PRICE RIGIDITY: MICROECONOMIC EVIDENCE AND MACROECONOMIC
IMPLICATIONS

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ABSTRACT

We review recent evidence on price rigidity from the macroeconomics literature, and discuss how this evidence is used to inform macroeconomic modeling. Sluggish price adjustment is a leading explanation for large effects of demand shocks on output and, in particular, the effects of monetary policy on output. A recent influx of data on individual prices has greatly deepened macroeconomists' understanding of individual price dynamics. However, the analysis of these new data raise a host of new empirical issues that have not traditionally been confronted by parsimonious macroeconomic models of price-setting. Simple statistics such as the frequency of price change may be misleading guides to the flexibility of the aggregate price level in a setting where temporary sales, product-churning, cross-sectional heterogeneity, and large idiosyncratic price movements play an important role. We discuss empirical evidence on these and other important features of micro price adjustment and ask how they affect the sluggishness of aggregate price adjustment and the economy's response to demand shocks.

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1 Introduction

A large empirical literature in macroeconomics has produced a diverse array of evidence supporting the notion that demand shocks have large effects on real output. One strand of this literature has focused on the effects of monetary shocks, documenting evidence for substantial "monetary non-neutrality" (see, e.g., Friedman and Schwartz, 1963; Christiano, Eichenbaum and Evans, 1999; Romer and Romer, 2004). Another strand has focused on the effects of shocks to government spending, documenting that they raise overall output substantially (see, e.g., Blanchard and Perotti, 2002; Ramey, 2011; Nakamura and Steinsson, 2012). A major challenge in macroeconomics has been how to explain these empirical findings. A large class of macroeconomic models in which the economy responds efficiently to shocks, implies that temporary demand shocks should have small effects on output and that monetary shocks, in particular, should have no effect on output.

A leading hypothesis for the large effects of demand shocks on output has been that prices (and wages) adjust sluggishly to changes in aggregate conditions. Consider a monetary shock. The efficient response to a doubling of the money supply is for all prices to double immediately and all real quantities to remain unchanged. This response relies on prices being very flexible. In that case, real interest rates and real output are completely divorced from movements in nominal interest rates and the money supply. However, if price adjustment is sluggish, a reduction of nominal interest rates by the central bank may translate into a reduction in real interest rates in the short run and thus increase output. In other words, sluggish price adjustment provides an explanation for the conventional wisdom that expansionary monetary policy increases output.

Fiscal stimulus is another potential source of variation in demand. If prices are flexible, a temporary increase in government spending results in a sharp rise in real interest rates. This "crowds out" private spending and implies that output increases only modestly. If, however, prices respond sluggishly to the stimulus (and the monetary authority doesn't make up for this by moving the nominal interest rate), increases in the real interest rate will be limited. This implies that the fall in private spending will be small and the overall effect of the stimulus will be to increase output substantially.

The same logic implies that sluggish price responses will mute the response of real interest rates to other aggregate shocks such as financial panics, increased uncertainty, bad news about future productivity or fluctuations in consumer sentiment (Keynes' "animal spirits"). By muting movements in real interest rates (and real wages), price rigidities imply that these shocks can result in substantial

variation in "aggregate demand." In this way, price rigidities greatly expand the role that these shocks can play in driving economic fluctuations.[1]

Many people's first reaction to the idea that major fluctuations in output—such as the Great Depression or the recession of 2007-2009—could be substantially a consequence of stickiness in prices and wages is that this doesn't sound plausible. But many types of economic disturbances call for sharp movements in real interest rates. Since price rigidities mute these movements, they imply that output can deviate substantially from its efficient level. Consider for instance the type of deleveraging shocks analyzed by Eggertsson and Krugman (2012) and Guerrieri and Lorenzoni (2011). An efficient response of the economy to such shocks calls for a sharp drop in real interest rates. However, if prices respond sluggishly and the nominal interest rate is constrained by its lower bound of zero, the real interest rate will be "stuck" at a level that is too high. In fact, rather than prices jumping down and beginning to rise (which would reduce the real interest rate), prices may fall gradually. This implies that real interest rates may actually rise, further exacerbating the initial shock. The substantial resulting deviation of the real interest rate from its "natural" or efficient level can lead to large inefficient drops in output.

In most models that feature sluggish responses of the aggregate price level, sticky prices and wages are not the whole story. A second key ingredient is coordination failures among price setters that lead prices to respond incompletely even when they change. Coordination failure among price setters arise when price changes are staggered and strategic complements——i.e., firm A's optimal price is increasing in firm B's optimal price. In that case, the first prices to change after an aggregate shock will not respond fully to the shock because other firms have not yet responded. This will in turn lead later firms to respond incompletely. The combination of nominal rigidities and coordination failures among price setters can generate long-lasting sluggishness of the aggregate price level and, therefore, large and long-lasting effects of demand shocks on output.

If the ultimate goal is to assess how sluggishly the aggregate price level responds to aggregate shocks, why not simply study this directly? Indeed a large literature has done just that. However, an important challenge faced by this literature has been to convincingly identify exogenous demand shocks (e.g., monetary shocks). Evidence on price rigidity at the micro level both helps bolster the case for sluggish price adjustment and helps us understand the mechanisms that give rise to this

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[1]Price rigidity also mutes the response of real interest rates in response to supply shocks and therefore changes the dynamic response of the economy to these shocks as well.

phenomenon. The idea behind using evidence of price rigidity is that it is unlikely that optimal prices are literally unchanged for long periods and then change abruptly by large amounts.

Business cycle models that feature nominal rigidities and coordination failure among price setters are often referred to as "New Keynesian." The behavior of these models—and therefore the policy conclusions they yield—depends critically on the assumptions made about price adjustment. For this reason, the characteristics of price adjustment have long been an important topic in empirical macroeconomics. Following the seminar work of Bils and Klenow (2004) this area has been especially active over the past decade and a great deal has been learned. In this article, we review the empirical literature on this issue with a focus on illustrating how its various strands help inform us about the extent to which micro price rigidity translates into sluggishness in the response of the aggregate price level to shocks. At the risk of oversimplifying a bit, the guiding question for each piece of empirical evidence will be: What does this piece of evidence imply about how sluggishly the overall price level responds to changes in aggregate conditions?

The paper proceeds as follows. In section 2, we lay out some basic facts about the frequency of price change in the U.S. economy. In section 3, we present a simple monetary model that helps explain why macroeconomists have had such a persistent interest in price rigidity by illustrating the close connection between price rigidity and economy's response to monetary (and other demand) shocks. The remainder of the paper delves into various features of price adjustment that complicate the relationship between the degree of micro price rigidity and and the responsiveness of the aggregate price level to shocks. Sections 4 and 5 discuss temporary sales and cross-sectional heterogeneity, both of which are first order issues in defining what we mean by "the" frequency of price change. Sections 6 and 7 discuss evidence that firms adjust their prices more frequently when their incentives to do so increase—in particular, in periods of high inflation—and investigate the implications of "menu cost" models that can capture this empirical regularity for the macroeconomic consequences of price rigidity. Sections 8 and 9 discuss seasonality in price adjustment and the hazard function of price adjustment. Section 10 discusses evidence on the relationship between inflation and price dispersion, a crucial determinant of the welfare costs of inflation in leading monetary models. Section 11 discusses the important role that coordination failures play in monetary models, and evidence on the strength of these coordination failures that has been gleaned from micro-price data. Section 12 concludes.

2 Basic Facts About Price Rigidity in Consumer Prices

We start by seeking an answer to a basic question: How often do prices change? Until recently, the empirical evidence on this issue was rather limited. Even though consumer prices are public information—one can simply walk into a store to observe them—large-scale datasets on micro-price data are, in practice, difficult to obtain. The conventional wisdom among researchers working on New Keynesian business cycle models in the 1990's and early 2000's was that prices changed roughly once a year. A common citation for this fact was Blinder et al.'s (1998) survey study of firm managers.[2] Bils and Klenow (2004) shattered this conventional wisdom by documenting that the median monthly frequency of price change in the micro data underlying the non-shelter component of the Consumer Price Index (CPI) in the U.S. in 1995-1997 was 21%, implying a median duration of price rigidity of only 4.3 months.

Over the past decade, the literature on price rigidity has grown dramatically as new sources of comprehensive price data have become available to academic researchers. Among the most important are the datasets underlying the Consumer Price Index, Producer Price Index and Import and Export Price Indexes, collected by the U.S. Bureau of Labor Statistics (BLS). Klenow and Kryvtsov (2008) and Nakamura and Steinsson (2008a) have analyzed in detail the micro data underlying the U.S. CPI for the period of 1988-2005.[3] Table 1 presents results on the frequency of price change from these papers.

In principle, it is straightforward to calculate the frequency of price change—one simply counts the number of price changes per unit time. In practice, this calculation is complicated by the presence of temporary sales, stockouts, product substitutions, and cross-sectional heterogeneity in the BLS data. In Table 1, we report statistics both for "posted prices"—i.e., raw prices including sales—and for "regular prices"—i.e., prices excluding sales. Regular prices are identified using a "sales flag" in the BLS data. For regular prices, Table 1 reports statistics both including price changes at the time of product substitutions and excluding such price changes.

We report both the expenditure weighted median and mean frequency of price change, as well as the median and mean "implied duration." The implied duration for a particular sector is defined

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[2] Other important early studies include Lach and Tsiddon (1992), Carlton (1986), Cecchetti (1986), and Kashyap (1995).

[3] See also, Hosken and Reiffen (2004, 2007), who analyze the prevalence and characteristics of temporary sales in the BLS CPI data.

as $d = -1/\ln(1-f)$, where f is the frequency of price change in that sector.[4] The median implied duration is the implied duration for the sector with the median frequency of price change, while the mean implied duration is defined as the expenditure weighted mean of the implied durations in different sectors.[5] Nakamura and Steinsson (2008a) report results for several different ways of treating observations that are missing due to sales and stockouts. The issue is that the frequency of price change may be larger or smaller over the course of these events than at other times. The statistics reported in Table 1 from Nakamura and Steinsson (2008a) estimate the frequency of price change over the course of these events using the price before and after the missing period.[6]

The results in Table 1 illustrate two important issues that arise when assessing price rigidity. First, the extent of price rigidity is highly sensitive to the treatment of temporary price discounts or "sales." For posted prices, the median implied duration is roughly 1.5 quarters, while for regular prices, it is roughly 3 quarters depending on the sample period and the treatment of substitutions. But why is it interesting to consider the frequency of price change excluding sales? Isn't a price change just a price change? The sensitivity of summary measures of price rigidity to the treatment of sales implies that these are first order questions, and recent work has shed a great deal of light on them. This work has developed several arguments, based on the special empirical characteristics of sales price changes, for why macro models aiming to characterize how sluggishly the overall price level responds to aggregate shocks should be calibrated to a frequency of price change substantially lower than that for posted prices. We discuss this work in section 4.

