

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 the 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 raises 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 in which 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.

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 & Schwartz 1963, Christiano et al. 1999, Romer & 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 & Perotti 2002, Ramey 2011, Nakamura & Steinsson 2011). 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 an increase in 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 does not 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, and fluctuations in consumer sentiment (Keynes's "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 does not sound plausible. But many types of economic disturbances call for sharp movements in real interest rates. Because 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 & Krugman (2012) and Guerrieri & 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

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

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 failures 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 in turn will 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 phenomenon. The idea behind viewing price rigidity as reflecting price adjustment frictions is that it is unlikely that optimal prices are literally unchanged for long periods and then change abruptly by large amounts, so such price patterns in the data must reflect the presence of some form of adjustment cost.

Business cycle models that feature nominal rigidities and coordination failures 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 seminal work of Bils & 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 oversimplification, the guiding question for each piece of empirical evidence is, what does it imply about how sluggishly the overall price level responds to changes in aggregate conditions?

The article proceeds as follows. Section 2 lays out some basic facts about the frequency of price change in the US economy. Section 3 presents a simple monetary model that helps explain macroeconomists' persistent interest in price rigidity by illustrating the close connection between price rigidity and the economy's response to monetary (and other demand) shocks. The remainder of the article delves into various features of price adjustment that complicate the relationship between the degree of micro price rigidity and the responsiveness of the aggregate price level to shocks. Sections 4 and 5 discuss temporary sales and cross-sectional heterogeneity, respectively, both of which are first-order issues in defining what we mean by "the" frequency of price change. Sections 6 and 7 present 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 examine seasonality in and the hazard function of price adjustment, respectively. 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 explores the important role that coordination failures play in monetary models and examines 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 data sets on micro price data are, in practice, difficult to obtain. The conventional wisdom among researchers working on New Keynesian business cycle models in the 1990s and early 2000s was that prices changed roughly once a year. A common citation for this was Blinder et al.'s (1998) survey study of firm managers. Other important early studies include Carlton (1986), Cecchetti (1986), Lach & Tsiddon (1992), and Kashyap (1995). Bils & Klenow (2004) shattered this conventional wisdom by documenting that the median monthly frequency of price change in the micro data underlying the nonshelter component of the US Consumer Price Index (CPI) 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 data sets underlying the CPI, Producer Price Index (PPI), and Import and Export Price Indexes, collected by the US Bureau of Labor Statistics (BLS). Klenow & Kryvtsov (2008) and Nakamura & Steinsson (2008) analyze in detail the micro data underlying the US CPI for the period 1988–2005. **Table 1** presents results on the frequency of price change from these papers. Readers are also referred to Hosken & Reiffen (2004, 2007), who analyze the prevalence and characteristics of temporary sales in the BLS CPI data.

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** presents 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 frequencies of price change, as well as the median and mean implied durations. The implied duration for a particular sector is defined as $d = -1/\ln(1 - f)$, where f is the frequency of price change in that sector.² The median implied duration is the implied duration for the sector with the median frequency of price change, whereas the mean implied duration is defined as the expenditure weighted mean of the implied durations in different sectors.³ Nakamura & Steinsson (2008) report results for several different ways of treating observations that are missing because of 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 & Steinsson (2008) estimate the frequency of price change over the course of these events using the price before and after the missing period.⁴

²A 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 $d = 1/\lambda = -1/\ln(1 - f)$.

³Why 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 data sets such as the BLS data, arising from products dropping out of the data set owing to product turnover and BLS resampling. Dropping censored spells would lead to biased duration estimates.

⁴Other methods considered in Nakamura & Steinsson (2008) 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 and higher implied durations.

