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ISSN 1657-7191 edición electrónica

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20
OCTUBRE DE 2008

Universidad de los Andes
Facultad de Economía

CEDE
Centro de Estudios sobre Desarrollo Económico

Serie Documentos Cede, 2008-20
ISSN 1657-7191

Octubre de 2008

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Edición, diseño de cubierta, preprensa y prensa digital:
Proceditor ltda.
Calle 1C No. 27 A – 01
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Teléfonos: 2204275, 220 4276, Fax: extensión 102
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Impreso en Colombia – Printed in Colombia

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Sticky Prices and Moderate Inflation

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September 18, 2008

Abstract

Recent evidence shows that there is great heterogeneity in the price setting frequency across sectors, and that those changing prices frequently do so even under low inflation. What happens to price setting strategies of sticky price goods under moderate inflation? We built a dataset of monthly newspaper and magazine prices for Colombia, for the period 1960-2005, an exceptional example of prolonged moderate inflation. Within this macroeconomic scenario, and the novel database, we study the frequency of price adjustment, the relative importance of time- and state-dependent theories, and their evolution as inflation declined from moderate rates to single digits.

Keywords: Sticky Prices, Moderate Inflation, Disinflation.
JEL classification: E31, E52, F41.

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†I appreciate helpful comments from Luis-Felipe Zanna, Munir Jalil, Marcela Eslava, André Lunardelli, and participants at the Universidad de los Andes/CEDE Seminar, the Lacea 2007 Annual Meeting and the XXXII SAE. Thanks also to Diana L. Lizarazo and Nicolás De Roux for their superb research assistance.

Precios rígidos e Inflación Moderada

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Septiembre 18, 2008

Resumen

La literatura reciente ha encontrado que hay una gran heterogeneidad en la frecuencia con que diferentes sectores de la economía ajustan sus precios. También ha mostrado que los sectores caracterizados por frecuentes cambios de precios lo hacen aún bajo inflaciones bajas y estables. ¿Qué pasa con las estrategias de cambios de precios de bienes con precios típicamente rígidos en escenarios de inflación moderada? Para responder la pregunta, construimos una base de datos de precios de periódicos y revistas para Colombia, 1960-2005, y aprovechamos el periodo de inflación moderada que caracterizó la vida macroeconómica de Colombia durante casi un cuarto de siglo. Con esos datos, se estudian la frecuencia de cambio de precios, la importancia relativa de las teorías de ajuste de precios estado- y tiempo- dependientes, y su evolución al caer la inflación de niveles moderados a bajos.

Palabras clave: Precios rígidos, inflación moderada, desinflación
Clasificación JEL: E31, E52, F41

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[†] Agradezco los comentarios de Luis-Felipe Zanna, Munir Jalil, Marcela Eslava, André Lunardelli, y de los participantes del Seminario CEDE de la Universidad de los Andes, la conferencia anual de Lacea (2007) y del Simposio de Análisis Económico XXXII (2007). Gracias también a Diana L. Lizarazo y Nicolás De Roux por su excelente trabajo como asistentes de investigación

1 Introduction

Nominal price rigidities have been at the heart of macroeconomic analysis since Keynes' seminal contribution (1936). Most modern macroeconomic models incorporate price stickiness as a natural assumption, one that does not require much justification. In many models, slow price adjustment is a key element in generating important results, such as the effectiveness of monetary shocks. The idea of price stickiness has even inspired a manifesto (Ball and Mankiw, 1994). Blanchard, in his recent "State of the Macro" paper (2008), claims that "[t]he study of nominal price and wage setting is one of the hot topics of research in macro today."

From an empirical standpoint, Cecchetti's (1986) study of newsstand magazine prices is one of the early papers with strong evidence supporting the idea that prices adjust infrequently. In a similar vein, Kashyap (1991) provides evidence of price stickiness through an examination of retail catalog prices. Evidence from Blinder et al. (1998) and Fabiani et al. (2005), based on interviews with business firms, also supports the idea of price stickiness. More recently, several studies have documented the price adjustment process in a variety of countries, using large micro datasets. Bils and Klenow (2004), using a dataset covering 70% of US consumer spending, find that half of the price spells (periods during which prices remain unchanged) last less than 4.3 months. This result suggests that price changes are more frequent than previously thought. This study has inspired similar efforts in several European countries. Taken as a whole, the evidence from the euro area is more benign as far as the slow price adjustment idea is concerned (Dhyne et al., 2005). The average duration of price spells ranges from 4 to 5 quarters, although, as Bils and Klenow report, there is substantial heterogeneity across goods. Finally, Nakamura and Steinsson (2007) provide evidence showing that excluding price changes associated with sales leads to an upward adjustment in the estimated median duration, to a range between 8 and 11 months for the prices underlying the US CPI.

The evidence in Bils and Klenow and the subsequent literature sparked by their contribution greatly enhanced our knowledge of nominal price adjustment patterns. Apart from the discussion as to whether the average or median duration of price spells is sticky or not, one of its main contributions is that it provided economists with quantifications regarding the great extent of heterogeneity of price adjustments across product categories.

In the theoretical literature, heterogeneity in price adjustments across product categories has started to be incorporated into otherwise standard macro models. For instance, the recent and important contributions by Carvalho (2006) and Carvalho and Schwartzman (2008) show that the heterogeneity of price stickiness plays a crucial role in determining the real effects of monetary shocks. In particular, heterogeneity causes monetary shocks to have larger real effects than one sector economies with an analogous frequency of price adjustments. These findings highlight the importance of not focusing exclusively on the mean of price change distributions; this recent research has made clear that understanding price adjustment strategies of the sticky and flexible extremes of price change distributions is very relevant.

Bils and Klenow (2004) have shown that goods at the flexible end of a price change distribution change prices often, even under low inflation scenarios. Where our knowledge remains weak is concerning the behavior of sticky prices, particularly at higher yet not extreme rates of inflation. Our paper is aimed at enhancing our knowledge precisely in this area—sticky prices under moderate inflation.

There are other reasons why economists are interested in achieving a better understanding of the price setting behavior of sticky price goods under alternative macroeconomic conditions. For instance, a strand of the literature highlights the fact that monetary policy should target measures of core inflation, which are made up mostly of sticky price goods, rather than targeting headline inflation, where flexible price goods have a heavy weight (see, for instance, the recent discussion in Mishkin, 2008). The reason for this is that a policy aimed at stabilizing the inflation of sticky price goods avoids unnecessary disallignments of sticky prices relative to flexible prices. In other words, such a strategy helps the economy come closer to the natural rate of output. This is yet another reason for enhancing our understanding of the price setting strategies of sticky price goods under alternative macroeconomic conditions.

Colombia, the case we focus on here, is an exceptional example of a country with a very prolonged moderate inflation scenario. Indeed, Colombia had inflation rates ranging between 15 and 30% for roughly a quarter of a century—since the early-'70s up until the late-'90s. From 1999 onward, Colombia successfully disinflated and was able to keep inflation in single digits (see Figure 1). In this paper, focusing on a single sticky price sector as explained below, we take advantage of this unique macroeconomic environment in order to study the frequency of price adjustments, the relative importance of the time and state dependent theories of price adjustment, and the manner in which they changed as inflation declined from the moderate range to its present low level.

