtained for the sample period covered by our dataset when the Schilling was the legal tender in Austria (1996–2001).[18] Overall, the results for the short sample are qualitatively similar to the long sample. The exception is that the frequency of a price change, conditional on the last price not being attractive price has a negative sign and neither conditional frequency is significant. The results for the longer sample are thus more in line with the theoretical model.

We now turn to the analysis of time-regular policies. We implement the model by looking at the determinants of the excess proportion of price changes taking place at the beginning of a year and, separately, at the beginning of a quarter; such behaviour will be called seasonal adjustment. Empirically, price changes in the Austrian data are, indeed, more frequent at the beginning of the year and, for some products, also at the beginning of a quarter (see Baumgartner et al., 2005).

According to the implications of the model, the same line of reasoning as for attractive prices applies to the share of price changes at the beginning of a period. Firms which have a flatter profit function will change their prices less frequently, by a larger amount and prefer a seasonal pattern of their price adjustment, i.e. have a larger proportion of price changes at the beginning of a period. Thus, in a large cross-sectional data set the share of price changes at the beginning of a period should be negatively related to the (conditional) frequency of price changes and positively to the average size of price changes. As in the regression for attractive prices, the average product-specific inflation and inflation volatility as well as the product group dummies and the share of sales prices have been included in the regression as additional control variables.

[18] The sample period form the introduction of the euro to the end of our sample (2002–2003) is too short to be analysed separately.

One important difference between the seasonal pattern and attractive prices is that, in some industries, firms tend to change prices together. For example, clothing stores hold simultaneous sales. This tendency to synchronize price changes needs to be controlled for so as to avoid spurious correlation between seasonal patterns and the conditional frequencies of adjustment. Therefore we include, on the right hand side of the regression, the synchronization index of price changes as defined by Fisher and Konieczny (2000). It summarizes, with a single number, the tendency of prices to be changed together. The index is defined as the ratio of sample standard deviation of the monthly proportion of price changes for a given product category to the standard deviation of the proportion under the assumption that price changes are perfectly synchronized.

The dependent variable in this regression is the ratio of the number of price changes taking place at the beginning of the period to the number of all price changes in that period, normalized to avoid it being bounded. Given that our data are monthly we adopt two definitions of a period: a year and a quarter. In yearly regressions we compute the ratio of the number of price changes in a January of any year to all price changes in the sample; in quarterly regressions we compute the ratio of the number of price changes in any January, April, July or September to the number of all changes in the sample. The (normalized) dependent variable is obtained by dividing the yearly (quarterly) statistics by the share of valid price observations at the beginning of the year (quarter). According to this definition, a number above 1 indicates that relatively more prices are changed at the beginning of the period than average. The resulting dependent variable is not bounded and OLS can be applied in the estimations. For a robustness check we also run equivalent regressions with the (log-odds ratio of the) non-normalized share of price changes at the beginning of a period as the dependent variable. The results are qualitatively very similar.

The regression results for seasonal price setting, shown in table 2 are also broadly consistent with model’s implications. Table 2 shows the results for period defined as a year (column 1) and period defined as a quarter (column 2). Of the two specifications, price setting at the beginning of a year is empirically more relevant (the mean of the dependent variable is 2.01, indicating that price changes in January are 101% more frequent than in the other months of the year) than price adjustment at the beginning of a quarter (with a mean dependent variable of 1.16). Therefore, we regard the first column in the table as our standard specification and treat the results for price setting at the beginning of a quarter as an additional specification for a robustness check.

Table 2: Explaining the Share of Price Changes at the Beginning of a Period (Year, Quarter)

Variable	Period = Year Coefficient	Period = Quarter Coefficient
Constant	0.591 ***	0.904 ***
Frequency cond. on seas (F_i^{seas})	-2.926 ***	-1.771 ***
Frequency cond. on not seas (F_i^{nseas})	-0.650	1.367 ***
Size of price changes$_i$	1.635	1.020 ***
Av. Inflation$_i$	0.592 ***	0.023
Stdv. Inflation$_i$	-0.039	-0.007
Group processed food	-0.117	0.011
Group energy	-0.293	0.102
Group industrial goods	-0.116	0.062 ***
Group services	0.552 ***	0.038
Share of sales prices$_i$	-1.371	-0.741 *
Synchronization of price changes$_i$	5.643 ***	0.676 ***
Number of observations	491	480
Adjusted R^2	0.458	0.221

*Notes: Estimation method is OLS; standard errors are computed using White's correction for heteroskedasticity; inflation is calculated as monthly changes in the corresponding product category's sub-index; the number of products included is lower than the maximum 517 because some variables are not defined for all products; *** denotes significance at the 1%, ** at the 5% and * at the 10% level.*