Table 1 Frequency of price change in consumer prices

	Median		Mean	
	Frequency (% per month)	Implied duration (months)	Frequency (% per month)	Implied duration (months)
Nakamura & Steinsson (2008)				
Regular prices (excluding substitutions 1988–1997)	11.9	7.9	18.9	10.8
Regular prices (excluding substitutions 1998–2005)	9.9	9.6	21.5	11.7
Regular prices (including substitutions 1988–1997)	13.0	7.2	20.7	9.0
Regular prices (including substitutions 1998–2005)	11.8	8.0	23.1	9.3
Posted prices (including substitutions 1998–2005)	20.5	4.4	27.7	7.7
Klenow & Kryvtsov (2008)				
Regular prices (including substitutions 1988–2005)	13.9	7.2	29.9	8.6
Posted prices (including substitutions 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 US Bureau of Labor Statistics (BLS) Consumer Price Index (CPI) micro data from 1988 to 2005. Regular prices exclude sales using a sales flag in the BLS data. Excluding substitutions denotes that substitutions are not counted as price changes. Including substitutions denotes that substitutions are counted as price changes. For the statistics from Nakamura & Steinsson (2008), 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. The median frequency 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 the ELIs using CPI expenditure weights. The within-ELI mean is weighted in the case of Klenow & Kryvtsov (2008) but not Nakamura & Steinsson (2008). The median implied duration is equal to $-1/\ln(1-f)$, where f is the median frequency of price change. The mean frequency denotes the weighted mean frequency of price change. The 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 the ELIs using CPI expenditure weights.

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, whereas for regular prices, it is roughly three quarters depending on the sample period and the treatment of substitutions.⁵ 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 frequencies of price change. For example, in Nakamura & Steinsson's (2008) results 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%, whereas

⁵For posted prices, the median frequencies of price change in Nakamura & Steinsson (2008) are close to those in Bils & Klenow (2004), whereas Klenow & Kryvtsov (2008) report higher frequencies of price change. Klenow & Kryvtsov (2008) note that these differences result from different samples (all cities versus top three cities) and different weights (category weights versus item weights).

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 monetary models. 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., the 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, whereas calibrating to the median frequency or the mean implied duration yields a better approximation to a full-fledged multisector model. We discuss this work in Section 5.

Certain product categories—in particular, durable goods—undergo frequent product turnover. For some of these 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 women’s 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 women’s dresses is 25.8% per month. It seems likely that most of the price adjustment for women’s 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.

As seen in **Table 1**, the median frequency of regular price change including product substitutions is roughly 1.5 percentage points higher than that excluding product substitutions.⁶ 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 United States have become available in recent years to academic researchers. Álvarez (2008) and Klenow & Malin (2011) tabulate studies using these data and present their estimates of the frequency of price change. An important part of this body of work was carried out within the context of the Inflation Persistence Network of the European Central Bank. The conclusions of this work are summarized in Álvarez et al. (2006) 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 product bar codes are scanned at the checkout aisle. A broad-based data set on supermarkets for the US economy is the SymphonyIRI research database. Also widely used is the Dominick’s Finer Foods database, available online from the Kilts Center for Marketing of the University of Chicago Booth School of Business (<http://research.chicagobooth.edu/marketing/databases/dominicks/index.aspx>). This data set 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 et al. (2011) and Gopinath et al. (2011). An important advantage of scanner data sets is that they often include information on the quantities 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. Nakamura et al. (2011) show that pricing policies differ a great deal across supermarket chains.

Recent studies also apply similar methods to broad-based BLS data sets on US producer prices both for domestic and internationally traded goods (Gopinath & Rigobon 2008, Nakamura &

⁶The CPI research database provides an imperfect measure of product turnover 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.

Steinsson 2008, Goldberg & Hellerstein 2011) as well as producer prices in other countries (see Álvarez 2008 and Klenow & Malin 2011, and references therein). 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 & Zerom 2010, Eichenbaum et al. 2011, Anderson et al. 2012b, Goldberg & Hellerstein 2013). 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 nonprice features that may vary over time (Carlton 1979).

Similarly, wage rigidity and price rigidity are closely intermingled, as 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. Numerous recent studies have examined issues similar to those described above using broad-based data on wages (e.g., Dickens et al. 2008, Barattieri et al. 2010, Le Bihan et al. 2012, Sigurdsson & Sigurdardottir 2012). It can be difficult to interpret data on both producer prices and wages because 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 a representation of 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 on which macroeconomists have focused, it is useful to introduce a 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 the length of time 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 of its own past value and the log price p_{it}^* set by firms that change their prices in period t :

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

Although few macroeconomists would argue for this model as a literal description of how firms set prices, the goal of this so-called 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.⁷

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 unimportant constant term to zero. Suppose that firms discount future profits at a rate β and face an isoelastic demand curve $y_{it} - y_t = -\theta(p_{it} - p_t)$, where y_{it} denotes log demand for product i , y_t

⁷For example, Woodford (2009) shows that the Calvo model can provide a good approximation to firms' pricing behavior when firms face information processing costs.

