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Micro Evidence on the Adjustment of Sticky-Price Goods: It's How Often, not How Much

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MICRO EVIDENCE ON THE ADJUSTMENT OF STICKY-PRICE GOODS: IT'S HOW OFTEN, NOT HOW MUCH

Lorenz Goette, Rudolf Minsch and Jean-Robert Tyran*

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We use a unique panel data set to analyze price setting in restaurants in Switzerland 1977-93, for items known to have sticky prices. The macroeconomic environment during this time period allows us to examine how firms adjust prices at low (0%) and fairly high (7%) inflation. Our results indicate that firms strongly react to inflation in the timing of their price adjustment: hazard of price changes is increasing with time and becomes steeper at higher inflation rates. However, we find little evidence that the amount by which they change the price responds to the inflation rate.

Keywords: Sticky prices, Inflation, Nominal inertia
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1 Introduction

Knowing the exact form of price stickiness and the mechanisms behind price adjustment is of critical importance in understanding the equilibrium dynamics of, e.g., monetary shocks to an economy (Dotsey, King, and Wolman 1999, Ball, Mankiw and Reis 2004). Recent access to the micro base of the CPI in several countries has produced a first wave of new results with respect to the extent of price stickiness. The results from two (sets of) studies from the US (Bils and Klenow 2004) and the Euro area (Dhyne et al. 2005) show pronounced heterogeneity in the extent of price stickiness across sectors. Non-processed and processed food items have relatively little price stickiness, while services exhibit more sticky prices.

These earlier studies make progress in understanding the price adjustment of goods with quick price adjustment. The aim of this study is to examine price adjustment of sticky price items. Two factors limit the earlier studies in addressing this issue. First, examining the behavior of firms with very sticky prices requires long panels of prices, but such data is scant (Bils and Klenow, 2004). Second, theories of price stickiness make specific predictions about how firms respond to changes in inflation. But the sample periods of earlier studies cover phases with insufficient variation in inflation, such as, e.g., the Euro area in recent years.

In this study, we use part of the CPI base of Switzerland between 1978 and 1993. We use data on prices from restaurants, a sector known from earlier studies for long periods of price stickiness. We choose highly standardized items that account for a large revenue share of the restaurants and can be tracked over a long time. The macroeconomic environment in Switzerland over this time period contains long stretches of low inflation, but also sharp rises to up to 7 percent. This allows us to examine how firms adjust prices of goods known for their price stickiness.

Observing long individual price series also allows us to directly *observe* the duration of price quotes and hazard rates of price changes, i.e. the probability of price adjustment conditional on a particular individual history of non-adjustment. In contrast, earlier studies (e.g., Bils and Klenow, 2004) calculate implied durations from the frequency of price changes, requiring assumptions about the shape of hazard rates. Recent models of price adjustment make specific predictions about the shape of hazard functions and how they change with inflation. For example, state-dependent pricing models predict *increasing* hazard functions (i.e. price-setters are more likely to adjust the longer they have not adjusted) while time-dependent theories predict flat hazard rates with spikes at regular intervals and no relation with inflation. Recent studies find decreasing hazards of price adjustment (e.g. Dhyne et al. 2005). But declining hazard functions cannot be taken as evidence against state-dependent pricing models unless one tightly controls for the types of goods: If the goods in a sample have inherently different pricing hazards (i.e., if one mixes "flexible-price" goods and "sticky-price" goods in the analysis), this can easily lead to downward-sloping hazards, even though each individual hazard function might be increasing. The reason is that the flexible-price goods exit at a higher frequency early on, leading to a downward-sloping overall hazard function. The items we consider in our study have similar baseline hazards, making the evidence on the slope of hazard functions more credible.

Our main findings are as follows. We find strong evidence that firms react to inflation with the timing of price adjustments. At higher inflation rates, we observe more frequent price adjustment. The effect is quantitatively large. The examination of the hazard functions is supportive of state-dependent pricing models: We find that the pricing hazard is upward sloping, as these models predict. We also find that it becomes steeper as inflation increases. However, the size of the price increases barely seems to respond to changes to the inflation rate, even when we control for different durations. Taken together, our results are suggestive of a model where the main choice variable is the *timing* of the price adjustment, and much less the size of it. Recent models that incorporate fairness concerns on the side of consumers (Rotemberg 2002, 2004) may account for this result.

2 Description of data

The Federal Office has collected our data for Statistics for the purpose of calculating the Consumer Price Index (CPI) in Switzerland. Data were collected quarterly (in the first two weeks of February, May, August, November) and show the price of a specific item in a particular restaurant. We analyze the prices of four items served in restaurants over 63 quarters in the range 1977:Q4 to 1993:Q2.[1] The five items under investigation are two beverages, a bottle of mineral water (300 ml) and a cup of coffee (served with cream), and three dishes, spaghetti with meat sauce (spaghetti Bolognese), a sirloin steak with a side dish (entrecôte café de Paris), and a sausage salad (cold dish).