As for the sticky price sector, we built a novel dataset based on newspaper and magazine prices; these constitute a traditional sample of goods

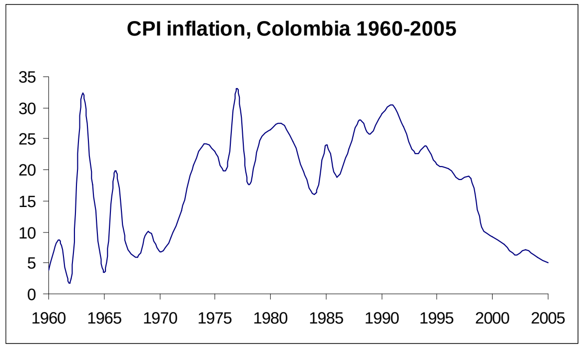

Figure 1: CPI Inflation in Colombia. 1960-2005.

that exhibit great nominal stickiness. The dataset consists of monthly prices over the period 1960-2005, and covers 32 of the main newspapers and magazines in Colombia. This dataset has several advantages. The long time series dimension offers the possibility of looking at the evolution of price setting behavior in a particular industry alongside varying macroeconomic conditions and it alleviates several econometric problems that become exacerbated when using shorter panel datasets. The monthly frequency of the collected data also represents an advantage. Indeed, the high frequency of our dataset allows for a better assessment of the time and state dependent characteristics of the price setting process; likewise, to explore the role of potential seasonal factors. Fisher and Konieczny (2003) highlight some additional desirable features of newspaper prices—newspapers are a perishable good, one usually made by a firm producing a single product. Although advertising might be an important part of revenues, it depends on circulation, which in turn should be a function of pricing decisions. Finally, discounts are rare and the physical costs of price changes are close to zero.

Our study is similar to one carried out by Gagnon (2007), who examines how price setting works across different inflation regimes. He uses micro data from Mexico for the period 1994-2004. His paper provides a much wider coverage of goods (over 200 product categories versus 32 goods in ours), while ours has a longer time-series dimension (45 years versus 10 years in his). Additionally, our inflation characteristics are different. His sample encompasses the tequila crisis–a large inflation hike followed by a quick disinflation—whereas ours involves an extended period (over a quarter of a century) of moderate inflation rates followed by a period of slow disinflation. Nonetheless, we view our effort as complementary to his, and discuss some of the differences and similarities between the two studies throughout the paper.

The question of how to introduce slow price adjustments into macroeconomic models has also interested economists. The literature has taken two broad approaches. The first consists of time dependent models, which usually assume an exogenous and staggered timing of price changes. Calvo (1983) and Taylor (1980, 1999) proposed the best known modeling strategies within this category.[1] The second category consists of state dependent models, which stress the idea that the timing of price adjustments is determined endogenously. The importance of the debate regarding which approach is best became apparent when Caplin and Spulber (1987) proposed a state dependent pricing model, wherein monetary policy becomes ineffective. Golosov and Lucas (2007) incorporate Caplin and Spulber's idea into a DSGE model, and (not surprisingly) find that monetary shocks do not have a large or persistent effect on economic activity. The debate between time and state dependent models offers a nice framework for organizing the discussion in our paper in a coherent way. With this purpose in mind, we will use this broad distinction extensively in this article.

In the end our main messages are as follows: We find that, despite studying a sample where inflation remains mostly in the 20 to 30% range, (sticky) prices continue to exhibit great nominal stickiness. Indeed, the related price spells last over a year on average, and more than a third of the price spells last over 12 months. When considering the determinants of price change patterns, we find that both time and state dependent elements have a statistically significant role, with the signs predicted by the corresponding theories. Nevertheless, the economic (quantitative) importance of the time dependent elements is much larger, and has increased with the recent disinflation. For instance, a typical time-dependent element—such as when a price completes a year without a change –has a large effect on the fraction of items that change price. To match the size of this effect, we would need an accumulated inflation rate (a typical state-dependent element) of almost 70% since the last price change.

We also find a positive time dependence—that is, the longer a price remains fixed, the greater the probability of observing a price change. In what constitutes an important contribution to the relevant literature, we show that the positive time dependence disappears if one does not control for the effect of inflation on the probability of a price change. (This is in addition

[1]With respect to this category, we might also include examples of endogenous time-dependent pricing rules (e.g., Bonomo and Carvalho, 2008).

to controlling for accumulated inflation since the last price change.) In other words, not controlling for the effect of inflation introduces a downward bias on the duration coefficient. This might help explain the puzzling negative time dependence found in several papers in the literature (e.g., Baumgartner et al., 2005, Fougere et al., 2005, Álvarez et al., 2005, and Dhyne et al., 2005).

The paper is organized as follows: Section 2 provides a description of the data. In section 3, we show that newspapers—despite the fact that we use a sample dominated by relatively high inflation rates— indeed constitute an industry characterized by a high degree of price stickiness. In section 4, we take a closer look at the relationship between frequency (of price changes), duration and the inflation rate; we explore the timing and the synchronization of price changes, and close the section by providing estimates of survival and hazard functions that further illustrate the characteristics of price changes. In section 5, we explore the determinants of price changes using a probabilistic model. This estimation allows us to better assess the relative importance of state and time dependent theories with respect to price change determination. It also allows us to better understand how the peculiar evolution of inflation in Colombia has shaped the price change mechanisms underlying sticky price goods. Section 6 concludes and highlights several policymaking implications.

2 The Data

We collect prices of 32 newspapers and magazines, on a monthly frequency, for the period January 1960 through December 2005. The prices collected are the ones printed on the covers of newspapers; we do not collect information regarding subscription prices. Not all publications existed in 1960 and we were not always able to collect all the potential data, particularly for the older issues of a few regional newspapers. The appendix provides a list of the publications and further details regarding our dataset.

We end up with 9778 observations, which we label p_{it}, that is, the price of publication i during period (month and year) t. As an illustration, Figure 2 plots the (log)prices for five of the main newspapers. Each plot includes a CPI index normalized to be equal to the price of *El Tiempo* (the main national newspaper) at the first observation of each plot.

Price data as the ones just described exhibit certain characteristics that need to be taken into account. In particular, price spells could be censored (price spells are periods during which the price remains unchanged). For instance, we do not know how long the prices collected in January 1960 (the

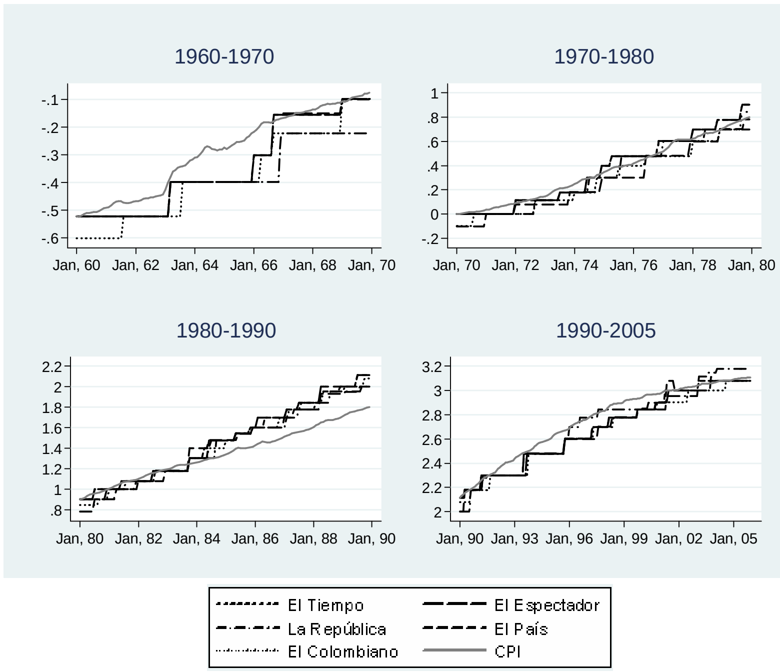

Figure 2: Log prices of five of the main publications and the (log) CPI.

first month we collected) were in place. Those price spells are thus said to be 'left' censored. Similarly, we do not know the duration of spells that were cut short at the end date of the sample –i.e., December 2005. Those observations are 'right' censored. If an observation is both right and left censored, it is said to be double censored.

