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Price-Setting in Mexico and the Real Effects of Monetary Shocks*

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Abstract: In this paper I use novel micro data underlying the Mexican CPI to establish stylized facts about prices in the Mexican economy. I then analyze the implications and consistency of the empirical results for the degree of monetary non-neutrality generated in both time and state-dependent pricing models. I find that the real effects of monetary shocks importantly depend on the type of nominal rigidity considered and on the treatment of sales in the statistics that are calibrated into the models.

Keywords: Price Micro Data, Price Rigidity, Menu Cost Models.

JEL Classification: E30, E31, E32.

Resumen: En este trabajo se utilizan nuevos micro datos del INPC para establecer hechos estilizados sobre la fijación de precios en la economía mexicana. Posteriormente, se analizan las implicaciones y la coherencia de los resultados empíricos para el grado de la no neutralidad monetaria generada por modelos tiempo y estado dependientes. Se encuentra que los efectos reales de los choques monetarios dependen en gran medida del tipo de rigidez nominal considerada y del tratamiento de las ofertas en los momentos utilizados para la calibración de los modelos.

Palabras Clave: Micro Datos de Precios, Rigideces Nominales, Modelos con Costos de Menú.

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1 Introduction

How important is the role of sticky prices as a transmission mechanism through which monetary shocks could influence the economy's real variables remains a central question in monetary economics. General equilibrium models that incorporate nominal rigidities, for example, by assuming a fixed probability of price change or a cost of adjusting prices, can generate short-run non-neutrality of monetary shocks since changes in the money supply are not matched one-by-one for changes in expected inflation.

In this setting, the evidence obtained from the data on individual firms' price-setting matters. First, because the use of a model that incorporates nominal rigidities is justified by the evidence in support of the relevance of this feature. And second, because the measured stickiness of prices is a key determinant for the answers this type of models get. If prices change relatively infrequently these models will predict large real effects of monetary shocks, whereas if prices change quite frequently the predicted real effects will be small. Because of this, the price-setting behavior observed in the data is important to properly incorporate micro-founded nominal rigidities into general equilibrium models that aim to quantify monetary non-neutrality.

Taking advantage of a novel price data set, the objective of this paper is to analyze these subjects in the context of the Mexican economy. I do so in two steps. First, using micro data from the Mexican Consumer Price Index (CPI), I establish stylized facts about prices in this economy. Second, I analyze the implications of the price-setting empirical results for the degree of monetary non-neutrality generated in the main types of sticky price models. I find that the real effects of monetary shocks importantly depend on the type of nominal rigidity considered and on the treatment of sales in the statistics that are calibrated into the models. This last result suggests that previous studies that considered posted prices (including sales) for sticky price models' calibration have underestimated monetary non-neutrality.

Although there are previous studies about price-setting in Latin American countries, these have faced different limitations in terms of the quality or in the sample coverage of the data.

This paper contributes to the empirical macroeconomics literature by presenting new evidence of the price-setting behavior in Mexico using high quality micro data of product-level price quotes underlying the Mexican CPI. In addition to reporting new estimates of the main price statistics, this paper exhibits the importance of sales in the current price-setting, a fact that had not been documented for Mexico, or for any other Latin American country using CPI micro data.

More specifically, unlike the data sets used in previous studies, the data used in this paper reports additional information about the price collection; particularly, about the presence of sales, stockouts, and product substitutions, features of the CPI micro data that have proved to be relevant for the price-setting empirical analysis in advanced economies (Nakamura and Steinsson, 2008), but previously not available for the case of Mexico. This paper is the first to use this novel micro data to analyze the price-setting behavior in this economy. The data set covers 6 and a half years spanning from June 2009 to December 2015.

In the first part of the paper I establish some stylized facts about prices in Mexico. I find that 17.0% (24.7%) of posted prices adjust each half-month period (month), while for regular prices (excluding sales) this number is 13.4% (20.6%). From the total number of price changes, on average, around 60% are increases. Additionally, the average duration of posted and regular prices is 18.1 and 19.9 semimonthly periods (9 and 10 months), respectively. With respect to the size of price changes, it is found that the magnitude of adjustments is large, more than 10%, while the semimonthly inflation rate is only 0.2%. The magnitude of decreases is larger than the magnitude of increases, even when sales are excluded. Moreover, the distribution of the size of price changes is bimodal, yet small price changes are also frequent. In both the frequency and the size of price adjustments, there is a considerable heterogeneity in the price-setting across goods and services.

Besides these results, I find that sales play an important role for posted prices' flexibility as 31.2% of posted price changes are due to sales. Furthermore, it is observed that from the total number of ending sales only 46.5% of them return to their previous regular level, whereas the 19.8 and 33.7% end at a higher and lower regular price, respectively. Sales'

price-setting exhibits considerable different features compared to the one observed for regular prices. For example, the magnitude of these price changes is 18%, almost the double of the one for regular prices. Additionally, sale prices last, on average, only 4 semimonthly periods (2 months). Lastly, the dynamic features of sales in the Mexican data suggest that this type of changes respond more to idiosyncratic considerations rather than to aggregate ones.

Finally, an analysis of price dynamics shows that the main source of variations in inflation is the size of price changes (the intensive margin) rather than the fraction of products that adjusts (the extensive margin). Although the frequency is only moderately correlated with inflation (0.12 and 0.18 correlation coefficient for posted and regular prices), a larger correlation is found for the frequency of increases, nevertheless the degree of correlation with inflation is stronger for the one of decreases. With respect to the size of price changes, its correlation with inflation is more than 0.9, while, in this case, the correlation for the size of increases is stronger than the one for decreases.

In the second half of the paper, I analyze the implications and consistency of the previous empirical results for both time and state-dependent models. To do so I consider the CalvoPlus model of [Nakamura and Steinsson \(2010\)](#). This model combines both time and state-dependent pricing by assuming that firms face, with some fixed probability, two types of menu costs: a low one and a high one. Given these assumptions, this model nests both [Calvo \(1983\)](#) and [Golosov and Lucas \(2007\)](#) models as special cases.

With respect to the monetary non-neutrality results, I find that these largely depend on the type of nominal rigidity considered. For the case of the Mexican economy, the monetary non-neutrality obtained with the [Calvo \(1983\)](#) model is more than 8 times larger than the one obtained with the [Golosov and Lucas \(2007\)](#) model. Furthermore, the CalvoPlus model results, which combine both time and state-dependent pricing, suggest that monetary non-neutrality is limited as monetary shocks account for around 3% of the observed Mexican business cycle.¹ Importantly, in the three models considered the price statistics used in the

¹Specifically, I compare the variance of real output generated by the model in response to calibrated monetary shocks with the variance of the Hodrick-Prescott (HP) filtered real GDP for Mexico. This measure of non-neutrality is used in [Nakamura and Steinsson \(2010\)](#).

calibrations have large implications for the results. More precisely, the real effects of monetary shocks obtained with regular prices are between 1.5 and 2 times larger than the ones obtained with the posted prices calibrations. These last results show the relevance of sales in posted prices' flexibility and its consequences on the quantitative results obtained with this type of price-setting models.

On the one hand, this paper is related to the empirical price-setting literature. During the last decade, there has been a growing number of studies that employ price micro data underlying official CPIs to uncover evidence on the existence and importance of nominal rigidities (Klenow and Kryvtsov, 2008; Nakamura and Steinsson, 2008). Price-setting studies have not been limited to data bases from official CPIs. Previous studies have analyzed micro data from producer (Nakamura and Steinsson, 2008), as well as export and import price indexes (Gopinath and Rigobon, 2008). Other alternative data sources that have been used are prices from scanner data collected by supermarkets or other large retailers (Midrigan, 2011), or scraped data collected by "scraping" prices from websites (Cavallo and Rigobon, 2011). Nevertheless, the latter data bases generally have a limited sample coverage, compared to the one of the CPIs. For an extensive survey about this literature see Klenow and Malin (2011) and Nakamura and Steinsson (2013).

Regarding the existing studies for Latin American countries, Medina, Rappoport, and Soto (2007) and Barros et al. (2012) analyzed the price-setting in Chile and Brazil, respectively, using CPI micro data. Furthermore, Chaumont et al. (2011) for Chile and Borraz and Zipitriá (2012) for Uruguay, employed high frequency scanner data from supermarkets to document new insights about retail pricing in emerging market economies. Finally, Cavallo (2013) used daily scraped data from large supermarkets in Argentina, Brazil, Chile and Colombia to compare price rigidity across developing countries. From the above, none of the CPI data bases previously used report if the product was on sale at the time it was quoted. Additionally, these data bases are at a monthly frequency. Regarding scanner and scraped data bases, although they have higher frequencies their sample coverage is limited to the goods sold at supermarkets and, from these, only Cavallo (2013) data bases identify the presence of

sales.²

For Mexico, three studies have used alternative data bases derived from the CPI micro data. [Gagnon \(2009\)](#) used monthly averages of the price quotes underlying the Mexican CPI, for the period from 1994 to 2002, to document new insights about the relation of inflation and the setting of individual prices during low and high inflation episodes. [Ysusi \(2010\)](#) employed the same data base to describe the price-setting for the years 2002 to 2009, a period during which the Mexican Central Bank, *Banco de México*, started to implement an inflation targeting policy. Finally, [Cortés, Murillo, and Ramos-Francia \(2012\)](#) used restricted access product-level price indexes, calculated at a semimonthly frequency, to document stylized facts about prices for the period from 2002 to 2011. They concluded that the patterns observed in prices are consistent with those of an economy with low and stable levels of inflation.

Compared to the above mentioned studies for the Mexican economy, my results for the frequency of price change show that prices adjust less often. The higher frequencies found in those studies could be explained by differences in the years and CPI basket covered, and by certain features of the data sets previously used that lead to an overestimation of the frequency of price change, a key statistic when analyzing the real effects of monetary shocks. These features, and their implications for the price statistics calculations, are discussed below. Moreover, because of data limitations, none of the previous studies for Mexico differentiate between posted and regular price changes. As documented in this paper, given the importance of sales for price flexibility, this difference turns out to be relevant for the recent price-setting behavior.

On the other hand, this paper is related to the literature of sticky price models. Two main types of pricing models have been developed to incorporate nominal rigidities. Time-dependent ones, originally proposed by [Calvo \(1983\)](#), assume that firm's timing of price

²Different empirical strategies can be implemented to identify sales when these are not reported in the data, for example, see [Nakamura and Steinsson \(2008\)](#). From the previous Latin American studies, [Barros et al. \(2012\)](#) showed results for regular prices by eliminating large v-shaped changes, while [Chaumont et al. \(2011\)](#) reported that after applying a similar filter posted and regular prices features were very similar, suggesting that this type of filters does not worked well in their data.

changes is fixed regardless of firms' incentives. Hence, in this type of models, firms choose the size of price changes but not the timing. In contrast, state-dependent models incorporate both decisions about the timing and magnitude of price adjustment into the firms' price-setting problem by assuming that there is a fixed cost of changing prices. This cost is called in the literature a "menu cost". In such a setup, the timing and the magnitude of price adjustments will depend on the state of the economy.

The primary interest of these models is to analyze whether firm level nominal rigidities have important macroeconomic implications for monetary non-neutrality. From the previous literature, the menu cost model developed by [Goloso and Lucas \(2007\)](#) was the first to introduce idiosyncratic firm level shocks. The inclusion of this type of shocks considerably improved the ability of the model to match the price statistics observed in the data. However, this model suffers from some inconsistencies, for example that it does not generate sufficient small price changes and predicts small real effects of monetary shocks, compared to the ones generated by a time-dependent models.

Different extensions from this model have followed to question the implications of monetary neutrality generated by menu costs. For example, [Nakamura and Steinsson \(2010\)](#) proposed a model with both intermediate inputs in production and heterogeneity across sectors, [Midrigan \(2011\)](#) introduced models with multi-product firms and [Kehoe and Midrigan \(2015\)](#) examined a model that generates temporary price changes. All these extensions yield increases in the monetary non-neutrality results. As well, the CalvoPlus model extension considered in this paper also doubles the degree of non-neutrality generated by the otherwise [Goloso and Lucas \(2007\)](#) model.

The paper proceeds as follows. In [Section 2](#) I describe the Mexican micro data of consumer prices used in this paper. [Section 3](#) presents the price statistics employed for the empirical analysis. Additionally, in this section, I discuss some issues that arise when calculating price statistics from the CPI micro data. [Section 4](#) documents the empirical findings of the paper regarding the price-setting behavior in the Mexican economy. The CalvoPlus model is described in [Section 5](#). The quantitative results of the real effects of monetary shocks gener-

ated by the three considered models are reported in [Section 6](#). Finally, [Section 7](#) concludes.

2 Mexican Micro Data of Consumer Prices

2.1 Description

This section describes the data set employed in the paper. I use a novel restricted access micro data of product-level price quotes underlying the Mexican CPI, the *Índice Nacional de Precios al Consumidor* (INPC).³ Inflation in Mexico is calculated two times a month with prices collected in 46 major cities and metropolitan areas of the country. Prices of food and travel services are quoted four times a month, while the rest of goods are quoted twice a month. Given that the majority of products in the CPI are quoted on the latter basis, the main results of the paper are reported at a semimonthly frequency.⁴ Compared to other countries' CPIs, commonly quoted on a monthly basis, the Mexican index has the advantage of having high frequency price quotes with nationwide statistical representativeness.

The goods and services in the CPI, from here on referred to as products, are classified into three aggregation levels: (1) item's variety, (2) the city of quotation, and (3) generic item categories. Generic items are broad consumption categories used to group individual products, for example "Carbonated drinks" or "Haircuts", while varieties apply to some generics with further disaggregated classifications. The expenditure weights that are used to calculate inflation are specific to the variety-city-generic level. For this paper, I use the expenditure weights introduced in April 2013 to calculate aggregate statistics from product-level prices.⁵

³Since the creation of the INPC in 1969, Banco de México was in charge of the measurement of inflation. In June 2011, the exclusive right to elaborate the national price indexes was given to the Mexican National Institute of Statistics, the *Instituto Nacional de Estadística y Geografía* (INEGI). To ensure consistency of the inflation measurement, no methodological changes occurred with this administrative change.

⁴Through the paper I refer to the goods and services that are quoted four times a month as weekly quoted products: total of 48 price quotes a year. For these products, semimonthly statistics were calculated considering the prices of the last week of each half-month period. Also, I refer to the products quoted twice a month as semimonthly quoted products: total of 24 price quotes a year.

⁵The Mexican CPI weights are derived from the Survey of Household's Income and Expenditures, *Encuesta Nacional de Ingreso y Gasto de los Hogares* (ENIGH). This ensures that generic item categories cover over 95% of Mexican households' expenditures (INEGI, 2013).

The micro data identifies quoted products at a detailed level. Each price quote has information about the product's brand, a description, and an outlet unique identification number (e.g. Generic item: Carbonated drinks; Brand: Coca-Cola; Description: Bottle of 3 liters, non-returnable; Outlet: 31272 in Mexico City). Additionally, the data used in this paper contain information about the price collection. In particular, it specifies if the product was on sale or was missing when the price was quoted.⁶ There is also a variable that identifies product substitutions and specifies the reason why they were substituted.

