# Sectoral Asymmetries in the Frequency of Price Changes and Exogenous Interest Shocks

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**Abstract**

This paper studies sectoral price responses to an exogenous interest rate shock in the United States. It has two main findings. First, an interest rate shock causes strong relative price effects as price responses to such a shock differ largely among sectors. Second, sectoral asymmetries in the frequency of price changes are only weakly associated with sectoral price responses. I show that a multi-sector model where sectors differ not only in the frequency of price changes but also in the structure of production costs is capable of explaining these two findings.

**Keywords:** Sectoral Heterogeneity in the Frequency of Price Changes; The Shocks to Monetary Policy.

**JEL Classification Numbers:** E13, E31, E32, E37, E47, E52, E58

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# Contents

# 1 Introduction

Heterogeneity in Price Flexibility and Monetary Policy Shocks This paper seeks to answer two questions: First, do shocks to monetary policy in the United States induce sectoral prices to exhibit common or divergent dynamics? Second, are such shocks the cause of different price dynamics in fast-adjusting sectors where prices change often, compared to the slow-adjusting sectors where prices change infrequently?

In regards to the first question, I find that monetary policy shocks in the United States lead to divergent sectoral price dynamics, which suggests there are relative price effects of the monetary shocks. Indeed, I find that while the price responses in some sectors to monetary shocks are muted, they are strongly positive or negative in others. This finding is in conformity with the finding in Balke & Wynne (2007). However, they surprisingly find that a contractionary shock to monetary policy preponderantly results in positive initial price responses. In contrast, I find that the initial price responses in sectors to such a shock are equally divided between negative and positive responses, resulting in the initial aggregate price responses staying muted.[1] The difference in the distribution of negative and positive price responses between Balke & Wynne (2007) and this paper can result from the fact that a measure of the output gap is missing in Balke & Wynne (2007). As argued in Giordani (2004), the absence of an output gap measure in a VAR model may result in "the price puzzle", which refers to the counter-intuitive finding that an unanticipated monetary tightening causes an increase in the price level. I use the capacity utilization rate as a measure of the output gap and find no evidence of the puzzle at the *disaggregated level* in the VAR model.

Next, I analyze the correlation between the frequency of price changes in a sector and its impulse response functions to a contractionary monetary policy shock over five years. I find the frequency of price changes in a sector does not play a decisive role in its price responses to monetary policy shocks as the correlations between the frequency of price changes and the price responses are weak and never significantly different from zero. This finding contrasts sharply with

[1] This finding is similar to the finding in Boivin, Giannoni & Mihov (2009).

that in Bils, Klenow & Kryvtsov (2003) who find that a higher frequency of price change for a given sector is associated with a higher price response when a contractionary monetary shock occurs. I show the assumption in Bils, Klenow & Kryvtsov (2003), that the isolated monetary shocks and sector-specific price shocks are orthogonal, is violated for a considerable number of sectors. The violation of such a critical assumption may drive the counter-intuitive finding in Bils, Klenow & Kryvtsov (2003).

Lastly, I attempt to develop a DSGE model to explain two important findings in this paper: the interest rate shock causes strong relative price effects, and there is a weak association between the impulse response functions of sectoral prices and the frequency of price changes in sectors. Three DSGE models are considered: The first model is the *one-sector model*, in which it is assumed that the frequency of price changes is the same among all sectors. The second model is the *multi-sector model with symmetric cost structure* in which sectors differ only in regards to the frequency of price changes. Lastly, the third model is the *multi-sector model with asymmetric cost structure* in which sectors differ not only in regards to the frequency of price changes, but also in their production costs' structure. I show that while the *one-sector model* can explain the second finding successfully, this model may not explain the strong relative price effects of the interest rate shock at the disaggregated level. Quite the opposite, the *multi-sector model with symmetric cost structure* is successful in explaining relative price effects. Yet, this model fails to account for the low correlations of the frequency of price changes with sectoral price responses over five years following an interest rate shock. The last model, the *multi-sector model with asymmetric cost structure*, on the other hand, successfully explains both of the aforementioned two findings of the empirical section. Therefore, I conclude this model outperforms the other two models.

The organization of the paper is as follows. Section 2 presents the emprical strategy for isolating monetary shocks in the United States and studies the impulse response functions of sectoral prices to such shocks. Section 3 develops three theoretical models and evaluates the success of these models in explaining the strong relative price effects of the interest rate shock and the weak correlations between the frequency of price changes and sectoral price responses. The last section concludes the discussion.

# 2 The Empirical Section

This section develops my empirical strategy for analyzing sectoral price responses following a contractionary shock to the federal funds rate. Before investigating how sectoral prices change following an exogenous interest shock, it is first useful to study the aggregate dynamics.