A second important issue that is illustrated by the results reported in Table 1 is the distinction between the mean and the median frequency of price change. Take the results of Nakamura and Steinsson (2008a) on the frequency of regular price changes including substitutions for the sample period 1998-2005. The median monthly frequency of regular price change is 11.8%, while the mean monthly frequency of regular price change is 23.1%. Again, this difference is first order for the measurement of how sluggishly the overall price level responds to aggregate shocks in conventional

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[4]A constant hazard λ of price change implies a monthly probability of a price change equal to $f = 1 - e^{-\lambda}$. This implies $\lambda = -\ln(1-f)$ and $d = 1/\lambda = -1/\ln(1-f)$.

[5]Why has this literature focused on frequency measures (and "implied duration" measures constructed by inverting the frequency) as opposed to direct duration measures? The primary explanation is the large number of censored price spells in datasets such as the BLS data, arising from products dropping out of the dataset due to product turnover and BLS resampling. Dropping censored spells would lead to biased duration estimates.

[6]Other methods considered in Nakamura and Steinsson (2008a) include using only contiguous price observations and carrying forward the old regular price through sales and stockouts. These methods yield somewhat lower frequencies of price change.

monetary models.[7] This difference arises because the distribution of the frequency of regular price change across products is highly skewed and begs the question: which summary measure of price rigidity—e.g., mean or median frequency of price change—should we focus on when calibrating a simple macro model? Recent work has argued that calibrating to the mean frequency is inappropriate, while calibrating to the median frequency or the mean implied duration yields a better approximation to a full fledged multi-sector model. We discuss this work in section 5.

Certain product categories—in particular, durable goods—undergo frequent product turnover. For some such goods, an important portion of price adjustment likely occurs not through price changes for a particular item but rather at the time of product turnover. For example, the frequency of price change for womens' dresses not counting product turnover is only 2.4% per month, which might suggest a duration of prices of over 40 months. However, the frequency of product turnover for womens' dresses is 25.8% per month. It seems likely that most of the adjustment of prices for womens' dresses occurs at times when retailers discontinue older dresses and replace them with new ones. The same is true (to a lesser extent) for many other product categories.

Table 1 reports statistics on the frequency of regular price change both including price changes at the time of product substitutions and excluding such price changes. The median frequency of regular price change including product substitutions is roughly 1.5 percentage points higher than excluding product substitutions.[8] However, price changes due to product substitutions may differ substantially in terms of their implications for the adjustment of the aggregate price level to shocks because their timing may be motivated to a much larger extent than for other price changes by factors other than a firm's desire to change its price—factors such as product development cycles and seasonality in demand. We discuss this in more detail in section 8.

Comprehensive data on consumer prices in a number of countries other than the U.S. have in recent years become available to academic researchers. Alvarez (2008) and Klenow and Malin (2011) tabulate studies using these data and their conclusions regarding the frequency of price change. An

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[7]The difference is even larger when we look at the results reported in Klenow and Kryvtsov (2008). The results of Klenow and Kryvtsov (2008) and Nakamura and Steinsson (2008a) are quite similar for the median frequency and implied duration of regular price changes. Larger differences arise for the other statistics. For posted prices, the median frequency of price change in Nakamura and Steinsson (2008a) are close to those in Bils and Klenow (2004), while Klenow and Kryvtsov (2008) report higher frequencies of price change. Klenow and Kryvtsov (2008) note that these differences are due to different samples (all cities versus top three cities) and different weights (category weights versus item weights).

[8]The CPI research database provides an imperfect measure of product introduction by providing an indicator for whether a product undergoes a "forced substitution." A forced substitution occurs if the BLS is forced to stop sampling a product because it becomes permanently unavailable.

important part of this body of work was carried out within the context of the Inflation Persistence Network (IPN) of the European Central Bank. The conclusions of this work are summarized in Álvarez et al. (2005b) and Dhyne et al. (2006).

Scanner data have provided new insights into high frequency price dynamics for consumer packaged goods. These data are often collected directly from supermarkets as products are "scanned" at the checkout aisle. A broad-based dataset on supermarkets for the U.S. economy is the IRI Research Database. Another widely used database is the *Dominick's Finer Foods* database, available online from the Kilts Center for Marketing of the University of Chicago Booth School of Business.[9] This dataset includes consumer prices, as well as a measure of wholesale costs, for a leading Chicago supermarket chain. Similar data for another supermarket chain are analyzed by Eichenbaum, Jaimovich, and Rebelo (2011) and Gopinath et al. (2011). An important advantage of scanner data sets is that they often include information on quantity sold as well as prices. A disadvantage of these data is their exclusive focus on consumer packaged goods, and in some cases a single retail outlet.[10]

Recent studies also apply similar methods to broad based BLS datasets on U.S. producer prices both for domestic and internationally traded goods (Nakamura and Steinsson, 2008a; Goldberg and Hellerstein, 2011; Gopinath and Rigobon, 2008) as well as producer prices in other countries (see Alvarez (2008) and Klenow and Malin (2011) for citations). This literature is less extensive than the literature on consumer prices mainly because producer price data are less readily available to researchers. Yet the retail sector accounts for only a small fraction of value added. A major goal of the literature on consumer prices is to indirectly help us understand the behavior of manufacturer prices. A small number of papers have studied the relationship between consumer and producer prices (e.g., Nakamura and Zeron, 2010; Goldberg and Hellerstein, 2012; Eichenbaum et al., 2011; Anderson et al., 2012) . These papers tend to find rapid pass-through of changes in producer prices to consumer prices. A complication with interpreting data on producer prices is that producer contracts often exhibit substantial non-price features that may be varied over time (Carlton, 1979).

Similarly, wage rigidity and price rigidity are closely intermingled, since wages are a primary source of costs for many firms. Prices are particularly rigid in the service sector, a phenomenon that should perhaps be viewed as indirect evidence for wage rigidity. A number of recent studies

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[9]http://research.chicagobooth.edu/marketing/databases/dominicks/index.aspx

[10]Nakamura, Nakamura, and Nakamura (2011) show that pricing policies differ a great deal across supermarket chains.

have studied issues similar to those described above using broad-based data on wages (Dickens et al., 2008; Barattieri et al., 2012; Le Bihan et al., 2012; Sigurdsson and Sigurdardottir, 2012). A complication with interpreting data on both producer prices and wages is that they often derive from long-term relationships. This implies that observed producer prices and wages in a given month may be installment payments on a long-term contract rather than representing marginal costs or benefits for the buyer or seller at that point in time (Barro, 1977; Hall, 1980).

3   A Simple Model of Monetary Non-Neutrality

To understand why price rigidity plays such a central role in the macroeconomics literature, as well as the particular features of price adjustment that macroeconomists have focused on, it is useful to introduce a very simple model of price adjustment and derive its implications for the adjustment of the aggregate price level and the effects of monetary shocks on the economy. We make the simplest possible assumption about the timing and frequency of price adjustment: for each firm, an opportunity to change its price arrives at random with probability $(1 - \alpha)$ (Calvo, 1983). This implies that the probability that a firm changes its price in a given period is independent of the shocks hitting the economy or how long it has been since this firm last changed its price. In this case, the log aggregate price level p_t in the economy will (up to a first order approximation) be a weighted average its own past value and the log price p_{it}^* set by firms that change their price in period t:

$$p_t = (1 - \alpha)p_{it}^* + \alpha p_{t-1}. \tag{1}$$

While few macroeconomists would argue for this model as a literal description of how firms set prices, the goal of this "Calvo model" is to provide a tractable model of price adjustment to be incorporated into general equilibrium business cycle models. The key question for empirical analysis is whether the Calvo model—despite its simplicity—can nevertheless provide an adequate approximation, at an aggregate level, to a more complex pricing process.[11]

Suppose that firms produce using a linear production technology with constant productivity and labor being the only variable input. This implies that marginal costs are proportional to wages: $mc_t = w_t$, where mc_t is log nominal marginal costs, w_t is the log nominal wage and we have set an

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[11]For example, Woodford (2009) shows that the Calvo model can provide a good approximation to firms' pricing behavior when firms face information processing costs.

unimportant constant term to zero. Suppose that firms discount future profits at a rate β and face an iso-elastic demand curve $y_{it} - y_t = -\theta(p_{it} - p_t)$, where y_{it} denotes log demand for product i, y_t denotes log real aggregate output, and p_{it} denotes the log price of product i. It is simple to show that given this setup (up to a linear approximation) firms will set a price that is a discounted average of the marginal costs they expect to prevail over the period that their price remains fixed:[12]

$$p_{it}^* = (1 - \alpha\beta) \sum_{j=0}^{\infty} (\alpha\beta)^j E_t mc_{t+j}. \tag{2}$$

Let m_t denote log nominal output:

$$m_t = y_t + p_t. \tag{3}$$

Assume for simplicity that the monetary authority varies the money supply or the nominal interest rate in such a way as to make log nominal output follow a random walk with drift:

$$m_t = \mu + m_{t-1} + \epsilon_t. \tag{4}$$

This specification for monetary policy is equivalent to a simple rule for the money supply if the "velocity" of money is constant.[13]

Suppose that households' utility functions are $\log C_t - L_t$, where C_t denotes consumption and L_t labor. This implies that households' labor supply is vertical and given by $w_t - p_t = c_t$. Combining this equation with $mc_t = w_t$ from above and using $m_t = y_t + p_t$ and $c_t = y_t$ yields $mc_t = m_t$. Now consider the special case where average growth in nominal output is zero ($\mu = 0$ in equation (4)). Since marginal costs equal aggregate nominal output, which follows a random walk, $E_t mc_{t+j} = m_t$ for all j. Using this to simplify equation (2) yields $p_{it}^* = m_t$.

Notice that this last result implies that a given firm's optimal price is independent of the prices set by other firms in the economy. (The optimal price is proportional to aggregate nominal output, which is exogenous.) Thus, in this simple model the coordination failure discussed in the introduction—where one firm changes its price by less than it otherwise would because other firms haven't changed their price—doesn't arise. The pricing decisions of firms in this simple model are said to be strategically independent—neither strategic complements, nor strategic substitutes.

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[12]See, e.g., Woodford (2003, ch. 3) and Gali (2008, ch. 3).

[13]More generally, the central bank can achieve this target path for nominal output in a broad class of monetary models by appropriately varying the nominal interest rate.

Combining $p_{it}^* = m_t$ and equation (1), we have that the evolution of the aggregate price level is governed by

$$p_t = (1-\alpha)m_t + \alpha p_{t-1}. \tag{5}$$

The dynamics of output and inflation in this economy are then governed by equations (3)–(5).