This novel data set has some important advantages over the ones used in previous empirical studies for the Mexican economy. On the one hand, [Gagnon \(2009\)](#) and [Ysusi \(2010\)](#) used publicly available data of monthly average prices, across the two or four price quotes in each month, of the products underlying the Mexican CPI. These prices are published every month in the Mexican Government's Official Gazette, the *Diario Oficial de la Federación* (DOF). The use of averages, instead of direct price quotes, complicates the inference of the results because changes in average prices are smaller and more frequent which results, as will be shown below, in a bias on the price-setting statistics.⁷

On the other hand, [Cortés, Murillo, and Ramos-Francia \(2012\)](#) used restricted access product-level price indexes, called relative prices (*precios relativos*), which are calculated twice a month for inflation measurement. The use of direct price quotes for the price-setting analysis is preferable to the use of relative prices for two reasons. First, because for the weekly quoted items these semimonthly indexes are calculated with the average of the two weekly prices, hence, a bias similar to the one present in the DOF data arises. And second, because the practices that are used to compute relative prices under stockout periods could lead to spurious small price changes.⁸ Finally, none of those data sets include information

⁶Sales are defined as non-conditional price discounts in terms of a minimum number of products bought or a determined form of payment ([INEGI, 2013](#)).

⁷For further discussion about the implications of using average prices for the price-setting analysis see [Gagnon \(2009\)](#). To control for the bias generated by the averaging of prices, a filter is employed in that paper to obtain monthly statistics from the DOF micro data.

⁸When a product is missing, there are two approaches to calculate the relative price: (1) impute the mean variation of the relative prices at the lowest available aggregation level, and (2) carry forward the last relative price observed. For the majority of items the first approach, which generates spurious small price changes, is used. For further details about the construction of relative prices for inflation measurement, see [INEGI \(2013\)](#).

about sales or missing values.

This paper is the first to use this novel database to analyze the price-setting dynamics in the Mexican economy. Although this micro data is only available since June 2009, as detailed below, it has richer information than the DOF or the relative prices data sets. Particularly, this data allows me to address the presence of sales and stockouts in the Mexican consumer prices, features that have proved to be relevant for the price-setting empirical analysis using CPI micro data (Nakamura and Steinsson, 2008). Additionally, this paper is the first to directly control for a source of spurious small price changes in the Mexican CPI micro data.⁹ The implications of these features for the empirical results are further discussed in Section 3 and Section 4.

2.2 Sample Coverage

The data set covers 6 and a half years spanning from June 2009 to December 2015. However, to control for price censoring (when the beginning of the price is not observed), for the empirical results of the paper I focus on the last five years of the data. As will be further discussed, censoring generates a downward bias on the price duration estimates since the longest spells are the ones more likely to be left out of the sample. Because of this, I restrict to price spells that ended with a price change that occurred in or after January 2011. For consistency, the rest of the price-setting analysis is done for this sample that goes from January 2011 to December 2015.¹⁰

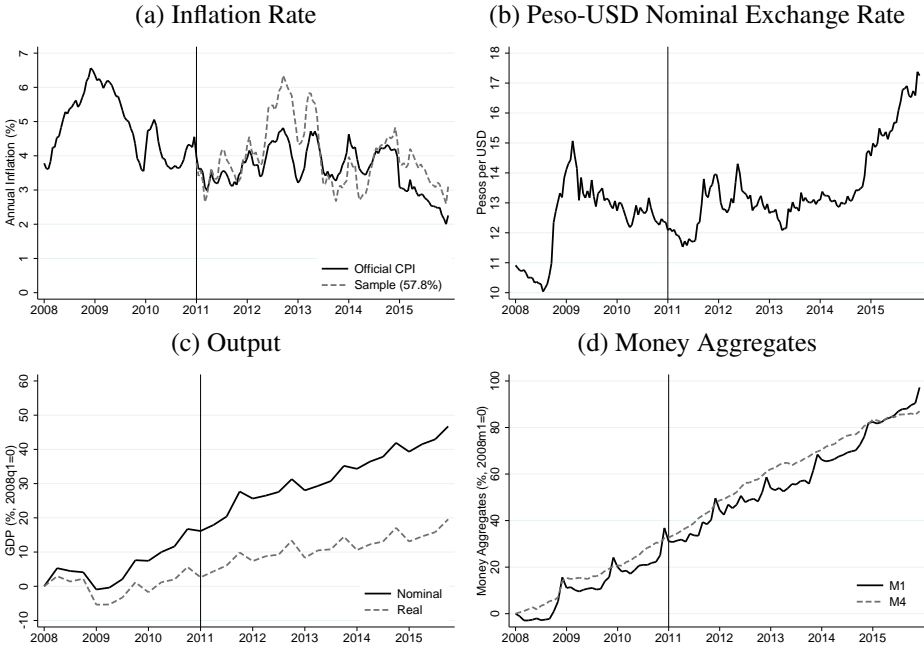
Figure I shows some relevant macroeconomic indicators for the time period covered by the data. During those years, annual inflation in Mexico was relatively stable averaging 3.6%,

⁹The prices of half of the generic items considered in the analysis are recorded in a unit measure (e.g. price per kilogram). In the cases where the product size differs from the unit measure, the posted price is converted as a price per unit and then recorded in the data (INEGI, 2013). This conversion practice, which leads to spurious small price changes, is also present in the DOF and relative prices data bases as those are obtained from the recorded price quotes. As described in Appendix A, in this paper, I address this issue by reconstructing the posted price using a conversion factor reported in the data.

¹⁰The selection of the time period, which is relevant for the duration calculations, has no implications for the rest of the price-setting statistics, such as the frequency and magnitude of price changes. Note that, restricting to this period, the price changes that occurred at the beginning of the sample could have followed from a spell with a maximum duration of 36 semimonthly periods (from July 2009 to December 2010).

a number close to the 3% target of Banco de México. Conversely, the Mexican peso-USD nominal exchange rate presented some important fluctuations during the period, and more recently have shown a depreciation pattern from the end of 2014 onward. Nonetheless, in a recent paper, [Kochen and Sámano \(2016\)](#) show that the exchange rate pass-through to consumer prices in the Mexican economy is low. In terms of economic activity, after the 2008-2009 global financial crisis, real output has been increasing at an annual rate of 2.8% during the sample period. Lastly, money aggregates have been growing at a relatively constant rate.

Figure I: MACROECONOMIC CONTEXT



SOURCE: Banco de México and INEGI.

NOTES: The vertical line indicates the sample starting period: January 2011.

With respect to the CPI basket coverage, out of the total 283 generic items, I restrict my analysis to 244. The main groups of items that were excluded from the analysis are: energy and government regulated fares, shelter, private education and tourism services. All-together the selected items represent 57.8% of the Mexican CPI basket, measured by household expenditure weights. This sample comprises more than 17 million weekly and semimonthly prices from 245,578 different products.

The generic items sample selection is as follows. First, I drop 20 generic items, with a participation of 30.3% in the CPI, whose prices require specific treatments for inflation measurement. For this group of items, called subsystems (*subsistemas*), the prices reported in the data are city aggregated price indexes, which if used in the analysis, would affect the results as spurious small price changes would arise.¹¹ Examples of these items are shelter, private education, and telecommunication services. Second, I exclude the items with regulated prices because their price dynamics reflect administrative considerations rather than those of the market. Given that most of energy and utilities items are regulated, I also drop natural and LP gas. Overall, a total of 11 items with an expenditure weight of 10.8% were excluded for the above said reasons. Third, 6 additional items, with a total weight of 0.23%, were excluded because of missing data before the December 2010 CPI basket revision.¹² Finally, I exclude air transportation and tourism services, with a joint weight of 0.82%, because of spurious small price changes' concerns.¹³

3 Price Statistics

This section introduces the price statistics that are used to characterize the price-setting dynamics in the Mexican economy. First, I present some inflation accounting definitions of the empirical macroeconomics literature. Second, I discuss some relevant issues that arise when calculating aggregate price statistics from the micro data. I make particular emphasis on the treatment of sales, stockouts and product substitutions when calculating the statistics.

¹¹This type of goods are analogous to the composite-good items of the United States CPI. [Eichenbaum et al. \(2014\)](#) show that considering the prices of these items generates a large share of spurious small price changes.

¹²In December 2010 there was a major basket revision that resulted in a reduction of generic item categories, from 315 to the current number of 283. For details about this revision see [INEGI \(2013\)](#).

¹³[Eichenbaum et al. \(2014\)](#) argued that airline fares in the United States CPI micro data are problematic for the price-setting analysis since the final consumer price is affected by a myriad of taxes and fees that could change the final price in small amounts. Additionally, tourism services could have problems related to non-transaction prices. For these reasons I exclude these items from the empirical analysis.

3.1 Definitions

I follow the price-setting empirical literature (Klenow and Kryvtsov, 2008) and define aggregate inflation (π_t) as the weighted sum of log price changes of individual products:

$$\pi_t = \sum_{s \in \Upsilon_t} \omega_t^s \Delta p_t^s \quad (1)$$

where ω_t^s is product s expenditure weight at time t .¹⁴ Υ_t is the set of products considered for the price-setting calculations at that time period. This set could change over time depending on the treatment of sales, missing values and item substitutions. $\Delta p_t^s = p_t^s - p_{t-1}^s$, where p_t^s is the price in logs of product s at time t .

This methodology to compute inflation differs from Mexico's official one which calculates inflation as the percentage change of Laspeyres indexes (INEGI, 2013). I, however, follow the literature standards to make my results comparable to previous studies. Panel (a) of Figure I presents the official CPI annual inflation rate and the one calculated using equation (1) for the 57.8% selected sample. Despite the methodology and sample differences, both series present very similar dynamics with a correlation coefficient of 0.73 over the period between January 2011 to December 2015.

Aggregate inflation can be decomposed as the product of two price statistics, the fraction of products that change their price (the extensive margin) and the average size of those price changes (the intensive margin):

$$\pi_t = \underbrace{\left(\sum_{s \in \Upsilon_t} \omega_t^s I_t^s \right)}_{fr_t} \underbrace{\left(\frac{\sum_{s \in \Upsilon_t} \omega_t^s \Delta p_t^s}{\sum_{s \in \Upsilon_t} \omega_t^s I_t^s} \right)}_{dp_t} \quad (2)$$

¹⁴Product-level weights are function of the expenditure weights at the variety, city and generic item level, to which the product belongs, and of the number of products selected at that aggregation level in each time period. Specifically: $\omega_t^s = (\omega^g \omega^{c,g} \omega^{v,c,g}) / (\#\Upsilon_t^{v,c,g})$, where ω^g are generic item weights, that sum to 1 across items, $\omega^{c,g}$ are city weights, that add to 1 across all cities for each generic, and $\omega^{v,c,g}$ are variety weights, that sum to 1 for each city and generic item. $\Upsilon_t^{v,c,g}$ is the set of all individual products at the (v, c, g) level considered at time t .

where I_t^s is an indicator variable equal to 1 if a price change has occurred and zero otherwise. The term fr_t is henceforth referred to as the aggregate frequency of price change and dp_t is the aggregate size of non-zero price changes, at time t . These statistics jointly characterize the inflation dynamics since they capture how often and for how much prices change. The statistics excluding the time subindex, e.g. fr and dp , are the average of those statistics across time.

As emphasized by [Gagnon \(2009\)](#), an additional decomposition that is also informative of the relationship between inflation and price changes is obtained by separating the frequency and the size between price increases and decreases:

$$\pi_t = \underbrace{\left(\sum_{s \in \Upsilon_t} \omega_t^s I_t^{s+} \right)}_{fr_t^+} \underbrace{\left(\frac{\sum_{s \in \Upsilon_t} \omega_t^s I_t^{s+} \Delta p_t^s}{\sum_{s \in \Upsilon_t} \omega_t^s I_t^{s+}} \right)}_{dp_t^+} + \underbrace{\left(\sum_{s \in \Upsilon_t} \omega_t^s I_t^{s-} \right)}_{fr_t^-} \underbrace{\left(\frac{\sum_{s \in \Upsilon_t} \omega_t^s I_t^{s-} \Delta p_t^s}{\sum_{s \in \Upsilon_t} \omega_t^s I_t^{s-}} \right)}_{dp_t^-} \quad (3)$$

where I_t^{s+} (I_t^{s-}) is an indicator variable equal to 1 if a price increase (decrease) has occurred and zero otherwise. The terms fr_t^+ and dp_t^+ (fr_t^- and dp_t^-) are the frequency and the size of price increases (decreases).

Before proceeding, it is convenient to identify the role of these statistics in time and state-dependent price-setting models. On the one hand, in the [Calvo \(1983\)](#) model it is assumed that the frequency of price change (fr_t) is exogenously fixed. Hence, in time-dependent models the only possible source of variation for inflation (π_t) has to come from the size of price changes (dp_t) of the adjusting firms. On the other hand, menu cost models incorporate the price adjustment decision into individual firms' price-setting problem. Consequently, in state-dependent models, aggregate inflation could vary from both changes in the frequency and the size of price changes.

Additional to the previous statistics, the duration of price spells is of interest since it is a direct measure of price stickiness. For this paper, I consider two empirical approaches to calculate the duration of prices. One approach, commonly used in the literature, is to estimate the duration implied by the frequency of price change. Assuming that price changes follow

an exponential distribution with a constant hazard rate λ , the probability of an adjustment in each time period is given by $fr = 1 - e^{-\lambda}$, which implies a mean duration of price spells equal to $dur^{imp} = 1/\lambda = -1/\log(1 - fr)$. However, as will be shown in [Section 4.2](#), the aggregation level at which the frequency is inverted has large consequences for implied duration calculations.¹⁵

The other approach is to calculate the duration of prices observed in the data. A well-known disadvantage of directly calculating the duration is spell censoring, i.e. when the beginning or the ending of the price is not observed. Spell censoring could bias the duration downwards since large spells are the ones more likely affected. As mentioned, to control for this bias, I restrict to a sample of uncensored spells that ended with a change in or after January 2011. With this sample, I calculate the average duration of price spells (dur) by aggregating prices' durations across spells and products, weighting each spell with the sum of its time-varying weights (w_t^s) along its duration:

$$dur = \frac{1}{T - \omega_{cens}} \sum_s \sum_{\tau} \left(\sum_{t_0^{\tau,s} \leq t \leq t_T^{\tau,s}} \omega_t^s \right) dur^{\tau,s} \quad (4)$$

where τ identifies the price spell, $t_0^{\tau,s}$ and $t_T^{\tau,s}$ are the beginning and the ending periods of spell τ , and $dur^{\tau,s}$ is the spell duration. The term $1/(T - \omega_{cens})$ corrects for the total number of time periods (T) and for the weight of the omitted censored price spells (ω_{cens}).¹⁶

3.2 Sales, Stockouts and Product Introduction

Although the calculations of the previous statistics may seem straightforward, there are some relevant issues to be considered. The presence of sales, stockouts and product substitutions are some features of consumer prices micro data that have direct implications for the statistics calculations. In this context, the empirical strategy chosen to deal with these fea-

¹⁵By Jensen's inequality, since $-1/\log(1 - x)$ is a convex function between 0 and 1, the average implied duration of lower aggregation levels will be larger than the one implied by the aggregate frequency.

¹⁶The statistic in equation (4) is what [Baharad and Eden \(2004\)](#) defined as "average duration per price", but incorporating time-varying weights on prices. This approach to calculate the average duration of prices is similar to the one used in [Klenow and Kryvtsov \(2008\)](#).

tures will depend on which types of price adjustments are considered relevant from a macro perspective and hence should be incorporated in the analysis.