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 (see, e.g., Woodford 2003, chapter 3; Galí 2008, chapter 3):

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

Let m_t denote log nominal output:

$$m_t = y_t + p_t. \quad (3)$$

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:

$$m_t = \mu + m_{t-1} + \epsilon_t. \quad (4)$$

This specification for monetary policy is equivalent to a simple rule for the money supply if the velocity of money is constant.⁸

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 $w_t - p_t = c_t$, where $c_t = \log C_t$. 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 in which average growth in nominal output is zero ($\mu = 0$ in Equation 4). Because 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 failures discussed in Section 1, in which one firm changes its price by less than it otherwise would because other firms have not changed their price, do not arise. The pricing decisions of firms in this simple model are said to be strategically independent—neither strategic complements nor strategic substitutes.

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

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

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.

⁸More 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.

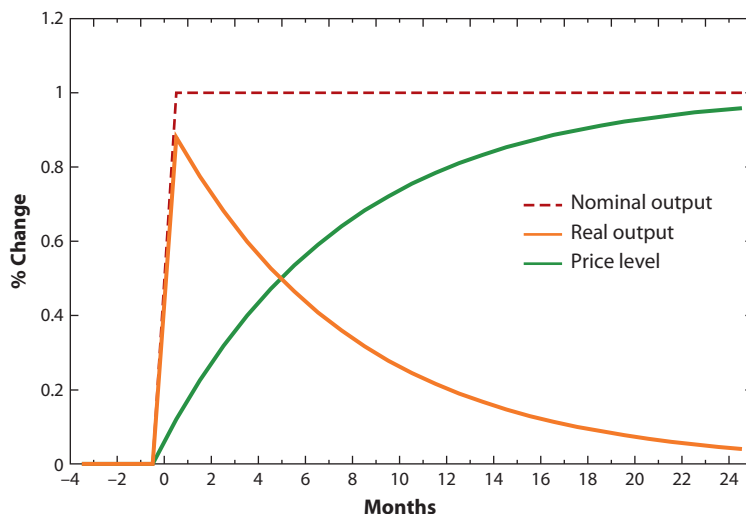


Figure 1

Response of real output and the price level to a one-time permanent shock to nominal aggregate demand in the Calvo model.

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 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 nine regular price changes over a roughly seven-year period. Which of these summary measures of

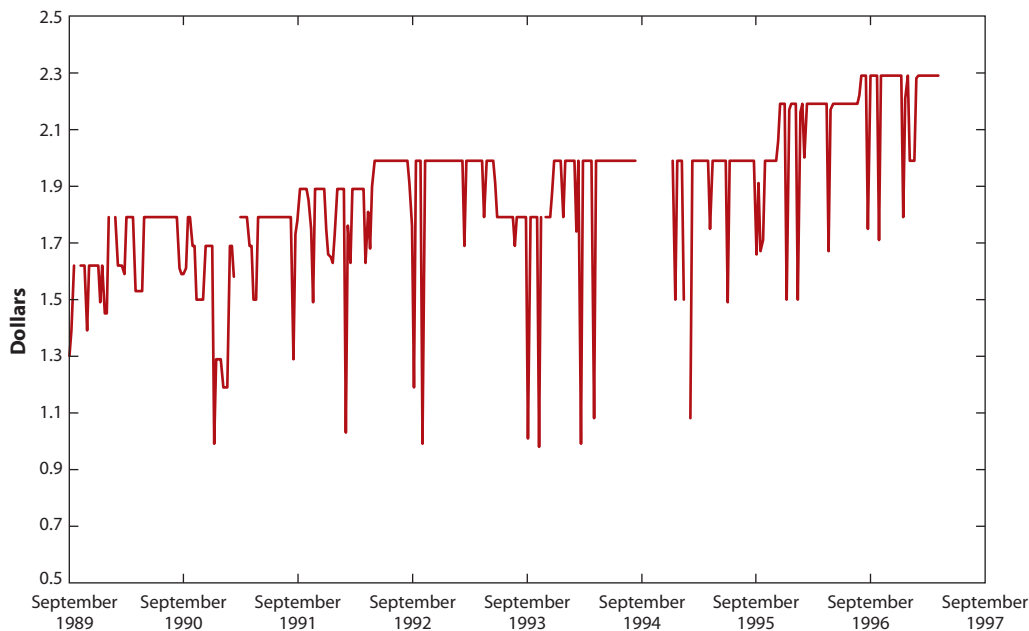


Figure 2

Price series of Nabisco Premium Saltines (16 oz) at a Dominick's Finer Foods store in Chicago.