Figure 1 presents the impulse response of nominal output, real output and the price level to a permanent unit shock to nominal output (starting from initial values of $y_{-1} = p_{-1} = 0$). Initially, most prices are stuck at their old level and the price level responds only partially to the change in nominal output. In the short run, thus, real output rises. Over time, more and more prices respond and real output falls back to its steady state level. It is easy to see from equations (3)–(5) that the response of real output is $y_t = \alpha^t$. In other words, the size of the boom in output at any point in time after the shock is simply equal to the fraction of firms that have not had an opportunity to change their prices since the shock occurred. All firms that have had such an opportunity have fully adjusted to the shock.

This illustrates that as the frequency of price change approaches one, the degree of monetary non-neutrality goes to zero, while monetary non-neutrality can be large and persistent if the amount of time between price changes is large. A simple measure of the amount of monetary non-neutrality in this model is the cumulative impulse response (CIR) of output—the sum of the response of output in all future periods (the area under the real output curve in Figure 1). In this simple model, the CIR of output is $1/(1-\alpha)$ and the CIR is proportional to the variance of real output. Another closely related measure is the half-life of the output response, $-\log 2/\log \alpha$. Using these measures, one can see that it will matter a great deal for the degree of monetary non-neutrality in this model whether the frequency of price change is 10% per month or 20% per month.

In this simple model, there is a clear link between the frequency of price change and the degree of sluggishness of the aggregate price level following a monetary shock. An analogous argument can be made for other demand shocks. The link between price rigidity and the aggregate economy's response to various shocks explains macroeconomists' persistent interest in the frequency of price adjustment. In the following sections, we will discuss how changing some of the critical assumptions of this simple model regarding the nature of price adjustment—e.g., allowing for temporary sales, cross-sectional heterogeneity, and endogenous timing of price changes—affects the speed of adjustment of the aggregate price level, and, in turn, the response of output to various economic disturbances.

4   Temporary Sales

Figure 2 plots a typical price series for a grocery product in the United States. This figure illustrates a central issue in thinking about price rigidity for consumer prices: Does this product have an essentially flexible price, or is its price highly rigid? On the one hand, the posted price for this product changes quite frequently. There are 117 changes in the posted price in 365 weeks. The posted price thus changes on average more than once a month. On the other hand, there are only 9 regular price changes over a roughly 7 year period. Which of these summary measures of price rigidity is more informative? Which should we use if we wish to calibrate the frequency of price change in the model in section 3?

One view is simply that "a price change is a price change," i.e., all price changes should be counted equally. However, Figure 2 also illustrates well that sales have very different empirical characteristics than regular price changes. While regular price changes are in most cases highly persistent, sales are highly transient.[14] In fact, in most cases, the posted price returns to its original value following a sale. Table 2 reports results from Nakamura and Steinsson (2008a) on the fraction of prices that return to the original regular price after one-period temporary sales in the four product categories of the BLS CPI data for which temporary sales are most prevalent. This fraction ranges from 60% to 86%.[15] Clearance sales are not included in these statistics because a new regular price is not observed after such sales. Nakamura and Steinsson (2008b) argue that clearance sales, like other types of sales, yield highly transient price changes.

This evidence strongly suggests that firms are not reoptimizing their prices based on all available new information when sales end. Furthermore, the empirical characteristics of sales price changes do not accord well with the simple model developed in section 3. This model and most other standard macroeconomics models do not yield "sale-like" price changes in which large price decreases are

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[14]Sales are identified either by direct measures such as a "sale flag" (as in the BLS data), or by using sale filters that identify certain price patters (such as "V-shaped" temporary discounts) as sales. While it is often said that looking at a price series it is easy to identify the regular price and the timing of sales, constructing a mechanical algorithm to do this is more challenging. Nakamura and Steinsson (2008a), Kehoe and Midrigan (2010), and Chahrour (2011) consider different complex sale filter algorithms that allow, for example, for a regular price change over the course of a sale and for the price to go to a new regular price after a sale. Such algorithms are used both by academics and by commercial data collectors such as IRI and AC Nielsen in identifying temporary sales.

[15]It is noticeable that the fraction of prices that return to the original price after a sale is negatively correlated with the frequency of regular price change across these categories. In fact, Table 2 shows that the probability that the price returns to its previous regular price can be explained with a frequency of regular price change over this period that is similar to the frequency of regular price change at other times (third data column). In addition, higher frequency data sets suggest that many sales are shorter than one month. This suggests that the estimates in Table 2 for the fraction of sales that return to the original price are downward biased.

quickly reversed.

To answer the question of how to treat sales in coming up with a summary measure of price rigidity, it is essential to understand how the distinct empirical characteristics of sales affect their macroeconomic implications. Several recent papers have attempted to develop more sophisticated models to capture the dynamics of prices displayed in Figure 2 and Table 2 and investigate their implications for the rate of adjustment of the aggregate price level, and the extent of monetary non-neutrality. These authors have also investigated the extent to which simpler models—like the one presented in section 3—generate approximately correct rates of price adjustment when they are calibrated to the frequency of price change including or excluding sales.

Kehoe and Midrigan (2010) build a menu cost model in which firms can either change their price permanently—i.e., change their regular price—or, at a lower cost, change their price temporarily— i.e., have a sale. They choose the parameters of their model to match moments such as the size and frequency of price changes, and the probability of return to the original price in the BLS CPI data. In their model, sales are simply temporary price changes, motivated by firms' desire to change their prices temporarily. The timing and magnitude of sales are fully responsive to the state of the economy and a large fraction of quantity sold is sold at sales prices. Nevertheless, sale price changes contribute little to the response of aggregate prices to monetary shocks. In their model, thus, the degree of monetary non-neutrality is close to the same as if sales price changes were completely absent. The key intuition is that because sale price changes are so transitory, they have a much smaller long-run impact on the aggregate price level "per price change" than do regular price changes.

Guimaraes and Sheedy (2011) introduce the idea that firms use sales to price discriminate between low and high price elasticity consumers into a macroeconomic business cycle model.[16] In their model—just as in the model of Kehoe and Midrigan (2010)—price flexibility associated with sales has a minimal effect on the degree of sluggishness of the aggregate price level in response to demand shocks (including monetary shocks). In Guimaraes and Sheedy (2011), this result arises not only because of the transitory nature of sales, but also because retailers have an incentive to avoid holding sales simultaneously—i.e., they have an incentive to stagger the timing of sales. This implies that

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[16]Sobel (1984) originally introduced the idea that sales might be due to price discrimination between customers with different price elasticities. Other important papers on sales in the industrial organizations (IO) literature include Varian (1980), Salop and Stiglitz (1982), Lazear (1986), Aguirregabiria (1999), Hendel and Nevo (2006), and Chevalier and Kashyap (2011). Hosken and Reiffen (2004) use BLS CPI data to evaluate the empirical implications of IO models of sales.

low sale prices "average out" across stores, reducing their effect on the aggregate price level.

An important point is that, in both Kehoe and Midrigan (2010) and Guimaraes and Sheedy (2011), even though a disproportionate fraction of goods are sold on sale, the regular price nevertheless continues to be the dominant factor in determining the trajectory of the aggregate price level and thus the response of aggregate output to demand shocks. These studies underscore the general lesson that a "price change is not just a price change" in determining how rapidly the aggregate price level reacts to macroeconomic shocks—the same frequency of price change may correspond to very different levels of responsiveness of the aggregate price level.

Eichenbaum, Jaimovich, and Rebelo (2011) analyze a scanner dataset from a large U.S. retailer and argue that it is useful to think of retail prices in terms of a "reference price"—defined as the modal price in a given quarter—and deviations from this reference price. They show that reference prices are quite sticky even though posted prices change on average more than once a month.[17] They develop a simple pricing rule that matches the behavior of prices in their data well. They then show that an economy in which prices are set according to this pricing rule generates a degree of monetary non-neutrality that is similar to that of a menu cost model calibrated so that the frequency of price change matches the frequency of reference price changes in their data, while a menu cost model calibrated to match the frequency of overall price changes yields much less monetary non-neutrality.

Another argument for why it may be appropriate to view sales as contributing less to the response of the aggregate price level to changes in macroeconomic conditions than an equal number of regular price changes is that firms' decisions to have sales may be "orthogonal to macro conditions" to a greater extent than their decisions regarding regular price changes. Anderson et al. (2012) analyze a unique dataset from a large U.S. retailer that explicitly identifies sales and regular prices. They show that regular prices react strongly to wholesale price movements and wholesale prices respond strongly to underlying costs, but the frequency and depth of sales is largely unresponsive to these shocks.[18] Coibion, Gorodnichenko, and Hong (2012) show that the frequency and size of sales falls when unemployment rates rise (i.e., changes in the behavior of sales raise rather than reduce prices in a recession). In contrast, Klenow and Willis (2007) show that in the BLS CPI data, the size of sales

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[17]Stevens (2011) reaches similar conclusions based on identifying infrequent breaks in pricing regimes in scanner data. She adapts the Kolmogorov-Smirnov test to identify changes in the distribution of prices over time. She finds that the typical pricing regime lasts 31 weeks and contains four distinct prices.

[18]The idea that sales may not respond to changes in macroeconomic conditions is suggestive of information costs, sticky information or rational inattention (Mankiw and Reis, 2002; Burstein, 2006; Woodford, 2009; Sims, 2011).

price changes is related to recent inflation in much the same way as the size of regular price changes. Klenow and Malin (2011) present evidence that sales do not fully wash out with cross-sectional aggregation in the BLS CPI data, but do substantially cancel out with quarterly time aggregation. More research is needed to assess the extent to which sales respond to macro conditions.

To sum up, three main reasons have been emphasized as potential reasons why temporary sales need to be treated separately in analyzing the responsiveness of the aggregate price level to various shocks. First, sales are highly transitory, limiting their effect on the long-run aggregate price level, even if their timing is fully responsive to macroeconomic shocks. Second, retailers may have an incentive to stagger the timing of sales, reducing their impact on the aggregate price level. Finally, sales may be on "autopilot," i.e., unresponsive to macroeconomic shocks.

5   Heterogeneity in the Frequency of Price Change

There is a huge amount of heterogeneity in the frequency of price change across sectors of the U.S. economy. Figure 3 illustrates this by plotting a histogram of the frequency of regular price change across different product categories in the CPI from Nakamura and Steinsson (2008a). While many service sectors have a frequency of price change below 5% per month, prices in some sectors—such as gasoline—change several times a month. A key feature of this distribution is that it is strongly right-skewed. It has a large mass at frequencies between 5-15% per month but then it has a long right tail with some products having a frequency of price change above 50% and a few close to 100%. As a consequence, the expenditure-weighted median frequency of regular price change across industries is about half the mean frequency of regular price change (see Table 1).