Our dataset is unique and features several desirable qualities. First, the observed prices are true transaction prices that can be fully flexible. In our data, there are no sales, rebates or special offers (in contrast to scanner data from supermarkets, see Chevalier, Kashyap and Rossi 2000). In addition, tips and taxes are fully included Switzerland. Restaurant prices are free to change at any time. In particular, there are no government regulations on restaurant prices, and there are no long-term contracts involved in restaurant prices under study.

Second, the items under study are core business items for the restaurant, and price-setters therefore have strong incentives to choose optimal prices. For example, spaghetti was the single most important item on the Swiss restaurants' menu in 1997 (Gastrosuisse 2001). According to the survey of the Swiss restaurant association Gastrosuisse, coffee and tea contributed 11.3 percent to the total sales of restaurants, mineral water 10.6 percent.

Third, the chosen goods have approximately constant quality attributes. To be able to study relatively long (in our sample 16 years) time-series is that quality attributes of the goods must be constant over time. Restaurants may differ with respect to location or ambiance. Therefore, even a physically identical product like a bottled mineral water may have different (cross-sectional) quality attributes when served in different restaurants. However, for our purposes, an important property of the data set is persistency of quality attributes over time. Fortunately, the Swiss federal office of statistics did a very careful job. For example, it started a new time series when the owner changed or when the restaurant underwent a major renovation. A

[1] Unfortunately it is not possible to analyze time periods after 1993 as the Federal Office of Statistics changed several aspects of the price collection procedure. In particular, new codes for the individual restaurants were introduced which makes relating pre- and post-1993Q2 data impractical.

further advantage of our data is that quality adjustments as a substitute for price adjustments are rather improbable.

Fourth, we have chosen the four items served in restaurants because they are standard examples for “sticky-price goods”. For example, Bils and Klenow (2004) find implied mean durations of unchanged prices for “beer, ale, other alcoholic beverages away from home” of 15.2 months, “distilled spirits away from home” 12.1 months and “breakfast or brunch” 11.4 months. In Europe, prices seem to be stickier in general than in the US (see Dhyne et al. 2005 for an overview). For example, in Italy, the implied mean duration of a “glass of beer in café” is 27.9 months; “a meal in a restaurant” is 17.9 months (Veronese et al. 2005, table A3.3). Our findings are comparable to those from other European countries. For example, we find a mean duration for mineral water of 23.7 months (5.92 quarters), and for spaghetti of 21.8 months (5.44 quarters).

Fifth, we have extraordinarily long individual price series (of up to 63 quarters). This allows us to study the adjustment of our sticky-price goods in detail at the individual level. For example, we can calculate real price erosion, durations of price quotes and hazard rates (see below for definitions) for an item in a particular restaurant. The length of our series allows us to observe durations of up to nine years. While there are other studies with long horizons, and studies with a larger cross-sectional dimension, our data set is much richer in the cross sectional dimension than studies with a comparable length of individual price series with constant quality attributes in low-inflation environments. For example, the study by Kashyap (1995) analyzes the behavior of three price-setters over a period of 35 years with a total of 12 goods (e.g. a duffel bag or chamois shirt).

Sixth, our observations come from an environment with low but variable inflation. The average annual inflation in Switzerland over the period under study was at 3.53%, and annualized quarterly inflation rates (to the corresponding quarter of the previous year) range from -0.1% to 7.2%. In contrast, the studies run in the context of the Inflation Persistence Network in several European countries (Dhyne et al. 2005) observe an environment with low (around 2%) and stable inflation in which there is less scope to study the effect of changes in inflation on pricing.[2]

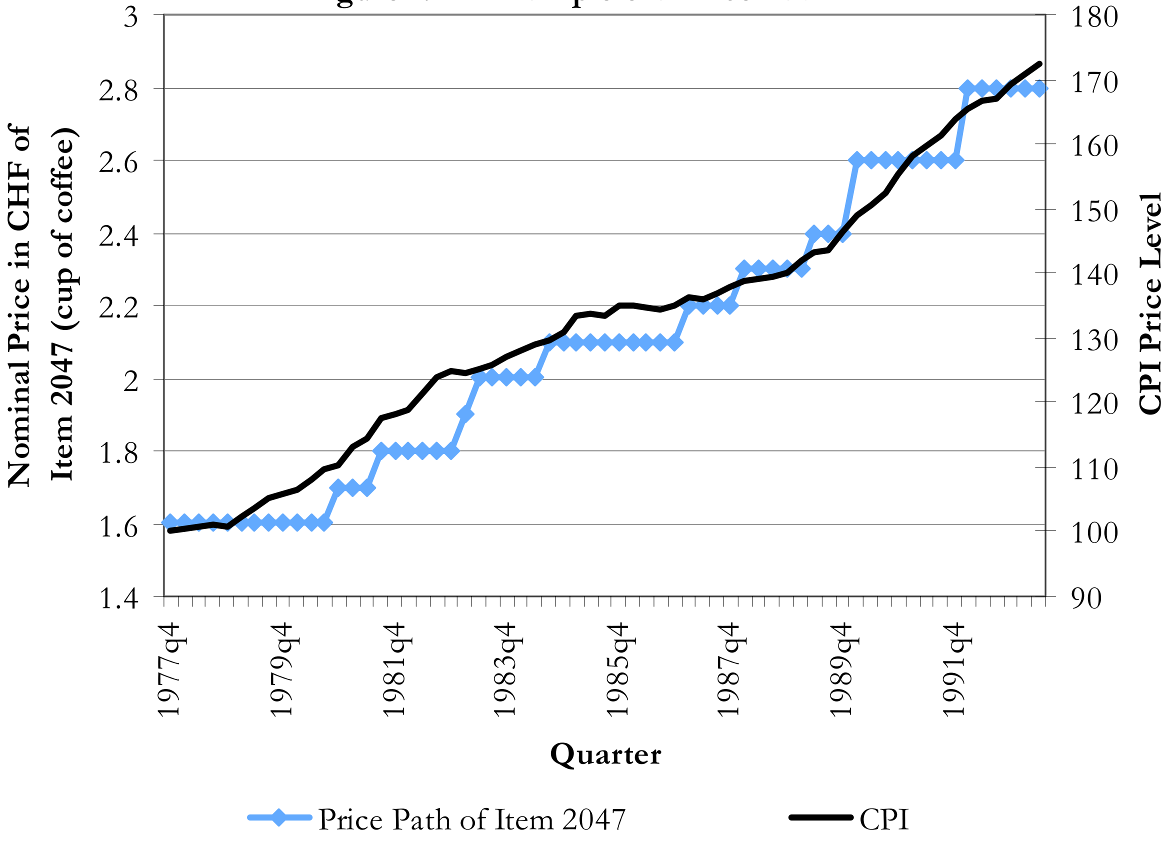
[2] The studies also have shorter time horizon (of 6 to 11 years) around the Euro-changeover.