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; in other words, all price changes should be counted equally. However, **Figure 2** also illustrates that sales have very different empirical characteristics than regular price changes do. Whereas regular price changes are in most cases highly persistent, sales are highly transient.⁹ In fact, in most cases, the posted price returns to its original value following a sale. **Table 2** reports results from Nakamura & Steinsson (2008) 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%.¹⁰ Clearance sales are not included in these statistics because a new regular price is not observed after such sales. Nakamura & Steinsson (2008, supplementary material) argue that clearance sales, like other types of sales, yield highly transient price changes.

⁹Sales are identified either by direct measures such as sales flags (as in the BLS data) or by sale filters that identify certain price patterns (such as V-shaped temporary discounts) as sales. Although it is often said that by looking at a price series, one can easily identify the regular price and the timing of sales, constructing a mechanical algorithm to do this is more challenging. Nakamura & Steinsson (2008), Kehoe & 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 ACNielsen to identify temporary sales.

¹⁰It 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 (the third data column). In addition, higher-frequency data sets indicate 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.

Table 2 Transience of temporary sales

	Fraction return after one-period sales	Frequency of regular price change	Frequency of price change during one-period sales	Average duration of 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. The first data column gives the median fraction of prices that return to their original level after one-period sales. The second is the median frequency of price changes excluding sales. The third lists the median monthly frequency of regular price change during sales that past 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 levels after one-period sales. The fourth data column gives the weighted average duration of sale periods in months. Data taken from Nakamura & Steinsson (2008).

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 quickly reversed.

To answer the question of how to treat sales in arriving at 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—such as 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 & Midrigan (2010) build a menu cost model in which firms can either change their prices permanently (i.e., change their regular price) or, at a lower cost, change their prices 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, sales price changes contribute little to the response of aggregate prices to monetary shocks. In their model, thus, the degree of monetary non-neutrality is similar to that in the case where sales price changes are completely absent. The key intuition is that because sales 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 & Sheedy (2011) develop the idea that firms use sales to price discriminate between low- and high-price elasticity consumers into a macroeconomic business cycle model.¹¹

¹¹Sobel (1984) introduced the idea that sales might result from price discrimination between customers with different price elasticities. Other important papers on sales in the industrial organization (IO) literature include Varian (1980), Salop & Stiglitz (1982), Lazear (1986), Aguirregabiria (1999), Hendel & Nevo (2006), and Chevalier & Kashyap (2011). Hosken & Reiffen (2004) use BLS CPI data to evaluate the empirical implications of IO models of sales.

In their model—just as in the model of Kehoe & 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 & 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 low sales prices average out across stores, reducing their effect on the aggregate price level.

An important point is that, in both Kehoe & Midrigan (2010) and Guimaraes & 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 et al. (2011) analyze a scanner data set from a large US 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.¹² 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. (2012b) analyze a unique data set from a large US 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 are largely unresponsive to these shocks.¹³ Coibion et al. (2012) and Anderson et al. (2012b) show that the frequency and size of sales fall when unemployment rates rise (i.e., changes in the behavior of sales raise rather than reduce prices in a recession). In contrast, Klenow & Willis (2007) demonstrate that in the BLS CPI data, the size of sales price changes is related to recent inflation in much the same way as the size of regular price changes. Klenow & 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.

In summary, there are three main 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

¹²Stevens (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.

¹³The idea that sales may not respond to changes in macroeconomic conditions is suggestive of information costs, sticky information, or rational inattention (Mankiw & Reis 2002, Burstein 2006, Woodford 2009, Sims 2011).

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 US economy. Figure 3 illustrates this in a histogram of the frequency of regular price change across different CPI product categories from Nakamura & Steinsson (2008). Whereas 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% and 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 in which half the prices adjust all the time (e.g., gasoline) and half hardly ever adjust (e.g., haircuts) or one in which all prices adjust half of the time? A related question is, if one wishes to approximate the behavior of the US economy using a model with homogeneous firms, should one calibrate the frequency of price change to the mean or median frequency of price