One problem with censoring is that long price spells are more likely to be censored than short price spells. Ignoring this censoring might, for instance, bias statistics for the duration of price spells and the frequency of price changes. The literature has developed several methods for dealing with some of the problems that arise because of censoring (e.g., Kiefer, 1988, and the papers summarized in Dhyne et al, 2005). We use several of these strategies in the paper. In any event, inasmuch as we use many years of monthly data, censoring is not likely to be determinant. Indeed, in our sample we have a total of 744 price spells. Of those, 688 (92.5%) are uncensored, 24 left censored, 31 right censored and one is double censored.

3 Are sticky prices sticky?

We begin by exploring whether or not newspaper prices are indeed sticky in a sample dominated by inflation rates above 20%. We do this by reporting the results for the frequency of price changes and the length of price spells.

3.1 Frequency

How frequently do we observe price changes in our sample? A direct measure of the frequency of price changes (F) results from dividing the number of price changes by the number of observations where a price change could have been observed (see the appendix for details). We call this frequency estimate "Frequency 1". We obtain a frequency of adjustment of 7.3% for the whole sample. To put this result in perspective, Bils and Klenow (2004) find that magazines in the US exhibit price changes 8.6% of months.

A potential problem with Frequency 1 is that it does not take into account the presence of censored observations. Such an omission might bias our estimates, as long spells are more likely to be censored. We thus calculate a second frequency measure, "Frequency 2", to correct for this bias, in a manner proposed by Kiefer, 1988 (see the appendix for details). The results suggest that, as expected, the effect of censored observations in our sample is small. Frequency 2 is 7.4%, which is very similar to Frequency 1, as described above.

3.2 Duration

The duration of price spells is the length of time during which a price remains unchanged. One indirect way to estimate the mean duration of price spells is to find the inverse of the frequency estimates. Under this indirect estimation method, duration (D) is defined as $D = 1/F$. Using the estimates of frequency described above, the mean duration of price spells are 13.7 and 13.5 for frequencies 1 and 2, respectively. That is, despite that for most of the period inflation remained above 20%, we still observe that prices remain unchanged for prolonged periods. Again, to put our results in perspective, Bils and Klenow (2004) find that the average duration of price spells for magazines in the US is 11 months.

Duration can also be estimated directly, by simply counting the number of months that each spell lasts. We call this estimate Duration 3, and report the means and medians in Table 1. In reporting the results, we separate them according to the kind of observed censoring. This is important, as the disadvantage of directly estimating average durations is that they do not deal with the censoring problem. Table 1 shows that the results are robust. For the noncensored observations, spells last on average a bit more than a year. Censored observations generally last longer, confirming our suspicion that long spells are more likely to be censored. The main point is that, despite having a sample where inflation fluctuated mostly between 20 and 30%, we still observe a high degree of price stickiness. In the next section, we take a closer look at the evolution and characteristics of this stickiness.

Duration 3 (1960 - 2005)			
	Mean	Median	# obs
D3 nc	12,5	11	688
D3 lc	16,8	11	24
D3 rc	23,1	23	31
D3 dc	28,0	28	1
D3 total	13,1	11	744

Notes: nc, lc, rc and dc are, respectively, non, left, right and double censored.

Table 1

4 A first look at the pattern of price changes

Having established the presence of high price stickiness, we begin to explore the characteristics of price changes and whether features proposed by the time and state dependent theories are present in our sample.

4.1 Frequency, Duration and Inflation

Sticky price models are built under the assumption that prices remain constant during long periods of time –i.e., that the price change frequency should be low. Why prices remain constant for prolonged periods of time is still a matter of debate among economists. Different theories propose alternative explanations consistent with peculiar frequency and duration patterns.

We begin by taking a broad look at the relationship of frequency and duration with inflation. Several state dependent models predict that at higher inflation rates, the frequency of price changes should be higher and their duration lower (e.g., Cecchetti, 1986). Gagnon (2007) finds evidence in favor of this idea, in the midst of the Mexican inflation crisis of the mid-90s. In Table 2, we split the frequency and duration estimates, described in the previous section, according to inflation characteristics. More specifically, we split the sample between low inflation periods (1967-1972 and 1999-2005) and moderate inflation periods (1973-1998). We define moderate inflation as corresponding to rates above 15%. Considering that with the exception of a couple of months, annual inflation remained below 30%, the definition of moderate inflation is roughly consistent with those proposed by Dornbusch and Fischer (1993) and Hofstetter (2008).[2] Low inflation corresponds to rates below 15%.

Table 2 shows that during the periods of low inflation, the frequency is between 4.5% and 6.5% whereas under moderate inflation, it jumps to almost 9%. By the same token, the spells under moderate inflation last just under a year, while lasting between 15.3 and 22.3 months during periods of low inflation. As predicted by the state dependent theories, prices adjust more frequently when inflation is higher.[3]

Frequency and Duration			
	Low Inflation		Moderate Inflation
	1967 - 1972	1999 - 2005	1973 - 1998
Frequency 1	4,5%	5,4%	8,7%
Frequency 2	4,8%	6,5%	8,8%
Duration 1	22,3	18,5	11,6
Duration 2	20,9	15,4	11,3
Duration 3nc	21,4	15,3	11,2

Table 2

Figure 3 reinforces the message: it depicts a scattering of the annual frequencies of price adjustments and the corresponding annual inflation rate. Here, the positive relationship is evident. The linear regression shows that an additional point of inflation, raises the frequency of price changes by 0.26%. To put the result in perspective, Gagnon (2007) shows that during high inflation periods in Mexico, a 1% increase in inflation, increases the frequency by 0.36.[4]

As the state dependent theories predict, inflation appears to explain part of the frequency of price changes. On the other hand, time dependent

[2] In carrying out this separation, we omit the data for 1960-1966. During that period, only a small fraction of the newspapers considered here existed. Additionally, inflation during these years jumped back and forth from low to moderate levels, and this preliminary description of the data would not control for these large variations.

[3] The Duration 3 sample allows us to check if the differences identified in Table 2 are statistically significant. Performing such a test (not reported here) shows that the two Duration 3 sample means, during low and moderate inflation rates, are statistically different at the 1% level.

[4] His sample, though, contains more heterogenous goods and much higher inflation rates than ours.

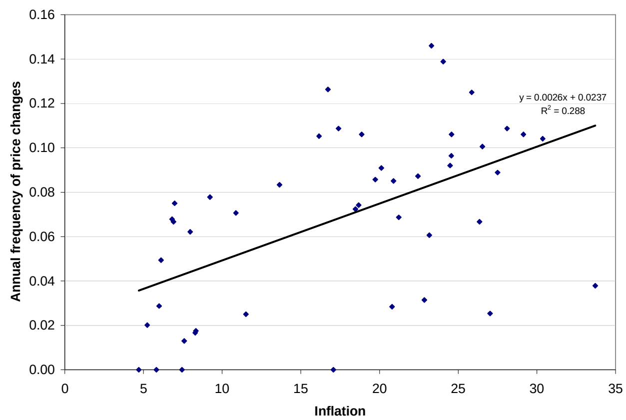

Figure 3: Annual frequency of price changes

theories a la Calvo or Taylor predict no relationship between these two variables. Here, price changes occur at fixed intervals (Taylor) or randomly (Calvo). Nevertheless, once we estimate a probabilistic model (in section 5), we show that inflation –despite being significant and retaining the sign suggested by Figure 3– plays a minor quantitative role in determining price changes, once we control for time dependent features.

4.2 The timing of price changes

Taylor's time dependent model calls for the occurrence of perfectly staggered price adjustments, say, once a year. Taylor's model is potentially consistent with a higher frequency of adjustment during a certain month, but only so long as the assumption of uniform staggering is relaxed. In other words, if price setters change prices once a year, but have a preferred month for taking this step, we might observe a higher frequency of price changes during that particular month. In Calvo's setup, where each period a constant and random fraction of firms receive a signal to change prices, we do not expect more frequent price changes during any particular month.