The simple model in section 3 assumes a common frequency of price adjustment for all firms in the economy. The huge amount of heterogeneity and skewness in the frequency of price change across products begs the question: How does this heterogeneity affect the speed at which the aggregate price level responds to shocks? In other words, will the price level respond more sluggishly to shocks in an economy where half the products adjust their prices all the time (say gasoline) and half hardly ever adjust (say haircuts) or one where all prices adjust half of the time? Or put slightly differently, if one wishes to approximate the behavior of the U.S. economy using a model with homogeneous firms, should one calibrate the frequency of price change to the mean or median frequency of price change in the data? As we discuss in section 2, the difference between these different summary measures is

14

first order.

Several authors have sought to answer these questions using detailed multi-sector models of the economy designed to incorporate cross-sectional heterogeneity in the frequency of price change. Bils and Klenow (2002) consider a multi-sector business cycle model with Taylor (1980) pricing and up to 30 sectors. The frequency of price change and expenditure weight for each sector is calibrated to match the empirical distribution of price rigidity. They find that a single-sector model with a frequency of price change roughly equal to the median frequency of price change in the data most closely matches the degree of monetary non-neutrality generated by their multi-sector model. This analysis motivates the focus on the median frequency of price change in Bils and Klenow (2004).

Carvalho (2006) focuses on the Calvo (1983) model of price setting—but also considers the Taylor (1980) model as well as several sticky information models. He incorporates strategic complementarity among price setters into his model and considers a broad class of processes for nominal aggregate demand. He shows that under this wide range of assumptions, the multi-sector version of his model that incorporates heterogeneity in the frequency of price change, generates much more monetary non-neutrality than the single-sector version in which the degree of price rigidity is calibrated to match the average frequency of price change.

In particular, Carvalho (2006) shows that in a multi-sector version of the model presented in section 3 (in the limiting case of no discounting), a single-sector model calibrated to the average duration of price spells matches the monetary non-neutrality generated by an underlying true multi-sector model. Table 1 reports that for consumer prices excluding sales in the U.S., the average duration is 9-12 months—much longer than the duration implied by the average frequency of price change and slightly longer than the duration implied by the median frequency of price change.[19]

Why does heterogeneity amplify the degree of monetary non-neutrality relative to a single-sector model calibrated to the average frequency of price change? The core intuition for this can be illustrated in a two sector version of the model presented in section 3 where the two sectors differ only in their degree of price rigidity. Suppose this economy is hit by a positive shock to nominal aggregate demand—i.e., a shock that raises all firms' optimal prices. Output jumps up when the shock occurs and then begins to fall back to steady state as firms adjust their prices (see Figure 1). Recall that at any point in time after the shock, the amount of monetary non-neutrality (the

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[19]The average duration, mean$(1/f)$, is larger than the duration implied by the average frequency, $1/\text{mean}(f)$, by Jensen's inequality since $1/x$ is a convex function.

increase in output) resulting from the shock in this model is equal to the fraction of firms that have not had a chance to change their price since the shock occurred.

If some firms have vastly higher frequencies of price change than others, they will change their prices several times before the other firms change their prices once. But all price changes after the first one for a particular firm are "wasted" in that they don't contribute to the adjustment of the aggregate price level to the shock since the firm has already adjusted to the shock.[20] If it were possible to re-allocate some price changes from the high frequency of price change sector to the low frequency of price change sector this would speed adjustment of the aggregate price level since a higher proportion of firms in the low frequency of price change sector have not adjusted to the shock.

A question that arises regarding the results of Bils and Klenow (2002) and Carvalho (2006) is whether they carry over to a setting where firms choose the timing of their price changes optimally subject to a menu cost. Nakamura and Steinsson (2010) address this question. We show that in an economy with low inflation and large price changes—like the U.S.—the timing of price changes is dominated by idiosyncratic shocks. This implies that the frequency of price change doesn't respond much to aggregate shocks and the effects of heterogeneity in price rigidity on the degree of monetary non-neutrality are similar to what they are in the Calvo and Taylor models. However, when inflation is high and price changes are relatively small, the menu cost model yields quite different results. In this case, aggregate shocks affect the frequency of price change more and this reduces the degree to which heterogeneity amplifies the amount of monetary non-neutrality. When we calibrate our model to the U.S. economy, we find that our multi-sector model generates a degree of monetary non-neutrality that is similar to that of a single sector model with a frequency of price change equal to the median frequency of price change in the data.

The huge amount of heterogeneity in the frequency of price change across sectors has testable implications regarding movements in relative prices and relative inflation rates across sectors. Other things equal, sectors in which prices change frequently should see a more rapid response of inflation relative to sectors with more sticky prices following an expansionary demand shock. Bouakez, Cardia, and Ruge-Murcia (2009a,b) use sectoral data to estimate a multi-sector DSGE model of the U.S. economy. Their estimates of the frequency of price change are highly correlated with Nakamura

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[20]Recall that the model in section 3 had no strategic complementarity. Firms therefore respond fully to the shock the first time they change their price after the shock occurs. (This argument also goes through in a model with strategic complementarity, but it involves slightly more steps.)

and Steinsson's (2008) sectoral estimates of the frequency of price change excluding sales. Other papers have found less support for this basic prediction of New Keynesian models. Bils, Klenow, and Kryvtsov (2003) find that the relative price of flexible price sectors falls after an expansionary monetary policy shock identified using structural VAR methods.[21] Mackowiak, Moench, and Wiederholt (2009) consider the response of sectoral prices to sectoral shocks. They find that there is little variation in the speed of the response of sectoral prices to such shocks between sectors with flexible prices and sectors with more sticky prices.

6   Inflation and the Frequency of Price Change

The simple model presented in section 3 makes the strong assumption that the timing of price adjustment is completely random and the frequency of price change is constant over time. Thus, firms do not respond to changes in economic conditions by changing the timing and frequency of price changes even though the incentives to change prices may have increased. The theoretical literature on price rigidity has emphasized that models that instead allow firms to choose the timing of price changes optimally, can yield vastly different conclusions about the speed of adjustment of the aggregate price level and the amount of monetary non-neutrality resulting from a given amount of micro price rigidity. We will discuss these models in more detail in section 7. These theoretical results motivate empirical work on the degree to which firms are more likely to change their prices when they have a stronger incentive to do so.

One way of studying this issue empirically is to investigate the extent to which the frequency of price change rises in periods of high inflation, since an increase in inflation raises the incentive firms have to change prices. Gagnon (2009) studies this question using data on price adjustment in Mexico over the period 1994 to 2002. Mexico experienced a serious currency crisis in December of 1994. Year-on-year inflation in Mexico rose from below 10% in the fall of 1994 to about 40% in the spring of 1995. Inflation then fell gradually to below 10% in 1999. Gagnon finds that at low inflation rates (below 10-15%) the frequency of price change comoves weakly with the inflation rate because movements in the frequency of price increases and movements in the frequency of price decreases offset each other. At higher inflation rates there are few price decreases and the frequency of price

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[21]A related result in Bils and Klenow (2004) is that sectors with a low frequency of price change do not have smaller innovations to inflation nor do they have more persistent inflation processes than sectors with a high frequency of price change as simple sticky-price models suggest they should.

change rises rapidly with inflation. Gagnon compares his empirical results with simulations of a menu cost model along the lines of Golosov and Lucas (2007). He finds that his menu cost model matches the variation in the frequency of price change very closely. His results provide strong support for the idea that firms respond to incentives in terms of the timing and frequency of their price changes.[22]

Alverez et al. (2011) study the same issues in an even more volatile setting: Argentina over the period 1988-1997. Argentina experienced a hyperinflation in 1989 and 1990, with inflation peaking at almost 5000%. A successful stabilization plan ended the hyperinflaton in 1991 and inflation fell quickly so that after 1992 there was virtual price stability. Alvarez et al. present theoretical results for a menu cost model showing that in the neighborhood of zero inflation, the frequency of price change should be approximately unresponsive to inflation, while at inflation rates that are high relative to the size of idiosyncratic shocks the elasticity of the frequency of price change with respect to inflation should be approximately 2/3. They then show empirically that both of these results hold in their Argentinian data. At low inflation rates (less than 10%) the frequency of price change is approximately uncorrelated with inflation, but at high inflation rates the elasticity of the frequency of price change with inflation is very close to 2/3 over a very large range. Again, the menu cost model seems to fit data on the frequency of price change remarkably well.[23]

The extent to which the frequency of price change varies with inflation has also been studied using the U.S. CPI microdata for the period 1988 to 2005 (Klenow and Kryvtsov, 2008; Nakamura and Steinsson, 2008a). However, the U.S. sample has the disadvantage of a low and stable inflation rate, which is not ideal for making inference about the relationship between the frequency of price change and inflation. Nevertheless, an interesting feature of the results of both Klenow and Kryvtsov's and our paper is that the frequency of price increases varies more with inflation than the frequency of price decreases. This asymmetry is more pronounced in our results— we find that the frequency of price decreases is largely unresponsive to the inflation.[24] We show that this feature of the data

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[22]Gagnon, Lopez-Salido, and Vincent (2012) argue that a very large proportion of the response of inflation to large shocks such as the collapse of the peso in Mexico in late 1994 and VAT increases in Mexico in April 1995 and January 2010 results from the "extensive margin" of price adjustment—i.e., the frequency rather than the size of price changes. Karadi and Reiff (2012) make a similar point for large VAT changes in Hungary. They argue that a model with leptokurtic idiosyncratic shocks along the lines of Midrigan (2011) is better able to capture this fact than the Golosov-Lucas model.

[23]Other papers that study related questions include Konieczny and Skrzypacz (2005) for Poland (1990-1996), Barros, Bonomo, Carvalho, and Matos (2009) for Brazil (1996-2008), and Wulfsberg (2010) for Norway (1975-2004).

[24]We focus on the median frequency of price change, while Klenow and Kryvtsov focus on the mean frequency of price change. We show that the difference between the time variation in the median and mean arises from the strong upward trends in the frequency of price change in gasoline and airplane tickets, which account for a small fraction of the economy but greatly influence the mean due to their high frequencies of price change. The relationship between

arises naturally in a menu cost model with idiosyncratic shocks and positive trend inflation. In such models, prices adjust only when they breach the lower or upper bound of an "inaction region." Positive inflation implies that the distribution of relative prices is asymmetric within the inaction region with many more prices bunched towards the lower bound than the upper bound. The bunching toward the lower bound implies that the frequency of price increases covaries more than the frequency of price decreases with shocks to inflation.

7 The Selection Effect

The theoretical literature on the macroeconomic effects of price rigidity has shown that price changes that are timed optimally by firms in response to macroeconomic shocks tend to lead to more rapid adjustment of the aggregate price level than price changes that are timed randomly. Golosov and Lucas (2007) refer to this as the "selection effect."