2.1 An example of a price series and some definitions

We use an example of a price series for one particular item to explain some of the concepts we use below. The individual price series of *figure 1* has also been selected because it is typical in many ways for the data set we analyze (see section B).

Figure 1 shows the nominal price of a cup of coffee served in one particular restaurant (left scale, item 2047 in our database) and of the price level in Switzerland (CPI, right scale) from the 4th quarter in 1977 to the 3rd quarter in 1993. The dots in the individual price series show price observations (63 quarters in total).

Figure 1: An Example of a Price Path



The first price observation in figure 1 for the cup of coffee in this restaurant is at a *nominal price* of CHF 1.60. In the 3 years following 1977Q4, the nominal price of item 2047 was held constant, and was then increased by CHF 0.10 which corresponds to a *relative price increase* of 6.25%. Over the entire period, the restaurant increased its nominal price 10 times, and the last nominal price we observe is at CHF 2.80. In our analysis, we chose to *fully censor* the individual price series. This means that we define the number of quarters between two observed nominal price changes as a *price spell*, and we call the length of the spell the *duration* of a nominal price quote. For example, the first three spells in figure 1 have a length of 3, 6, and 1 quarters.[3] Full censoring implies that the observations before the first and after the last observed price change are not included in the analysis. In the example, this means that the first 12 quarters and the last 6 quarters are not included in the analysis (in the whole sample, the mean is 30.9 quarters, and 25% of the items can be observed for more than 53 quarters).

We calculate the *real price erosion* for each item as the percentage by which the real price eroded over the duration. For example, the real price eroded by 3.71% over the first spell (i.e., the 3 quarters from 1980Q4 to 1981Q2) while the nominal price was held constant at CHF 1.70 but the CPI rose from 110.38 to 114.49 (i.e., rose by 3.71%).

2.2 Descriptive Statistics

Item 2047 shown in *figure 1* is typical in a number of ways for the dataset, as can be seen from *table 1*, where we show the price growth for the different products in our sample. For example, the nominal price of this item increases from a level of CHF 1.60 in 1977Q4 to CHF 2.80 in 1993 and the average nominal price for coffee was CHF 1.93 over the entire period. The total increase for this item is 75.0% (or, 3.61% annual rate), which is very similar to the average increase of all coffees in our sample (annual rate 3.79%). Table 1 also shows that the different products behave quite similarly in terms of average price growth. An important aspect of our data is that prices in our sample on average increased more than the CPI.

[3] We implicitly assume that prices only change once a quarter if they change at all. We also assume that the timing of changes is the same across all quarters. While this might be a problem if price spells were typically shorter than one period, the issue becomes negligible if prices are sticky for more than a quarter. In 8.3 percent of all cases, we observe price changes in two consecutive periods.

Table 1: Average nominal prices and price changes

	Number of Observations	Mean nominal price 1977:4 – 1993:2	Avg. nominal price change (annualized, 1977:4 – 1993:2)
Coffee	11,007	1.93	3.79%
Mineral Water	12,511	2.39	3.73%
Spaghetti	7,475	8.67	4.37%
Steak	7,870	21.99	3.31%
Sausage Salad	7,126	7.59	4.82%
Total	45,989	--	4.06%

Table 2 shows the average duration of a price spell (in quarters) and the average price erosion over the time period over which the price was held constant. In line with the evidence from earlier studies, we find a substantial amount of price rigidity in our data. Prices are kept constant for approximately six quarters, with only small differences across the products. Over these periods, the real prices eroded by approximately 5 percent, again with only minor differences between products.