## 2.1 Aggregate Dynamics after an Exogenous Shock in the Federal Funds Rate

In Tugan (2013), the following VAR model is considered,

$$\Upsilon_t = B_0 + \sum_{k=1}^{k_{max}} B_k \Upsilon_{t-k} + A_0 \mathcal{E}_t \tag{2.1}$$

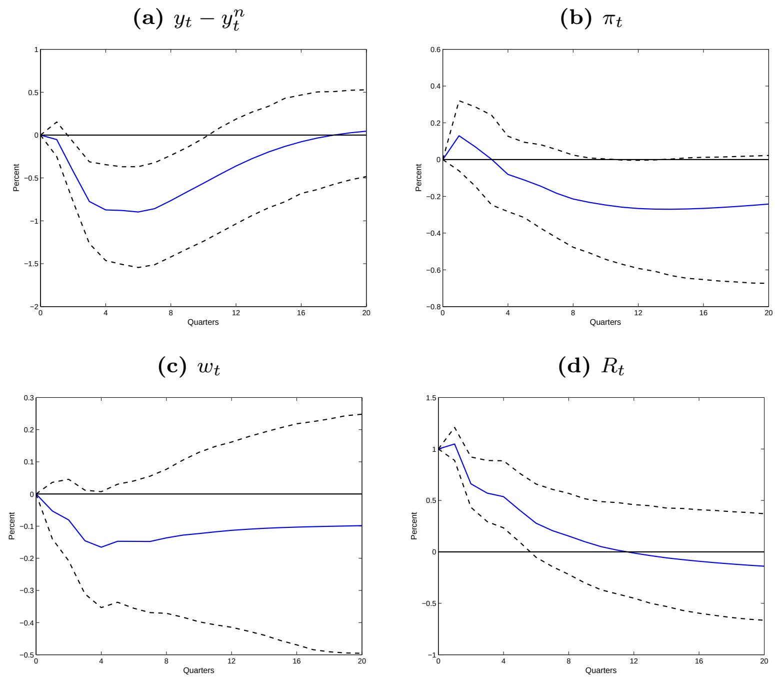
where structural shocks and the number of lags included are denoted by $\mathcal{E}_t$ and $k_{max}$, respectively. $A_0$ stands for the contemporaneous response matrix of the variables to these shocks. Lastly, $\Upsilon_t$ denotes the vector of variables contained in the VAR and is given as:

$$\Upsilon_t = [y_t - y_t^n, \pi_t, w_t, R_t] \tag{2.2}$$

I again consider the VAR model in (2.1) to study the aggregate dynamics in this paper. The variables included in $\Upsilon_t$ are the capacity utilization rate in manufacturing as a measure of the output gap $(y_t - \hat{y}_t^n)$, annualized inflation $(\pi_t)$, the real wage $(w_t)$ and the federal funds rate $(R_t)$ (See Tugan (2013) for a detailed explanation of the variables). The VAR is quarterly and contains four lags of each variable. The sample spans the period of 1959Q1-2013Q1. The ordering of the variables in the VAR system implies that the capacity utilization, inflation and the real wage respond to the monetary policy shock with one quarter lag. This assumption is standard in the literature.

Figure 1 displays the impulse responses of the variables contained in the VAR system to an unanticipated 1% rise in the federal funds rate together with the 95% error bands estimated with the method

**Figure 1:** *The VAR-Based Impulse Responses of Aggregate Variables to Monetary Shocks*




**Note:** In the figure, the solid line indicates the estimated point-wise impulse responses. The area between the dashed lines shows the 95% confidence interval estimated with the method suggested by Sims & Zha (1999).

proposed by Sims & Zha (1999). Regarding the effect of the expansionary policy shock, as shown in the figure:

- The point estimates suggest the output gap stays below its pre-shock level for four years after the shock. In addition, in the first two years, the 95% confidence bands indicate, the fall in the output gap is statistically significant with a trough occurring after about one and half years following the shock.

- Inflation is above its pre-shock level in the early periods following the shock but rises over time below the pre-shock level. The trough realizes after about three years according to the estimated impulse responses. It is notable that my finding of a delayed effect from the interest shock on inflation is a general finding for the United States economy as noted by Woodford (2003).

- The real wage falls after the shock, suggesting nominal wages fall relative to nominal prices following the expansionary shock.

- Lastly, the federal funds rate remains above its pre-shock level for about three years following the shock according to its point impulse response estimates.

## 2.2 Sectoral Price Responses after Interest Shocks

To study sectoral price responses following an unanticipated 1% increase in the federal funds rate, I consider again the same VAR in (2.1) but add the annualized percentage change in a sectoral price to the vector of variables($\Upsilon_t$) in (2.2). This variable is denoted by $\pi_{it}$ and is measured as $4 \times (lnp_{it} - lnp_{it-1})$.

$$\Upsilon_t = \left[ \begin{array}{ccccc} y_t - y_t^n, & \pi_t, & \pi_{it}, & w_t, & R_t \end{array} \right] \tag{2.3}$$

Two identifying assumptions for isolating exogenous interest shocks in (2.3) are worth mentioning. First, the Federal Reserve observes sectoral price movements before setting the federal funds rate. Second, there is at least a quarter lag in the response of sectoral prices to the federal funds rate shocks. The latter is consistent with with the aforementioned assumption that there is a quarter delay in the aggregate price level's response to the federal funds rate shock. Were the sectoral prices assumed to respond contemporaneously to the shock while the aggregate price level was not, the analysis would be internally inconsistent.