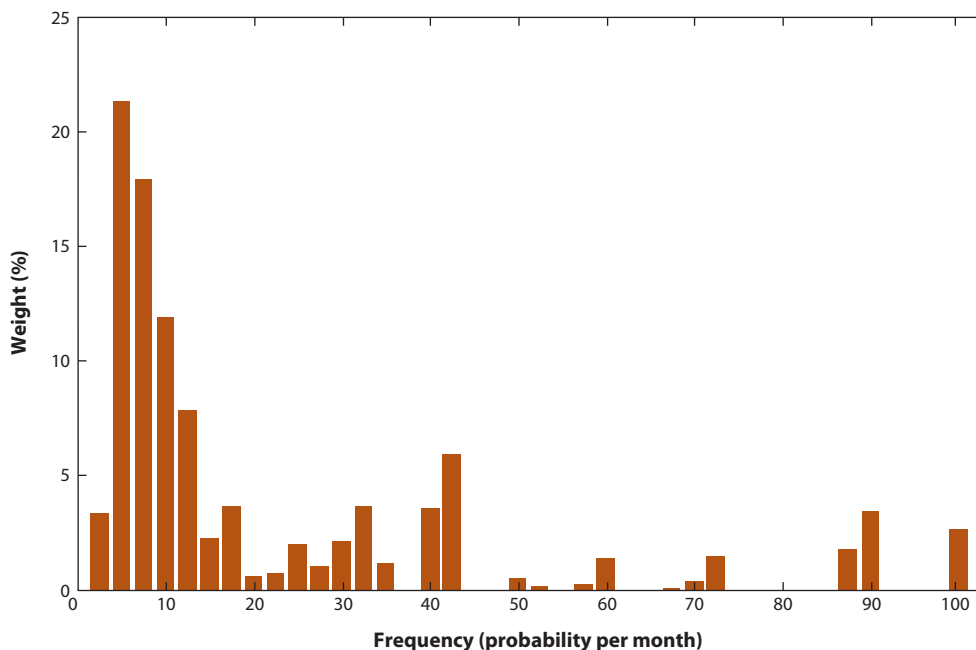


Figure 3

The expenditure weighted distribution of the frequency of regular price change (percent per month) across product categories (entry-level items) in the US Consumer Price Index (CPI) for the period 1998–2005. Data taken from Nakamura & Steinsson (2008).

change in the data? As we discuss in Section 2, the difference between these different summary measures is first order.

Several authors have sought to answer these questions using detailed multisector models of the economy designed to incorporate cross-sectional heterogeneity in the frequency of price change. Bills & Klenow (2002) consider a multisector 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 multisector model. This analysis motivates the focus on the median frequency of price change in Bills & 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 multisector version of his model that incorporates heterogeneity in the frequency of price change generates much more monetary non-neutrality than does 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 multisector 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 multisector model. **Table 1** reports that for consumer prices excluding sales in the United States, 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.¹⁴

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 in which 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 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 prices 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 do not contribute to the adjustment of the aggregate price level to the shock as the firm has already adjusted to the shock.¹⁵ If it were possible to reallocate some price changes from the sector with the high frequency of price change to that with the low frequency of price change, this would speed the adjustment of the aggregate price level because a higher proportion of firms in the low frequency of price change sector have not adjusted to the shock.

¹⁴The average duration, $\text{mean}(1/f)$, is larger than the duration implied by the average frequency, $1/\text{mean}(f)$, by Jensen's inequality because $1/x$ is a convex function.

¹⁵We note that the model in Section 3 has no strategic complementarity. Firms therefore respond fully to the shock the first time they change their prices after the shock occurs. (This argument also goes through in a model with strategic complementarity, but it involves slightly more steps.)

A question that arises regarding the results of *Bils & Klenow (2002)* and *Carvalho (2006)* is whether they carry over to a setting in which firms choose the timing of their price changes optimally subject to a menu cost. *Nakamura & Steinsson (2010)* address this question by showing that in an economy with low inflation and large price changes (e.g., the United States), the timing of price changes is dominated by idiosyncratic shocks. This implies that the frequency of price change does not 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, which reduces the degree to which heterogeneity amplifies the amount of monetary non-neutrality. When we calibrate our model to the US economy, we find that our multisector 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 et al. (2009a,b)* use sectoral data to estimate a multisector DSGE (dynamic stochastic general equilibrium) model of the US economy. Their estimates of the frequency of price change are highly correlated with *Nakamura & 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 et al. (2003)* find that the relative price of flexible price sectors falls after an expansionary monetary policy shock identified using structural vector autoregression (VAR) methods.¹⁶ *Mackowiak et al. (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 that 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 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, as 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 1994–2002. Mexico experienced a serious currency crisis in December

¹⁶A related result in *Bils & 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 do sectors with a high frequency of price change, as simple sticky-price models suggest they should.