Observing a higher frequency of price changes during a specific month could also be consistent with certain state dependent theories. In Colombia, the conventional wisdom holds that many price changes take place in January. Part of the explanation lies in the fact that minimum wages and other government controlled prices (such as utilities) are traditionally raised in January. Such behavior implies potentially higher costs for firms, which

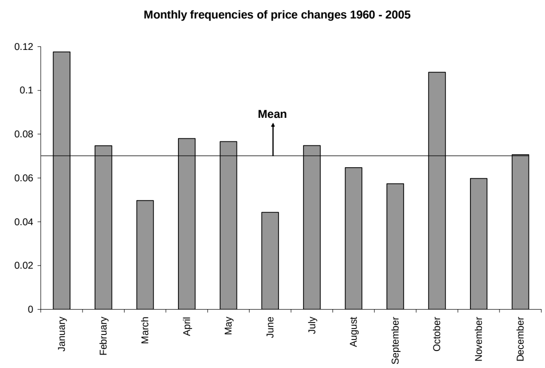

Figure 4: Monthly frequencies of price changes.

can translate into a higher frequency of price changes for that month. In a similar sense, under Rotemberg's (2006) optic, it could be that price changes in January are more palatable to consumers; they perceive that costs are now higher and are therefore willing to accept raised prices. This might therefore be a preferred strategy for price setters. In this regard, this particular state dependent model and Taylor's time dependent model (without perfect staggering) would lead to observationally equivalent predictions.

In Figure 4, we plot the frequencies of price changes across months. The bars in the plot indicate the number of price changes as a percentage of the total number of observations available each month. We observe a peak in January, suggesting a preference for changing prices that month. There is another peak, though slightly smaller, in October. March and June are two notable troughs, while the rest of the months are closer to the mean frequency. As mentioned above, the peak in January could be consistent with Rotemberg's state dependent model. As for the time dependent theories, the results might be interpreted as evidence of time dependent strategies a la Taylor, only with imperfect staggering. In other words, prices remain unchanged for a long period of time (say, a year), and price changes exhibit a relatively high degree of synchronization across goods (i.e., many price changes take place in January). We explore how plausible this idea is in the following subsection by estimating the degree of synchronization of price changes.

4.3 Synchronization

Following Fischer and Konieczny (2000), we estimate a synchronization index, calculated as the ratio of the actual standard deviation of frequencies to the maximum theoretical standard deviation. In case of perfect synchronization, the index takes a value of 1, whereas where perfect staggering exists, the ratio takes de value of 0 (see the appendix for details). The index can also be interpreted as a method of moments estimator providing the share of firms in the economy whose price changes are perfectly synchronized (Dias et al., 2005).

Inasmuch our panel of newspapers is not balanced, we calculate the ratio for several balanced partitions of the data. In general, the longer the sample being looked at, the fewer the number of publications that can be followed to form a balanced panel. In Table 3, we report the results, together with the number of newspapers included in each calculation.

Synchronization						
				Low π		**Moderate π**
1960-05	1970-05	1970-98	1986-00	1967-72	1999-05	1973-98
0,56 (4)	0,50 (8)	0,51 (10)	0,39 (21)	0,61 (7)	0,45 (7)	0,55 (7)

Note: Numbers in parentheses report the number of publications included in each panel.

Table 3

In short, the results in Table 3 suggest that close to one half of price changes are synchronized.[5] The broad separation between low and moderate inflation eras gives no clear prediction as to the impact of inflation on the synchronization rates. Compared to the synchronization results for the Euro area (e.g., Dhyne et al, 2005), where the statistic varies between 0.13 and 0.48, our results suggest a high degree of synchronization. Of course, in making this comparison, we should keep in mind that we are focusing on a specific type of product. In any case, the high degree of synchronization in price changes can reconcile time dependent pricing with our evidence; firms could use time dependent price setting strategies with imperfect staggering (synchronized).

An alternative way of assessing the relative importance of time and state dependent models, related to the degree of synchronization, was recently proposed by Klenow and Kryvtsov (2005). They show that for any given month, inflation can be broken down into the product of the fraction of items that change prices and the average price change. In time dependent models, all variance in inflation should be explained by the second term (the

[5] We tried other partitions of the data and obtained similar results.

intensive margin). Why? In time dependent models, a constant fraction of firms adjust prices each period –i.e., the variance in the first term should be zero. Looking at a very large dataset of the US CPI, they find that the intensive margin explains 95% of inflation variance, and interpret this result as evidence in favor of time dependent models.

Calculating these margins in our sample is feasible, but the results should be interpreted with caution. Indeed, although we use a large time series, what matters for this index is the cross-section dimensions of the dataset. In this respect we only have 32 newspapers.[6] Hence, we interpret this evidence as indicative of the direction in which things have moved in keeping with the disinflation, but do not make much of the size of the index itself. The results are, nevertheless, very interesting. For the moderate inflation period (1973 to 1998) the time dependent features (intensive margin) account for 15% of the 'newspapers' inflation variance. By way of contrast, for the period 1999 to 2005, when inflation was kept in single digits (though was still greater than 5%), the time dependent features explain 30% of inflation variance. As expected, the importance of time dependent models increases as the level of inflation decreases. This result is consistent with Gagnon's (2007): he finds that time dependent elements dominate in Mexico if inflation is low (below 10 to 15%) while state dependent features gain relevance when inflation is high (above 10 to 15%).

Finally, we should stress that, according to the evidence in Gagnon (2007), the intensive margin explains most of the inflation variance for those goods exhibiting a high frequency of price changes. The reason is that those goods change price frequently, independently of the inflation rate. By contrast, there is more room for sticky price goods to adjust to a higher inflation rate by changing prices more frequently. Accordingly, since we are focusing on a sticky price good, we should not necessarily expect the intensive margin to explain close to 100% of the inflation variance.

4.4 The Hazard and Survival Functions

The Hazard Function: Hazard functions, in our context, investigate the probability of a price change conditional on the elapsed duration of a price spell. They report the rate at which a price spell will be completed at duration t, assuming that it lasted until t. They also provide information on the empirical relevance of price setting theories. For instance, Calvo's (1983) model assumes that every period a certain fraction of firms adjusts

[6] In the early years of the sample, the number of newpapers is actually much lower.

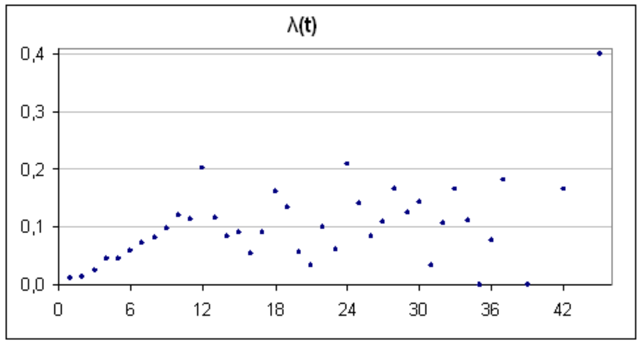

Figure 5: Hazard Function

its prices. More specifically, price changes follow a Poisson distribution, implying a constant hazard function. Alternatively, Taylor's (1980, 1999) modelling strategy assumes that prices are set for a fixed period of time. In other words, price changes only occur when a fixed duration is completed. If such is the case, we should observe peaks at said durations in the hazard function. A richer set-up –truncated Calvo models– combining both ideas, was proposed by Dotsey et al., (1999). Here an increasing hazard function is expected –i.e., the probability of a price change increases as the spell becomes longer.

We report Kaplan-Meier estimates of this function, $\lambda(t)$, in Figure 5 (see the appendix for details).[7] The hazard function exhibits positive time duration up through period 12 –i.e., within this range, the probability of a price change increases with time. In period 12 the function jumps, a fact consistent with Taylor's idea that firms set prices for a fixed period of time (in this case, for one year). We also observe peaks at months 18 and 24 –i.e., the rate at which spells conclude also jumps after having remained a year-and-a-half or two years in place. After periods 12, 18 and 24, we observe periods of negative time dependence ($d\lambda(t)/dt < 0$); if a firm has not changed its price in months 12, 18 or 24, then it becomes less likely that it will do so in the following months. Finally, since the accuracy of the estimator declines with longer durations –as there are fewer observations available for inference– we do not bother to interpret the longer durations in the hazard function.