Table 2: Average duration of price quotations (number of quarters) and real price erosion

	Duration		Erosion (relative to CPI)	
	Mean (s.d.)	Maximum	Mean	Maximum
Coffee	5.38 (3.27)	19	5.09%	20.94%
Mineral Water	5.92 (4.24)	38	5.76%	35.06%
Spaghetti	5.44 (4.23)	30	5.47%	27.69%
Steak	5.01 (3.83)	33	5.03%	26.71%
Wurst Salad	5.43 (4.15)	34	5.53%	34.02%
Total	5.46 (3.92)		5.36%	

We found no evidence of an attrition bias. In particular, the items starting at 1977Q4 but ending before 1993Q3 do not have systematically different distributions of price changes than the items extending through the entire sample (Mann-Whitney tests, $p < 0.05$ for each good).

How do the descriptive statistics compare to the literature? In our data set, 42 percent of all spells involve erosion of real prices of 5 percent or more before nominal prices adjusted, and 8.5 percent of all spells involve erosion of real prices of 10 percent or more. However, these rates of real erosion are much smaller than those found in Cecchetti (1986). He reports (p. 426) that the average inflation since the last nominal price change almost always exceeds 10 percent, and values of 20 percent are quite common. A possible reason why we observe less extreme price stickiness than Cecchetti could be that our items are core business items, while the revenues from newsstand sales may be relatively unimportant for the large printing houses.

The descriptive statistics of our data are also similar to Kashyap (1995). He finds an average duration of 4.9 quarters, and s.d. of 5 quarters (Kashyap 1995: 252). He finds that catalog prices in the U.S. typically are unchanged for more than a year and that there is considerable heterogeneity in the average duration.

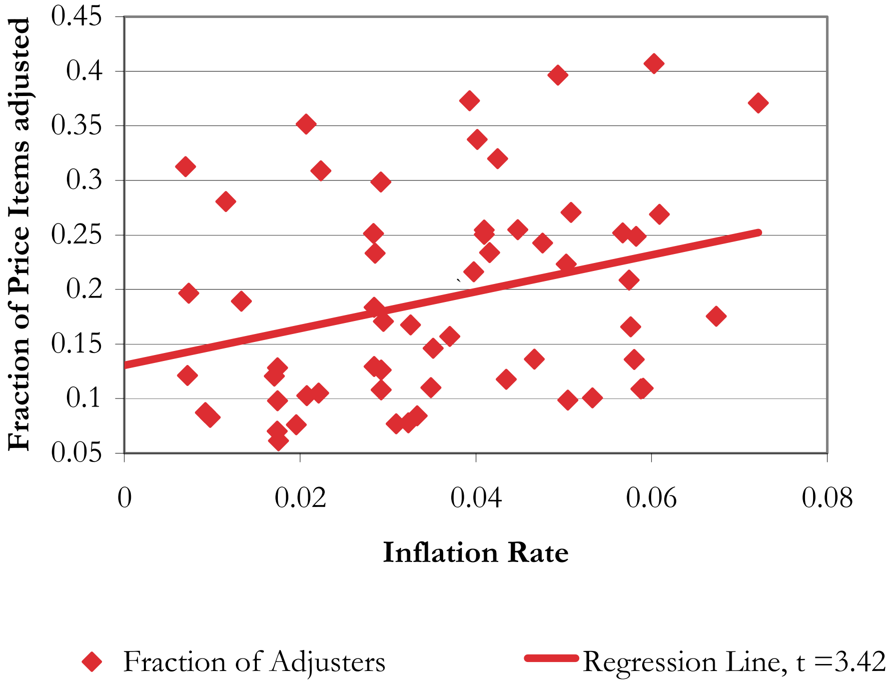
3 Evidence on Price Adjustment

3.1 The Frequency and Size of Price Adjustments

Figures 2 and *3* display the basic pattern of price adjustment in our data. *Figure 2* plots the fraction of items in the sample that adjusted prices against the annualized quarterly inflation rates. As can be seen from this figure, there is substantial variation in the fraction of price adjustments, but in general, a higher inflation rate leads to more frequent price adjustment. The impact of inflation on the frequency of adjustment is quite large and statistically significant (as in all the figures below, we use robust standard errors to calculate the *t*-statistic): An increase in the inflation rate from 1 percent to 6 percent leads to a more than 50 percent increase in the adjustment frequency, from 15 percentage points to almost 25 percentage points.

In stark contrast, *Figure 3* displays the relationship between nominal price changes and the inflation rate for each quarter. The red squares in Figure 3 mark the average price increases of the price adjusters in each quarter. The figure shows a weak relationship that is just barely statistically significant: The coefficient of a regression of the average price change in a given quarter on the inflation rate is 0.18. This means that raising the inflation rate from 1 percent to 6 raises the nominal price increase from 6 percent to approximately 7 percent (see upper regression line in *figure 3*).