1994. Year-on-year inflation in Mexico rose from below 10% in the fall of 1994 to approximately 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 change rises rapidly with inflation. Gagnon compares his empirical results with simulations of a menu cost model along the lines of Golosov & 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.¹⁷

Álvarez 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 5,000%. A successful stabilization plan ended the hyperinflation in 1991, and inflation fell quickly so that after 1992 there was virtual price stability. Álvarez 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 two-thirds. They then show empirically that both 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 two-thirds over a very large range. Again, the menu cost model seems to fit data on the frequency of price change remarkably well. Other papers that study related questions include Konieczny & Skrzypacz (2005) for Poland (1990–1996), Barros et al. (2009) for Brazil (1996–2008), and Wulfsberg (2010) for Norway (1975–2004).

The extent to which the frequency of price change varies with inflation has also been studied using the US CPI micro data for the period 1988–2005 (Klenow & Kryvtsov 2008, Nakamura & Steinsson 2008). However, the US 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 & Kryvtsov and Nakamura & Steinsson is that the frequency of price increases varies more with inflation than does 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 inflation.¹⁸ We show that this feature of the data 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 toward the lower bound than toward the upper bound. The bunching toward the lower bound implies that the frequency of price increases covaries more than does the frequency of price decreases with shocks to inflation.

¹⁷Gagnon et al. (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 value-added tax (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 & 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 can.

¹⁸We focus on the median frequency of price change, whereas Klenow & 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 owing to their high frequencies of price change. The relationship between the frequency of price change and inflation documented for the median good holds individually in all sectors with a substantial degree of price rigidity.

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 do price changes that are timed randomly. Golosov & 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 & 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 & Weiss 1977, 1983); in other words, they will wait to change their price until their relative price has fallen to a trigger level s , and at that point they will 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 regardless of 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 thus will 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 are selected to be the firms that most need to change their price (i.e., 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 (for related work, see Caballero & Engel 1991, 1993; Caplin & Leahy 1991, 1997; Danziger 1999; see Head et al. 2012 for a completely different, search-based, argument for why micro price rigidity may be entirely divorced from sluggishness of the aggregate price level).

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 in which aggregate shocks have no effect on which firms and how many firms change their prices. The Caplin-Spulber model is the opposite extreme case in which 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 the real world is located on the spectrum between the Calvo model and the Caplin-Spulber model.

A key feature of the data that is fundamentally inconsistent with the simple models discussed above is the large size of price changes. Klenow & Kryvtsov (2008) show that the average absolute size of price changes for US consumer prices is very large—roughly 10%. A related fact is that approximately 40% of regular price changes are price decreases. In the simple models discussed above, firms are reacting only to aggregate inflation when they change prices. Because inflation is almost always positive, these models imply that almost all price changes should be price increases. Furthermore, as Golosov & Lucas (2007) point out, with inflation of approximately 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 & 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.¹⁹ 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.”

Midrigan (2011) argues, however, that this result is quite sensitive to the distribution of idiosyncratic shocks. In an extreme example, 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 toward a weaker selection effect than Golosov & Lucas’ model implies. First, he 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, although 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 (see Kashyap 1995 and Lach & Tsiddon 1996 for earlier evidence on small price changes).²⁰ Finally, Midrigan 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.²¹ 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 data sets, it is necessary to pool multiple products to obtain an estimate of the distribution of the size of price changes across 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 & 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 nonadjusters—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

¹⁹The 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 US supermarket chain to be 0.7% of revenue. Nakamura & Zerom (2010) estimate the costs of changing prices for coffee manufacturers to be 0.23% of revenue.

²⁰Klenow & Kryvtsov (2008) find a large number of small price changes in the US CPI data. More recently, Eichenbaum et al. (2012) argue that measurement problems may be behind many of the small price changes observed in the type of data used by Midrigan (2011). Also, Anderson et al. (2012a) find very few small regular price changes in data from a large US retailer.

²¹More recently, Bhattarai & Schoenle (2012) present related evidence from US 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. Álvarez & Lippi (2012) explore the implications of such economies of scope in price setting in more detail.