Survivor function: $S(t)$, reports the probability of a price spell surviv-

[7] Note that these Kaplan Meier estimates allow for censoring (e.g., Kiefer 1988).

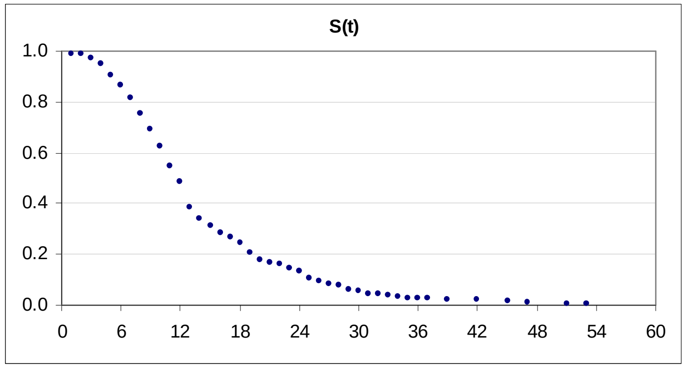

Figure 6: Survivor Function

ing until time t. Put another way, for each t, the function gives the fraction of price spells with a duration of t months or more. The survivor function can also provide information about price setting strategies, as discussed above in connection with the hazard function. For instance, a marked drop at 12 months would indicate that once-a-year price changes are a relevant strategy.

We report Kaplan-Meier estimates for this function in Figure 6 (see the appendix for details). The survivor function decreases up through month 12. After month 12, we observe a sharper decline which is consistent with the findings for the hazard function. Again, we interpret this as meaning that a large proportion of firms change prices after they have remained unaltered for a year. The survivor function then becomes flatter; by month 18, more than 20% of the price spells have still survived. Again, this finding signals that price stickiness remains relevant, despite of the moderate inflation scenarios predominant in our sample. Finally, we also observe declines after months 18 and 24, a finding consistent both with the results for the hazard function discussed above and with Taylor's proposal concerning time dependent price-setting strategies.

The survivor and hazard functions illustrate important features of price setting strategies, but these results do not control for other factors such as inflation. In the next section, we explore the determinants of price change probabilities controlling for other relevant factors.

5 The Determinants of Price Changes

In this section, we estimate a model that studies the determinants of the probability of price changes. The empirical strategy involves estimating a logit model where the dependent variable is a dummy, one that takes the value of 1 if there is a price change during the month in question for a certain publication, and is 0 otherwise. The set of determinants is based on the results discussed until now. Dummy variables exist for durations of 6, 12, 18 and 24 months; these are intended to capture the time dependent elements we identified in the hazard and survivor functions. Monthly dummies are also included, inasmuch as, based on our findings above, some months seem more prone to price changes. Here we intent to quantify the economic importance of these features. State dependent factors, such as the inflation level and the size of the last price change, are also part of the determinants. The accumulated inflation since the last price change is also included as a determinant of price changes. We also explore if, after controlling for all these factors, the duration (that is, the number of months since the last price change) affects the probability of a price change. The hazard function seems to point to a positive time dependence—that is, the longer a price remains constant, the more likely it is to change. The appendix provides details on the construction of the different variables.[8]

The main results are presented in Table 4. Aside from the variables just described, all regressions include fixed effects.[9] We report both the marginal effects and the odd ratios. In both cases, we present a baseline specification inclusive of all the variables listed above, along with the fixed effects. We also report several alternative specifications in order to assess the robustness of our findings.[10] The main initial discussion focuses on the baseline regressions. At the end of the section, we explore their robustness. We begin interpreting our results with the state dependent variables.

State dependent variables: As expected, increases in annual inflation raise the likelihood of observing a price change corroborating more formally our evidence as summarized in Figure 3. Nevertheless, the economic impor-

[8] Several papers in this literature explore the role of VAT changes on the probability of a price change. This is not relevant in our sample, as the goods we analyze do not pay VAT taxes in Colombia.

[9] With our sample, the model cannot be estimated with time effects. Given that the cross-section dimension of the sample is small, we have several periods without price changes; a time dummy would then "perfectly predict" the outcome and the estimation is not possible.

[10] We also estimated probit models. The results, not reported in the paper, are very similar to the logits.

tance of the coefficient is rather small, more so if the model is evaluated at low inflation rates—that is, the currently relevant rates. For instance, the predicted probability of observing a price change if the model is evaluated at an inflation rate of 4.5% (the observed inflation at the end of 2006) is 2.5%.[11] If inflation were to increase by 2 percentage points, then the predicted probability of observing a price change would only increase to 2.9%.

In panel A of Figure 7, we plot the probability of price changes under different inflation rates; the rest of the variables are evaluated at the mean. Inflation affects the probability of a price change in an exponential way, but the impact is rather flat and small when inflation remains within the single digit range. Despite being highly significant, the quantitative impact of inflation shocks on the probability of price changes is small.