Figure 2: The Frequency of Price Adjustment and Inflation, 1978 - 1993 (quarterly data)



These results are similar to what was found in other studies: For example, Lach and Tsiddon (1992) find that despite large differences in the inflation rate in their sample, the amount by which prices are typically adjusted does not vary much. Similar findings are reported in Kashyap (1995) and, more recently, Ratfai (2004). For example, Cecchetti (1986) reports that the average price increase in his sample of magazines rose from 23.5% (in the 1960s) to 25.3% (in the 1970s) while the average CPI inflation rose from 2.4% to 7.1%. The implied slope of 0.38 in Cecchetti's data[4] is slightly larger than ours (0.18).

As pointed out in Rotemberg (2004), this is somewhat puzzling from the point of view of a state-dependent pricing model (Sheshinsky and Weiss, 1977; Dotsey, King and Wolman, 1999, Dotsey and Wolman, 2004). In these models, firms face an administrative cost of

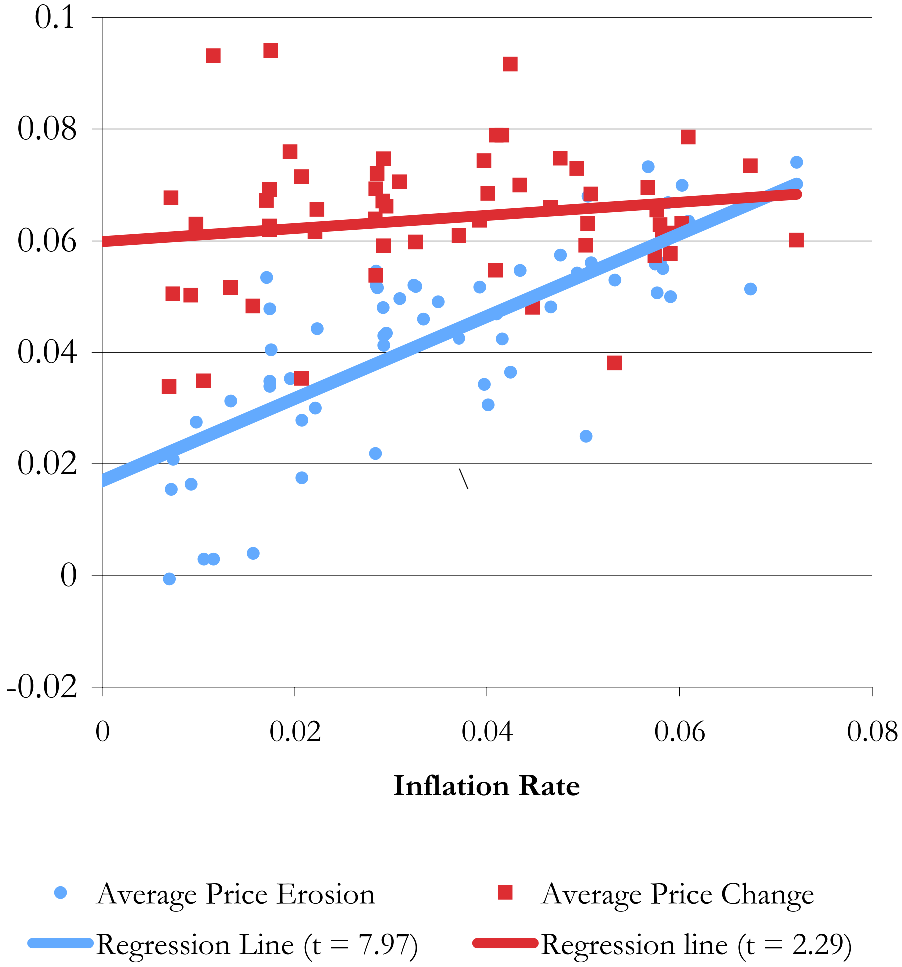
[4] This number is taken from Rotemberg (2004: 36).

changing the nominal price. They set a (S,s) price band within which they let the real price of their good erode and then reset the price. In the face of higher inflation, it is optimal to widen the (S,s) band, and – as a consequence – raise prices by more when adjusting. This is a fairly robust property of this class of models. Only under further restrictions on the profit function (see Dotsey and King, 2004, for a discussion) is it also true that this implies more frequent price adjustment. Our results, and others, seem to suggest that firms leave their (S,s) band approximately constant, and increase their prices more often when inflation is high.

Our data allows looking at this issue from a different angle: If it is true that firms use approximately the same (S,s) band irrespective of inflation, it should also be true that the real price erosion at adjustment is approximately invariant to inflation, because the two are equivalent according to the logic of (S,s) models. On the other hand, if the firms' main concern is the timing of a price change, then we should see a stronger relationship between real price erosion and inflation than between price changes and inflation.

More specifically, if we regress the real price erosion since the last price adjustment on inflation in the quarter of adjustment, the coefficient should be approximately equal to 0.18, the result from the regression of the price changes. Yet, the two approaches yield quite different results. The blue circles in *Figure 3* display the mean price erosion of the price adjusters in each quarter. It is puzzling that the relationship between price erosion at adjustment and inflation is quite steep (the regression coefficient is 0.78). The slope of the regression line is significantly greater than for price changes ($p < 0.001$), but significantly less than one ($p < 0.01$). These results suggest that the choice variable of the firms is much less how much to adjust their prices, but rather when to adjust.