The crucial result is that the sign on both conditional frequencies, i.e. if the last price change was at the beginning of a year (F_i^{seas}) or was not at the beginning of a year (F_i^{nseas}) is negative, as predicted by the model. The coefficient on F_i^{seas} which, as argued before, is more relevant in terms of the theoretical model, is significant at the 1% level. In other words, in markets where prices are changed infrequently, a large proportion of these changes take place in January. Note that in the regression we control for the synchronization of price setting. While the index is not a perfect control[19], the inclusion of the index in the regression reduces the likelihood that the negative sign is due to some markets being characterized by yearly price changes in January only.

[19] It leaves several degrees of freedom as it summarizes, with just a single number, the monthly pattern in the proportion of price changes.

The coefficient on the size of price changes has the expected positive sign (recall that, given the inflation rate, the size of price changes is inversely related to the frequency of adjustment) but the effect is only marginally significant (at the 11% level). An unexpected result is that the average product specific inflation has a positive effect on the share of price changes at the beginning of the year. This is at odds with the theoretical model, which implies that the higher is inflation, the more frequent are price adjustments and the less likely are firms to adjust prices at predefined dates. The coefficient on inflation volatility is negative, as expected, but the effect is not significant. Only services show a significantly higher share of price changes at the beginning of the year than the reference group (unprocessed food), which is related to the fact that many service prices in Austria are regularly changed in January (see Baumgartner et al. (2005)). The commercial practice of sales and temporary promotions is obviously not an important determinant of seasonal price setting in January: the coefficient on the sales variable is negative but not significant. Finally, the coefficient on the synchronization variable is positive and significant at the 1% level. This indicates that in markets where firms synchronize price changes, adjustment in January is frequent.

The regression results for the quarterly pattern of adjustment, shown in the second column of table 2, are similar. The results are qualitatively equivalent to those in the second column with a few exceptions. The frequency of price changes conditional on the last price change not at the beginning of a quarter (F_i^{nseas}) has a positive sign and group effects are somewhat different. Significance patterns are also a bit different. Although some results of this specification (e.g. for the size of price changes and average inflation) are more in line with the theoretical model, it is not our preferred specification as its fit measured by an adjusted R^2 of 0.22 is much lower than in the previous regression; this is not surprising given the quarterly seasonal pattern in price adjustment is much weaker than the yearly pattern.

To sum up, the regression results for both, the share of attractive prices and the share of price changes at the beginning of a period, support the cross-sectional implications of the model developed in the previous section: in markets which are characterized by a low adjustment frequency (independent of the adjustment to attractive prices), large price changes and lower average inflation, we find a high share of attractive prices as implied by the model. And in markets with low adjustment frequency (independent of the seasonal adjustment), large price changes and a higher synchronization of price changes, the share of price changes at the beginning of a year (and a quarter) is high. The only result that is not consistent with the model and cannot readily be explained with other common price-setting practices is the positive relation between average inflation and the share of price changes at the beginning of the year. But all other results are broadly consistent with the model and/or can be rationalized by the stylized facts of price setting in

Austria. Thus, we conclude that our theory is clearly supported by the cross-sectional relations in the Austrian micro CPI data.

4. Conclusions and Extensions

Regular adjustment is ubiquitous in many environments, yet the reasons for such behaviour have not received much attention. In this paper we make a small step towards explaining the incidence of regular adjustment. It is attributed to the heterogeneity in adjustment costs over time and/or levels and the heterogeneity in the shape of the benefit function across policymakers.

The empirical results obtained from a large Austrian data set are consistent with the model. As the benefit function heterogeneity is not observable, we show that the model implies a negative relationship between the average frequency of adjustment and the incidence of regular policies. We treat adjustment frequency as an instrument and find that firms which change prices infrequently often choose regular policies, by setting attractive prices or by adjusting at the beginning of a year or of a quarter.

An alternative source of differences is heterogeneity in price adjustment costs across policymakers as in Dotsey, King and Wolman (1999). In their model firms are otherwise identical but the costs of changing nominal prices differ across firms. These differences lead, in turn, to different frequencies of price changes. With adjustment costs heterogeneity across policymakers, the implied correlation between the frequency of adjustment and the incidence of regular policies is positive. The reason is that, whenever its adjustment cost is low and so the firm changes prices often, the profit function is flat over the variation of the real price and the firm is likely to choose a regular policy. Therefore our empirical results do not support the joint hypothesis that adjustment costs vary across time (or levels) and across firms.