Figure 2 shows the impulse responses of the *sectoral* and *aggregate price levels* to an unanticipated 1% increase in the federal funds rate.[2] Sectoral price level impulse responses represent the price level responses for 125 Personal Consumption Expenditure (PCE) categories for which I have an estimate of the frequency of price changes.

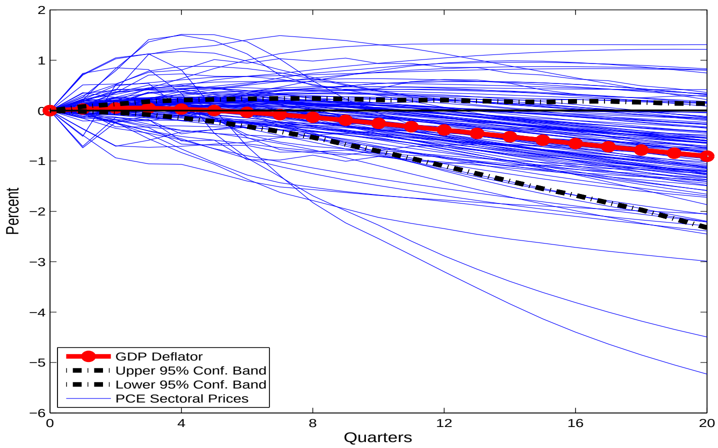
A crucial finding in Figure 2 is that an unanticipated change in the federal funds rate produces relative price effects in the United States. The existence of such effects requires only the lowest and highest sectoral price level responses to differ significantly. A stronger condition is met in Figure 2. Indeed, not only the highest sectoral price level responses differ radically from the lowest ones, but they are also outside the 95% confidence bands for the impulse responses of the aggregate price level.

It is also notable that initially, positive and negative sectoral responses are distributed evenly. This results in an initial muted response of the aggregate price level. Following this phase, the sectoral prices' responses are predominantly negative. As a consequence, the aggregate price level shows a decline following the initial phase. It is worth mentioning that these findings are in conformity with the findings in Boivin, Giannoni & Mihov (2009) who use the factor augmented vector autoregression (FAVAR) approach to study the sectoral price responses to a federal funds rate shock. They advocate their method by showing that a contractionary interest rate shock results in a fall in most sectoral prices and that there is no evidence of a "price puzzle" when the FAVAR approach is used. However, Hanson (2004) finds that the "price puzzle" is mainly associated with the 1959-1979 sample period, and that evidence of a "price puzzle" is weak during the 1976-2005 period which Boivin, Giannoni & Mihov (2009) consider. Whether the FAVAR approach alleviates the puzzle or not is uncertain when the data sample is extended back to 1959. It is notable that even if my data sample covers the period in which Hanson (2004) finds strong evidence of the puzzle, my results do not indicate a "price puzzle".

How strong is the association between sectoral price level responses

[2] Since both $\pi_t$ and $\pi_{it}$ are measured as four times the difference in the log of price levels between two periods, to obtain impulse responses for sectoral and aggregate price levels ( denoted with $lnp_{it}$ and $lnP_t$, respectively), cumulative impulse responses for $\pi_t$ and $\pi_{it}$ are obtained, which are then scaled down by four.

**Figure 2:** *The Impulse Responses of the Price Levels of PCE Categories to an Unanticipated 1% Increase in the Federal Funds Rate Shocks*




**Note:** In the figure, the thick solid line marked with circles shows the aggregate price responses following an unanticipated 1% increase in the federal funds rate whose 95% confidence intervals are marked by the thick dashed lines. The thin solid lines, on the other hand, display the sectoral impulse responses to the same shock.

in Figure 2 and sectoral frequency of price changes? Is a higher frequency of price change in a sector is associated with a higher or a lower price level response in periods? To answer these questions, I first define the frequency of price changes in sectors and describe my data on the frequency of price changes. The frequency refers to the percentage of firms that adjust their prices in a quarter. Our monthly frequency of price changes data for the United States comes from Nakamura & Steinsson (2008). In contrast to the frequency of price changes in Bils

& Klenow (2004), where only the frequency of price changes including sales are reported, using the frequency of price changes data in Nakamura & Steinsson (2008) has the advantage that the frequencies in sectors are reported for both non-sales price changes and price changes including sales. Since Nakamura & Steinsson (2008) find that the frequency of price changes including sales in some sectors differs radically from that of non-sales price changes to a great margin, it is important to check the robustness of my results in this paper for non-sales price changes and price changes including sales.