“uncertainty shocks”—variation in the variance of idiosyncratic shocks—and finds that this model can match the countercyclicality 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, however, match up well with standard estimates of the variability of plant-level productivity, materials costs, and wages (e.g., evidence from the US Annual Survey of Manufacturers suggests that the annual volatility of total factor productivity is in the neighborhood of 10%). An alternative view is that idiosyncratic shocks stand in for other, unmodeled sources of variation in firms’ desired prices. Burstein & 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 result from 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 suggest that the process for regular price changes and that for temporary discounts may be largely disconnected from each other (Eichenbaum et al. 2011, Anderson et al. 2012b). 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, whereas the timing, depth, and frequency of sales are determined by intertemporal price discrimination, advertising, and shifts in fashion and tastes.

8. SEASONALITY IN THE FREQUENCY OF PRICE CHANGE

In Section 6, we present 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 & Steinsson (2008) 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 results from a spike in the frequency of price change in January. Álvarez et al. (2006) 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. Alternatively, 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 as to why. It may be a result of 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 & Tenreyro (2007) show that the response of output to monetary shocks identified using a structural VAR is larger for shocks that occur in the first and second quarters than for shocks that occur in the third and fourth quarters.

Nakamura & Steinsson (2010) document 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 (e.g., the fall and spring clothing seasons in apparel) that are largely orthogonal to a firm's desire to change its price. Although 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, although 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 is 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, conversely, 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 in which inflation is the only motive for price adjustment, a firm's 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, as transient idiosyncratic shocks tend to flatten the hazard function (see Nakamura & Steinsson 2008 for a discussion of this issue).²² In addition, the inaction region is narrower in good times because more is at stake in getting the price right. This also tends to flatten the hazard function (Klenow & 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

²²The 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 the price back to its previous level.

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 & Kryvtsov (2008) and Nakamura & Steinsson (2008) study the hazard function of price change for US CPI data. Both papers seek to control for cross-sectional heterogeneity in the frequency of price change. Klenow & 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.²³

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 price 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 & 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 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 US supermarket chain, Eichenbaum et al. (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 levels when reference prices change. Campbell & 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 et al. (2012) carry out a similar exercise with the SymphonyIRI marketing data set. 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

²³One argument for using a random effects model is that hazard function models with fixed effects suffer from the incidental parameters problem, leading to biased coefficient estimates.

point at which 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 & Woodford 1997, Woodford 2003).

Burstein & Hellwig (2008) compare the welfare costs of inflation arising from price dispersion in the Calvo model with those in 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, whereas 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.²⁴

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 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 IO economists because it provides evidence regarding 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 are uncorrelated with the rate of inflation.

Much of the literature on this topic has focused not on price dispersion but rather on the dispersion of inflation rates for individual prices within a category or the dispersion of inflation rates across categories. This concept has often been referred to as relative price variation. We, 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 & Elwertowski 1976, Parks 1978, Fischer 1981, Debelle & Lamont 1997). Van Hoomissen (1988) and Lach & Tsiddon (1992) consider the relationship between category-level inflation rates and inflation dispersion at the product level within the category. Both these papers use Israeli data. Van Hoomissen considers 13 narrow product categories over the period 1971–1984, whereas Lach & Tsiddon consider 26 narrow product categories over the period 1978–1984. 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 estimates a negative quadratic term, suggesting that price dispersion rises less steeply at high rates of inflation.

Álvarez et al. (2011) present theoretical results for a menu cost model, showing that price 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, whereas 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 one-third. 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 1989–1990 and subsequent stabilization, Álvarez et al. show that the elasticity of price dispersion in the data lines up well with the

²⁴Kiley (2000) and Levin et al. (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.

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 one-third for homogeneous goods but is 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.²⁵

11. COORDINATION FAILURES 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 prices 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 failures among price setters. In such models, the firms that have an opportunity to change their prices 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 failures can be powerful amplification mechanisms for the aggregate effects of nominal rigidities. They imply that the sluggishness of the aggregate price level lasts well beyond the point at which all prices have changed at least once. Coordination failures among price setters can thus imply that the macro sluggishness of prices is much greater than the micro rigidity of prices.

Coordination failures among price setters have 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 & Tsiddon 1996). In the Taylor (1980) and Calvo (1983) models, price changes are assumed to be staggered. In Golosov & 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 failures 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 prices soon after an aggregate shock will respond incompletely to the shock because other firms have not yet changed their prices.²⁶ 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 prices that period.