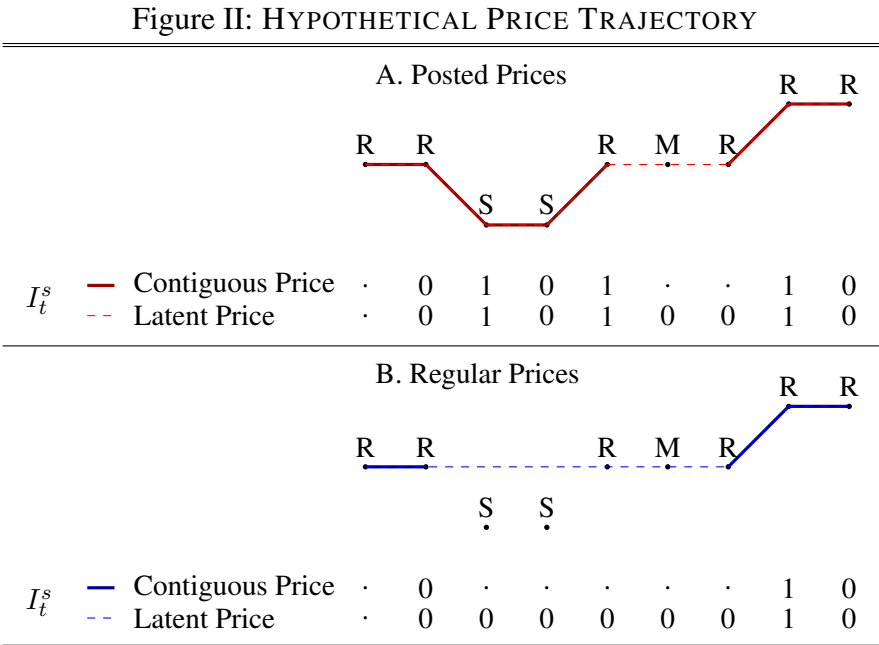
This issue is particularly important with respect to sales as it is still an open question for macroeconomists if sale-related price changes respond to aggregate shocks or rather they reflect entirely idiosyncratic forces (Klenow and Malin, 2011). For example, Nakamura and Steinsson (2008) pointed out that some types of sales are orthogonal to the business cycle, e.g. end-of-season sales in apparel, which suggests the convenience of removing them to analyze the price-setting dynamics. Yet, Kryvtsov and Vincent (2015) recently documented that in the United Kingdom (U.K.) and United States (U.S.) CPI micro data there is evidence that the frequency of sales does respond to macroeconomic conditions, thus, we should care for this type of price changes.

Given these considerations, I distinguish between two classes of price changes: (1) posted and (2) regular, where the latter exclude all sale-related price adjustments. To filter out sale-related changes, identified using the sale flag reported in the data, I consider two empirical strategies. The first is to calculate price changes considering only contiguous regular observations. The second, widely used in the empirical literature, is to construct regular “latent” prices that carry forward the last observed regular price during sale periods. For the treatment of stockouts, analogous strategies are considered: calculate posted and regular prices considering only contiguous observations, and construct latent prices that carry forward the last observed posted or regular price during missing periods. For all the empirical results of the paper, latent prices were constructed using observations at the lowest quoting frequency available (weekly or semimonthly, depending on the product) carrying forward the last observed price during a maximum period of 5 months.¹⁷

Figure II exemplifies the contiguous and latent price strategies for a hypothetical price trajectory. Panel A shows that for posted prices, the price change indicator variable (I_t^s) will differ between strategies only during and after stockouts. Additionally, Panel B shows these

¹⁷If there is no observed price after this period the latent price is left missing as in the contiguous price strategy. This maximum period of 5 months is also used in Nakamura and Steinsson (2008).

strategies for regular prices. Now, since sales are treated as missing regular prices, the contiguous and latent strategies will be distinct during and after both sales and missing periods. This example shows the trade-off between the two empirical strategies. On the one hand, the contiguous price strategy has the advantage of not making further assumptions about the behavior of prices during unobserved periods. On the other hand, however, this strategy implies a loss of information about the price-setting that follows from sales or stockouts periods. As will be shown below, in the Mexican data, excluding this type of changes has non-negligible implications for regular prices' statistics.



NOTES: Hypothetical 9 period price trajectory. R denotes regular price, S sale, and M denotes missing value. I_t^s is an indicator variable equal to 1 if a price change has occurred and zero otherwise. Regular Prices denotes prices excluding sales. Contiguous Price refers to the strategy that calculates price changes with contiguous observations only. Latent Price denotes the strategy that carries forward the last observed price during stockouts and/or sales.

Panel A of [Table I](#) presents the main aggregate statistics considering these two strategies. For the case of posted prices, both strategies yield very similar numbers. Meanwhile, for regular prices, although the frequency and duration statistics are broadly equal, the annual inflation rate obtained with contiguous observations is 2 percentage points (p.p.) higher than the one obtained with latent prices. This result is explained by the price-setting that follows

from sales in the Mexican data. From the total, 46.5% of sales return to their previous level while 33.7 and 19.8% end at a lower and higher regular price, respectively. Hence, restricting to contiguous observations will exclude a considerable share of regular price changes at the end of sales which, as mentioned, are mainly decreases and consequently inflation will be higher. Given these results, in order to include this type of regular price changes and because the other statistics are very similar across strategies, for the rest of the empirical analysis I report price statistics calculated with latent prices. The results considering only contiguous prices are available upon request.

A final issue to be considered refers to price changes associated with product substitutions. In this respect, it can be argued that price adjustments that arise from the introduction of new products should be considered for the price-setting analysis. To examine this, I identify new product introductions using a variable in the data that reports the cause of product substitutions. The major groups of items that have this type of substitutions are household durables, apparel, and transportation goods. Panel B of [Table I](#) reports the aggregate price statistics including price adjustments that result from the introduction of new products. In general, the inclusion of this type of price changes does not affect the main aggregate statistics. Because of this, and given that for inflation measurement none of the price changes related to product substitutions are considered, for the rest of the analysis I focus on price changes of identical products only.

4 Empirical Evidence from the Micro Data

4.1 Frequency of Price Change

This section presents the empirical findings of the paper about the price-setting behavior in the Mexican economy. I start by answering the question about how often prices change. The main statistics related to the frequency of price change are reported in [Table II](#). For both posted and regular prices, the table reports the semimonthly frequency of price change (fr), the fraction of changes that are increases (fr^+/fr), and the median across generic items' semimonthly frequencies ($Md\ fr$). For posted prices, additional statistics related to sales

Table I: PRICE-SETTING UNDER DIFFERENT STRATEGIES

	Posted Prices		Regular Prices	
	Contiguous	Latent	Contiguous	Latent
A. Identical Products				
Annual π	3.6	3.9	6.1	4.1
π	0.2	0.2	0.3	0.2
fr	17.1	17.0	13.4	13.4
dp	0.9	1.0	1.9	1.3
dur	17.5	18.1	19.8	19.9
B. Including New Product Substitutions				
Annual π	3.6	4.1	6.1	4.3
π	0.2	0.2	0.3	0.2
fr	17.1	17.1	13.4	13.5
dp	0.9	1.0	1.9	1.3
dur	17.5	18.1	19.8	20.0

SOURCE: Banco de México and INEGI.

NOTES: With the exception of the annual inflation rate (π), all statistics are in semimonthly frequency. All the statistics reported were calculated for each semimonthly period and then averaged across time. Regular Prices denotes prices excluding sales. fr denotes the semimonthly frequency of price change. dp is the log size of price changes times 100. dur is the duration of prices. Contiguous refers to the strategy that considers contiguous observations only. Latent denotes the strategy that carries forward the last observed price during stockouts and/or sales.

are also presented: the fraction of sale-related price changes ($Sales/fr$)¹⁸, the percentage of ending sales that return to their previous regular level (Return), and the percentage of prices tagged as sales from the total semimonthly observations (Perc.). All statistics reported were calculated for each semimonthly period first, and then aggregated across time.

For the total sample, the table shows that, on average, 17% of posted prices adjust each half-month period. From the adjusting prices, 57.6% are price increases and 31.2% are sale-related changes. At the end of sales only 46.5% return to their previous regular level, while 19.8 and 33.7% end at a higher and at a lower regular price, respectively.¹⁹ On average,

¹⁸I define sale-related price changes as all adjustments in which the current (p_t^s) or the last period price (p_{t-1}^s) was tagged as a sale. Note that, using this definition, in the example presented in Figure II both price changes at the beginning and at the end of the sale period will be accounted as sale-related ones.

¹⁹As a comparison, Klenow and Kryvtsov (2008) reported that about 60% of sales in the U.S. CPI return to their previous regular level.

around 6.8% of all posted prices are tagged as sales each half-month period. Overall, these results highlight the relevance of sales for the frequency of posted price change in the Mexican economy. When sales are excluded, using latent prices as described in [Section 3.2](#), the semimonthly frequency of regular price change is reduced to 13.4%.²⁰ Of the total regular price changes, on average, the share of increases is 60.4%.

Besides these results, the median frequency of price change across consumption categories has also been of interest as it is a statistic less affected by products with a high frequency of price change and, hence, it is a good measure for the degree of price stickiness under heterogeneity. For the Mexican micro data, the table shows that the median across generic items' frequencies, for posted prices, is 10.9% which is considerably smaller than the mean frequency of 17%. Meanwhile, for regular prices, the median frequency across items is 7.7%, also below the 13.4% average.

Furthermore, [Table II](#) presents statistics for generic items' major groups.²¹ The frequencies of price change across groups indicate that there is a large heterogeneity in terms of how often prices change, which was suggested by the median frequency results. This heterogeneity is particularly evident when comparing services to the rest of the goods. For example, considering posted prices, the semimonthly frequency for services is 3.7%, a fifth of the aggregate one, and close to 85% of price changes are increases. In terms of goods, the groups that display higher frequencies are unprocessed food and household goods with 45.2 and 26.3% of prices adjusting each half-month period, respectively. The goods with the lowest ones are apparel and recreation goods with semimonthly frequencies of 6.8 and 8.6%, respectively.

Part of the heterogeneity observed in posted prices is accounted by the prevalence of sales across groups. For example, while in household durables almost 60% of prices are sale-

²⁰The fact that 31.2% of posted price changes are due to sales does not necessarily imply that the frequency of regular price change should be smaller than the posted one in that percentage. Discrepancies could arise when sales do not return to their previous regular level, which in the case of the Mexican data occurs in 53.5% of the cases, because that type of end-of-sale adjustments are counted as both sale-related and regular price changes.

²¹These groups were constructed from both inflation measurement and household expenditure official classifications and are similar to the ones considered by [Nakamura and Steinsson \(2008\)](#) for the U.S. CPI.

Table II: SEMIMONTHLY FREQUENCY OF PRICE CHANGE BY MAJOR GROUP

Major Group	Weight	Posted Prices						Regular Prices		
		Md fr	fr	fr^+/fr	Sales			Md fr	fr	fr^+/fr
					Sales/ fr	Return	Perc.			
Processed Food	14.7	12.8	14.9	58.2	36.4	51.4	6.7	9.6	11.2	61.7
Unprocessed Food	8.4	40.3	45.2	53.6	21.4	36.1	8.1	33.9	39.6	54.4
Household Goods	2.4	29.6	26.3	54.4	46.6	42.8	14.0	19.8	17.3	57.7
Household Durables	1.7	19.0	18.4	54.7	59.6	53.9	19.3	9.8	10.0	61.5
Apparel	5.3	6.7	6.8	61.6	47.1	75.3	6.3	3.9	3.9	71.5
Transportation Goods	3.4	18.9	17.8	63.4	18.2	37.4	7.2	17.4	16.1	66.0
Recreation Goods	1.2	7.1	8.6	59.9	37.9	44.6	6.1	5.0	6.2	64.7
Health and P. Care Goods	5.3	25.3	24.2	54.3	49.2	45.1	15.3	15.9	15.4	57.7
Services	15.5	3.8	3.7	84.3	8.1	51.8	1.2	3.6	3.5	86.8
Total Sample	57.8	10.9	17.0	57.6	31.2	46.5	6.8	7.7	13.4	60.4

SOURCE: Banco de México and INEGI.

NOTES: Regular Prices denotes prices excluding sales. All the statistics reported were calculated for each semimonthly period and then averaged across time. The frequency statistics are reported in percentage per half-month period. fr denotes the aggregate mean frequency of price change. Md fr is the median of generic item level frequencies. fr^+/fr denotes the percentage of price changes that are price increases calculated as the mean frequency of increases over the mean frequency of price change. Sales/ fr denotes the percentage of sale-related price changes. Return is the percentage of ending sales that return to their previous regular price. Perc. is the percentage of observations tagged with a sale flag.

Table III: MONTHLY FREQUENCY OF PRICE CHANGE IN CPI MICRO DATA COMPARISON WITH PREVIOUS STUDIES

Country	Paper	Sample Period	Posted Prices		Regular Prices	
			<i>fr</i>	Md <i>fr</i>	<i>fr</i>	Md <i>fr</i>
Brazil	Barros et al. (2012)	1996:03-2008:12	37.2	40.6		
Chile	Medina, Rappoport, and Soto (2007)	1999:01-2005:07	46.1	33.3		
Euro Area	Dhyne et al. (2006)	1996:01-2001:01	15.1			
France	Alvarez, Le Bihan, and Lippi (2014)	2003:04-2011:04	17.1		14.7	
Israel	Baharad and Eden (2004)	1991:01-1992:12	24.5	21.4		
Mexico	Gagnon (2009)	1994:01-2004:12	29.4			
Mexico	Ysusi (2010)	2002:07-2009:12	35.0			
Mexico	Cortés, Murillo, and Ramos-Francia (2012)	2002:06-2011:12	29.8			
Mexico	Kochen (this paper)	2011:01-2015:12	24.7	18.2	20.6	14.1
Norway	Wulfsberg (2009)	1997:01-2004:12	21.9	14.3	21.3	13.0
Portugal	Dias, Dias, and Neves (2004)	1997:01-2001:01	22.0	11.7		
United Kingdom	Kryvtsov and Vincent (2015)	1996:02-2013:09	17.2		14.6	
United States	Klenow and Kryvtsov (2008)	1988:02-2005:01	36.2	27.3	29.9	13.9
United States	Nakamura and Steinsson (2008)	1998:01-2005:12	26.5	19.4	21.1	8.7

SOURCE: Klenow and Malin (2011) and own research.

NOTES: Regular Prices denotes prices excluding sales. All the frequency statistics are reported in percentage per month. *fr* denotes the aggregate mean frequency of price change. Md *fr* is the median of disaggregated consumption categories level frequencies.

related ones, for services this number is 8.1%. Additionally, for this last group, on average, only 1.2% of prices are on sale each period while for household durables this figure is 19.3%. Among goods, after durables, the groups with the largest share of sale-related price changes are household goods, apparel, and health and personal care goods with around 50% of posted price changes due to sales. Once sales are excluded, the frequency of price change seems more homogeneous. Particularly, the previous mentioned groups have smaller frequencies of regular price change closer to the mean.

In addition to the semimonthly statistics, [Table III](#) reports the monthly frequency of price change calculated using the last (weekly or semimonthly) price of the month. Considering monthly observations, the frequency of price change obtained for posted and regular prices is 24.7 and 20.6%, respectively.²² The table also presents the monthly frequency of price change for different studies across countries using CPI micro data. Although discrepancies in methodology, sample coverage, and expenditure weights make comparisons difficult, it is informative to compare aggregate measures of price stickiness across countries and studies. However, few of the papers listed distinguish between posted and regular prices and, excluding this paper, all of them are for advanced economies.