The frequency of price changes is estimated for *entry level items* (ELIs) of CPI in Nakamura & Steinsson (2008). However, for many components of CPI, price series for the disaggregated sectors are not available for earlier periods, prior to 1970s. In contrast, sectoral price indexes can be obtained from 1959 for the bulk of the Bureau of Economic Analysis' personal consumption expenditure (PCE) categories. For this reason, I use the PCE categories for estimation. However, since the frequency of price changes are not readily available for the PCE categories, the categories are mapped to the components of the CPI index in my analysis. To map ELIs in the CPI with the PCE categories, I use the mapping that Andrea Tambalotti made available.[3] If a PCE category is matched with only one component of CPI, the frequency of price changes in that PCE category is taken as that of the CPI component. If there are multiple ELIs that map with a single PCE category, the frequency of price changes in this PCE category is measured as the weighted average of the frequency of price changes in these CPI components, with the weights given as the 2000 CPI expenditures of the ELIs reported by Nakamura & Steinsson (2008).

Lastly, the frequencies of price adjustment within ELIs in Nakamura & Steinsson (2008) are reported as *monthly* percentages. The quarterly frequency of price changes in ELIs are estimated by using the method described in Tugan (2013). It is notable that the quarterly frequencies are found to differ substantially among the mapped PCE categories. The frequencies range from as low as 6.7% in intracity mass transit to 100% in net purchases of used motor vehicles.

Now, I study the association of the frequency of price changes with the impulse responses of sectoral prices to interest rate shocks. First,

[3] I am grateful to Andrea Tambalotti for sharing his mapping with me.

the impulse responses of sectoral prices to an unanticipated 1% federal funds rate shock have been obtained as in Figure 2. Next, for each period, the correlation between the frequency of price changes in sectors and their impulse responses is estimated. Figure 3 demonstrates these correlations and the corresponding 95% confidence bands. The correlations reveal that if prices change more frequently in a sector, it is more likely that prices in that sector increase after a contractionary interest rate shock during the first year following the shock. This finding holds whether or not the frequency of price changes includes non-sales price changes (See Panel (a) and Panel (b) of Figure 3). After one year following a contractionary monetary shock, on the other hand, a higher frequency of price change in a sector is associated with a lower impulse response. It is notable that the correlations of the impulse responses with the frequency of non-sales price changes and with the frequency of price changes *including sales* are rather similar. Lastly, the fact that the value of zero is contained in the confidence intervals for the correlations indicates that one may not reject the hypothesis of no correlation between the frequency of price changes and the impulse responses of sectoral prices following a contractionary federal funds rate shock. In other words, the price responses after a contractionary shock in sectors are only weakly associated with the fraction of firms in the sector that change their prices in a quarter.

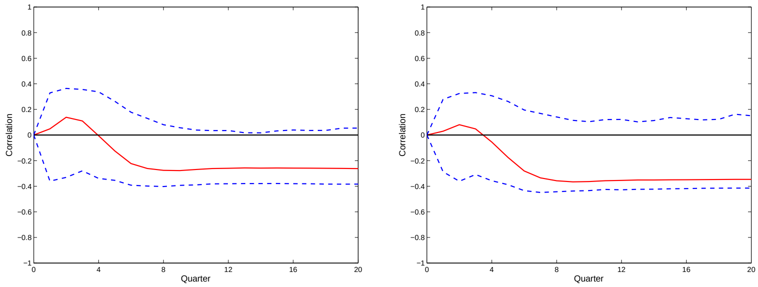
These findings contrast with those in Bils, Klenow & Kryvtsov (2003) who find that there is an anomaly in the relative price movements following an unanticipated change in the federal funds rate. Indeed, the price of the flexible-price category rises *significantly* relative to that of the sticky-price category in the first eight months following a contractionary interest shock. They reason there are two possible explanations for this finding: either the sticky price models are incapable of explaining relative price movements following the exogenous monetary shocks or else inferred monetary shocks are not orthogonal to persistent price shocks in the flexible- and sticky-price categories. I discuss the Bils, Klenow & Kryvtsov (2003) model in detail in Appendix A and offer an explanation for the contrasting findings in this paper and theirs.