To illustrate the selection effect, it is helpful to start with a highly stylized model due to Caplin and Spulber (1987). Time is continuous. Nominal output follows a continuous time version of equation (4)—i.e., a Brownian motion with drift. The distribution of the shock to this Brownian motion is bounded below in such a way that nominal aggregate demand always increases. Firms face a fixed cost of changing their price but choose the timing of price changes optimally. These assumptions imply that firms will adopt an Ss policy (Sheshinski and Weiss, 1977, 1983), i.e., they will wait to change their price until their relative price has fallen to a trigger level s and at that point raise their relative price to a target level S. Finally, assume that the initial distribution of relative prices is uniform between s and S.

In this setting, nominal shocks have no effect on real output no matter how infrequently individual prices change. To see this, consider a short interval of time over which nominal output increases by Δm. Firms with relative prices smaller than $s + \Delta m$ at the beginning of this interval hit the lower trigger and raise their prices by $S - s$. These firms represent $\Delta m/(S - s)$ fraction of all firms. Their price changes will thus yield an increase in the price level equal to Δm, which implies $\Delta y = \Delta m - \Delta p = 0$. Intuitively, the firms that change their prices are not selected at random as in the Calvo model, but rather selected to be the firms that most need to change their price—i.e.,

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the frequency of price change and inflation documented for the median good holds individually in all sectors with a substantial degree of price rigidity.

those firms that have the largest "pent up" desire to change their price. This implies that these firms change their price by much more than if they were selected randomly and the price level therefore responds much more strongly to the nominal shock.[25]

The difference in conclusions regarding monetary non-neutrality between the Calvo model and the Caplin-Spulber model is striking. However, both models are extreme cases. The Calvo model is the extreme case where aggregate shocks have no effect on which firms and how many firms change their prices. The Caplin-Spulber model is the opposite extreme case where aggregate shocks are the only determinant of which firms and how many firms change their prices. More recent work has explored intermediate models and used empirical evidence on the characteristics of micro price adjustment to calibrate these models with the goal of assessing where on the spectrum between the Calvo model and the Caplin-Spulber model the real world is.

A key feature of the data that is fundamentally inconsistent with the simple models we have discussed so far is the large size of price changes. Klenow and Kryvtsov (2008) show that the average absolute size of price changes for U.S. consumer prices is very large—roughly 10%. A related fact is that about 40% of regular price changes are price decreases. In the simple models we have discussed, firms are reacting only to aggregate inflation when they change prices. Since inflation is almost always positive, these models imply that almost all price changes should be price increases. Furthermore, as Golosov and Lucas (2007) pointed out, with inflation of about 2.5% per year and firms changing prices every 4-8 months, the size of price changes should on average be much smaller than 10%.

Golosov and Lucas (2007) interpret these empirical findings as providing evidence for large, highly transitory, idiosyncratic shocks to firms' costs. They consider a model with fixed costs of price adjustment and a combination of aggregate shocks and idiosyncratic shocks calibrated to match the large size of price changes.[26] Their main conclusion is that their realistically calibrated menu cost model still yields a very strong selection effect. The selection effect reduces the degree of monetary non-neutrality by a factor of six relative to the Calvo model. They conclude that realistically modeled price rigidity yields monetary non-neutrality that is "small and transient."

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[25]For related work see Caballero and Engel (1991, 1993), Caplin and Leahy (1991, 1997), and Danziger (1999). See Head et al. (2012) for a completely different—search based—argument for why micro price rigidity may be completely divorced from sluggishness of the aggregate price level.

[26]The fixed cost of changing prices in their model amounts to 0.5% of revenue. This lines up well with empirical estimates of the costs of changing prices. Levy et al. (1997) estimate the costs of changing prices for a U.S. super market chain to be 0.7% of revenue. Nakamura and Zerom (2010) estimate the costs of changing prices for coffee manufacturers to be 0.23% of revenue.

Midrigan (2011) argues, however, that this result is quite sensitive to the distribution of idiosyncratic shocks. To take an extreme case, if idiosyncratic shocks take values of either zero or a very large number, then the model can be parameterized such that firms adjust their prices only when they are hit by an idiosyncratic shock. This largely eliminates the selection effect and yields a model that is quite similar to the Calvo model.

Midrigan (2011) presents several pieces of empirical evidence from the Dominick's scanner data that point towards a weaker selection effect than Golosov and Lucas' model implies. First, Midrigan shows that the distribution of the size of price changes in the Dominick's data is quite dispersed, whereas the Golosov-Lucas model implies a distribution of the size of price changes that is very concentrated around the upper and lower bounds of the inaction region. Second, while the average size of price changes is large, there are many small price changes; whereas the Golosov-Lucas model implies that firms "selected" to make a price change have a strong incentive to do so, and therefore change their prices by substantial amounts.[27] Finally, Midrigan (2011) argues there is substantial coordination in the timing of price changes within categories.

Midrigan (2011) makes two changes to the Golosov-Lucas model so as to be able to match the facts about the distribution of the size of price changes he documents. First, he assumes a leptokurtic distribution of shocks (shocks with larger kurtosis than the normal distribution). Second, he assumes "returns to scale" in price adjustment—if a firm chooses to change the price of one of its products, it can change the price of another product for free.[28] He then shows that the selection effect is small in his model and the degree of monetary non-neutrality generated by his model is only slightly smaller than in the Calvo model.

It is important to note that interpreting the empirical evidence on the size distribution of price changes is complicated by the potential role of heterogeneity across products. Given the short time-series of prices available for a given product in both the BLS data and scanner datasets, it is necessary to pool multiple products to obtain an estimate of the distribution of the size of price changes across

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[27]See Kashyap (1995) and Lach and Tsiddon (1996) for earlier evidence on small price changes. Klenow and Kryvtsov (2008) find a large number of small price changes in the U.S. CPI data. More recently, Eichenbaum, Jaimovich, Rebelo, and Smith (2012) have argued that measurement problems may be behind many of the small price changes observed in the type of data used by Midrigan (2011). Also, Anderson, Jaimovich, and Simester (2012) find very few small regular price changes in data from a large U.S. retailer.

[28]More recently, Bhattarai and Schoenle (2012) present related evidence from U.S. PPI data. They document that firms for which a larger number of products are sampled in the PPI database change prices more often and by smaller amounts, which they explain by returns to scale in price adjustment. Alvarez and Lippi (2012) explore the implications of such economies of scope in the price setting in more detail.

products. If the parameters of the repricing rule differ across products, this could potentially generate the wide distribution of absolute sizes observed in the data—even if the true distribution for each product was as in the Golosov-Lucas model.

Berger and Vavra (2011) use the BLS CPI data to show that the distribution of the size of price changes— the distribution of $\log(P_{it}/P_{it-1})$ across firms i at time t excluding non-adjusters—becomes more dispersed in recessions. They argue that standard models of price adjustment are inconsistent with this fact. Vavra (2012) augments a standard menu cost model to include "uncertainty shocks"— variation in the variance of idiosyncratic shocks—and finds that this model can match the counter-cyclicality of price change dispersion. Vavra's model implies that monetary non-neutrality falls in recessions because prices become more flexible.

Taken literally, the models described above assume that the large amount of variability in retail prices arises from large, unobserved idiosyncratic shocks to firms' productivity. This does not, how-ever, match up well with standard estimates of the variability of plant-level productivity, materials costs and wages.[29] An alternative view is that idiosyncratic shocks are a stand-in for other, unmod-eled sources of variation in firms' desired prices. Burstein and Hellwig (2007) emphasize that large price changes could also arise from idiosyncratic shocks to consumer demand. Nakamura (2008) shows that for grocery products, only a small fraction of price variation is common across products and outlets suggesting that most price changes are not responding to cost or demand shocks but rather are due to price discrimination or dynamic pricing strategies.

A prominent feature of price adjustment that is clearly at odds with all the menu cost models discussed in this section is frequent temporary sales that revert back to exactly the prior regular price. So, do temporary sales imply that we should discard the menu cost model? Not necessarily. The menu cost model matches a variety of empirical facts about *regular* price changes. The menu cost model is therefore potentially a good model for thinking about regular price changes. Both institutional evidence regarding price setting at large retailers and statistical evidence based on data from such retailers suggests that the process for regular price changes and temporary discounts may be largely disconnected from each other (Anderson et al., 2012; Eichenbaum et al., 2011). It may therefore be that regular price changes are well described by a menu cost model with the underlying desired price governed by traditional cost and demand factors, while the timing, depth and frequency

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[29]For example, evidence from the U.S. Annual Survey of Manufacturers suggests that the annual volatility of TFP is in the neighborhood of 10%.

of sales is determined by intertemporal price discrimination, advertising, and shifts in fashion and tastes.

8   Seasonality in the Frequency of Price Change

In section 6, we presented evidence that firms react to increases in the incentive to change prices by increasing the frequency of price change. This evidence suggests that the timing of at least some fraction of price changes is chosen purposefully. However, there is also considerable evidence suggesting that the timing of some price changes follows a regular schedule. Nakamura and Steinsson (2008a) document considerable seasonality in price setting in the United States. For consumer prices, they find that the median frequency of regular price change is 11.1% in the first quarter and then falls monotonically to 8.4% in the fourth quarter. They also find that the frequency of price change across months spikes in the first month of each quarter and that price increases play a disproportionate role in generating these patterns. For producer prices, the degree of seasonality is similar but quite a bit larger. The frequency of price change in the first quarter is 15.9% and falls to only 8.2% in the fourth quarter. For producer prices, most of the seasonality is due to a spike in the frequency of price change in January. Alvarez et al. (2005b) find similar patterns for consumer prices in the Euro Area.

One interpretation of this seasonality is that it provides evidence that the timing of some price changes follows a regular schedule as in the model of Taylor (1980). However, this is not the only possible interpretation. An alternative interpretation is that there may be seasonality in cost changes. For example, it may be that wages are more likely to change in January than in other months of the year. If this is the correct explanation, it, of course, begs the question: Why are wages more likely to change in January? This, in turn, may be due to regular schedules playing an important role in the timing of wage changes. A possible consequence of this pronounced seasonality in price adjustment is that monetary non-neutrality might be larger for shocks that occur early in the year than for shocks that occur later in the year. Indeed, Olivei and Tenreyro (2007) show that the response of output to monetary shocks identified using a structural VAR are larger for shocks that occur in Q1 and Q2 than for shocks that occur in Q3 and Q4.

Nakamura and Steinsson (2010) document very pronounced seasonality in product turnover for both apparel and transportation goods. They argue that this suggests that the timing of product

turnover is likely to be motivated primarily by factors such as development cycles and changes in consumer tastes—for example, the fall and spring clothing seasons in apparel—that are largely orthogonal to a firm's desire to change its price. While the introduction of the new spring clothing line may be a good opportunity for a firm to adjust its price, this type of new product introduction does not occur *because* of the firm's desire to adjust its price. That is, while price changes are likely to occur when new products are introduced, new products are not introduced because the old products were mispriced. If the timing of product substitutions are less "selected," it may be appropriate to model product substitutions not as optimally timed price changes such as those that arise in a pure menu cost model but rather as price changes without any selection effect such as those that arise in the Calvo or Taylor models.