Figure 3: Price Changes, Price Erosion and Inflation Rate



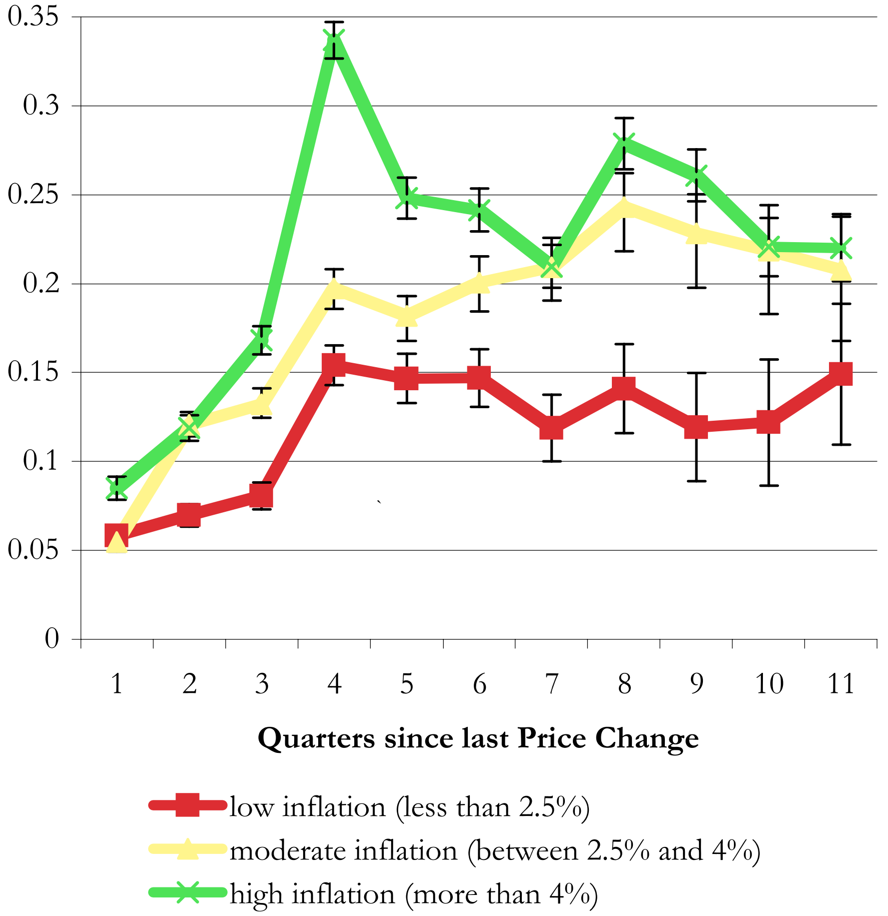
3.2 The Hazard of Price Changes

A second way to examine price adjustment that uses the special features of our data set is to examine the hazard of a price change, i.e. the probability of changing the price conditional on not having changed the price for *k* periods. This is particularly interesting if there is a large fraction of observations with very long periods of price fixity, as we observe them in our data set. *Figure 4* displays the hazard rates for three different inflationary environments. They were chosen such that each contains about a third of the quarters of our sample periods. Consider first the hazard function for very low (< 2%) inflation rates. There are two defining characteristics. Initially, the hazard of changing the price before one year is quite low: Less than 20 percent of the items see their price changed in the first three quarters. At the fourth quarter, the hazard of the price change then jumps up and stays roughly constant. Thus, at low inflation rates, we observe a weakly increasing hazard function.

Moving to the price hazard at intermediate levels of inflation, we find that the hazard in the early durations is somewhat higher. The more marked difference, however, is in the long durations: After six quarters, there is a strong increase in the probability of adjusting the price. Overall, the upward slope in the hazard function becomes more pronounced, compared to the episodes with very low inflation. As the standard error bars already indicate, the shift in the hazard function is statistically highly significant.

Moving to even higher inflation (> 4%), the hazard of price changes becomes even steeper. However, we observe the formation of a strong peak at a yearly frequency of price changes, but also a higher hazard of prices changes thereafter. A formal statistical test rejects the equality of the implied survivor functions during the three inflation episodes (log-rank test of equality of the survivor function, Chi-squared(2) = 234.6, $p < 0.001$).

Figure 4: The Hazard of Price Changes and Inflation (1979 - 1993, +/- s.e. of estimate)



Three features are noteworthy about this pattern: First, the pattern is consistent with the intuition from a state-dependent pricing model: Models of this class predict that the hazard of a price change should be increasing with time, and the hazard should be higher at higher inflation. We find that the hazard of a price change is roughly increasing in the duration, and higher inflation rates raise the hazard for all durations, but more so for longer durations.

Second, our results differ sharply from those obtained using much broader sets of goods as in, for example, the research conducted by the Inflation Persistence Network of the European Central Banks (see Dhyne et al., 2005, for an overview). Typical results from these studies show a *falling* hazard of price adjustment, which could be taken as evidence against state-dependent models. However, falling hazards can also be generated if the underlying goods have very different (possibly increasing) hazards of adjustment.[5] The fact that we find increasing hazard rates for fairly narrowly defined goods is at least suggestive that heterogeneity between goods is driving the results in studies that use a broad range of goods.