In the next section, I aim to explain the empirical findings in this section with the three DSGE models. In the first model (*one-sector model*), the fraction of firms that may change their prices in a quar-

**Figure 3:** *Correlations of $\lambda_i$ with the Impulse Responses of $P_i$ to an Unanticipated 1% Increase in $R_t$*

**(a)** *With the Frequency of Non-Sales Price Changes*

**(b)** *With the Frequency of Price Changes Including Sales*




**Note:** The solid lines display the correlations of frequency of price changes in a sector with the impulse responses of sectoral prices to an unanticipated 1% increase in the federal funds rate at each quarter. The 95% confidence intervals for these correlations are shown with dotted lines and are estimated using the block-bootstrap method explained in Appendix B.

ter in all sectors *after an interest rate shock* are assumed to be the same in all sectors. In this model, the frequency of price changes in the economy is approximated by the median frequency of price changes in all sectors. It is notable that the assumption of the same frequency of price changes among sectors in the *one-sector model* does not necessarily contradict with the finding in Bils & Klenow (2004) and Nakamura & Steinsson (2008) that the distribution of frequency of price changes among sectors is wide in the United States. Indeed, while the frequency of price changes may differ largely among sectors for *sector-specific* and *other types of shocks*, they are the same for *an interest rate shock*. In the second model (*multi-sector model with a* *symmetric cost structure*), sectors are allowed to differ only in the frequency of price changes after *an interest rate shock*. In the last model, (the *multi-sector model with an asymmetric cost structure*), sectors differ not only in the frequency of price changes but also in the cost structure. The performance of these models in explaining the weak association of the frequency of price changes with impulse responses of sectoral prices after an interest rate shock in the economy is then evaluated. The findings are in favor of the *multi-sector model with an asymmetric cost structure*.

# 3 Theoretical Models

In this section, I consider a variant of the theoretical model in Tugan (2013). Since the model environment is discussed in detail in Tugan (2013), only a summary of main features of the model and the dynamic equations needed to solve the DSGE models are stated here:

- Price setting is staggered along the lines of Calvo (1983).
- Wage setting is staggered along the lines of Erceg, Henderson & Levin (2000).
- When the optimization signal is not received by firms and workers, wages and prices are set according to the backward-looking indexation rule.
- There is habit persistence in consumption.
- It is assumed that consumption decisions are made and prices and wages are set one period before observing the interest rate shocks.
- Firms are obliged to pay their wage bill in advance. As a consequence, when the monetary authority decides to introduce an unanticipated increase in the interest rate, real marginal costs may rise despite a fall in output accompanying the contractionary shock.
- The one- and multi-sector models differ only in the assumption regarding the frequency of price changes after an interest rate

shock. In the one-sector model, the frequency of price changes in all sectors is assumed to be homogenous. In the multi-sector model, on the other hand, there is a heterogeneity in price setting among sectors. As a matter of fact, in some sectors, prices change more frequently than in others.

## 3.1 The Structural Equations in the Models

Now, I state the main equations of the model. Let the hat over variables, $\hat{R}_t$, $\pi_{t+1}$ and $\varphi$ denote the log-deviation of the variables from their corresponding steady states; the nominal interest rate; inflation in prices; and, the intertemporal elasticity of substitution, respectively. The IS equation is given by:

$$E_{t-1}\hat{x}_t = E_{t-1}\hat{x}_{t+1} - \varphi E_{t-1}\Big(\hat{R}_t - \pi_{t+1}\Big) \tag{3.1}$$

where $\hat{x}_t$ is defined as:

$$\hat{x}_t = \bigg(\Big(\hat{y}_t - b\hat{y}_{t-1}\Big) - b\beta E_t\Big(\hat{y}_{t+1} - b\hat{y}_t\Big)\bigg) \tag{3.2}$$

In (3.2), $\hat{y}_t$, $b$ and $\beta$ denote the log-change in output; the habit formation parameter; and, the discount factor, respectively.

The second equation in the models is the wage inflation equation ($\pi_t^w$):

$$\pi_t^w - \gamma_w \pi_{t-1} = \xi_w E_{t-1}\bigg(\omega_w \hat{y}_t + \varphi^{-1}\hat{x}_t - \hat{w}_t\bigg) + \beta E_{t-1}\bigg(\pi_{t+1}^w - \gamma_w \pi_t\bigg) \tag{3.3}$$

In (3.3), $\omega_w$ and $\gamma_w$ denote the elasticity of real wages paid for the number of hours worked with respect to output changes for a constant marginal utility of real income and the backward-looking indexation parameter in wages, respectively. Lastly, letting $1 - \alpha_w$, $\sigma_{\mathcal{H}}$ and $\theta_w$ denote the probability of receiving a wage change signal by a differentiated labor type; the Frisch-elasticity of labor; and, the wage elasticity of substitution among differentiated labor types, respectively, the wage stickiness parameter $\xi_w$ in (3.3) can be written as:

$$\xi_w = \frac{(1-\alpha_w)}{\alpha_w}\frac{1-\alpha_w\beta}{(1+\theta_w\sigma_{\mathcal{H}}^{-1})}$$

The third equation in the models is the monetary policy rule. The monetary authority is assumed to control the interest rate and implements the following Taylor Rule to stabilize the economy:

$$R_t = \rho_R \times R_{t-1} + [a_\pi \pi_t + a_y(y_t - y_t^n)] - \rho_R \times [a_\pi \pi_{t-1} + a_y(y_{t-1} - y_{t-1}^n)] + \epsilon_t \tag{3.4}$$

where $y_t^n$ and $\epsilon_t$ denote the potential output and the shock in monetary policy, respectively. The calibrated values for $\rho_R$, $a_\pi$ and $a_y$ are given as 0.92, 1.24 and 0.33, respectively. These calibrated values are based on Rudebusch (2002).