	Baseline		Robustness							
	Marginal Effects	Odd-ratios	Marginal Effects				Odd-Ratios			
Annual π	0.0032***	1.063***		0.0029***	0.0039***	0.0016***		1.055***	1.077***	1.029***
	(0.00044)	(0.0091)		(0.00044)	(0.0003)	(0.0003)		(0.0089)	(0.0066)	(0.0059)
Size of last nominal price change	-0.0014***	0.97***	-0.0012***		-0.0013***	-0.0014***	0.98***		0.98***	0.98***
	(0.00023)	(0.0043)	(0.00024)		(0.00023)	(0.00023)	(0.0042)		(0.0043)	(0.0042)
Accumulated π (since last price change)	0.00066**	1.012**	0.0024***	0.00058*		0.0022***	1.044***	1.011*		1.04***
	(0.00031)	(0.0059)	(0.00023)	(0.00032)		(0.00016)	(0.0045)	(0.0059)		(0.0031)
Duration	0.0032***	1.063***	-0.000031	0.0031***	0.0043***		0.99	1.06***	1.085***	
	(0.00058)	(0.012)	(0.00044)	(0.0006)	(0.00028)		(0.0078)	(0.012)	(0.0062)	
January	0.035**	1.71***	0.033**	0.036**	0.036**	0.032**	1.64***	1.71***	1.74***	1.64***
	(0.014)	(0.32)	(0.015)	(0.015)	(0.015)	(0.014)	(0.3)	(0.31)	(0.32)	(0.3)
February	0.0049	1.094	0.00032	0.0057	0.0068	0.0002	1.006	1.107	1.131	1.004
	(0.011)	(0.22)	(0.011)	(0.012)	(0.012)	(0.011)	(0.20)	(0.22)	(0.23)	(0.20)
March	-0.023***	0.584**	-0.028***	-0.023***	-0.022***	-0.028***	0.529***	0.595**	0.610**	0.523***
	(0.0082)	(0.14)	(0.008)	(0.0084)	(0.0084)	(0.0078)	(0.12)	(0.14)	(0.14)	(0.12)
April	0.0016	1.031	-0.0047	0.0014	0.0046	-0.005	0.916	1.026	1.088	0.909
	(0.011)	(0.21)	(0.01)	(0.011)	(0.011)	(0.01)	(0.18)	(0.2)	(0.22)	(0.18)
May	0.0024	1.046	-0.0039	0.0019	0.0054	-0.0041	0.93	1.036	1.103	0.926
	(0.011)	(0.21)	(0.011)	(0.011)	(0.011)	(0.01)	(0.19)	(0.21)	(0.22)	(0.18)
June	-0.024***	0.58**	-0.029***	-0.023***	-0.022***	-0.028***	0.523***	0.595**	0.61**	0.519***
	(0.0081)	(0.13)	(0.0079)	(0.0084)	(0.0084)	(0.0077)	(0.12)	(0.14)	(0.14)	(0.12)
July	0.0011	1.02	-0.0036	0.00075	0.0033	-0.0037	0.936	1.014	1.064	0.932
	(0.011)	(0.2)	(0.011)	(0.011)	(0.011)	(0.01)	(0.19)	(0.2)	(0.21)	(0.19)
August	-0.01	0.81	-0.013	-0.011	-0.0089	-0.013	0.767	0.807	0.835	0.763
	(0.0094)	(0.17)	(0.0095)	(0.0096)	(0.0096)	(0.0093)	(0.16)	(0.17)	(0.17)	(0.16)
September	-0.018**	0.677*	-0.021**	-0.017*	-0.017*	-0.021**	0.644**	0.698*	0.695*	0.643**
	(0.0086)	(0.15)	(0.0087)	(0.0089)	(0.0087)	(0.0086)	(0.14)	(0.15)	(0.15)	(0.14)
October	0.022*	1.432*	0.021	0.022*	0.022*	0.021	1.393*	1.426*	1.449**	1.402*
	(0.013)	(0.27)	(0.013)	(0.013)	(0.013)	(0.013)	(0.26)	(0.26)	(0.27)	(0.26)
November	-0.0084	0.844	-0.0092	-0.0097	-0.0082	-0.0091	0.837	0.827	0.848	0.837
	(0.0098)	(0.18)	(0.01)	(0.0098)	(0.0097)	(0.01)	(0.17)	(0.17)	(0.18)	(0.17)
6 month Dummy	-0.0023	0.957	-0.0011	0.0014	-0.0026	-0.0048	0.979	1.025	0.95	0.913
	(0.0096)	(0.18)	(0.01)	(0.01)	(0.0095)	(0.0095)	(0.18)	(0.19)	(0.18)	(0.17)
12 month Dummy	0.102***	3.245***	0.101***	0.102***	0.104***	0.103***	3.116***	3.18***	3.302***	3.185***
	(0.019)	(0.48)	(0.019)	(0.019)	(0.019)	(0.019)	(0.46)	(0.46)	(0.49)	(0.47)
18 month Dummy	0.053**	2.074***	0.052**	0.049**	0.055**	0.061***	1.995***	1.978***	2.111***	2.2***
	(0.022)	(0.47)	(0.022)	(0.021)	(0.022)	(0.023)	(0.45)	(0.45)	(0.48)	(0.5)
24 month Dummy	0.061**	2.251***	0.061**	0.054*	0.061**	0.081**	2.18***	2.058***	2.255***	2.631***
	(0.029)	(0.63)	(0.03)	(0.028)	(0.029)	(0.034)	(0.62)	(0.58)	(0.63)	(0.75)
Predicted probability	0.0561		0.059	0.0577	0.0558	0.058				
Observations	9147		9147	9246	9147	9147				
Log likelihood	-2189,3		-2215,8	-2239,2	-2191,5	-2203,1				
Restricted log likelihood	-2408,8		-2408,8	-2443,3	-2408,8	-2408,8				
LR chi2	439.12		386.02	408.33	434.61	411.39				
Prob > chi2	0,00		0,00	0,00	0,00	0,00				
Pseudo R2	0,09		0,08	0,08	0,09	0,09				

† Notes: All regressions include fixed effects. ***significant at 1%, **significant at 5%, *significant at 10%. Standard errors in parentheses. Marginal effects are evaluated at the mean. For Dummy Variables they should be interpreted as a discrete change of the dummy variable from 0 to 1. Mean values of the first four independent variables are, in order, 17.98, 22.01, 13.1, 9.48. They are calculated with the baseline sample. The sample is larger when we omit the lagged price change, but mean values remain virtually identical.

Table 4

[11] The rest of the variables are evaluated at the mean.

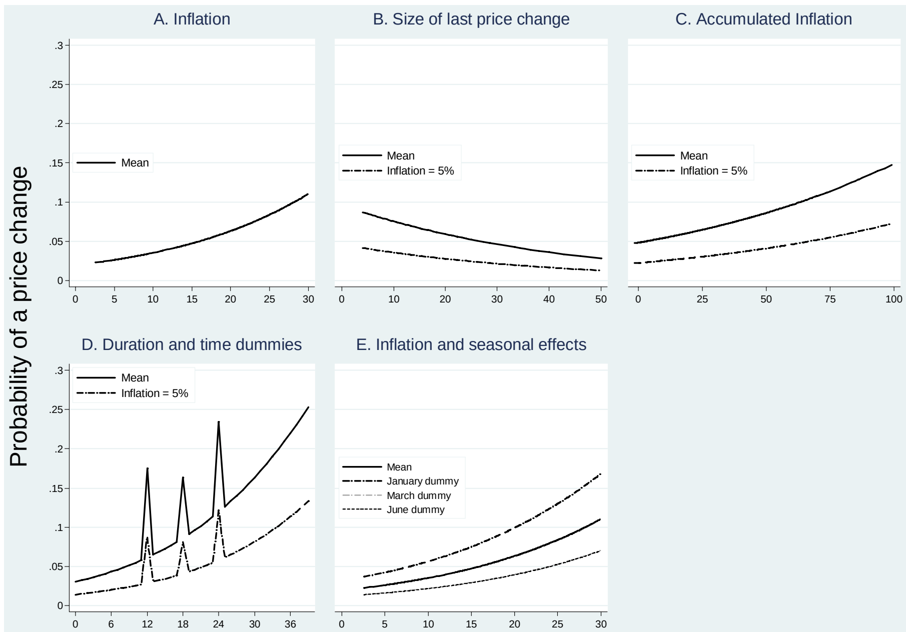

Figure 7: Probability of price changes.

The size of the last nominal price change is always significant and has the expected negative sign; if a firm chooses a large nominal price change, the probability of observing a price change decreases thereafter. The marginal effect, evaluated at the mean, shows that a 1% increase in the size of the last nominal price change will reduce the probability of a price change by 0.14%, a tiny number considering that the predicted probability of a price change is close to 6%. Panel B in Figure 7 plots how the probability changes at different levels of lagged price changes. We report the results on the one hand with all the variables evaluated at the mean and, on the other, with annual inflation set at 5% (with all remaining variables still evaluated at the mean). The latter scenario allows us to look at the results while focusing on an inflation rate close to the relevant current rate. In this case, it becomes latent how flat the curve is—i.e., changes in past price adjustments have little impact on the probability of price changes.

The accumulated inflation since the last price change is also significant and exhibits the expected positive sign. The more the price of a newspaper is eroded by inflation, the larger the chance of observing a price change. Again though, the effect is small—increasing the accumulated inflation by

10% points (away from the mean) increases the predicted probability of a price change by 0.7%. In Panel C in Figure 7, we illustrate the role of accumulated inflation. Note that up until a relatively high accumulated rate (say, 50%), the curve remains quite flat—i.e., the impact of the variable on the probability of a price change is quantitatively small. This effect is reinforced if we set inflation equal to 5%.

Given the context, one is tempted to ask the question, why we include both the 'inflation rate' and the 'accumulated inflation rate' as separate determinants of the probability of a price change. On the one hand, the *accumulated inflation* (since the last price change) captures the deviation from the optimal price in a standard Ss setting. On the other hand, the *inflation rate* might capture Rotemberg's hypothesis: if inflation is high, a price change becomes more palatable to consumers, independently of the accumulated inflation since the last price change. This explains why we included both variables and why they show up with positive and significant coefficients. As we later discuss in connection with the robustness tests, omitting either of the variables can bias other coefficients.

Time dependent elements: Duration has a positive and significant coefficient; the longer a price is in place, the larger the chance of observing a price change, after controlling for several factors. Of course, the impact of duration has to be considered along with the role of the time dummies for 6, 12, 18 and 24 months.