For posted prices, in general, Euro Area countries display low price change frequencies, while Latin American countries are among those with the highest ones. Countries like Israel and the U.S. have a monthly frequency of price change around 25%. My calculations for Mexico, that 24.7% of the prices change each month, is below the average frequency across the studies presented. Considering regular prices, the monthly frequency I document for Mexico (20.6%) is similar to the results of [Nakamura and Steinsson \(2008\)](#) for the U.S. (21.1%). Yet, large differences arise when comparing the median frequency of regular prices.

Although from the frequency of posted and regular prices the exact percentage of sale-related price changes cannot be directly obtained, comparing these two statistics gives some

²²Note that if the probability of price change was fixed, as assumed in time-dependent models, the monthly frequency implied from semimonthly estimates would be $fr + (1 - fr)fr$. The results presented in [Table II](#) would imply a monthly frequency of 31.1 and 25% for posted and regular prices, respectively. The discrepancies between those and the frequencies obtained using the last price of the month suggest that the fixed frequency assumption may not be adequate for the Mexican data.

insights about the relevance of sales in the price-setting. For example, [Table III](#) shows that for France and the U.K. the frequency of regular prices is around 0.85 times the one for posted prices. In Norway, this number is 0.97 suggesting that sales do not play an important role in that country's price-setting. For the U.S. the share of regular price changes that adjust each month is around 0.80 times the share of posted prices. My calculations for Mexico imply that the monthly frequency of regular prices is 0.83 times the one calculated for posted prices.

Finally, compared to the previous evidence for Mexico I find that prices adjust less often. Using data from the DOF for the period between 1994 to 2004, [Gagnon \(2009\)](#) documented a monthly frequency of posted price change of 29.4%. With that same data for the years from 2002 to 2009 [Ysusi \(2010\)](#) calculated that 35% of prices adjust each month. Finally, using relative prices between 2002 and 2011, [Cortés, Murillo, and Ramos-Francia \(2012\)](#) obtained a frequency of 29.8%. The higher frequencies found in those studies could be explained by the differences in the years and CPI basket covered, or by the characteristics of the data sets used, which, as mentioned in [Section 2.1](#), could bias the frequency upwards.

To analyze the extent of this, for my sample, I compute the monthly frequency of posted price change that would be obtained with those data sets. Using monthly averages, as the ones reported in the DOF micro data, the monthly frequency rises to 33.8%, 9.1 p.p. higher than the one calculated with direct price quotes. This number is close to that reported by [Ysusi \(2010\)](#), which suggests that the large frequency found in that study is accounted for the averaging of prices as that paper did not filter for that as in [Gagnon \(2009\)](#).²³ Additionally, when I calculate the frequency of relative prices, which are also reported in my data, I obtain that 27.1% of prices adjust each month, 2.4 p.p. higher than the baseline. Overall, these results show the importance of using the direct price quotes for the price-setting analysis because otherwise biased calculations about the rigidity of prices, a key statistic when analyzing the real effects of monetary shocks, could arise.

²³The upward bias in the monthly frequency due to the averaging on prices is explicitly recognized by the author ([Ysusi, 2010](#), pp.5).

Table IV: DURATION UNDER DIFFERENT APPROACHES

	Posted Prices		Regular Prices	
	Mean	Median	Mean	Median
Implied Duration				
Aggregate fr	5.4	-	7.0	-
Generic Item fr	15.7	8.0	18.7	11.5
Product fr	19.0	10.0	21.1	12.6
Duration of Price Spells				
Uncensored	18.1	11.0	19.9	14.0
W/ Left Censored	21.8	13.0	23.7	16.0

SOURCE: Banco de México and INEGI.

NOTES: Regular Prices denotes prices excluding sales. All duration results are reported in semimonthly periods. Implied Duration is calculated as $-\log(1 - fr)$, where fr is the frequency of price change. For generic item and product-level results implied durations are obtained at each aggregation level, using that level average frequency, and then the mean or median duration is obtained across them. Duration of Price Spells is calculated as presented in equation (4). Uncensored denotes the calculation considering non-censored spells only. W/ Left Censored includes both uncensored and left censored spells.

4.2 Duration of Prices

The duration of prices has an inverse relation to the frequency of change: the more often a product price adjusts, the shorter its price spells will be. This statistic is also of interest as it is an immediate measure for the degree of price stickiness. Table IV presents different estimations for the duration of prices using the two empirical approaches described in Section 3.1. With respect to the implied duration approach, as mentioned, the aggregation level at which the frequency is inverted has large implications for the results. Specifically, for posted prices, the duration implied by the aggregate frequency is just 5.4 half-months (2.7 months), whereas the average across the implied durations of generic item and product-level frequencies is 15.7 and 19 semimonthly periods (7.8 and 9.5 months), respectively. The table also reports the median of the implied durations across these two lower aggregation levels. Particularly, the median duration implied by product-level frequencies is 10 and 12.6 semimonthly periods (5 and 6.3 months) for posted and regular prices, respectively.

On the other hand, the mean duration of uncensored price spells, calculated as in equation

Table V: DURATION OF PRICE SPELLS BY MAJOR GROUP

Major Group	Weight	Posted Prices				Regular Prices	
		<i>dur</i>	Md <i>dur</i>	Sales		<i>dur</i>	Md <i>dur</i>
				<i>dur</i>	Md <i>dur</i>		
Processed Food	14.7	19.3	14.0	3.7	2.0	21.5	18.0
Unprocessed Food	8.4	6.8	3.0	2.1	1.0	7.8	4.0
Household Goods	2.4	10.9	5.0	3.3	2.0	13.9	8.0
Household Durables	1.7	10.0	6.0	4.9	4.0	14.2	10.0
Apparel	5.3	23.7	18.0	6.3	4.0	29.8	23.0
Transportation Goods	3.4	8.7	6.0	5.9	5.0	9.3	7.0
Recreation Goods	1.2	23.1	17.0	6.4	4.0	25.2	19.0
Health and P. Care Goods	5.3	11.0	6.0	3.9	3.0	14.4	10.0
Services	15.5	33.3	27.0	9.5	7.0	33.7	27.0
Total Sample	57.8	18.1	11.0	4.0	2.0	19.9	14.0

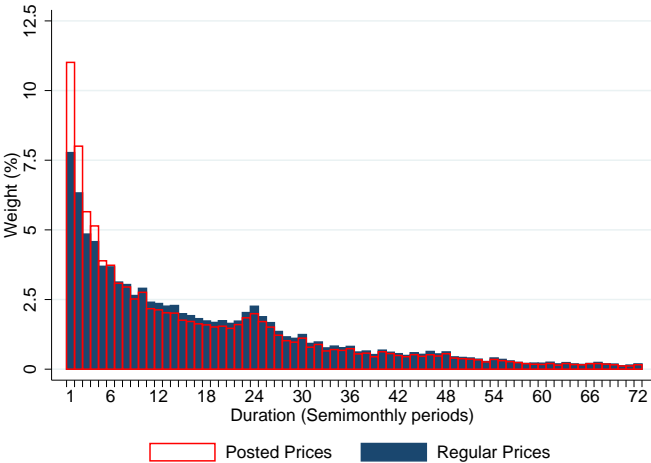
SOURCE: Banco de México and INEGI.

NOTES: Regular Prices denotes prices excluding sales. All duration results are reported in semimonthly periods. *dur* (Md *dur*) denotes the mean (median) duration of uncensored price spells calculated as in equation (4).

(4), is 18.1 and 19.9 semimonthly periods (9 and 10 months) for posted and regular prices, respectively. The median across uncensored spells is 11 and 14 semimonthly periods (5.5 and 7 months), respectively. As discussed, a disadvantage of this approach is the downward bias generated by spell censoring. To assess the degree of this bias on the previous results, [Table IV](#) also reports the duration of prices including left censored spells, which circumvents this type of censoring by assuming that the first observed price was a new price. Including these spells, the mean and median duration, for both posted and regular prices, increases almost 4 and 2 half-month periods, respectively. These results show that, even when restricting to uncensored spells that ended with a change in or after January 2011, the duration obtained with uncensored spells will be downwards biased in more than 1.5 months. Nevertheless, these calculations are still larger than the ones implied by generic item level frequencies and are in line with the ones obtained using product-level frequencies. Given these results, for the remainder of the paper I focus on the duration of uncensored spells as it seems the most robust statistic without making assumptions about the distribution of price changes or about spells' censoring. Yet, the results should be interpreted considering this downward bias.

In addition, [Table V](#) shows the duration of prices by items' major groups. The table presents the mean and median duration of posted, sale, and regular uncensored price spells. For the total sample, the table reports that the average duration of a sale in the Mexican data is 4 half-month periods (2 months), while its median duration is 2 (1). Consistent with the results for the frequency of price change, [Table V](#) shows a large heterogeneity in the length of prices. Now, the groups that reported the smallest frequencies are the ones with larger durations. For example, for posted prices, the mean duration of services is 33.3 semimonthly periods (13.5 months), while for apparel and recreation goods the mean duration is around a year. The goods with the shortest price spells are unprocessed food and transportation goods, with a mean duration of 6.8 and 8.7 half-month periods (3.4 and 4.3 months), respectively.

Figure III: DISTRIBUTION OF THE DURATION OF PRICE SPELLS



SOURCE: Banco de México and INEGI.
 NOTES: Weighted histogram of the duration of uncensored price spells across time and product-level items. Regular Prices denotes prices excluding sales. Bin size is 1 half-month period.

4.2.1 The Distribution of the Duration of Prices

One advantage of measuring the duration of price spells directly from the data is that it allows a distributional analysis. [Figure III](#) presents the distribution of the duration of price spells ($dur^{\tau,s}$) across products and time, using the spell-level weights defined in equation (4), for both posted and regular prices. In line with the results about the duration of sales, the

figure shows that the distribution of posted prices has a larger share of spells of less than 2 months, compared to the one of regular prices. Moreover, for posted prices over 11% of the price spells last only one half-month period, while when sale-related changes are excluded this number is around 7%. The distribution of the duration of prices has an important right bias, with a considerable share of prices lasting 1 year (24 half-month periods) or more. Opposite to the results for the frequency, a consequence of the right bias in the distribution is that the median duration across spells is smaller than the mean duration, as was previously documented in [Table IV](#).

4.3 Size and Magnitude of Price Changes

The size of non-zero price changes is the other statistic that, together with the frequency of price change, characterizes inflation dynamics. [Table VI](#) presents the empirical results for the size of price changes across items' major groups. For both posted and regular prices, the table presents the size of price changes (dp), the magnitude of changes given by the size of price changes in absolute value ($|dp|$), and the magnitude of price increases ($|dp^+|$) and decreases ($|dp^-|$). For posted prices, additionally, the magnitude of sale-related price changes is also presented. Note that through the paper I refer to the size of price changes, calculated as the log-difference times 100, as percentage changes.

[Table VI](#) shows that, on average, the magnitude of posted price changes is 12.4%. From these, the magnitude of sale-related price changes averages 18.0%. In turn, the magnitude of regular changes is 10.1%. These results show that the magnitude of price changes is large, particularly compared to the semimonthly inflation rate of 0.2%. Moreover, the average size of price changes, which averages increases and decreases, is considerably smaller than the magnitude (1.0 and 1.3% for posted and regular prices, respectively). This fact suggests the importance of idiosyncratic shocks to marginal costs as it reflects that in each time period there are multiple changes, of an average magnitude of 10%, which cancel out to a 1% average. It could be argued, however, that the low average size reported in [Table VI](#) may be a result of the averaging across time. To address this concern, the table reports this statistic taking the absolute value of dp_t at each semimonthly period before averaging across time (dp

Table VI: SIZE AND MAGNITUDE OF PRICE CHANGES BY MAJOR GROUP

Major Group	Weight	Posted Prices						Regular Prices				
		dp at t	dp	$ dp $	$ dp^+ $	$ dp^- $	Sales	dp at t	dp	$ dp $	$ dp^+ $	$ dp^- $
							$ dp $					
Processed Food	14.7	1.3	1.2	9.4	9.1	9.8	13.5	1.7	1.6	7.4	7.2	7.5
Unprocessed Food	8.4	2.3	0.5	16.0	15.3	16.6	24.0	2.4	0.5	14.0	13.3	14.7
Household Goods	2.4	1.3	0.5	11.3	10.9	11.7	15.6	1.3	1.0	7.8	7.6	8.0
Household Durables	1.7	1.5	0.6	11.7	11.2	12.2	14.3	1.9	1.5	8.1	7.8	8.6
Apparel	5.3	3.8	1.7	16.2	14.4	18.6	24.8	3.3	3.2	8.9	8.4	9.9
Transportation Goods	3.4	1.0	0.7	3.8	3.6	4.4	7.2	1.1	0.8	3.2	3.1	3.5
Recreation Goods	1.2	2.2	1.4	11.6	10.9	12.8	18.1	2.6	2.2	8.1	7.9	8.2
Health and P. Care Goods	5.3	1.2	0.5	12.5	11.9	13.1	17.5	1.2	0.9	8.1	7.7	8.5
Services	15.5	4.3	4.1	7.8	7.1	11.1	20.6	4.6	4.5	6.6	6.4	7.7
Total Sample	57.8	1.2	1.0	12.4	11.6	13.5	18.0	1.6	1.3	10.1	9.4	11.1

SOURCE: Banco de México and INEGI.

NOTES: Regular Prices denotes prices excluding sales. With the exception of dp at t , all the statistics reported were calculated for each semimonthly period and then averaged across time. dp denotes the average size of price changes. $|dp|$ is the magnitude of price changes. $|dp^+|$ and $|dp^-|$ are the magnitude of price increases and decreases, respectively. dp at t is the average size of price changes at each half-month period taking the absolute value of each period before averaging across time.

at t). The resulting averages are only slightly above the previous numbers, which show that the cause of the small average size is intra-period heterogeneity. Similar evidence has been cited in previous papers, such as [Golosov and Lucas \(2007\)](#), as a motivation for the use of idiosyncratic productivity shocks in price-setting models.²⁴

On the other hand, with respect to the magnitude conditional on the direction of the change, the magnitude of decreases is larger than the one of increases, even when sales are excluded (11.1 and 9.4% for the case of regular prices). A state-dependent model with trend inflation also generates this result. As will be described below, in a menu cost model firms' decision variable is its real price deflated by the aggregate price level. Under trend inflation, there will be relatively fewer small price decreases because, when the desired real price of a firm is below the current one, there could be cases in which it is optimal to leave the nominal price unchanged, and not pay the menu cost, letting aggregate inflation to reduce the real price. Note that this mechanism does not apply for price increases. For this reason, in a menu cost model with trend inflation price decreases will be larger in magnitude than price increases.

Finally, the results in [Table VI](#) suggest that price-setting is also heterogeneous in terms of the size of price changes. Considering posted prices, between the major groups, unprocessed food and apparel are the groups with the largest price changes (average magnitudes of more than 15%). In contrast, the groups with smallest price changes are transportation goods and services with changes of less than 4 and 8%, respectively. Group heterogeneity is substantially reduced considering regular prices, which shows that the presence of sales across groups plays an important role in explaining the above said differences. The groups with the largest sales are apparel and unprocessed food, with sale-related changes of around 25%.

²⁴The evidence that the magnitude of price changes is, on average, large has been consistently documented in previous studies ([Klenow and Malin, 2011](#)).