## 3.2 The Econometric Method

Since the number of sectors for which the frequency of price changes data is available is quite large, it is impractical to solve the DSGE models by considering each individual sector. To circumvent this problem, as in Tugan (2013), I reduce the number of sectors in the model to 10 by including each sector in one of the percentiles of the frequency of price changes and approximating the frequency of price changes in a sector with the median frequency in the percentile group where that sector is contained. It is notable that only the frequency of price changes *including sales* are considered in calibration. Since the calibration of $f_k$, $\alpha_{pk}$ and some other parameters of the models are extensively discussed in Tugan (2013), I skip describing the calibration method here and only discuss the estimation method for the free parameters of the models. Let $\mathcal{P}$ denote the vector of free parameters to be estimated. In the models, $\mathcal{P}$ contains 7 parameters:

$$\mathcal{P} = [\varphi, \sigma_{\mathcal{H}}, \theta_p, \theta_w, b, \gamma_p, \gamma_w]$$

where $\theta_p$ and $\gamma_p$ denote the price elasticity of substitution for sectoral goods and the backward-looking indexation parameter in prices, respectively.[4]

$$\hat{\mathcal{P}}(\hat{A}_T) = \arg\min_{\mathcal{P}} (\hat{h}_T - f(\mathcal{P}))' \hat{A}_T' \hat{A}_T (\hat{h}_T - f(\mathcal{P})) \tag{3.5}$$

[4]See (3.1) for the definition of $\varphi$, (3.2) for the definition of $b$, and (3.3) for the definitions of $\sigma_{\mathcal{H}}$, $\theta_w$ and $\gamma_w$.

where $\hat{A}_T$ and $f(\mathcal{P})$ show the weighting matrix and the model-based impulse responses and correlations for a given parameter vector $\mathcal{P}$. Lastly, $\hat{h}_T$ stands for the vector of estimated VAR-based impulse responses and correlations and is given by:

$$\hat{h}_T = \left| \ \mathcal{C}^{y_t - y_t^n}_{1,20}, \quad \mathcal{C}^{\pi_t}_{1,20}, \quad \mathcal{C}^{w_t}_{1,20}, \quad \mathcal{C}^{R_t}_{1,20} \quad \rho_{\mathcal{C}^{lnp_{it}}_{1,20}, \lambda_i} \ \right|' \tag{3.6}$$

where $\mathcal{C}^{Z}_{1,20}$ denotes the impulse responses of the variable $Z$ to an unanticipated 1% rise in the federal funds rate between the $1^{th}$ and $20^{th}$ quarters following the shock as shown in Figure 1, and $\rho_{\mathcal{C}^{lnp_{it}}_{1,20}, \lambda_i}$ represents the correlations of the frequency of price changes in sectors($\lambda_i$) with the impulse responses of sectoral prices between the $1^{th}$ and $20^{th}$ quarters ($\mathcal{C}^{lnp_{it}}_{1,20}$) as in Figure 3. Lastly, $T$ stands for the sample size of the data used to estimate the VAR-based impulse responses. As a weighting matrix, I use the diagonal matrix whose diagonal elements are given by the inverse of standard errors of each term in $\hat{h}_T$. This matrix ensures more precisely estimated VAR-based correlations and impulse responses have larger weights when choosing parameters in (3.5).

## 3.3 Results

Before presening the aggregate and disaggregated model-based dynamics following an unanticipated 1% increase in the federal funds rate and comparing the outcomes in the models with the VAR-based dynamics, I first report the structural parameter estimates in the models in Table 1. Woodford (2003) shows that the backward-looking indexation rule in prices ($\gamma_p$) and the habit persistence in consumption ($b$) induce hump-shaped dynamics after a monetary shock. Hence, high estimates for such parameters in Table 1 can be related to the hump-shaped dynamics of consumption in Figure 1. Low estimates for the intertemporal elasticity of substitution ($\varphi$) and the Frisch-elasticity of labor ($\sigma_H$) are consistent with the estimates reported in Hall (1988) and Boldrin, Christiano & Fisher (2001). The estimated value for the backward-looking indexation rule in wages ($\gamma_w$) is close to its upper theoretical limit of one, as Christiano, Eichenbaum & Evans (2005) assume. Lastly, the calibrated values for the price elasticity