9 The Hazard Function of Price Adjustment

The hazard function of price change can, in principle, be a powerful way of distinguishing between alternative models of price adjustment. The hazard that a price spell will end after a certain number of periods is defined as the probability that the price spell will end after that number of periods given that it has survived to that point. Intuitively, the hazard function answers the question: are prices that have recently changed more likely than others to change again? Or is it the case that prices become more likely to change the longer they have remained unchanged? If prices become more likely to change the longer they have remained unchanged, the hazard function is upward sloping. If, on the other hand, price adjustments tend to be clumped together (as in the case of temporary sales) then the hazard function is downward sloping.

In a simple setting where inflation is the only motive for price adjustment, a firms' incentive to adjust its price will grow over time as its price drifts away from its optimal level. If firms in such an environment can optimally choose the timing of price changes subject to a fixed cost, the hazard function of price change will be upward sloping. In contrast, in the Calvo model described in section 3, the probability of price adjustment is constant and the hazard function is flat regardless of a firm's incentives to adjust its price. This sharp dichotomy breaks down, however, when idiosyncratic shocks are added to the model. In this case, menu cost models can give rise to a multitude of different shapes for the hazard function of price change, since transient idiosyncratic shocks tend to flatten

the hazard function.[30] In addition, the inaction region is narrower in good times since more is at stake in getting the price right. This also tends to flatten the hazard function (Klenow and Kryvtsov, 2008).

A major empirical issue in the analysis of hazard functions of price adjustment is how to account for cross-sectional heterogeneity in the frequency of price change. It is well known in the literature on duration models that estimating hazard functions based on pooled data from many heterogeneous products can generate downward sloping hazard functions even when the true hazard function for any individual product is flat or upward sloping (e.g., Kiefer, 1988). The logic is that conditioning on a price having survived for a longer period of time will naturally tend to restrict the sample to products with a lower frequency of price change, leading to a downward sloping hazard function.

Klenow and Kryvtsov (2008) and Nakamura and Steinsson (2008a) study the hazard function of price change for U.S. CPI data. Both papers seek to control for cross-sectional heterogeneity in the frequency of price change. Klenow and Kryvtsov do this by allowing for fixed effects in the hazard function for goods in different deciles of the distribution of regular price changes. In our paper, we estimate separate hazard functions for different major product groups and, in addition, allow for good specific random effects in the hazard.[31]

For price changes including sales, we estimate sharply downward sloping hazard functions. This reflects the "bunching" of price changes associated with temporary sales. For regular prices changes, we estimate hazard functions that are somewhat downward sloping for the first few months and then largely flat except for a spike at 12 months for services prices. Klenow and Kryvtsov estimate flat hazard functions for regular price changes except for a spike at 12 months. We also estimate hazard functions for producer prices. These have a similar shape to the hazard function of regular price change for consumer prices, except that the spike in the hazard function at 12 months is larger and occurs in all product categories. The spike at 12 months in these estimated hazard functions supports the notion that some price changes occur on a fixed schedule as in Taylor (1980). However, failure of the empirical methods to fully purge the effects of heterogeneity is a potential concern regarding evidence of downward-sloping hazard functions. Also, the construction of the "regular

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[30]See Nakamura and Steinsson (2008a) for a discussion of this issue. The reason why idiosyncratic shocks flatten the hazard function is that they give rise to temporary price changes that are quickly reversed. Such price changes occur when the idiosyncratic shock is large enough that it is worthwhile for the firm to change its price temporarily to an "abnormal" level even though it realizes that it will soon have to change it back.

[31]One argument for using a random effects model is that hazard function models with fixed effects suffer from the incidental parameters problem.

price" variable in these studies is imperfect, implying that evidence of a downward-sloping hazard function could arise partly from missed "sales."

The menu cost model has stronger implications for the hazard of price changes as a function of the deviation of the current price from its optimal level than for the hazard of price changes as a function of time. In particular, a central implication of menu cost models is that prices should be more likely to change the further they are from their optimal level. Several recent papers have attempted to test this prediction using various observable proxies for a firm's optimal price. Using scanner data from a U.S. supermarket chain, Eichenbaum, Jaimovich, and Rebelo (2011) show that the prices (in particular the reference prices discussed in section 4) are more likely to change when the product's markup over marginal cost strays from its average level, and that markups tend to revert back to their average level when reference prices change. Campbell and Eden (2010) find that the probability of a price change is elevated when a store's price substantially differs from the average of other stores' prices for the same product. Gagnon, Lopez-Salido, and Vincent (2012) carry out a similar exercise with the SymphonyIRI Marketing dataset. They find similar results even when they filter out one-month sales or consider reference prices.

10   Price Dispersion and the Welfare Costs of Inflation

When prices are sticky and price setting is staggered across firms, inflation will cause dispersion in the relative price of otherwise identical goods. Consumers will respond to this price dispersion by increasing relative demand for the products with low relative prices at any given point in time. If production costs are increasing in output at the product level, price dispersion for identical products will, thus, inefficiently cause products with relatively low prices to be demanded to the point where their marginal cost of production is higher than that of relatively high priced goods. Most recent work on optimal monetary policy has modeled the costs of inflation as arising through this price dispersion channel (Rotemberg and Woodford, 1997; Woodford, 2003).

Burstein and Hellwig (2008) compare the welfare costs of inflation arising from price dispersion in the Calvo model and a menu cost model. They show that these welfare costs are an order of magnitude larger in the Calvo model than in the menu cost model. The reason is that in the Calvo model, prices can become arbitrarily out of line without this triggering a price change, while in the menu cost model, a large enough deviation of a price from its optimal value will trigger a price

change. Price dispersion for identical products thus increases much more sharply with inflation in the Calvo model than in menu cost models.[32]

This discussion suggests that documenting the empirical relationship between inflation and the degree of price dispersion for identical products is interesting both because it sheds light on the welfare costs of inflation and also because it may help distinguish between competing models of price adjustment. Traditionally, the relationship between price dispersion and inflation has also been of interest to industrial organization economists because it sheds light on the extent of product market imperfections.

In practice, researchers are not able to compare the prices of products that are completely identical. Instead, price dispersion is usually analyzed either across products within a narrow category or across categories. The identifying assumption being made is then that the differences in optimal prices across products within a category or between categories is uncorrelated with the rate of inflation.

Much of the literature on this topic has focused not on price dispersion but rather on dispersion of inflation rates for individual prices within a category or dispersion in inflation rates across categories. This concept has often been referred to as relative price variation. We will, however, refer to this as inflation dispersion. A number of papers starting with Mills (1927) have shown that inflation dispersion across product categories increases when aggregate inflation increases (see also Glejser, 1965; Vining and Elwertowski, 1976; Parks, 1978; Fischer, 1981; Debelle and Lamont, 1997). Van Hoomissen (1988) and Lach and Tsiddon (1992) consider the relationship between category level inflation rates and inflation dispersion at the product level within the category. Both of these papers use Israeli data. Van Hoomissen (1988) considers 13 narrow product categories over the period 1971-1984, while Lach and Tsiddon (1992) consider 26 narrow product categories over the period 1978-84. Inflation in Israel was high and variable over this period. Both papers find that intramarket inflation dispersion increases significantly with the rate of category level inflation. Van Hoomissen (1988) estimates a negative quadratic term suggesting that price dispersion rises less steeply at high rates of inflation.

Alverez et al. (2011) present theoretical results for a menu cost model showing that price

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[32]Kiley (2000) and Levin, Onatski, Williams, and Williams (2005) compare the welfare costs of inflation in the Calvo model with those in the Taylor model and find that the welfare costs in the Calvo model are much larger for similar reasons.

dispersion—defined as the standard deviation of log prices across products within a narrow product category—should be approximately unresponsive to inflation in the neighborhood of zero inflation, while at inflation rates that are high relative to the size of idiosyncratic shocks the elasticity of price dispersion with respect to inflation should be approximately 1/3. This contrasts sharply with the Calvo model in section 3, in which the distribution of relative prices scales one-for-one with the rate of inflation. Using micro-price data from Argentina over the period 1988-1997—a period that includes the hyperinflation of 1989-90 and subsequent stabilization—Alvarez et al. show that the elasticity of price dispersion in the data lines up well with the predictions of the menu cost model they analyze over a huge range of inflation rates. Price dispersion is unresponsive to inflation at low inflation rates. For high inflation rates, the elasticity of price dispersion with inflation is close to 1/3 for homogeneous goods, but smaller for heterogeneous goods. They conclude that welfare costs of price dispersion emphasized in the monetary economics literature are likely to be empirically relevant only for high inflation rates.[33]

11 Coordination Failure in Price Setting

A particularly simple feature of the model presented in section 3 is that firms respond fully to an aggregate shock the first time they have an opportunity to change their price after the shock. This feature implies that the response of the aggregate price level to a shock is complete once all prices have changed at least once, and, in particular, the effects of monetary shocks on real variables are limited to the time until prices have changed once. A prominent feature of many richer monetary models is that nominal rigidities are combined with other assumptions that lead to coordination failure among price setters. In such models, the firms that have an opportunity to change their price in a particular period adjust only partially to the shock because their optimal prices depend on the prices of other firms in the economy, which have not yet changed. Such coordination failure can be a powerful amplification mechanism for the aggregate effects of nominal rigidities. They have the implication that the sluggishness of the aggregate price level lasts well beyond the point when all prices have changed at least once. Coordination failure among price setters can, thus, imply that the macro sluggishness of prices is much greater than the micro rigidity of prices.

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[33]Reinsdorf (1994) studies the relationship between inflation and price dispersion for individual goods in the U.S. over the period 1980-1982 using micro-data underlying the U.S. CPI. In contrast to other studies, he finds a negative relationship between inflation and price dispersion.

Coordination failure among price setters has two essential ingredients. First, the timing of pricing decisions needs to be staggered. The empirical literature on price adjustment suggests that there is a large amount of staggering of price changes across price setters in the economy (see, e.g., Lach and Tsiddon, 1996). In the Taylor (1980) and Calvo (1983) models, price changes are assumed to be staggered. In Golosov and Lucas (2007) and the subsequent literature that uses menu cost models, staggering arises because the timing of price changes is largely determined by idiosyncratic shocks.