One source of strategic complementarity is intermediate inputs (Basu 1995). Consider a modified version of the model in Section 3, in which 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

²⁵Reinsdorf (1994) studies the relationship between inflation and price dispersion for individual goods in the United States over the period 1980–1982 using microdata underlying the US CPI. In contrast to other studies, he finds a negative relationship between inflation and price dispersion.

²⁶Strategic complementarity in price setting is closely related to Ball & Romer's (1990) notion of "real rigidities." It is also closely related to Taylor's (1980) notion of the "contract multiplier." Readers are referred to Cooper & John (1988) for a general discussion of strategic complementarity.

simplicity that Q_{it} is a constant elasticity of substitution (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 costs also depend on the aggregate price level, as 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 goods in the economy. This implies that firms' pricing decisions are strategic complements. If price setting is staggered, firms that have an opportunity to change their prices 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 the assumption of counterfactually low frequencies of price change.

The importance of coordination failures as amplification mechanisms 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 because it entails measuring how much prices respond to a given increase in costs rather than just whether they respond at all. Researchers using microdata on price adjustment have made some progress in this regard.

Gopinath et al. (2010) study the response of prices to changes in exchange rates at the micro level using BLS data on US 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 & Haller (2012) and Burstein & 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 et al. (2010) document, however, that prices respond much more in the long run to changes in the exchange rate. A related fact documented by Gopinath & 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.²⁷

However, evidence from microdata has also been used to level important critiques against models with staggered price changes and strong strategic complementarities in pricing. Bils et al. (2012) emphasize that strategic complementarity in price setting reduces volatility and increases the persistence of inflation. Rather than adjusting fully to an aggregate shock the first time it has an

²⁷An important challenge in interpreting this evidence is how to account for the endogeneity of exchange rate movements.

opportunity to change its price after the shock, a firm will respond with a sequence of smaller price changes in the same direction. Bils et al. 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.²⁸ 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 & 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 & 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 a small amount when it changes its prices as many of its competitors are not adjusting. Klenow & Willis (2006) point out that this implication is inconsistent with the empirical evidence on the large size of price changes. Golosov & Lucas (2007) and Burstein & Hellwig (2007) show that this same issue exists 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 model 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 (see Ball & Romer 1990, Kimball 1995, Nakamura & Steinsson 2010). 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 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,²⁹

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

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.

²⁸For firms that change their prices in period t , the reset price is simply their current price. For firms that do not change their prices, Bils et al. (2012) index the reset price to the rate of reset price inflation of those firms that change their prices. In a Calvo model without strategic complementarity, reset price inflation will equal the change in nominal aggregate demand, whereas actual inflation will lag behind as in **Figure 1** owing to price rigidity. This makes reset price inflation particularly useful for assessing the degree of strategic complementarity in price setting.

²⁹This 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 $M_t/P_t = 1$.

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 United States, inflation is small relative to observed price changes, so the size of price changes is determined primarily by the magnitude of idiosyncratic shocks. The response of a firm's price to an idiosyncratic shock is given by $\partial p_i / \partial A_i = -\Pi_{13} / \Pi_{11}$, 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 to the Klenow-Willis critique because lowering Π_{12} does not affect the firm's response to idiosyncratic shocks and therefore has little effect on the size of price changes.³⁰

12. CONCLUDING REMARKS

Sluggish adjustment of the aggregate price level plays a key role in determining how monetary shocks, as well as other fluctuations in 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 microfounded 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 in which temporary sales, product churning, cross-sectional heterogeneity, and retailer-manufacturer interactions play important roles. The recent empirical literature on price rigidity reviewed above 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-six years ago, Rotemberg (1987) wrote a wonderful survey article in the *NBER Macroeconomics Annual*, “The New Keynesian Microfoundations.” Broadly speaking, the topic of that survey is 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 past 26 years? Our sense is that we clearly have. The cutting-edge research 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. But the basic structure is similar today.

In contrast, the empirical evidence on which applied macroeconomists around 1987 were basing their models was extremely limited 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

³⁰Gertler & Leahy (2008) and Carvalho & Lee (2011) consider another form of strategic complementarity that is not subject to the Klenow-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_i/P_t and M_t/P_t .

industrial prices were the cutting edge empirically. Most of the empirical questions we focus on above are not even mentioned, presumably because no evidence regarding these questions existed at the time. Most of the models Rotemberg discusses are simple toy models. For the most part, they do not 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 past 26 years.

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