**Table 1:** *Estimates of Structural Parameters*

| | (a)<br>One-Sector Model | (b)<br>Multi-Sector Model<br>With *Symmetric* Cost Structure |
|---|---|---|
| $\varphi$ | 0.012 | 0.025 |
| $\sigma_{\mathcal{H}}$ | 0.740 | 0.001 |
| $\theta_p$ | 147.215 | 8.723 |
| $\theta_w$ | 19.970 | 6.030 |
| $b$ | 0.800 | 0.641 |
| $\gamma_p$ | 0.999 | 0.915 |
| $\gamma_w$ | 0.999 | 0.157 |
| Obj. Func. | 117.27 | 246.02 |

**Note:** *Obj. Func.* indicates the estimated value for the minimization problem discussed in (3.5). A lower value of *Obj. Func.* indicates a more successful model for accounting for aggregate dynamics and correlations between the frequency of price changes and sectoral price dynamics after an unanticipated 1% increase in the federal funds rate.

of substitution for sectoral goods ($\theta_p$) and the wage elasticity of substitution among differentiated labor types ($\theta_w$) in the literature vary enormously. Our estimates lie within the range of those calibrated values.

### 3.3.1 Aggregate Dynamics

Figure 4 and Figure 5 show the impulse responses of the output gap, inflation, the real wage and the federal funds rate over five years after a 1% contractionary shock in the federal funds rate in the *one-sector* *model* and the *multi-sector model with symmetric cost structure.* It is evident from these figures that the impulse responses of the aggregate variables are similar in these models. Output is constant on impact following the shock, by construction. Starting with the first period, output falls. This emanates from the fact that a higher interest rate reduces consumption by making saving more desirable. The *one-sector model* fails to account for increased inflation following the shock. The fall in real wage in the models is the product of two factors. First, a fall in output lowers nominal wage demand. To explain this, note that less effort is needed when output falls, and since the disutility from working is a convex function of effort, workers lowers their nominal wage if firms demand less effort. Second, an increase in prices contributes to a fall in real wage in earlier periods. Excessive fall in real wage in the *multi-sector model with symmetric cost structure* can be accounted for by the second factor.

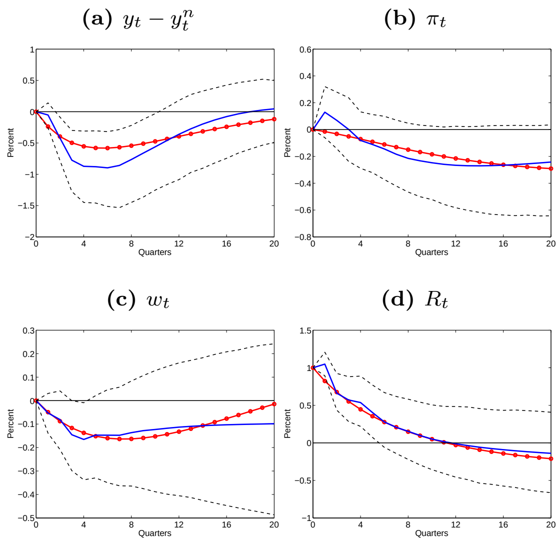
### 3.3.2 Model- and VAR-Based Correlations

The dynamics displayed in Figure 4 and Figure 5 reveal the DSGE models have similar predictions regarding the impulse responses of aggregate variables. In what they differ quite substantially is their prediction of the correlations of the frequency of price changes with the impulse responses of sectoral prices ($\rho_{\mathcal{C}_{1,20}^{lnp_{it}},\lambda_i}$).

Figure 6 shows the correlations of the frequency of price changes ($\lambda_i$) with the impulse responses of sectoral prices after an unanticipated 1% increase in the federal funds rate ($\mathcal{C}_{1,20}^{lnp_{it}}$). In the *one-sector model*, the correlations have to be zero by definition since all sectors have the same price response to the interest rate shock.[5] In the *multi-sector model with symmetric cost structure*, since it is assumed that sectoral prices are unresponsive to the shock on impact, the impact correlation is zero as illustrated in Panel (b) of the figure. Following

[5]It is notable that in the one-sector model, it is assumed that the *measured frequency of price changes* differs among sectors for sector-specific and aggregate shocks, except the federal funds rate shock. For the federal funds rate shock, on the other hand, it is assumed that the fraction of firms in all sectors that change their prices are the same. This implies when a contractionary interest rate shock occurs, prices in all sectors respond in the same way. Consequently, by construction, the correlations between *the measured frequency of price changes* and $\mathcal{C}_{1,20}^{lnp_{it}}$ are equal to zero in the *one-sector model.*