4.3.1 The Distribution of the Size of Price Changes

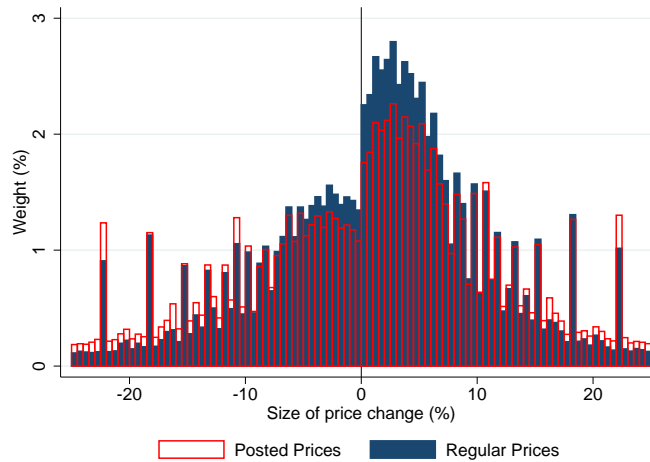
Additional to the average size, recent literature has also focused on the complete distribution of the size of non-zero price changes. This distribution is of interest since competing theories of price stickiness have different implications for its shape, particularly on the number of modes (Cavallo and Rigobon, 2011). For example, in time-dependent models, such as Calvo (1983), the distribution of price changes will inherit the distribution of idiosyncratic shocks. Therefore, under the assumption of normally distributed productivity shocks, the distribution of the size of price changes will have a unimodal shape centered at zero (Panel a, Figure VI).

In contrast, in state-dependent models the menu cost hypothesis has as an implication that small price changes are unlikely to occur. This is because, in this type of models, firms will only change their price when the desired price is far from the current one, and hence the benefit of adjusting is larger than the menu cost. Consequently, when a firm adjusts, the price change should be of a considerable magnitude. In particular, the Golosov and Lucas (2007) menu cost model generates a bimodal distribution with no small price changes (Panel b, Figure VI).

Because of these implications, the distribution of price changes found in the data is useful to compare across competing nominal rigidity theories. In the empirical literature, however, there is mixed evidence about the shape of this distribution. For example, Klenow and Kryvtsov (2008) documented that, for the U.S. CPI micro data, the distribution of price changes is close to unimodal centered at zero with a large share of small price changes. In contrast, Cavallo and Rigobon (2011) used scraped data from different supermarkets across 22 countries and concluded that unimodality is rejected in close to 2/3 of the retailers they analyzed.

For the case of Mexico, Figure IV shows the distribution of the size of non-zero price changes across products and time, for both posted and regular prices, using the product-level weights of equation (1). The histogram shows that, in the Mexican data, the distribution of

Figure IV: DISTRIBUTION OF THE SIZE OF PRICE CHANGES



SOURCE: Banco de México and INEGI.

NOTES: Weighted histogram of the size of non-zero price changes across time and product-level items. The size of price changes is the log size times 100. Regular Prices denotes prices excluding sales. The figure plots the price changes distribution in the $|\pm 25\%$ range only. Bin size is 0.5%.

price changes is bimodal centered at zero, with the right side mode having considerably more weight. The difference in the relative weight between the mode of increases and decreases could be explained by the existence of downward rigidity in some products and by the inflation level prevalent in Mexico. The figure also suggests that there is a considerable mass of small price changes, particularly for regular prices. Finally, the importance of sales for the size of price changes is also manifested in the overlapped histograms for posted and regular prices. Particularly, the distribution of posted prices, compared to the one of regular prices, exhibits a considerably lower weight in the modes close to zero while it shows fatter tails starting at price changes larger than 10%.

4.3.2 How Frequent Are Small Price Changes?

The bimodal shape of the distribution of the size of price changes in state-dependent models is a consequence of infrequent small price changes. Given this implication of the menu cost hypothesis, evidence of small price changes in the data has been used to criticize this type of price-setting models. In a recent paper, [Eichenbaum et al. \(2014\)](#) reassessed the previous evidence for the U.S. and argued that the majority of small price changes observed, in both CPI and scanner data, are usually generated by measurement errors. Given this ev-

idence, I analyzed in detail potential measurement problems in the Mexican data that could affect the empirical results. Particularly, as described in [Subsection 2.2](#), because of spurious small price changes' considerations, some generic items were excluded from the analysis.

Still, there is an additional source of spurious small price changes in the Mexican CPI micro data. For 113 of the 244 in-sample generic items, prices are recorded in a unit measure, e.g. price per kilogram. Whenever a product has a different size, the posted price is converted to the common unit ([INEGI, 2013](#)). As a consequence of this practice, spurious small price changes are likely to occur. To address this particular problem, I reconstruct posted prices using a variable in the data that reports the conversion factor used to adjust the original price to the unit measure. Additionally, I correct for all price changes of less than 1 cent, which are clearly due to a measurement error. The implications of these measurement errors and further details of these corrections are presented in [Appendix A](#). All the results presented in the paper were calculated with the corrected prices.²⁵

[Table VII](#) presents an accounting exercise that analyzes how frequent small price changes are in the Mexican price-setting. To characterize small price changes, I consider three thresholds commonly used in the literature: price changes of less than 1, 2.5 and 5%. For both posted and regular prices, the table reports the percentage of price changes that are smaller, in absolute value, than the above said thresholds ($\#/fr$). Plus, for the 5% threshold, the share of changes which are increases (Pos.) is also presented. For posted prices smaller than 5%, the share of sale-related adjustments (Sales) is reported as well.

From the total sample of posted price changes, 5.8% are smaller than 1%, 15.7% are smaller than 2.5%, and 32.1% are smaller than 5%. Within posted price changes smaller than 5%, 62.3% are increases and 14.7% are due to sales. Meanwhile, for regular prices, these numbers are larger: 7.3, 19.4 and 39.1%, for price changes smaller than 1, 2.5 and 5%, respectively. In this last group, on average, 63.7% of the changes are increases.²⁶ These

²⁵Spurious small price changes have implications for the frequency of small price changes and for the shape of the size of price changes' distribution. Nonetheless, its impact on the aggregate statistics is negligible: for both posted and regular prices, the frequency increases in around 0.2 p.p., while the size of price changes is reduced in 0.1 p.p.

²⁶Note that, for both posted and regular prices, the share of increases smaller than 5% is larger than the one

differences in the prevalence of small price changes between posted and regular prices were expected given the large magnitude of sales previously documented. Considering any of these thresholds, the frequency of small price changes presented in [Table VII](#) is in the middle between the previous evidence of small price changes of [Klenow and Kryvtsov \(2008\)](#) and [Eichenbaum et al. \(2014\)](#) for the U.S. CPI.²⁷ These results suggest that the prevalence of small price changes in Mexico is moderate.

[Table VII](#) also presents these results for the items' major groups. For example, transportation goods is the group with the largest share of small price changes regardless of the threshold considered. Furthermore, for this group close to 80% of all price changes, in absolute value, are of less than 5%, which could be explained by the fact that a price change of a non-negligible amount of money in this type of items is a small share of the total cost. On the other hand, the group with the smallest share of small price changes is unprocessed food, consistent with the large magnitude of price changes previously documented for this group.

Overall, the previous results for the size of price changes in Mexico do not represent a conclusive evidence in favor of one of the sticky price theories in particular. Still, evidence that price decreases are larger than increases, the bimodal shape of the size of price changes distribution, and the moderate frequency of small size price changes are facts that are in line with the menu cost hypothesis. But, as previously discussed, the [Golosov and Lucas \(2007\)](#) menu cost model will not generate sufficient small price changes as observed in the data. Nevertheless, a menu cost model with, for example, multi-product firms ([Midrigan, 2011](#)) can do a better work matching the empirical distribution of the size of price changes. In the following sections, I consider the CalvoPlus model of [Nakamura and Steinsson \(2010\)](#) which combines both time and state-dependent pricing by assuming that, with some fixed proba-

reported in [Table II](#), which considers all price changes. These results are in line with [Chen et al. \(2008\)](#) who document, using scanner data from the U.S., that small price increases are more frequent than decreases. This evidence seems consistent with the consumer inattentiveness hypothesis, which predicts that retailers could gain by implementing small price increases and, consequently, will be more frequent than decreases.

²⁷Specifically, [Eichenbaum et al. \(2014\)](#) calculated that, correcting for problematic small price changes, the share of price changes of less than 1, 2.5 and 5% in absolute value is: 3.6, 10.5, and 24.4% for posted prices, and 5.0, 13.8 and 32.2% for regular prices. These numbers are considerably smaller than the calculations of [Klenow and Kryvtsov \(2008\)](#) of 11.3, 23.4, and 39.8% for posted prices and 12.1, 25.4, and 44.3% for regular prices.

bility, firms receive the opportunity to adjust its price at a relatively lower cost. As will be shown, this assumption allows this model to generate the share of small price changes of the data.²⁸

4.4 Price Dynamics

This final subsection of the empirical results describes the aggregate dynamics of prices across the sample. The analysis of price-setting over time is of interest because in the presence of nominal rigidities price-setters have dynamic decision problems and hence the dynamic features found in the data are also helpful to distinguish between different price-setting theories (Klenow and Malin, 2011). Figure V presents semimonthly time series of the main price statistics. For easiness, these series are presented for posted prices only. The ones of regular prices present very similar dynamics.

On the one hand, Panel (a) presents the times series of inflation (π_t) and the frequency of price change (fr_t), i.e. inflation's extensive margin. Compared to the semimonthly inflation rate, the frequency of price change seems to be more volatile period to period, although fluctuating close to its 17.0% mean. The relation between these two series is relatively weak, with a correlation coefficient of only 0.12. When sales are excluded, this correlation is a bit larger at 0.18. Furthermore, the panel shows that in the first half-month period of 2014, as a response to the tax reform of that year, the share of adjusting prices rose to 27.5%.²⁹ The response of the frequency of price change to aggregate shocks, such as fiscal policy ones, has been interpreted in previous literature as supporting evidence of state-dependent pricing policies (Álvarez and Hernando, 2006).

On the other hand, Panel (b) analyzes the relation between inflation and the size of price change (dp_t), which captures how much of the inflation variation is explained by the intensive margin. The panel shows that both series comove almost perfectly. Specifically, the

²⁸For further discussion about small price changes and their aggregate implications in menu cost and other price-setting models see Midrigan (2011) and Alvarez, Le Bihan, and Lippi (2014), respectively.

²⁹The Mexican tax reform of January 2014 implied an increase of different taxes: (1) the preference rate of the Value Added Tax (VAT) in the frontier zones was eliminated to match the level of the rest of country (an increase of 11 to 16%); (2) an additional tax to sugary drinks and other products with high caloric density was implemented; and (3) the inclusion to the VAT of some goods and services, which were previously exempted.

Table VII: FREQUENCY OF SMALL PRICE CHANGES BY MAJOR GROUP

Major Group	Weight	Posted Prices					Regular Prices			
		$ \Delta p < 1\%$	$ \Delta p < 2.5\%$	$ \Delta p < 5\%$		$ \Delta p < 1\%$	$ \Delta p < 2.5\%$	$ \Delta p < 5\%$		
		<i>#/fr</i>	<i>#/fr</i>	<i>#/fr</i>	Pos.	Sales	<i>#/fr</i>	<i>#/fr</i>	<i>#/fr</i>	Pos.
Processed Food	14.7	7.0	17.7	36.4	59.5	18.7	9.2	22.7	45.3	61.0
Unprocessed Food	8.4	2.0	8.6	22.1	58.3	10.2	2.3	9.9	25.2	58.8
Household Goods	2.4	6.1	15.6	31.3	58.3	21.0	9.0	23.0	44.8	59.3
Household Durables	1.7	4.7	12.0	27.2	59.5	33.9	8.8	21.0	42.7	63.4
Apparel	5.3	3.5	9.4	22.6	72.5	6.2	6.2	16.5	39.2	72.9
Transportation Goods	3.4	24.7	51.0	76.5	63.4	13.8	27.2	54.7	80.1	65.6
Recreation Goods	1.2	6.9	16.4	31.7	63.8	13.1	9.7	22.8	43.6	63.9
Health and P. Care Goods	5.3	7.0	16.3	31.0	58.5	21.9	10.7	24.5	45.5	59.5
Services	15.5	4.6	21.7	48.4	86.8	2.0	5.0	23.2	51.4	87.0
Total Sample	57.8	5.8	15.7	32.1	62.3	14.7	7.3	19.4	39.1	63.7

SOURCE: Banco de México and INEGI.

NOTES: Regular prices denotes prices excluding sales. All the statistics reported were calculated for each semimonthly period and then averaged across time. $|\Delta p| < x\%$ denotes the set of price changes, in absolute value, of less than $x\%$, where x is 1, 2.5 and 5, respectively. *#/fr* is the percentage of observations in the set with respect to the total number of price changes. Pos. (Sales) denotes the percentage of price increases (sale-related price changes) in the set.

correlation coefficient between these two is 0.99 (0.98 excluding sales), around 8 times the correlation between the frequency and inflation. The size of price changes time series shows larger deviations than inflation, and its exhibits almost no correlation with the frequency (correlation coefficient of 0.01, and 0.02 for regular prices). Overall, these results indicate that inflation is almost completely accounted for by the size of price changes and not by the share of adjusting products.

To further analyze the importance of extensive and intensive margins for inflation dynamics, [Klenow and Kryvtsov \(2008\)](#) proposed the following variance decomposition:

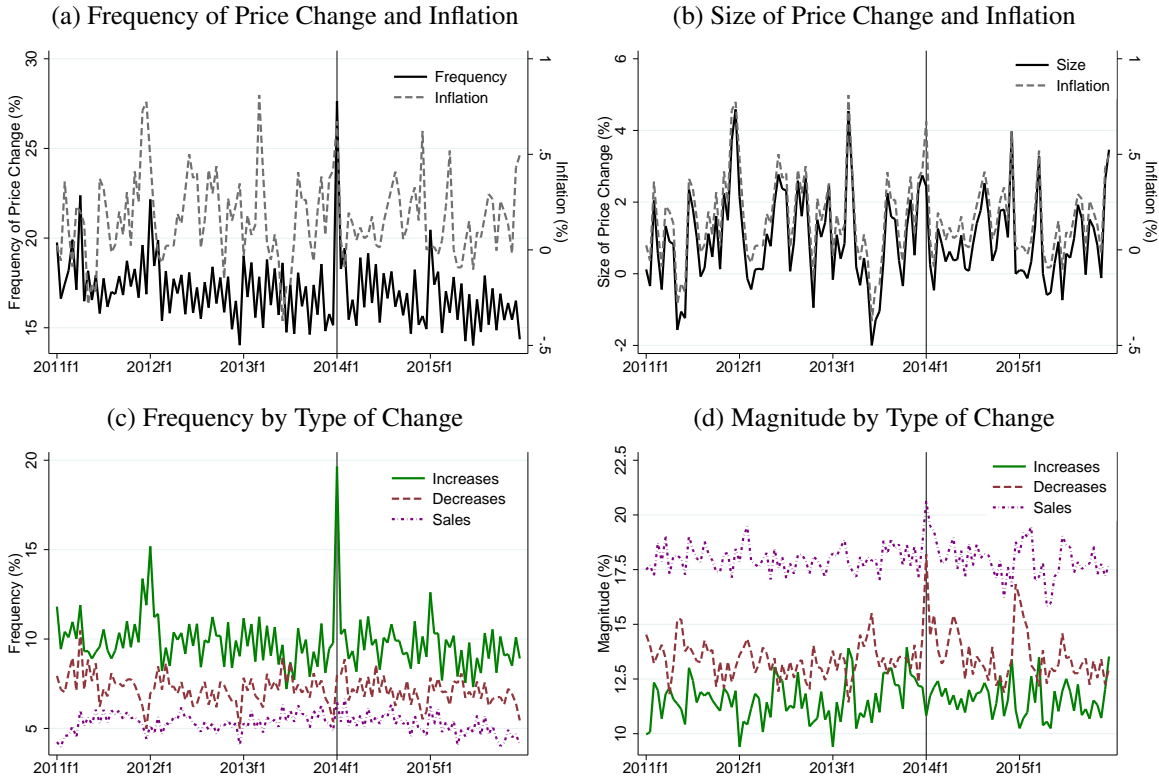
$$\mathbb{V}(\pi_t) = \underbrace{\mathbb{V}(dp_t) fr^2}_{IM} + \underbrace{\mathbb{V}(fr_t) dp^2 + 2 fr dp \mathbb{C}(fr_t, dp_t)}_{EM} + \mathcal{O}_t \quad (5)$$

where IM and EM denote the variance of inflation accounted for by the intensive and extensive margin, respectively. This decomposition is obtained by taking the variance of first-order Taylor series expansion of π_t , expressed as in equation (2), around the frequency and size sample means (fr and dp). \mathcal{O}_t are higher order terms from the Taylor expansion that are functions of fr_t .