Looking for now solely at the coefficients of the time dummies, we note that 12, 18 and 24 month dummies are all significant and have large coefficients. Not surprisingly, given the stylized facts uncovered in the previous section, the 12 month dummy is by far the most important one. Indeed, when a spell reaches the 12 month mark, our results show that the probability of a price change more than triples (odd ratio of 3.2). 18 and 24 month dummies also have a large impact according to our estimates. Each more than doubles the probability of a price change, showing that time dependent elements play a quantitatively large role, even after controlling for several state dependent variables. We should also note that the 6 month dummy turns out to be insignificant.

The joint effect of duration and the time dummies is depicted in Panel D in Figure 7; the impact of additional months on the probability of a price change increases exponentially and exhibits peaks at months 12, 18 and 24. The curve gains steepness much later if inflation is set at 5%. It is worth noting that the probability of a price change at month 24 is larger than at month 12. Nevertheless, this is a conditional probability—i.e., conditional on having already gone 24 months without a price change, a newspaper is

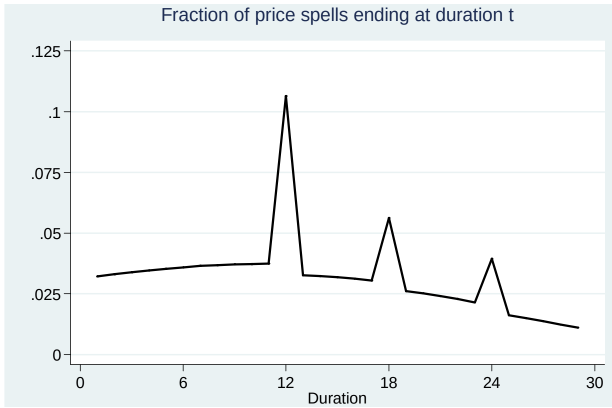

Figure 8: Fraction of price spells ending at duration t.

more likely to observe a price change than it would after 12 months. In Figure 8, we plot the same probabilities presented in Panel D, only now multiplied by the proportion of spells surviving each period. In other words, Figure 8 depicts the predicted fraction of firms changing prices for each duration. Note that here it becomes evident that going one year without a change in price is the instrumental time dependent feature present in our data.

As for the monthly dummies, we should note that the omitted month is December; therefore, the coefficients should be interpreted relative to this month. The results are consistent with our previous findings. Firms seem to be more prone to change prices in January, and less so in March and June. Those effects are highly significant and quantitatively important. For instance, the January dummy has an odds ratio of 1.7—i.e., when the month in question arrives, the odds of observing a price change increase by 1.7 times relative to what it was in December. The role played by this seasonal feature is depicted in Panel E in Figure 7.

How big are the relative effects? A way to assess the relative importance of some of the significant effects highlighted above is to estimate the 'relative effects' implied by the empirical model. More specifically, what we do is the following: The peak at month 12 in Figure 8 shows that the fraction of firms changing prices when the spell becomes a year old is 10.6%. We estimate for other variables what level would imply an identical predicted price change probability, with the rest of the variables being evaluated at their mean values. The results suggest that the time dependent features—

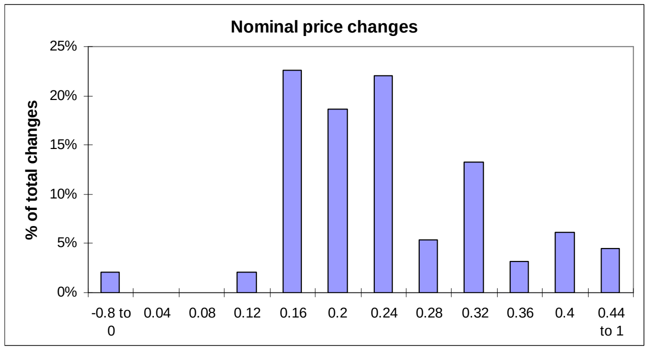

Figure 9: Histogram of price changes.

represented by a spell of a year during which prices remained unchanged—are quantitatively very important in determining the probability of price changes.

We find that an inflation rate of 29.4% implies a predicted probability of a price change of 10.6. That is, the same fraction of firms that change prices after a spell of one year, would also change prices if inflation goes up to 29.4%. In Colombia, after seven straight years of single digit inflation rates, the likelihood of inflation approaching 30% seems highly unlikely. Analogously, we would need an accumulated inflation since the last price change of 69.1% to match the impact of the 12-month factor. Both statistics highlight the importance of a simple 12 month rule (the traditional time dependent element) vis-à-vis inflation-related, state-dependent variables in determining the probability of a price change.

Similarly, the last price change should be -4.6% to have the same impact as the 12-month factor. In other words, we would require a negative lagged price adjustment. For the sake of perspective regarding this result, we note that the average price change in the baseline estimation is 22%. A lagged price change of -4.6% is highly unlikely in our sample. Indeed, as Figure 9 shows, we rarely observe nominal price decreases in our sample.

Robustness: In Table 4, in addition to the baseline estimations, we report several alternative specifications. Three important lessons emerge from this robustness exercises in Table 4:

(i) Not surprisingly, if inflation is dropped from the estimation, the coefficient on accumulated inflation becomes larger. Alternatively, if accumulated inflation is not included, the coefficient on inflation also becomes larger. The intuition is clear: the different effects that we capture by separating the two variables are forcibly captured by the one when we drop the other, thus making the respective coefficient larger. This result is important, as it suggests that similar estimations, common in the literature, when not including both variables (most only include accumulated inflation), might be overestimating the coefficient.

(ii) Accumulated inflation and duration are obviously positively related. Thus, it is not surprising that once one of these variables is dropped from the estimation, the coefficient for the remaining variable becomes larger.

(iii) Finally, we showed in the previous section that the frequency of price changes increased with inflation. In other words, other things being equal, duration and inflation should move in opposite directions. With this in mind, it should also not come as a surprise that if either inflation or duration is dropped from the estimation, then the remaining variable becomes less relevant in explaining the probability of a price change. This result is important for the literature—if inflation is omitted, estimations might fail to find the positive time dependence, not because it is absent in the price setting strategies, but rather because of a downward bias in the duration coefficient, induced by the omitted variable. The literature has usually stressed that the surprising negative time dependence obtained in several studies might likely be explained by the use of heterogeneous goods. For instance, Klenow and Kryvtsov (2005), in spite of obtaining declining hazards in their sample, reported that disaggregating the data into product categories yields mostly flat hazards, with spikes at 12 months. Our results suggest that not controlling for inflation might be a complementary reason for explaining negative time dependence.

6 Concluding remarks

Sticky prices remain sticky even under moderate inflation scenarios. Indeed, price spells last over a year on average, and more than a third of price spells last over 12 months. We also observe that the prices become stickier with disinflation. This has potentially important consequences for policymakers. Given that flexible prices adjust frequently even under low inflation (Bils and Klenow, 2004), the fact that sticky prices become stickier with disinflation, as we have shown here, increases heterogeneity in the frequency of price adjustments as inflation declines. As Carvalho (2006) and Carvalho and Schwartzman (2008) show, monetary shocks tend to have larger and more persistent real effects in heterogeneous economies than in one-sector economies with the same frequency of adjustments. This suggests that disinflation costs should be inversely related to inflation peaks, or more generally, to recent inflation history. Recent evidence—for example, Hofstetter (2008) and Senda and Smith (2008)—shows that disinflation costs measured using sacrifice ratios are indeed inversely related to the recent inflation history.[12]

When examining the determinants of price change patterns, we find that both time and state dependent elements play statistically significant roles, with the signs predicted by the corresponding theories. Nevertheless, the economic (quantitative) importance of time dependent elements is much larger, and has increased with the recent disinflation. For instance, a typical time-dependent element—such as when a price completes a year without a change —has a large effect on the fraction of items that change prices. To match the size of this effect, we would need an accumulated inflation rate change (a typical state-dependent element) of almost 70% since the last price.