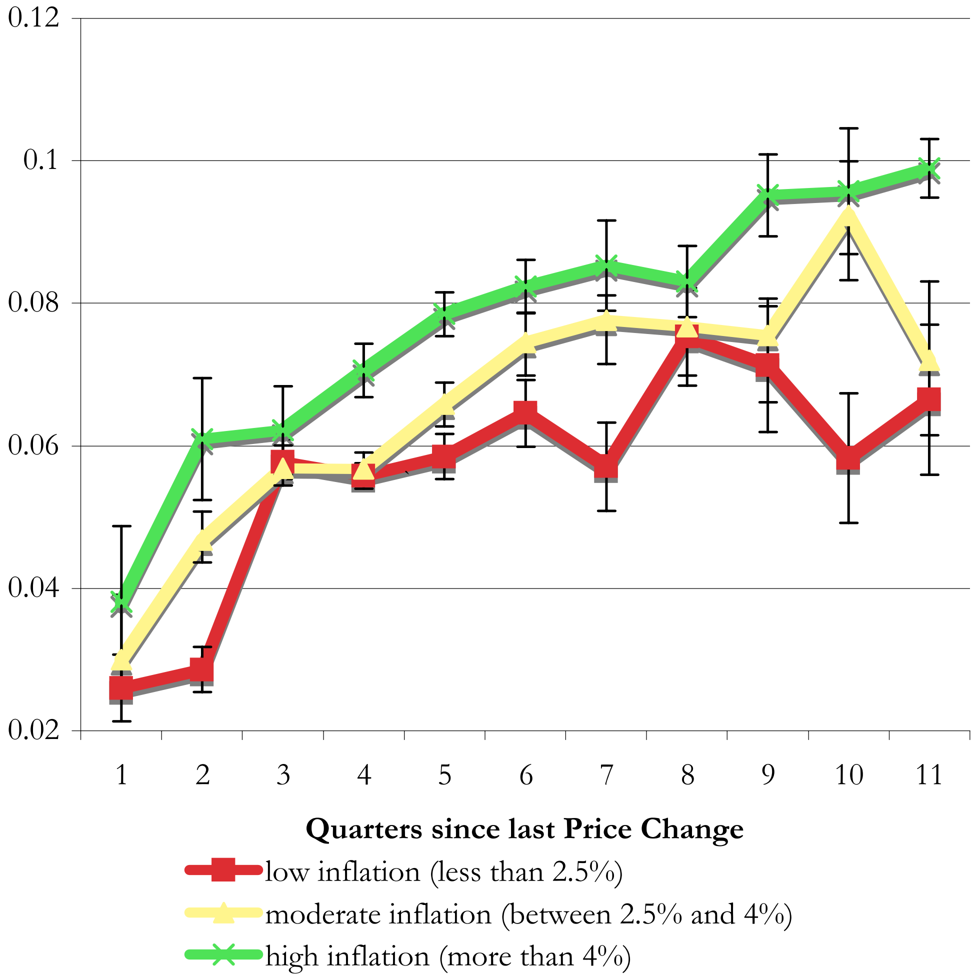
Third, our results suggest that different inflation rates affect the hazard rate of price changes in quite different ways. Thus, it may not be surprising that the relationship between the fraction of adjusters and inflation is not particularly tight (*Figure 2*). However, it affects strongly what vintage of prices will be adjusted. If we run an ANOVA of the duration since the last price change on inflation, product and producer dummies, we can explain 20 percent of the variance in the durations at adjustment. Interestingly, the lion's share (76% of the explained variance) is explained by inflation, and only 3% by the type of product, and 20% by the identity of the firm.

3.2 The Size of Nominal Price Changes Revisited

The previous analysis suggests a potential confound when examining the size of price increases at different inflation rates: If inflation induces firms to adjust prices earlier, this may affect the size of the price increase. To control for this, we examine the size of the price increases in the three inflation environments conditional on the duration of price changes. The results are displayed *Figure 5*. Conditioning on the duration makes the impact of inflation on price changes somewhat bigger, though not by much. With the exception of the very early durations, the typical difference in price increases between the low-inflation and high-inflation period at a given duration is about two percent. This is confirmed in a regression of individual price changes on inflation, a set of dummy variables for products and firms, and, most importantly, duration. The coefficient on the inflation rate is again small: it is 0.31 (t = 4.68, adjusted for clustering on quarters). Thus, raising the inflation from 1 percent to 6 percent raises the average price increase by 1.55 percentage points.

[5] Intuitively, this can happen if different products have very different baseline hazards. In early periods, a high frequency of price adjustments will be observed, mainly driven by the fast-adjusting goods. Later periods will show a low hazard of price changes, thus leading to a falling hazard rate, but only because none of the fast-adjusting goods are left in the later periods. This heterogeneity can swamp any increase the hazard rates might display for each good separately: This result is well-known in labor economics, in particular in the context of the hazard rate from exiting unemployment. See, e.g., Lancaster (1990).