Using this decomposition, the share of inflation variance accounted by IM is 97.2%. For regular prices, this share is only slightly smaller at 92.1%. These results show that the frequency of price change, which is the key endogenous variable that differentiates state and time-dependent models, is a relatively unimportant source of variations in inflation. Analogous results were obtained by [Klenow and Kryvtsov \(2008\)](#) for the U.S. and by [Cortés, Murillo, and Ramos-Francia \(2012\)](#) for the case of Mexico. On the contrary, [Gagnon \(2009\)](#) documented that during the high and volatile inflation period between 1994 to 2002 in Mexico, EM accounted for a much larger share of inflation variance (IM share of 41.4%). My results are in line with the above mentioned studies, which concluded that in a context of low and stable inflation, such as the one observed through the analyzed sample, the IM dominates the variance of inflation.

Besides the previous results, panels (c) and (d) of [Figure V](#) present the dynamics of the

Figure V: PRICE DYNAMICS



SOURCE: Banco de México and INEGI.

NOTES: Semimonthly time series for posted prices statistics. The vertical line indicate the tax changes of January 2014 in Mexico.

frequency and magnitude of price change distinguishing by the type of the adjustment. Panel (c) shows the time series of the frequency of increases (fr_t^+), decreases (fr_t^-), and sale-related changes. These series show that most of the variation in the frequency of price change, presented in panel (a), is explained by the frequency of increases. With respect to its relation with inflation, the correlation coefficient with the frequency of increases is 0.47, whereas with the frequency of decreases is -0.57. Interestingly, the degree of correlation with inflation is still stronger for decreases even when sales are excluded (correlations of 0.53 and -0.61 for regular increases and decreases, respectively). The frequency of changes due to sales is the one that exhibits the least variation across the sample and also the weakest relation with inflation (correlation coefficient of -0.09).

Panel (d) presents the time series of the magnitude of price increases (dp_t^+), decreases

($|dp_t^-|$), and sale-related changes. The series of the magnitude of increases and decreases shows a similar degree of variation across time. Consistent with the results of [Section 4.3](#), the magnitude of price decreases is above the one of increases in almost all the periods of the sample. Regarding its relation with the inflation rate, in this case, it is stronger for increases than for decreases (correlation coefficient of 0.47 and -0.39 for increases and decreases, respectively). The ordering in correlations is preserved also when sales are excluded (correlation of 0.59 and -0.53, respectively). Finally, the magnitude of sales, which is the largest of the three across the sample, exhibits even less variations across time around its 18.0% average. This last result, together with the frequency dynamics, suggests that sales in the Mexican data may be responding more to idiosyncratic forces rather than aggregate ones, as was argued by [Nakamura and Steinsson \(2008\)](#) for the case of the U.S.

Overall, as a conclusion of this section, the behavior of prices across time shows that the intensive margin (the size of price changes), rather than the extensive one (the fraction of products that adjust), is the main driver behind inflation dynamics. This fact suggests that time-dependent models hold well in this dynamic aspect of data because, as mentioned, in this type of models the variance of inflation has to come from the intensive margin only. Nonetheless, as discussed in detail in [Klenow and Kryvtsov \(2008\)](#), a similar result can be obtained with state-dependent models that incorporate relatively large idiosyncratic shocks, such as [Goloso and Lucas \(2007\)](#).

In this paper, rather than focusing on one of the pricing models in particular, I analyze the implications of the previous empirical results in both time and state-dependent type of models. To do so I consider the CalvoPlus model of [Nakamura and Steinsson \(2010\)](#). This model combines both time and state-dependent pricing by assuming that firms face, with some fixed probability, two types of menu costs: a low one and a high one. Given this assumption, as will be described below, this model nests both [Calvo \(1983\)](#) and [Goloso and Lucas \(2007\)](#) models as special cases.

5 The CalvoPlus Model

5.1 Setup

The model consists of a single sector economy with complete markets in which individual firms are subject to an aggregate shock on inflation and to an idiosyncratic productivity shock. The economy has three types of agents: (1) a representative household, (2) a continuum of monopolistically competitive firms, and (3) a monetary authority that exogenously sets the money supply. For the rest of the section, I denote $s^t = (s_0, \dots, s_t)$ as the history of events up to period t , and S^t the set of possible histories given the initial realization s_0 .

5.1.1 The Representative Household

The household's preferences are defined over leisure and a continuum of imperfectly substitutable goods indexed by z , denoting the producing firm. The household maximizes expected discounted utility given by:

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t [U (C(s^t), L(s^t))] \quad (6)$$

where \mathbb{E}_0 denotes the expectations operator conditional on information known at time $t = 0$, β is the intertemporal discount factor, $C(s^t)$ is the household's composite consumption good, and $L(s^t)$ denotes the household's labor supply.

The composite consumption good is given by a Dixit-Stiglitz index formed by differentiated goods:

$$C(s^t) = \left[\int_0^1 c(z, s^t)^{\frac{\theta-1}{\theta}} dz \right]^{\frac{\theta}{\theta-1}} \quad (7)$$

where $c(z, s^t)$ denotes the household's consumption of good z at state s^t . The parameter $\theta > 1$ denotes the elasticity of substitution between differentiated goods.

The household's income is composed by labor income and by firms' profits, which are owned by the household. Additionally, a complete set of Arrow securities are traded in this economy. Given these assumptions, the representative household budget constraint, at each

history s^t , can be written as:

$$P(s^t)C(s^t) + \mathbb{E}_t [D_{t,t+1}(s^{t+1})B(s^{t+1})] \leq B(s^t) + W(s^t)L(s^t) + \int_0^1 \Pi(z, s^t)dz \quad (8)$$

where $P(s^t)$ represents the aggregate price index, \mathbb{E}_t denotes the expectations operator conditional on information known at time t , $B(s^{t+1})$ is the state contingent payoffs of the portfolio of financial assets purchased by the household in period t and sold in period $t+1$, $D_{t,t+1}(s^{t+1})$ is the stochastic discount factor that prices these payoffs in period t , $W(s^t)$ is the wage rate in the economy, and $\Pi(z, s^t)$ denotes the profits of firm z in state s^t .

Each period, the household must decide how much of differentiated products z to consume. Given some aggregate expenditure level $C(s^t)$, this decision is solved by the household expenditure minimization problem. From the first order condition (FOC) of this problem, it can be solved for the household demand of each differentiated product z :

$$c(z, s^t) = \left[\frac{p(z, s^t)}{P(s^t)} \right]^{-\theta} C(s^t) \quad (9)$$

where $P(s^t)$, the lagrange multiplier of the minimization problem, is the price level given by:

$$P(s^t) = \left[\int_0^1 p(z, s^t)^{1-\theta} dz \right]^{\frac{1}{1-\theta}} \quad (10)$$

After solving for the household demand of each differentiated product z , the household maximization problem can be solved in terms of $C(s^t)$. Under regular assumptions, the FOCs of the maximization problem yield the leisure-consumption and Euler equations:

$$\frac{W(s^t)}{P(s^t)} = \frac{-U_L(s^t)}{U_C(s^t)} \quad (11)$$

$$D_{t,t+1}(s^{t+1}) = \beta \frac{U_C(s^{t+1})}{U_C(s^t)} \frac{P(s^t)}{P(s^{t+1})} \quad (12)$$

where $U_L(s^t)$ and $U_C(s^t)$ denote the marginal utility of labor and consumption, respectively.

Equation (11) describes the household labor supply, whereas equation (12) captures the relation between the discount factor and the sequence of consumption.

5.1.2 Firms

There is a continuum of firms of mass 1 which specialize in the production of each differentiated product z . Firms have a linear production function on labor and face a stochastic productivity shock:

$$y(z, s^t) = a(z, s^t) l(z, s^t) \quad (13)$$

where, for firm z at state s^t , $y(z, s^t)$ denotes the output, $l(z, s^t)$ the amount of labor used for production, and $a(z, s^t)$ is the idiosyncratic productivity shock. It is assumed that this shock, which is independent across firms, evolves according to an autoregressive (AR) process in logs:

$$\log a(z, s^t) = \rho \log a(z, s^{t-1}) + \epsilon(z, s^t) \quad (14)$$

where $\epsilon(z, s^t) \sim N(0, \sigma_\epsilon^2)$ are independent identically distributed.

Each firm z maximizes the value of its expected profits, discounted according to the price of the Arrow-Debreu contingent claims, subject to its production function and the behavior of aggregate variables:

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \left[\prod_{j=1}^t D_{j-1,j}(s^j) \right] \Pi(z, s^t) \quad (15)$$

where firms' profits in each state s^t are given by:

$$\Pi(z, s^t) = p(z, s^t) y(z, s^t) - W(s^t) l(z, s^t) - \chi(z, s^t) W(s^t) I(z, s^t) \quad (16)$$

In this last equation, $I(z, s^t)$ is an indicator function equal to one if the firm changes its price and zero otherwise. It is assumed that if the firm decides to adjust its price ($I(z, s^t) = 1$), it must hire $\chi(z, s^t)$ additional units of labor. This fixed cost of changing prices is what is called in the literature a “menu cost”. The key feature of the CalvoPlus model is that, with some fixed probability, firms receive an opportunity to change their prices at a relatively low

cost, whereas otherwise they face a high cost of changing prices. Specifically:

$$\chi(z, s^t) = \begin{cases} \chi_L, & \text{with Pr} = (1 - \alpha) \\ \chi_H, & \text{with Pr} = \alpha \end{cases} \quad \text{where } \chi_L < \chi_H, \alpha \in [0, 1] \quad (17)$$

Given the latter assumptions, this model has the appealing feature that it nests both [Calvo \(1983\)](#) and [Golosov and Lucas \(2007\)](#) models as particular cases. In the former one, it is assumed that firms are only allowed to change their price with some fixed probability, which can be obtained in this setup by assuming that with probability $(1 - \alpha)$ firms can change their price at no cost ($\chi_L=0$), while otherwise price changes are infinitely costly ($\chi_H=\infty$). The latter model is obtained by assuming that $\alpha = 1$ and, hence, there is a single cost of changing prices in the economy.

5.1.3 Monetary Authority

Money is introduced by assuming that aggregate nominal spending must be equal to the money stock in the economy:

$$\int_0^1 p(z, s^t) c(z, s^t) dz = P(s^t) C(s^t) = M(s^t)$$

where $M(s^t)$ is the money supply in state s^t .

It is assumed that the monetary authority exogenously sets a path for the money supply, which follows a random walk with drift in logs:

$$\log M(s^t) = \mu + \log M(s^{t-1}) + \eta(s^t) \quad (18)$$

where $\eta(s^t) \sim N(0, \sigma_\eta^2)$ are independent identically distributed random variables.

5.2 Equilibrium

5.2.1 Definition

An equilibrium in this economy is a set of contingent plans for the household $C(s^t)$, $L(s^t)$, and $B(s^{t+1})$; a set of contingent prices and allocations for firms $\{p(z, s^t)\}_z$, $\{y(z, s^t)\}_z$ and $\{l(z, s^t)\}_z$; and a set of contingent aggregate prices $W(s^t)$, $P(s^t)$ and $D_{t,t+1}(s^{t+1})$, such that:

- (i) Taking prices as given, the household's contingent allocations solve the household utility maximization problem.
- (ii) Given their productivity level and aggregate variables, contingent prices and allocations for each firm z solve its profit maximization problem.
- (iii) Clearing conditions for labor, goods, and money markets are satisfied.

5.2.2 Computing the Equilibrium

For the equilibrium computation, I assume preferences of the form: $U(C(s^t), L(s^t)) = \log C(s^t) - \omega L(s^t)$, where the parameter ω determines the disutility of labor. This specification, which is standard in the menu cost literature, ensures that the nominal wage, $W(s^t)$, is proportional to nominal spending, $P(s^t)C(s^t)$, and thus also proportional to the money supply, $M(s^t)$. The equilibrium is solved by value function iteration of firms' profit maximization problem expressed in real terms.

Firms maximize real profits, selecting the optimal price $p(z, s^t)$, subject to the productivity shock, the money supply shock, and the rest of the aggregate variables. Given the previous assumptions, the state space of the firms' maximization problem is infinite dimensional since the evolution of the endogenous aggregate variables depends on the joint distribution of all firms' prices and productivity levels. To make the problem tractable, following [Krusell and Smith \(1998\)](#), it is assumed that firms perceive the evolution of the price level as being a function of a small number of moments of this distribution. Specifically, firms perceive the following law of motion:

$$\frac{P(s^t)}{P(s^{t-1})} = \Gamma \left(\frac{M(s^t)}{P(s^{t-1})} \right) \quad (19)$$

Nakamura and Steinsson (2010) showed that, in this model, forecasting the aggregate price level using this single variable turns out to be highly accurate.

Given these assumptions, firm z 's maximization problem can be written recursively in the form of the following Bellman equation:

$$V \left(a(z, s^t), \chi(z, s^t), \frac{p(z, s^{t-1})}{P(s^t)}, \frac{M(s^t)}{P(s^t)} \right) = \max_{p(z, s^t)} \left\{ \Pi^R(z, s^t) + \mathbb{E}_t \left[D_{t,t+1}(s^{t+1}) V \left(a(z, s^{t+1}), \chi(z, s^{t+1}), \frac{p(z, s^t)}{P(s^{t+1})}, \frac{M(s^{t+1})}{P(s^{t+1})} \right) \right] \right\} \quad (20)$$

where $V(\cdot)$ is firm z 's value function. The state variables of the recursive problem are: (1) the realization of the idiosyncratic productivity shock $a(z, s^t)$, (2) the menu cost faced $\chi(z, s^t)$, (3) the price carried forward from the last period in current real terms $p(z, s^{t-1})/P(s^t)$, and (4) the aggregate real consumption, which is equal to the money supply divided by the price level $M(s^t)/P(s^t)$. $\Pi^R(z, s^t)$ denotes the firms' real profits, which are function of the contemporary state variables and the firm's decision to adjust its price.

The main challenge of this dynamic programming problem is to find the function Γ and the stationary distribution across states such that individual firms' price-setting, given by the policy function F of the firms' recursive problem, is consistent with the aggregate inflation implied by Γ . To solve for the equilibrium I follow the iterative procedure proposed by Nakamura and Steinsson (2010). The procedure iterates over Γ and F over a finite grid for the state variables until they are consistent with a stationary distribution. This procedure is described in detail in Appendix B. The transition probability matrices for the stochastic variables $a(z, s^t)$ and $M(s^t)$ are approximated using the method proposed by Tauchen (1986).