Why are regular policies important? Policymakers who adopt regular adjustment reduce their flexibility. The understanding of the costs and benefits of flexibility is not only of intrinsic importance to these policymakers but is also important for more general considerations. For example, monetary policy is more effective when nominal price adjustments are regular.

One way of viewing state-contingent (as opposed to regular) adjustment is that it provides the option of flexibility, at the cost of higher adjustment costs. This may result in hysteresis. The value of the option is lower under low and more stable inflation. Imagine that there is a setup cost of switching to regular adjustment, for example the expense on the organization of the work flow. A period of monetary stability may lead firms to switch to regular policies and, once the sunk cost has been paid, even when monetary stability falls, some firms may not abandon regular policies.

Finally, the effect of low inflation on the effectiveness of monetary policy depends on the source of the stability. If the reason inflation has been low and stable in recent years is mostly due to monetary policy, then we can expect greater incidence of regular policies and increased monetary effectiveness. On the other hand, assume inflation is low because of increasing competition. This raises demand elasticity and, so, by increasing the concavity of profit functions, may lower the incidence of regular price adjustments and so reduce the effectiveness of monetary policy.

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Appendix

A.I Definition of Attractive Prices for the Schilling Period (1996–2001)

Attractive prices are defined for prices ranges in order to take account of different attractive prices at different price levels: from 0 to 10 Austrian Schillings (ATS) all prices ending at x.00, x.50 and x.90 ATS, from 10 to 100 ATS all prices ending at xx0.00, xx5.00 and xx.90 ATS, from 100 to 1,000 ATS prices ending at xx0.00, xx5.00 and xx9.00 and xxx.90 ATS and so on. An equivalent rule has been defined to identify attractive prices in euro after the cash changeover (2001–2003). Table A1 shows an excerpt of a list of attractive prices for the Schilling case. In order to give a complete list of attractive prices, the table would continue to the right and to the bottom. The extension to the right would show multiples of 10 and 100 of the last four columns.

Table A1: Attractive Prices for the Schilling Period

below 1	1-9.99	10-99.99		100-999.99			1000-9999.99			
					105.00	see col. P		1050.00	see col. Q	see col. R
0.50	1.00	10.00	10.90	100.00	109.00	100.90	1000.00	1090.00	1009.00	1009.90
					115.00	101.90		1150.00	1019.00	1019.90
0.90			11.90	110.00	119.00	102.90	1100.00	1190.00	1029.00	1029.90
					125.00	103.90		1250.00	1039.00	1039.90
			12.90	120.00	129.00	104.90	1200.00	1290.00	1049.00	1049.90
					135.00	105.90		1350.00	1059.00	1059.90
			13.90	130.00	139.00	106.90	1300.00	1390.00	1069.00	1069.90
					145.00	107.90		1450.00	1079.00	1079.90
			14.90	140.00	149.00	108.90	1400.00	1490.00	1089.00	1089.90
					155.00	109.90		1550.00	1099.00	1099.90
	1.50	15.00	15.90	150.00	159.00	110.90	1500.00	1590.00	1109.00	1109.90
					165.00	111.90		1650.00	1119.00	1119.90
			16.90	160.00	169.00	112.90	1600.00	1690.00	1129.00	1129.90
					175.00	113.90		1750.00	1139.00	1139.90
			17.90	170.00	179.00	114.90	1700.00	1790.00	1149.00	1149.90
					185.00	115.90		1850.00	1159.00	1159.90
			18.90	180.00	189.00	116.90	1800.00	1890.00	1169.00	1169.90
					195.00	117.90		1950.00	1179.00	1179.90
	1.90	19.00	19.90	190.00	199.00	118.90	1900.00	1990.00	1189.00	1189.90
					205.00	119.90		2050.00	1199.00	1199.90
	2.00	20.00	20.90	200.00	209.00	120.90	2000.00	2090.00	1209.00	1209.90
					215.00	121.90		2150.00	1219.00	1219.90
			21.90	210.00	219.00	122.90	2100.00	2190.00	1229.00	1229.90
					225.00	123.90		2250.00	1239.00	1239.90
			22.90	220.00	229.00	124.90	2200.00	2290.00	1249.00	1249.90
					235.00	125.90		2350.00	1259.00	1259.90
			23.90	230.00	239.00	126.90	2300.00	2390.00	1269.00	1269.90
					245.00	127.90		2450.00	1279.00	1279.90
			24.90	240.00	249.00	128.90	2400.00	2490.00	1289.00	1289.90
					255.00	129.90		2550.00	1299.00	1299.90
	2.50	25.00	25.90	250.00	259.00	130.90	2500.00	2590.00	1309.00	1309.90