Figure 5: Mean Price Changes and Inflation (+/- s.e. of estimate)



4 Conclusions

This study examines the price adjustment of goods with sticky prices. To this end, we use a part of the CPI micro base from Switzerland from a sector that other studies have shown to have sticky prices: Prices of food items and drinks in restaurants (Bils and Klenow, 2004; Dhyne et al., 2005). The long time span covered by our data set (1978 – 1993) and the variability in inflation rates over this period (inflation varied between 0 and 7 percent) provide an ideal macroeconomic environment to examine how relatively sticky prices respond to different inflation environments.

Our main result is that for sticky price goods, it is mainly the frequency of price changes that varies over time, but not the size of the price adjustment. When we compare the frequency of price changes per quarter in our sample, we find that at one percent inflation, only about 15% of the items' prices are changed per quarter. When inflation is 6 percent, about 25% of the goods prices change every quarter. However, there is almost no difference in the amount of the price increase. When we consider the same five-percent increase in inflation, the average price increase changes from 6 percent to 7 percent.

As noted in Rotemberg (2004), this is only consistent with state-dependent pricing models if their (S,s) band is almost fixed. But if this is true, it implies that the real price erosion should also be approximately invariant to the inflation rate. When we examine how price erosion changes with inflation, we find that this does not hold: During high inflation, firms let their prices erode by more, but they do not raise prices by more than during low-inflation periods. This is again supportive of the view that the main variable over which the firms optimize is the frequency of adjustment.

A closer examination of the probability to change the price conditional on the duration since the last change reveals new interesting results: We find that the hazard of price changes is increasing, i.e., the longer it has been constant, the more likely it becomes that the price changes. We also find a strong and distinct impact of inflation on the price change hazard. Low inflation rates mainly increase the hazard at long durations, while higher inflation rates raise the hazard throughout. These findings are again consistent with optimizing models of price adjustment (Dotsey and King, 2004).

It is all the more puzzling why we find little evidence that firms take inflation into account when adjusting prices, and it is beyond the scope of this paper to explain this. A possible explanation is that for some low-priced goods, the firms are constrained by divisibility problems, thus effectively eliminating the size of the price increase as a choice variable. A model that shares this property is Rotemberg (2004). In his model, consumers are "disappointed" when they observe a price increase.

References

Ball, L., Mankiw, G., Reis, R. (2004): Monetary Policies for Inattentive Economics, *Journal of Monetary Economics* forthcoming.

Bils, M. and Klenow, P.J. (2004): Some Evidence on the Importance of Sticky Prices. *Journal of Political Economy* 112: 947-85.

Cecchetti, S.G. (1986): The Frequency of Price Adjustment: A Study of the Newsstand Prices of Magazines. *Journal of Econometrics* 31(3): 251-74.

Chevalier, J.A., Kashyap, A.K. and Rossi, P.E. (2003): Why Don't Prices Rise During Periods of Peak Demand? Evidence from Scanner Data. *American Economic Review* 93(1): 15-37.

Dhyne, E., Alvarez, L.J., Le Bihan, H., Veronese, G. Dias, D., Hoffman, J., Jonker, N., Lünnemann, P., Rumler, F. and Vilmunen, J. (2004): Price Setting in the Euro Area: Some Stylized Facts from Individual Consumer Price Data. European Bank Working paper no. 524, September 2005.

Dotsey, M., King, R.G. (2004): Monetary Discretion, Pricing Complementarity and Dynamic Multiple Equilibria. Forthcoming in: *Quartery Journal of Economics*.

Dotsey, M., King, R.G. and Wolman, A.L. (1999): State-Dependent Pricing and the Dynamics of Money and Output. *Quarterly Journal of Economics* 114(2): 655-90.

Eden, B. (2001): Inflation and Price Adjustment: An Analysis of Micro Data. *Review of Economic Dynamics* 4(3): 607-36.

Kashyap, A.K. (1995): Sticky Prices: New Evidence from Retail Catalogs. *Quarterly Journal of Economics* 110(1): 245-74.

Lach, S. and Tsiddon, D. (1992): The Behavior of Prices and Inflation: An Empirical Analysis of Disaggregated Data. *Journal of Political Economy* 100(2): 349-89.

Lach, S. and Tsiddon, D. (1996): Staggering and Synchronization in Price-setting: Evidence from Multiproduct Firms. *American Economic Review* 86(5): 1175-96.

Lancaster, T. (1990), *The Econometric Analysis of Duration Data*, Cambridge University Press: Cambridge, MA.

Ratfai, A. (2004): The Frequency and Size of Price Adjustment: Microeconomic Evidence. *Managerial and Decision Economics*, forthcoming.

Rotemberg, J.J. (2002): Customer Anger at Price Increases, Time Variation in the Frequency of Price changes and Monetary Policy. NBER Working paper no. 9320.

Rotemberg, J.J. (2004): Fair Pricing. NBER Working paper no. 10915 (November 2004).

Sheshinsky, E. and Weiss, Y. (1977): Inflation and Costs of Price Adjustment, *Review of Economic Studies* 44(2): 287 – 303.