5.3 Calibration

Table VIII reports the benchmark parameters calibrated outside the model, which will remain constant across the models considered in the next section. Given that the price statistics from the micro data are in semimonthly periods, I calibrate the model in that frequency. For the intertemporal discount I consider an annual discount value of 0.96, this implies a semi-monthly discount factor of $\beta = 0.96^{1/24}$. With respect to the elasticity of demand θ , that is of interest as it determines firms' markups³⁰, there is no clear consensus of what an appropriate value for this parameter is. For example, Golosov and Lucas (2007) used a value of 7, while Nakamura and Steinsson (2010) a value of 4. For this paper, I follow the latter calibration as it implies a markup of 33.3%, a number in line with the empirical evidence for Mexico.³¹

Table VIII: MODEL BENCHMARK PARAMETERS

Discount Factor	$\beta = 0.96^{1/24}$
Elasticity of Demand	$\theta = 4$
Steady-State Labor Supply	$L_{ss} = 1/3$
Disutility of Labor	$\omega = 2.25$
Money Supply Growth Rate Mean	$\mu = 0.0013$
Money Supply Growth Rate Std. Deviation	$\sigma_{\eta} = 0.0045$

The value of ω is set such that the labor supply in the steady state, L_{ss} , is equal to 1/3. This assumption for the labor supply value is standard in the literature. On the other hand, the calibration of the money supply process follows Nakamura and Steinsson (2010). Since the model does not incorporate a secular trend in economic activity, μ is set equal to the mean growth rate of nominal GDP less the mean growth rate of real GDP (presented in Figure I), while σ_{η} is equal to the standard deviation (s.d.) of nominal GDP growth.

The key parameters that determine the price-setting dynamics generated by the CalvoPlus model are: the probability of facing a low menu cost ($1 - \alpha$), the value of the low and high

³⁰In the flexible price economy, i.e. zero menu cost, the firm's optimal price is set as a constant markup, given by $\frac{\theta}{\theta-1}$, over marginal costs.

³¹Castañeda-Sabido and Mulato (2006) estimated an average markup of 31 percent for different industrial activities in Mexico.

cost of adjusting prices (χ_L and χ_H), and the idiosyncratic productivity shock parameters (ρ and σ_ϵ). As will be described below, for the quantitative results of the paper, these parameters are jointly calibrated so that the model matches a selected group of price-setting moments of the data.

6 Monetary Non-Neutrality Quantitative Results

This section analyzes the implications of the paper’s empirical results on the degree of monetary non-neutrality in the main types of sticky price models, namely in the [Calvo \(1983\)](#) and the [Golosov and Lucas \(2007\)](#) models. Additionally, a calibration for the full CalvoPlus model, targeted such that it can match the share of small price changes of the data, is also considered. Given the evidence about the relevance of sales for the price-setting dynamics, quantitative results considering both posted and regular prices calibrations are presented. It will be shown that the inclusion of sales has large implications in the monetary non-neutrality results as posted and regular prices possess different degrees of price stickiness that are calibrated into the model.

Through the section, monetary non-neutrality is calculated as the variance of output when the model is simulated with purely nominal aggregate shocks.³² To calculate this variance, I generate 15,000 realizations of the exogenous variable η_t to construct a time series for the money supply and then simulate the economy considering the model’s equilibrium functions. I drop the first 1,500 observations and then calculate the variance of the resulting output time series. To interpret the results of this exercise, I relate the variance of output generated by nominal shocks in the model to the business cycle fluctuations observed in the HP-filtered Mexican real GDP. Finally, I compare my results to the previous evidence of [Nakamura and Steinsson \(2010\)](#) for the case of the U.S. economy.³³

³²Other measures of monetary non-neutrality, such as the cumulative impulse response to a one time monetary shock, yield similar results. These alternative results are available upon request.

³³Note that this type of cross-country comparison, analyzing relative numbers of non-neutrality given the business cycle fluctuations, controls for the different size of shocks calibrated for each economy.

Table IX: CALIBRATED PARAMETERS AND MODEL FIT

	Posted Prices				Regular Prices			
	Data	Calvo	Golosov-Lucas	CalvoPlus	Data	Calvo	Golosov-Lucas	CalvoPlus
A. Calibrated Parameters								
χ_L	-	0.000	-	0.000	-	0.000	-	0.000
χ_H	-	0.743	0.010	0.015	-	0.641	0.009	0.019
Relative χ_H	-	37.96%	0.51%	0.79%	-	26.04%	0.35%	0.77%
ρ	-	0.788	0.636	0.524	-	0.783	0.786	0.806
$\sigma\epsilon$	-	0.140	0.066	0.070	-	0.129	0.047	0.047
$(1 - \alpha)$	-	0.170	0.000	0.088	-	0.134	0.000	0.080
B. Targeted Moments (%)								
fr	17.0	17.1	17.0	17.0	13.4	13.4	13.4	13.4
fr^+/fr	57.6	52.1	55.2	55.3	60.4	53.3	56.2	55.4
$ dp $	12.4	10.5	12.4	9.1	10.1	8.6	10.1	7.3
$\#/fr$ at $ \Delta p < 5\%$	32.1	32.2	0.0	33.0	39.1	39.2	0.0	39.0
$\#/fr$ at χ_L	-	99.0	0.0	51.6	-	99.0	0.0	59.2
C. Additional Moments (%)								
dp	1.0	0.7	0.8	0.8	1.3	1.0	1.0	0.9
$ dp^+ $	11.6	10.8	11.9	8.9	9.4	9.0	9.8	7.5
$ dp^- $	13.5	10.2	12.9	9.3	11.1	8.2	10.4	7.1

NOTES: Regular Prices denotes prices excluding sales. χ_L (χ_H) is the low (high) menu cost. Relative χ_H refers to the high menu cost as a fraction of the flexible price steady state semimonthly revenues times the probability of adjusting: fr (χ_H/Y_{ss}). ρ and $\sigma\epsilon$ are the autoregressive coefficient and the standard deviation of the idiosyncratic productivity shock. $(1 - \alpha)$ is the probability at which firms can change their price at χ_L . fr denotes the frequency of price change. fr^+/fr denotes percentage of price changes that are price increases. $|dp|$ is the magnitude of price changes. $|\Delta p| < 5\%$ denotes the set of price changes, in absolute value, of less than 5%. $\#/fr$ is the percentage of observations in the set with respect to the total number of price changes. dp is the average size of price changes. $|dp^+|$ and $|dp^-|$ are the magnitude of price increases and decreases, respectively.

6.1 The Calvo Model

The [Calvo \(1983\)](#) model is one of the most widely used in applied monetary economics. In this model, nominal rigidities are introduced by assuming that firms are only allowed to adjust their price with some fixed probability, regardless of their incentives. For this reason, in this type of models, price changes are said to be time-dependent. As discussed in [Nakamura and Steinsson \(2010\)](#), this assumption makes the model to be highly tractable, however, it runs into severe complications in the presence of large idiosyncratic shocks or a modest amount of aggregate inflation. Intuitively, as a result of this extreme nominal rigidity, a firm could accumulate such a large number of idiosyncratic and/or aggregate shocks that its desire to adjust its price may become so large that it would be preferable to shut down rather than continue producing at its present price.

To avoid this type of cases, I consider an approximation to this model in which firms, besides being allowed to change their price for free at a fixed probability, can always adjust but at a considerably high cost. This assumption guarantees that almost all price changes are time-dependent, precisely 99% of them, while allowing the model to incorporate large idiosyncratic shocks and aggregate inflation, which are relevant ingredients so that the model could better approximate the price-setting moments observed in the data.

Panel A of [Table IX](#) presents the parameters used in this paper such that the CalvoPlus model approximates the time-dependent pricing of the [Calvo \(1983\)](#) model and the data moments of both posted and regular prices. The calibration strategy is as follows. First, I set the value and probability of the low menu cost equal to zero and to the observed frequency of price change. Second, I jointly calibrate the value of the high menu cost and the idiosyncratic shock parameters such that 99% of the model's prices adjust at the low menu cost ($\#/fr$ at χ_L), while matching the frequency of increases and the share of price changes less than 5% in the data. Besides the main parameters, the table reports the expected high menu cost, given by the probability of adjusting, as a fraction of semimonthly revenues (Relative χ_H). The high menu cost which induces that only 1% of adjustments will be at this value is around 38 and 26% of the firms' revenues for posted and regular prices, respectively. Additionally, it is

worth mentioning that, the idiosyncratic shock s.d. calibrated such that the model approximates the share of small price changes in the data is around 30 times the one calculated for the aggregate money supply.

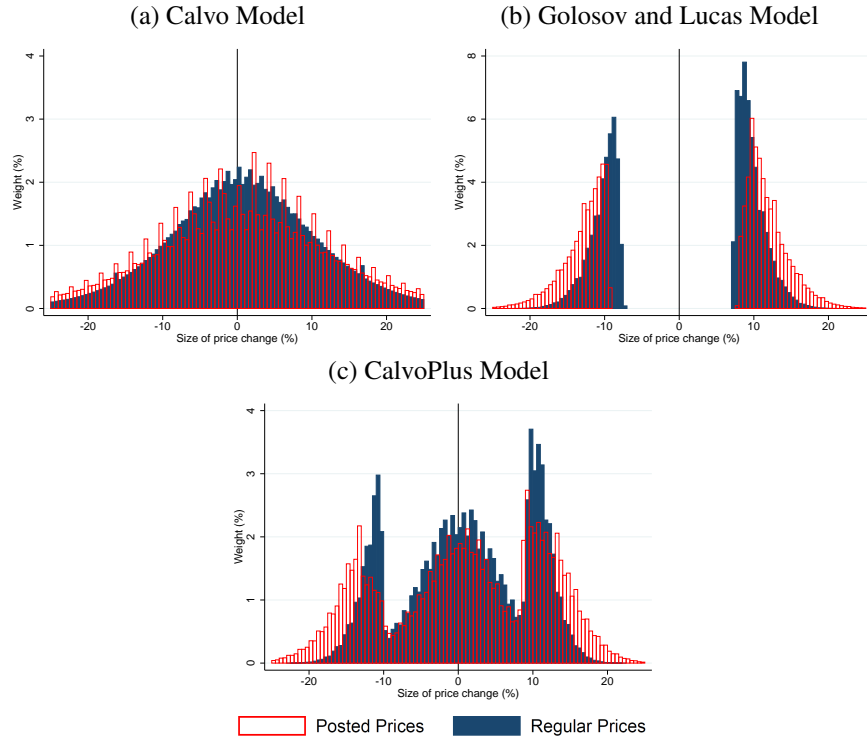
The model fit to the data is presented in Panel B. This panel shows that the model makes a relatively good adjustment to the observed price-setting. Nonetheless, despite large idiosyncratic shocks, the magnitude of price changes is around 2 p.p. below to the one in the data. Furthermore, this model does not possess an endogenous mechanism that generates that price decreases are larger than increases. For both posted and regular prices calibrations, contrary to what is observed in the data, the model's magnitude of increases is larger than the one of decreases. Additional to these moments, Panel (a) of [Figure VI](#) presents the complete distribution of price changes that is generated by the model. As discussed in [Section 4.3](#), this distribution inherits the shape of idiosyncratic shocks, and consequently, given the previous assumptions, exhibits a unimodal shape centered at zero. Given the larger size of price changes due to sales, the distribution of the model calibrated to posted prices exhibits a larger s.d. and fatter tails.

The results of the real effects of monetary shocks generated by the model are presented in [Table X](#). The table shows that monetary shocks in the [Calvo \(1983\)](#) model calibrated to posted prices account for close to 8% of the Mexican business cycle. Whereas, when this model is calibrated to the price statistics excluding sales, it generates fluctuations on output of the size of 14.7% of the observed business cycle, i.e. a result almost two times the one obtained with posted prices. These results highlight the importance of sales for the price-setting and the implications of the increased price flexibility on the monetary non-neutrality results of this model.

6.2 The Golosov and Lucas Model

On the other hand, the [Golosov and Lucas \(2007\)](#) model was the first to introduce idiosyncratic productivity shocks to an otherwise standard menu cost model. The introduction of this type of shocks makes that this model does a very good job matching the moments observed

Figure VI: DISTRIBUTION OF THE SIZE OF PRICE CHANGES BY PRICE-SETTING MODEL



NOTES: Histogram of the size of price changes in the stationary distribution generated by the Calvo (1983), Golosov and Lucas (2007), and Nakamura and Steinsson (2010) CalvoPlus model. The size of price changes is the log size times 100. Regular Prices denotes prices excluding sales. The figures plot the price changes distribution in the $|25\%|$ range only. Bin size is 0.5%.

in the data. Nevertheless, contradicting the results obtained with time-dependent models, this model predicts that nominal rigidities due to menu costs yield “small and transient” monetary non-neutrality results. In this section I analyze the results of this model calibrated to the Mexican economy.

Panel A of Table IX reports the model calibration for both posted and regular prices. This model is a particular case of the CalvoPlus model by assuming that the probability of a low menu cost is equal to zero and hence all price changes are made at a single cost. In this case, the high menu cost and the idiosyncratic shock parameters are jointly calibrated such that the model matches: the frequency of price change, the share of increases, and the magnitude of adjustments in the data. Similar to the previous results, the s.d. of the idiosyncratic shocks

Table X: MONETARY NON-NEUTRALITY AND THE BUSINESS CYCLE

	Posted Prices		Regular Prices	
	$\mathbb{V}(C)$	Fract. Total	$\mathbb{V}(C)$	Fract. Total
A. Non-Neutrality Results: Mexico				
Mexico GDP 1960-2015	4.374	100	4.374	100
Calvo	0.347	7.9	0.645	14.7
Golosov-Lucas	0.043	1.0	0.065	1.5
CalvoPlus W/ Small Price Changes	0.096	2.2	0.146	3.3
B. Nakamura and Steinsson (2010) Non-Neutrality Results: US				
U.S. GDP 1947-2005			2.720	100
Golosov-Lucas			0.055	2.0
CalvoPlus W/ Product Substitutions			0.173	6.4

SOURCE: OECD.Stats and [Nakamura and Steinsson \(2010\)](#).

NOTES: Regular Prices denotes prices excluding sales. For the data, $\mathbb{V}(C)$ is the variance of HP-filtered real GDP. For the model, $\mathbb{V}(C)$ is the variance of output when the model is simulated with purely nominal aggregate shocks. All the variance numbers reported are multiplied by 10^4 .

needed so that the model matches the magnitude of price changes are also large compared to the one of the monetary shock (around 15 and 10 times larger, for posted and regular prices). [Golosoov and Lucas \(2007\)](#) emphasized that this relatively large size of the idiosyncratic shock is crucial to generate a substantial number of price changes in this type of models. Despite the magnitude of these shocks, the calibrated menu costs are very reasonable in both calibrations. The table reports, for posted and regular prices, an expected menu cost of 0.51 and 0.35% of the firms' semimonthly revenue, respectively.³⁴

With respect to the model fit, unsurprisingly, the table's Panel B shows that the model does a very good job matching the above mentioned price statistics. This ability of the [Golosoov and Lucas \(2007\)](#) menu cost model to endogenously match these micro data moments is one of the main reasons that justify its use. However, as previously discussed, because of the menu cost assumption this model does not generates sufficient small price changes. In fact, for both calibrations there are no price changes smaller than 5%. This result is also shown in

³⁴These numbers are consistent with the menu costs' empirical evidence. Using store level data for supermarkets in the U.S. [Levy et al. \(1997\)](#) documented that menu cots are about 0.7% of firms' revenue.

the bimodal distribution of the size of price changes generated by the model as presented in Panel (b) of [Figure VI](#).