The second ingredient needed for coordination failure in price setting to amplify the sluggishness of the response of prices to shocks is strategic complementarity among price setters. If the pricing decisions of price setters are strategic complements—i.e., firm A's optimal price is lower the lower is firm B's price—and price setting is staggered, firms that have an opportunity to change their price soon after an aggregate shock will respond incompletely to the shock because other firms have not yet changed their prices.[34] The incomplete response of these early responders will in turn cause firms that change their prices later to respond by less, and so on. In this way, the combination of staggered price adjustment and strategic complementarity in price setting can lead to coordination failure where each cohort of price changers holds back because of the inaction of those firms that are not changing their price that period.

One source of strategic complementarity is intermediate inputs (Basu, 1995). Consider a modified version of the model in section 3, where the production function of firms is $y_{it} = AL_{it}^{a}Q_{it}^{1-a}$, where Q_{it} denotes an aggregate of material inputs used by the firm. Suppose for simplicity that Q_{it} is a CES aggregator of all the individual goods produced in the economy and that the value consumers place on these products for consumption purposes can be expressed by the same CES aggregator. In this case, the logarithm of the marginal cost of firm i is

$$mc_{it} = am_t + (1 - a)p_t,$$

where we have set an unimportant constant to zero. Firms' marginal costs depend on nominal aggregate demand as in the model in section 3 because the nominal wage is proportional to nominal aggregate demand. Now, however, firm' marginal cost also depend on the aggregate price level, since this is the cost of the materials firms use in production. Because the prices of firms' inputs are the prices of other goods in the economy, firms' marginal costs depend directly on the prices of the other

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[34]Strategic complementarity in price setting is closely related to Ball and Romer's (1990) notion of "real rigidities." It is also closely related to Taylor's (1980) notion of the "contract multiplier." See Cooper and John (1988) for a general discussion of strategic complementarity.

goods in the economy. This implies that firm pricing decisions are strategic complements. If price setting is staggered, firms that have an opportunity to change their price soon after an increase in nominal aggregate demand will moderate their price response because the prices of many of their inputs have not yet changed.

Several other sources of strategic complementarity are often also incorporated into monetary models. Prominent examples include demand structures in which the elasticity of demand is an increasing function of a firm's relative price (Kimball, 1995) and heterogeneous factor markets (Woodford, 2003, 2005; Altig et al., 2011). These features help structural monetary models match the long lasting effects of monetary shocks estimated in structural VARs without having to resort to assuming counter-factually low frequencies of price change.

The importance of coordination failures as an amplification mechanism for the effects of demand shocks in monetary models raises the issue of how to empirically discipline the degree of coordination failure in price setting. Measuring the degree of coordination failure in price setting is more difficult than measuring the degree of rigidity of micro prices since it entails measuring how much prices respond to a given increase in costs rather than just whether they respond at all. Researchers using micro-data on price adjustment have made some progress in this regard.

Gopinath, Itskhoki, and Rigobon (2010) study the response of prices to changes in exchange rates at the micro level using BLS data on U.S. import and export prices. They condition on times when prices change and find that firms adjust prices by only 0.25% for each 1% change in the cumulative exchange rate since the last price change. Fitzgerald and Haller (2012) and Burstein and Jaimovich (2009) perform similar exercises using official Irish micro data and American scanner data, respectively, and find even lower estimates of the initial price response to an exchange rate change. Gopinath, Itskhoki, and Rigobon (2010) document, however, that prices respond much more in the long-run to changes in the exchange rate. A related fact documented by Gopinath and Itskhoki (2010) is that the prices of imported goods continue to respond to changes in the exchange rate from before the product's previous price change. Furthermore, they find that pass-through of exchange rate changes into prices is higher for movements in the trade-weighted exchange rate—to which a firm's competitors can also be expected to respond—than to movements in bilateral exchange rates. These findings are consistent with models featuring strong strategic complementarities.[35]

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[35]An important challenge in interpreting this evidence is how to account for the endogeneity of exchange rate movements.

However, evidence from micro-data has also been used to level important critiques against models with staggered price changes and strong strategic complementarities in pricing. Bils, Klenow, and Malin (2012) emphasize that strategic complementarity in price setting reduces volatility and increases persistence of inflation. Rather than adjusting fully to an aggregate shock the first time it has an opportunity to change its price after the shock, a firm will respond with a sequence of smaller price changes in the same direction. Bils, Klenow, and Malin (2012) use BLS micro data on consumer prices to assess this prediction. They analyze "reset price inflation," which is meant to measure inflation in firms' desired prices.[36] They find that reset price inflation is quite volatile and has a serial correlation close to zero. They argue that this is hard to reconcile with the presence of large amounts of strategic complementarity.

Klenow and Willis (2006) develop a second critique of monetary models with large amounts of strategic complementarity. They show that a model with a substantial amount of demand-side strategic complementarity of the type emphasized by Kimball (1995) yields price changes that are much smaller than in the data unless idiosyncratic shocks are assumed to be massive and menu costs implausibly large. The demand system Klenow and Willis use is one in which firms are loathe to choose prices that deviate too far from those of their competitors. This implies that each firm chooses to change its price by only small amount when it changes its prices since many of its competitors are not adjusting. Klenow and Willis (2006) point out that this implication is inconsistent with the empirical evidence on the large size of price changes. Golosov and Lucas (2007) and Burstein and Hellwig (2007) show that this same issue arises in models in which strategic complementarity arises because of diminishing returns to scale in production.

What is the scope of this challenge? Whether this critique applies to a particular form of strategic complementarity turns out to depend on whether the strategic complementarity is of the "macro" or "micro" variety. This can be illustrated simply for the class of models for which the firm's profit function can be written as $\Pi(p_{it}/P_t, M_t/P_t, A_t)$, where p_{it} denotes firm i's price, P_t denotes the aggregate price level, M_t denotes nominal aggregate demand, and A_t is a vector of exogenous shocks.[37] In this case, firm i's desired price when prices are flexible is given by $\Pi_1(p_{it}/P_t, M_t/P_t, A_t) = 0$, where

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[36]For firms that change their price in period t, the reset price is simply their current price. For firms that do not change their price, Bils et al. index the reset price to the rate of reset price inflation of those firms that change their price. In a Calvo model without strategic complementarity, reset price inflation will equal the change in nominal aggregate demand, while actual inflation will lag behind as in Figure 1 due to price rigidity. This makes reset price inflation particularly useful for assessing the degree of strategic complementarity in price setting.

[37]See Nakamura and Steinsson (2010), Ball and Romer (1990) and Kimball (1995).

the subscript denotes a partial derivative. Consider for simplicity a steady state in which $p_{it}/P_t = 1$ and normalize M_t such that $M_t/P_t = 1$. In this case,[38]

$$\frac{\partial p_{it}}{\partial P_t} = 1 + \frac{\Pi_{12}}{\Pi_{11}}.$$

Recall that, by definition, p_{it} and P_t are strategic complements if $\partial p_{it}/\partial P_t > 0$. Pricing decisions are, therefore, strategic complements if $-\Pi_{12}/\Pi_{11} < 1$. Most macroeconomic models imply that $\Pi_{11} < 0$ and $\Pi_{12} > 0$. The degree of strategic complementarity therefore rises as either $-\Pi_{11}$ rises or Π_{12} falls.

Raising $-\Pi_{11}$ raises the extent of "micro" strategic complementarity. It lowers a firm's incentive to continue raising its price as its price rises relative to those of its competitors. Mechanisms that affect this channel include the curvature of demand faced by the firm as well as fixed factors or factors with firm-specific adjustment costs (e.g., firm specific capital). Lowering $-\Pi_{12}$ raises the extent of "macro" strategic complementarity. It reduces a firm's incentive to increase its prices in response to an increase in aggregate demand. Mechanisms that give rise to macro strategic complementarity include real wage rigidity and sticky intermediate inputs.

The Klenow-Willis critique points out that there is a link between strategic complementarities and the size of price changes. In economies such as the U.S., inflation is small relative to observed price changes, so the size of price changes is determined primarily by the magnitude of idiosyncratic factors. The response of a firm's price to an idiosyncratic shock is given by, $\partial p_t/\partial A_t = -\Pi_{13}/\Pi_{11}$, so increasing $-\Pi_{11}$ will both raise the degree of strategic complementarity and mute the response of the firm's desired price to other variables such as idiosyncratic shocks, leading to smaller price changes. Large amounts of "micro" strategic complementarities are, therefore, hard to reconcile with the large observed size of price changes. In contrast, "macro" forms of strategic complementarities are immune from the Klenow-Willis critique, since lowering Π_{12} does not affect the firm's response to idiosyncratic shocks, and therefore has little effect on the size of price changes.[39]

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[38]This equation is derived by differentiating $\Pi_1(p_{it}/P_t, M_t/P_t, A_t) = 0$ with respect to P_t and using $p_{it}/P_t = 1$ and $M_t/P_t = 1$.

[39]Gertler and Leahy (2008) and Carvalho and Lee (2011) consider another form of strategic complementarity that is not subject to the Klenow and Willis critique. In their models, strategic complementarity arises because labor markets are sector-specific and firm price setting is synchronized across firms within a sector. This does not fit into the framework discussed above because in these papers the profit function depends on the sectoral price index in addition to the variables p_{it}/P_t and M_t/P_t.

12   Concluding Remarks

Sluggish adjustment of the aggregate price level plays a key role in determining how monetary shocks, as well as other fluctuation is aggregate demand, affect the economy. Macroeconomists have long taken a wide range of approaches to acquiring empirical evidence regarding inflation dynamics—from estimating autoregressive models of aggregate data to studying the price adjustment behavior of individual goods. One goal of this literature has been to inform the development of micro-founded macroeconomic models that can be used to answer policy questions.

The recent influx of data on individual prices allows for a much more broad-based analysis of individual price dynamics than was previously possible. However, the new data also raise a host of new empirical issues that have not traditionally been confronted by parsimonious macroeconomic models of price-setting. Simple statistics such as the frequency of price change may be misleading guides to the flexibility of the aggregate price level in a setting where temporary sales, product-churning, cross-sectional heterogeneity, and retailer-manufacturer interactions play an important role. The recent empirical literature on price rigidity that we have surveyed has focused on determining which features of price adjustment at the microeconomic level are most important in determining the flexibility of the aggregate price level and attempting to quantify these features using broad-based data on prices.

Twenty-five years ago, Julio Rotemberg wrote a wonderful survey article published in the *NBER Macroeconomics Annual* on "The New Keynesian Microfoundations" (Rotemberg, 1987). Broadly speaking, the topic of this survey was the same as the topic of the present survey. Rotemberg's survey is therefore an interesting point of comparison for answering the question: Have we made progress in the last 25 years? Our sense is that we clearly have. The cutting edge research that Rotemberg surveyed was in the process of illustrating many of the basic theoretical channels that still play a vital role in the models we use today. This literature was developing tractable ways to model the key phenomena needed to capture nominal rigidities. Work along these lines has continued and some important theoretical ideas have been added to the mix since. But the basic structure is similar today.