We also find a positive time dependence—that is, the longer a price remains fixed, the greater the probability of observing a price change. In what constitutes an important contribution to the relevant literature, we show that the estimated positive time dependence disappears if one does not control for the effect of inflation on the probability of a price change. (This is in addition to controlling for accumulated inflation since the last price change.) In other words, not controlling for the effect of inflation introduces a downward bias on the duration coefficient. This might help explain the puzzling negative time dependence found in several papers in the literature (e.g., Baumgartner et al., 2005, Fougere et al., 2005, Álvarez et al., 2005, and Dhyne et al., 2005).

[12] The idea that time dependent rules are endogenous and that this might have important consequences for disinflation costs has also been recently explored from a theoretical standpoint by Bonomo and Carvalho (2008). They show that the interaction between the endogeneity of time-dependent rules and imperfect credibility increases the output costs of disinflation.

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Appendix:

Frequency 1:

$$F1 = \frac{\sum\limits_{i=1}^{n}\sum\limits_{t=2}^{T} d1_{it}}{\sum\limits_{i=1}^{n}\sum\limits_{t=2}^{T} d2_{it}}$$

with n = number of newspapers; T = number of periods; $d1_{it} = 1$ if the price of publication i changed in t, otherwise it is 0. $d2_{it} = 1$if product i in t was also observed in t-1, otherwise it is 0. In simple terms, F1 is number of changes divided by number of observations where we could have detected a price change.

Frequency 2:

$$F2 = \frac{n_{nc}+n_{lc}}{n_{nc}T_{nc}+n_{lc}T_{lc}+n_{lc}T_{rc}+n_{lc}\frac{n_{dc}}{n_{rc}}T_{dc}}$$

with n_{nc}, n_{lc}, n_{rc} and n_{dc} the number of uncensored, left, right and double censored price spells, respectively; call the average spell in each category T_{nc}, T_{lc}, T_{rc} and T_{dc}.

Duration: Indirect estimates: Di=1/Fi, for i = 1, 2. Direct estimates: $D3_j = \sum\limits_{j} Duration/\#Spells(j)$, with j = nc, lc, rc and dc.

Synchronization: $Sync = SD/SDMax$ with $SD = \sqrt{\frac{1}{T-1}\sum\limits_{t=2}^{T}(F_t - F)^2}$ and $SDMax = \sqrt{F(1-F)}$.

Survivor and hazard functions: Kaplan-Meier (KM) estimates of survivor and hazard functions. With observed spell lengths $t_1 ... t_k$, the survivor function is $\hat{S}(t_j) = \prod\limits_{j|t_j \leqslant t}(\frac{n_j - d_j}{n_j})$

where n_j is number of price spell at risk of exhibiting a price change at time t_j and d_j is the number of price changes at t_j. The hazard function $\hat{\lambda}(t_j) = d_j/n_j$, i.e., the rate at which spells are completed after duration t. The integrated hazard function is $\hat{\Lambda}(t_j) = \sum_{i \leqslant j} \hat{\lambda}(t_i)$

Construction of the variables used in the logit model:

Annual Inflation: for each month t, $\pi_t = (CPI_t/CPI_{t-12} - 1) * 100$

Size of last nominal price change: Calculated as a symmetric growth rate, i.e., $\Delta P = \frac{P_{ti} - P_{t-1i}}{(P_{ti} + P_{t-1i})/2} * 100$

Accumulated Inflation: If a price change occurred in period T, then for each period following T and up to the next price change, the accumulated inflation is $(CPI_{T+t}/CPI_T - 1) * 100$, for $t = 1$...next price change.

Duration: If a price change occurred in period T, then for each period following T and up to the next price change, Duration is $D_{T+t} = t$ for $t = 1$...next price change.

Newspapers included:

Publication	City of publication	Dates Included	Periodicity of publication	Comments
El Tiempo	Bogotá	Jan 60 - Dec 05	Daily	-
El Espectador	Bogotá	Jan 60 - Aug 01	Daily	Since September '01 it´s a weekly publication. From October '81 to May '84 it had a special edition on Mondays and it´s price was different from that of the rest of the week. In this period we took the price of Tuesday´s edition.
La República	Bogotá	Jan 60 - Dec 05	Daily	Monday´s edition is different from those of the rest of the week and has a different price. We took the price of Tuesday´s edition. From July '85 to June '87, the newspaper had different prices for Bogotá; in this period we took the price of Bogota.
El País	Cali	Jan 67 - Dec 05	Daily	-
El Heraldo	Barranquilla	June 78 - Dec 05	Daily	-
El Universal	Cartagena	Sept 83 - Dec 05	Daily	-
El Colombiano	Medellín	Jan 60 - Dec 05	Daily	-
La Patria	Manizales	May 67 - Dec 05	Daily	-
La Tarde	Pereira	Sept 83 - Dec 05	Daily	-
Vanguardia Liberal	Bucaramanga	Jan 73 - Dec 05	Daily	-
El Espacio	Bogotá	July 65 - Dec 05	Daily	The first entry of the sample corresponds to the first edition of the newspaper.
El Nuevo Siglo	Bogotá	Jan 64 - Dec 05	Daily	Before January '64 the price of the newspaper was not printed, that´s why we don´t have the prices for those dates.
El Diario de Occidente	Cali	Jan 66 - Oct 01	Daily	Since November '01 the newspaper is free and it isn´t included in the sample. The first entry of the sample corresponds to the first edition of the newspaper.
El Mundo	Medellín	April 79 - Dec 05	Daily	-
La Libertad	Barranquilla	June 85 - Dec 01 Sept 03 - Dec 05	Daily	-
La Opinión	Cúcuta	March 85 - Dec 05	Daily	-
El Nuevo Día	Ibagué	Dec 92 - Dec 05	Daily	The first entry of the sample corresponds to the first edition of the newspaper.
El Informador	Santa Marta	Jan 80 - Dec 05	Daily	-
El Diario del Sur	Pasto	March 85 - Dec 88 Aug 91 - Dec 05	Daily	From February '87 to December '88 we took the price of Sunday´s edition.
La Nación	Neiva	June 94 - Dec 05	Daily	Data for the period of November '97 to September '98 were provided by the management of the newspaper.
Crónica del Quindío	Armenia	May 96 - Dec 05	Daily	-
El Liberal	Popayán	Jan 79 - Dec 05	Daily	Up to December '96 the newspaper didn´t circulate on Monday. In this period the price corresponds to Tuesday´s edition.
El Pilón	Valledupar	March 96 - Aug 99 Feb 03 - Dec 05	Daily	-
El Meridiano de Córdoba	Montería	July 95 - Dec 05	Daily	-
El Diario del Otún	Pereira	Feb 82 - Dec 05	Daily	The first entry of the sample corresponds to the first edition of the newspaper.
Portafolio (semanal)	Bogotá	Sept 93 - Oct 97	Weekly	The first entry of the sample corresponds to the first edition of the newspaper.
Portafolio (diario)	Bogotá	Nov 97 - Dec 05	Daily	In '97 Portafolio became a daily publication. The first entry of the sample corresponds to the first daily edition of the newspaper.
Cambio	Bogotá	Aug 98 - Dec 05	Weekly	Before August '98 the magazine circulated with the name Cambio 16. The first entry of the sample corresponds to the first edition of the magazine with the name Cambio.
Cromos	Bogotá	Jan 60 - Dec 05	Weekly	-
Semana	Bogotá	May 82 - Dec 05	Weekly	The first entry of the sample corresponds to the first edition of the magazine.
Aló	Bogotá	Feb 87 - Dec 05	Semi-monthly	The first entry of the sample corresponds to the first edition of the magazine.
Dinero	Bogotá	May 93 - Dec 05	Semi-monthly	The first entry of the sample corresponds to the first edition of the magazine.