The monetary non-neutrality results of the menu cost model are presented in [Table X](#). For posted prices, the variance of output generated by nominal shocks represents only 1% of the Mexican business cycle, whereas for regular prices this number is 1.5%, i.e. 1.5 times higher. [Goloso and Lucas \(2007\)](#) pointed out that to obtain a good match between theory and data sales must either be removed or explicitly incorporated to the model. Although some recent papers have developed menu cost models that generate sale-type (or temporary) price changes ([Kryvtsov and Vincent, 2015](#); [Kehoe and Midrigan, 2015](#)), the standard practice in the literature is to calibrate models to regular prices. In this vein, my quantitative results calibrated to regular prices should be considered as the baseline of the paper. Nonetheless, the differences I found between posted and regular prices highlight the importance of filtering out sales to calculate moments from the data that are used to calibrate this type of models that aim to quantify the real effects of monetary shocks.

Overall, the menu cost model results indicate that the response of output to a monetary shock is limited. A similar conclusion is made by [Goloso and Lucas \(2007\)](#) for the U.S. economy using this model. Specifically, considering [Nakamura and Steinsson \(2010\)](#) results, as reported in [Table X](#), this model accounts for only 2% of that country business cycle, suggesting that the degree of monetary non-neutrality in these economies is very similar.³⁵ The low non-neutrality generated is explained by the selection effect present in menu costs models. This effect refers to the fact that, in these models, the firms that adjust their price as a response to the monetary shock are not selected at random, but rather are the firms whose prices are most out of line. Because of this effect, standard menu cost models generate small responses of output to monetary shocks.

³⁵[Goloso and Lucas \(2007\)](#) considered the cumulative impulse response of output to a one time monetary shock as a measure of monetary non-neutrality. They calculated that, given a shock of 1.25% in the money supply, the response of output in the first period is around 0.5%, and the shock dissipates in less than 3 months.

6.3 The CalvoPlus Model with Small Price Changes

Finally, I present the results for the full CalvoPlus model, which combines both time and state-dependent pricing policies. For this case, I consider a calibration that introduces small price changes into the otherwise [Goloso and Lucas \(2007\)](#) menu cost model by assuming that, with a fixed probability, firms are allowed to adjust their price at a zero menu cost. This assumption is a reduced form of the multiproduct firms model of [Midrigan \(2011\)](#), which has, as an appealing feature, economies of scope in price changes.³⁶ This mechanism to generate small price changes in menu cost models has also been used in [Vavra \(2014\)](#). In [Nakamura and Steinsson \(2010\)](#), the CalvoPlus model is calibrated so that the probability of receiving a zero menu cost is equal to the frequency of product substitution in the U.S. CPI. Their calibration aims to capture the idea that, when a new product is introduced, firms receive the opportunity to set its price at a relative lower menu cost.

The calibrated parameters for the CalvoPlus model with small price changes are reported in Panel A of [Table IX](#). For this model, I set the low menu cost equal to zero and jointly calibrate the remaining four parameters such that the model matches: the frequency of price change, the fraction of price increases, the magnitude of price changes, and the percentage of adjustments of less than 5%. The panel shows that, in both calibrations, the probability of receiving a free menu cost is smaller than the frequency of price change considered for the [Calvo \(1983\)](#) model (0.088 and 0.080, for posted and regular prices). Additionally, the calibrated high menu costs are moderate, although larger than in the [Goloso and Lucas \(2007\)](#) model.

With respect to the model fit, Panel B shows that the model does a good approximation to the targeted moments with exception of the magnitude of price changes. This result was expected given the introduction of small price changes into the model. On the other hand, the distribution of the size of price changes is presented in Panel (c) of [Figure VI](#). The resulting

³⁶In [Midrigan \(2011\)](#) multiproduct firms model, firms pay a single menu cost for changing the price of multiple products. This economies of scope assumption in prices adjustments allows the model to generate small price changes since once the menu cost is paid to change the price of a particular product, the price of other products will also be adjusted, possibly in small magnitudes.

histogram combines both time and state-dependent pricing distributions. The price changes of less than, say, 10% are adjustments of firms that received the opportunity of a free price change, whereas the larger ones exhibit the bimodal shape characteristic of menu cost models. Even though this distribution is still far from the empirical one presented in [Figure IV](#), this allows the CalvoPlus model to generate the share of small price changes observed in the data.

[Table X](#) shows that the variance of output generated by nominal shocks in this model accounts for 2.2 and 3.3% of the Mexican business cycle for posted and regular prices, respectively. These numbers are more than two times larger than the ones generated by the [Goloso and Lucas \(2007\)](#) model, but still fell short of the results obtained with the [Calvo \(1983\)](#) model. It should be noticed that the increase in monetary non-neutrality is a result of the time-dependent pricing that is being introduced by the CalvoPlus assumption, as it generates that a share of adjusting firms is chosen at random and consequently muting the selection effect present in menu cost models.

Panel B of [Table IX](#) shows that, in the calibrated CalvoPlus model, around 50 and 60% of price changes are made at the low menu cost, considering posted and regular prices. The previous numbers show that, even though the share of time-dependent pricing introduced is in between time and state-dependent models, the non-neutrality results are closer to the ones obtained with the [Goloso and Lucas \(2007\)](#) model. This result is consistent with [Nakamura and Steinsson \(2010\)](#), who showed that the real effects of monetary shocks generated by the [Calvo \(1983\)](#) model are quite sensitive to even a modest amount of state-dependent pricing. In that paper, the calibration to the frequency of product substitution yields that around 75% of price changes are made at the low menu cost. This higher proportion of time-dependent changes is the explanation of the larger non-neutrality found with the CalvoPlus model for the U.S. economy, as shown in Panel B of [Table X](#).

In sum, the quantitative results presented in this section show that for both time and state-dependent pricing models the statistics that are calibrated into the model have large implications for the non-neutrality results. Specifically, the real effects of monetary shocks obtained

with regular prices are between 1.5 and 2 times larger than to the ones generated with the posted prices calibrations. Additionally, these results largely depend on the type of nominal rigidity considered. For the case of the Mexican economy, the monetary non-neutrality obtained with the [Calvo \(1983\)](#) model is more than 8 times larger than the one obtained with the [Golosov and Lucas \(2007\)](#) model. Finally, the calibration of the CalvoPlus model, which combines both time and state-dependent pricing, suggests that monetary non-neutrality is limited close to the results obtained with a menu cost model.

7 Conclusions

The nature of price-setting has important implications for a range of issues in macroeconomics. Particularly, for models that incorporate some nominal rigidity the measured stickiness of prices is critical for the quantitative results any type of these models get. Therefore, the evidence obtained from the data on individual firms' price-setting is relevant to properly incorporate micro-founded nominal rigidities into price-setting models that aim to quantify monetary non-neutrality.

This paper contributes to the empirical macroeconomics literature by providing new evidence of the price-setting behavior in Mexico using product-level micro data underlying the Mexican CPI. I documented that 17.0% (24.7%) of posted prices adjust each half-month period (month), while for regular prices this number is 13.4% (20.6%). From these price changes, around 60% are increases. A new fact not previously documented for Mexico is that sales account for around 31.2% of price changes, which reflects the importance of sales for the aggregate level of price flexibility.

Furthermore, this paper documents that there is a large heterogeneity in the price-setting. I also found that the size of price changes are, on average, large and the distribution of non-zero price changes is bimodal centered at zero, with the right side mode having considerable more weight. Yet, there is a moderate amount of small price changes that are not explained by measurement error nor spurious small price changes. Finally, an analysis of the aggregate dynamics of the price statistics shows that the main source of variations in inflation is due

to the size of price changes (the intensive margin) rather than to the fraction of products that adjust (the extensive margin).

In light of this evidence, I analyze the implications and consistency of the empirical results for both time and state-dependent pricing models using the CalvoPlus model of [Nakamura and Steinsson \(2010\)](#). This model combines both time and state-dependent pricing by assuming that firms face, with some fixed probability, a low and a high menu cost. As a result, this model nests both [Calvo \(1983\)](#) and [Golosov and Lucas \(2007\)](#) models as special cases.

The quantitative exercises show that, for both types of models considered, the statistics that are calibrated into the model have large implications for monetary non-neutrality. Specifically, the real effects of monetary shocks obtained with regular prices are between 1.5 and 2 times larger than the ones obtained with the posted prices. These results highlight the importance of differentiating between the moments of posted and regular prices observed in the data that are calibrated into general equilibrium models. Additionally, the results largely depend on the type of nominal rigidity that is assumed. Considering the calibration to regular prices, the monetary non-neutrality obtained with the [Calvo \(1983\)](#) model account for 14.7% of the Mexican business cycle, while in the [Golosov and Lucas \(2007\)](#) model this number is only 1%. Finally, the CalvoPlus model results suggest that monetary non-neutrality is limited as monetary shocks account for around 3% of the observed business cycle.

The results presented in this paper suggest the importance of developing a richer model that can account for other features in data in order to analyze the real effects of monetary shocks in a broader framework. For example, one feature of the data, which have proved to be relevant in both time and state-dependent models ([Carvalho, 2006](#); [Nakamura and Steinsson, 2010](#)), is heterogeneity in the price-setting. Additionally, the relevance of sales in the Mexican price-setting suggests that a model with temporary price changes could be also of interest. Finally, a model that combines nominal rigidities with a small open economy framework could have a more realistic representation of the Mexican economy. The study of these extensions, and its applications, is left for future research.

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A Correcting for Spurious Small Price Changes

Measurement problems in the CPI micro data that generate spurious small price changes could bias the evidence for the prevalence of small price changes and the aggregate price-setting statistics. For the case of the Mexican CPI, the practice of registering prices in a unit size, e.g. price per kilogram, could generate a considerable number of spurious small price changes. From the 244 in-sample items 113 have prices reported in a unit size. Overall, these items represent 50.9% of the sample, measured by household expenditure weights.

To correct for this source of spurious small price changes I employ a variable that reports the factor used to convert the posted price to the per unit size price. For example, for the generic Carbonated drinks, whose price is reported in liters, a bottle of 600 milliliters has a conversion factor of $1/0.6$, i.e. 1.6667. With *cfact* I reconstruct the posted price of each product *s* as:

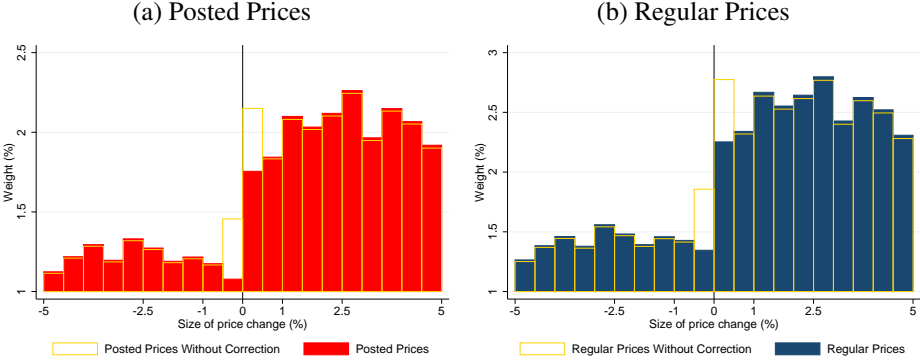
$$P_t^s = \frac{\lfloor 100 \text{ Rep}P_t^s / \text{cfact} \rfloor}{100}$$

where P denotes the corrected price, $\text{Rep}P$ the reported price, *cfact* the conversion factor, and $\lfloor \cdot \rfloor$ denotes the round function. The round function together with the operations for 100 are required to correct for the possible loss of precision in the data because of conversion factors like $1/0.6$. With this operation all the prices are rounded to 2 decimals (cents). As an additional correction, I eliminate all price changes of less than 1 cent since these changes are clearly due to measurement error or loss of precision in the data. Finally, I correct single price changes of 1 cent within an otherwise constant price spell, for example: 45, 45.01, 45.

Panels (a) and (b) of [Figure A.I](#) show the implications of these corrections for the size of price changes of posted and regular prices in the $[-5, 5]$ interval. The solid bars are the same distributions presented in [Figure IV](#), whereas the distributions in gold present prices without corrections. The difference between these histograms shows that, once correcting for these problems, a considerable amount of spurious small price changes is reduced. Nonetheless, for the aggregate statistics the consequence of these spurious small price changes is negligible: for both posted and regular prices the frequency increases in around 0.2 p.p., while the

size of price changes is reduced in 0.1 p.p.

Figure A.I: SPURIOUS SMALL PRICE CHANGES



SOURCE: Banco de México and INEGI.

NOTES: Weighted histogram of the size of non-zero price changes across time and product-level items. Regular Prices denotes prices excluding sales. The size of price changes is the log size times 100. The figures plot the price changes distribution in the $|5\%|$ range only. Bin size is 0.5%.

B Iterative Procedure to Solve for the Equilibrium

The equilibrium of the menu cost model is solved by value function iteration of the firm recursive problem following the iterative procedure proposed by Nakamura and Steinsson (2010). Specifically, to find the equilibrium functions and the stationary distribution that are consistent with an equilibrium I implement the following procedure:

- 1) Specify a finite grid of points for the state variables $(a, p_{-1}/P, M/P)$.
- 2) Propose a step function Γ on the M/P grid.
- 3) Given the proposed function Γ , solve for the firm's policy function F by value function iteration on the grid.
- 4) Check if Γ and F are consistent with an stationary distribution:
 - 4.1) Calculate the stationary distribution of the economy, call it Q , over $(a, p_{-1}/P, M/P)$ implied by the proposed Γ and F using the following algorithm:
 - 4.1.0) Start with an initial uniform distribution over the firm's states: $Q(a, p_{-1}/P, M/P)$.
 - 4.1.1) Map $Q(a, p_{-1}/P, M/P)$ into $Q(a, p/P, M/P)$ using the policy function F .
 - 4.1.2) Map $Q(a, p/P, M/P)$ into $Q(a_{+1}, p/P, M/P)$ using the transition probability for the productivity process.
 - 4.1.3) Map $Q(a_{+1}, p/P, M/P)$ into $Q(a_{+1}, p/P, M_{+1}/P)$ using the transition probability of the monetary supply.
 - 4.1.4) Map $Q(a_{+1}, p/P, M_{+1}/P)$ into $Q(a_{+1}, p/P_{+1}, M_{+1}/P_{+1})$ using the law of motion of the aggregate price level Γ .
 - 4.1.5) Check whether $\|Q(a_{+1}, p/P_{+1}, M_{+1}/P_{+1}) - Q(a, p_{-1}/P, M/P)\| < \xi$, where $\|\cdot\|$ denotes the sup-norm. If so, stop. If not, update the distribution Q and restart step 4.1.1.

- 4.2) Use the stationary distribution Q in states $(a_{+1}, p/P, M_{+1}/P)$, the law of motion Γ , and equation (10), to calculate the difference between the aggregate price level implied by the firms' policy function, call it P_{+1} , and the price level implied by Γ , call it $P(\Gamma)_{+1}$.
- 4.3) Check whether $\|P_{+1} - P(\Gamma)_{+1}\| < \xi$. If so, Γ and F are consistent.
- 5) If Γ and F are consistent these are equilibrium functions consistent with the stationary distribution Q . If not, update Γ and return to 3.