In constrast, the empirical evidence that applied macroeconomists around 1987 were basing their models on was extremely primitive relative to the evidence that has been amassed today. Cecchetti's (1986) study of the newsstand prices of magazines and Carlton's (1986) study of industrial prices

were the cutting edge empirically. Most of the empirical questions we have focused on in our survey are not even mentioned, presumably because no evidence regarding these questions existed at the time. Most of the models Rotemberg discussed are simple "toy" models. For the most part, they don't even try to be quantitative. How could they have given that there was little evidence on which to base the types of detailed assumptions needed for a quantitative model? Here the field has been transformed. We are still a long way away from having a definitive description of the relevant aspects of price adjustment. But we have certainly narrowed the playing field substantially over the last 25 years.

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Table 1
Frequency of Price Change in Consumer Prices

	Median		Mean	
	Freq.	Impl. Dur.	Freq.	Impl. Dur.
Panel A: Nakamura and Steinsson (2008):				
Regular Prices (Excl. Subs. 1988-1997)	11.9	7.9	18.9	10.8
Regular Prices (Excl. Subs. 1998-2005)	9.9	9.6	21.5	11.7
Regular Prices (Incl. Subs. 1988-1997)	13.0	7.2	20.7	9.0
Regular Prices (Incl. Subs. 1998-2005)	11.8	8.0	23.1	9.3
Posted Prices (Incl. Subs. 1998-2005)	20.5	4.4	27.7	7.7
Panel B: Klenow and Kryvtsov (2008):				
Regular Prices (Incl. Subs. 1988-2005)	13.9	7.2	29.9	8.6
Posted Prices (Incl. Subs. 1988-2005)	27.3	3.7	36.2	6.8

All frequencies are reported in percent per month. Implied durations are reported in months. These statistics are based on BLS CPI micro data from 1988-2005. "Regular Prices" exclude sales using a sales flag in the BLS data. "Excl. Subs." denotes that substitutions not counted as price changes. "Incl. Subs." denotes that substitutions are counted as price changes. For the statistics from our 2008 paper, we take the case referred to as "Estimate frequency of price change during stockouts and sales." "Posted Prices" are the raw prices in the BLS data including sales. "Median Freq." denotes the weighted median frequency of price change. It is calculated by first calculating the mean frequency of price change for each Entry Level Item (ELI) in the BLS data and then taking a weighted median across ELI's using CPI expenditure weights. The within ELI mean is weighed in the case of Klenow and Kryvtsov but not Nakamura and Steinsson. The "Median Implied Duration" is equal to $-1/\ln(1-f)$, where f is the median frequency of price change. "Mean Frequency" denotes the weighted mean frequency of price change. "Mean Implied Duration" is calculated by first calculating the implied duration for each ELI as $-1/\ln(1-f)$, where f is the frequency of price change for a particular ELI and then taking a weighted mean across ELI's using CPI expenditure weights.

Table 2
Transicence of Temporary Sales

	Frac. Return After One Period Sales	Freq. Reg. Price Ch.	Freq. Price Ch. During One Period Sales	Av. Dur. Sales
Processed Food	78.5	10.5	11.4	2.0
Unprocessed Food	60.0	25.0	22.5	1.8
Household Furnishings	78.2	6.0	11.6	2.3
Apparel	86.3	3.6	7.1	2.1

The sample period is 1998-2005. "Frac. Return After One Period Sales" denotes the median fraction of prices that return to their original level after one period sales. "Freq. Reg. Price Ch." denotes the median frequency of price changes excluding sales. "Freq. Price Ch. During One Period Sales" denotes the median monthly frequency of regular price change during sales that last one month. The monthly frequency is calculated as $1-(1-f)^{0.5}$ where f is the fraction of prices that return to their original level after one period sales. "Av. Dur. Sales" denotes the weighted average duration of sale periods in months.
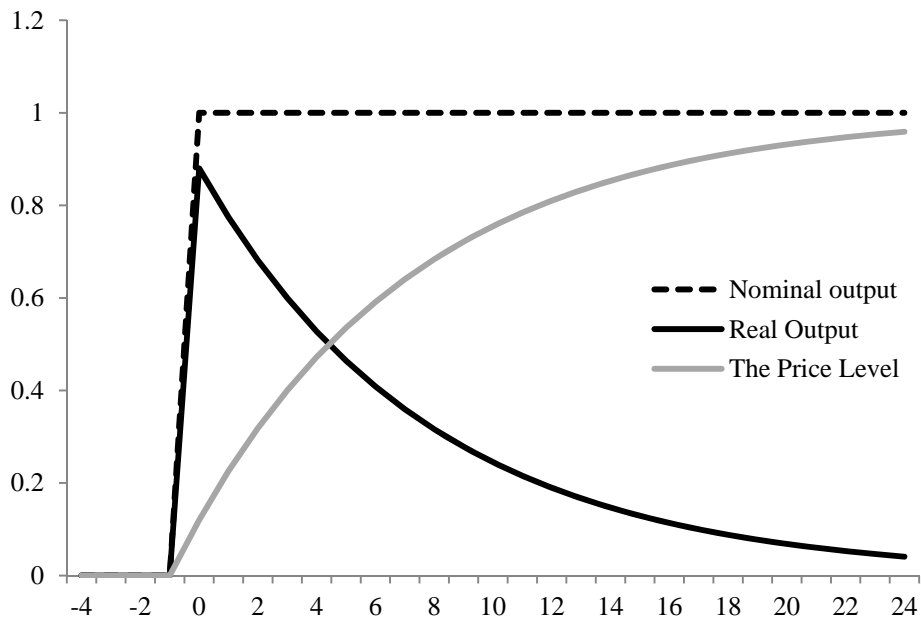
Figure 1
Response of Real Output and the Price Level to a One-Time Permanent Shock
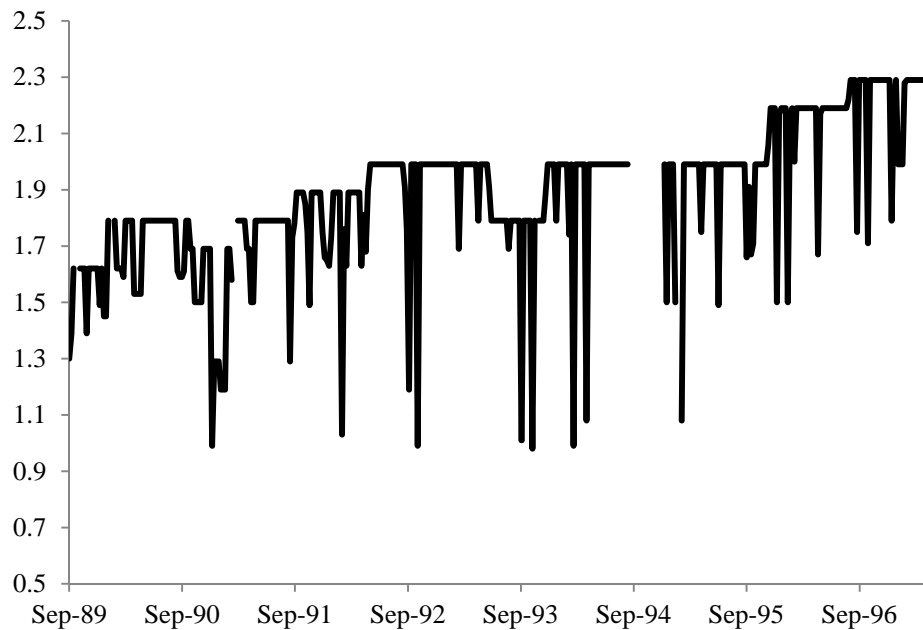to Nominal Aggregate Demand

Figure 2
Price of Nabisco Premium Saltines 16 oz. at a Dominick's Finer Foods Store in Chicago
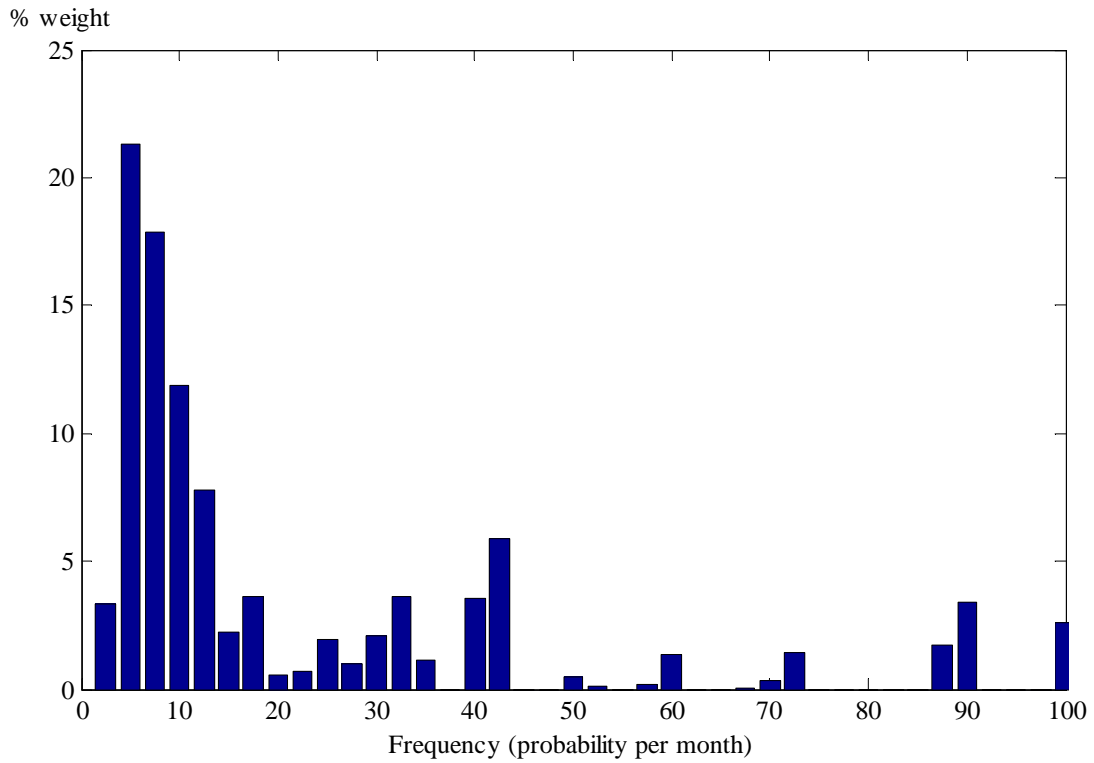
Figure 3
The Distribution of the Frequency of Price Change for U.S. Consumer Prices

Source: Nakamura and Steinsson (2008). The figure shows the expenditure weighted distribution of the frequency of regular price changes (percent per month) across product categories (ELI's) in the U.S. CPI for the period 1998-2005.