**Figure 4:** *Impulse Responses to an Unanticipated 1% Rise in* $R_t$
*(One-Sector Model)*

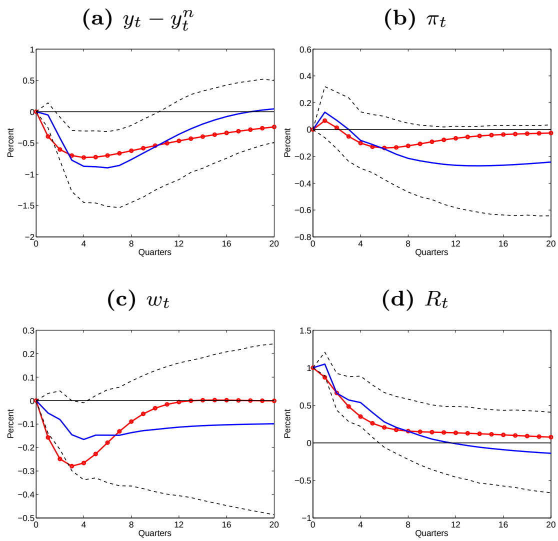



**Note**: The solid lines in panels show the VAR-based impulse responses and the area between dotted lines indicate the 95% confidence intervals estimated with the method suggested by Sims & Zha (1999). The solid lines marked with circles represent the dynamic responses of the variables as predicted by the model.

the impact period, the correlations are positive for two quarters. This is unconventional as it implies that a tightening of monetary policy shock leads to a higher price response in sectors where firms change prices frequently than sectors where firms change prices infrequently. Starting in the third quarter following the contractionary shock, the correlations become negative. This suggests that a higher frequency of price change in a sector is associated with a lower price response in these periods.

It is notable that compared to the *multi-sector model with sym-*

**Figure 5:** *Impulse Responses to an Unanticipated 1% Rise in* $R_t$
*(Multi-Sector Model with* Symmetrical *Cost Structure)*

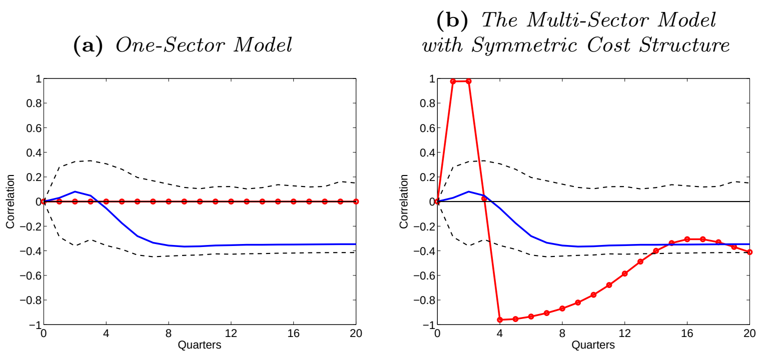



**Note**: The solid lines in panels show the VAR-based impulse responses and the area between dotted lines indicate the 95% confidence intervals estimated with the method suggested by Sims & Zha (1999). The solid lines marked with circles represent the dynamic responses of the variables as predicted by the model.

*metric cost structure*, the one-sector model performs much better in explaining the correlations. However, this model may not explain the rich set of sectoral price dynamics following the interest rate shock displayed in Figure 2. The *multi-sector model with symmetric cost structure*, on the other hand, can explain the wide distribution of the responses of sectoral prices to such a shock; yet, this model is unable to explain the correlations.

Why does the *multi-sector model with symmetric cost structure* fail to explain $\rho_{\mathcal{C}_{1,20}^{lnp_{it}},\lambda_i}$? To answer this question, I now detail the price-

**Figure 6:** *Model- and VAR-Based Correlations of* $\lambda_i$ *with* $\mathcal{C}_{1,20}^{lnp_{it}}$




**Note**: The solid lines show the VAR-based correlations and the area between dotted lines indicate the 95% confidence intervals estimated with the method suggested by Sims & Zha (1999). The solid lines marked with circles represent the dynamic responses of the variables as predicted by the model.

setting behavior of the firms in each model. It is assumed that firms set prices before observing shocks to the interest rate. In each sector, firms optimize their prices only when a price-change signal is received. The fraction of firms which receive this signal is different in each sector and is given by $\lambda_i = 1 - \alpha_{ip}$ for the sector $i$. It is well known that this fraction is equal to the probability of receiving a price change signal in each period in the Calvo (1983) model.

Let $p_{it-1}(i')$ and $P_t$ denote the last period price of the good produced by the firm $i'$ in *Sector* $i$ in the last period and the price of the composite consumption good, respectively. When no such signal is received, firms are assumed to set their prices according to the following partial adjustment backward-looking indexation rule:

$$\tilde{p}_{it}(i') = p_{it-1}(i') \left( \frac{P_{t-1}}{P_{t-2}} \right)^{\gamma_p} \tag{3.7}$$

where the tilde over $p$ denotes the price set according to the backward-looking indexation rule and $\gamma_p$ shows the backward-looking indexation parameter. If $0 < \gamma_p < 1$, there is partial backward-looking indexation in the economy.

When a firm is capable of setting an optimal price, it sets $p_{it}(i')^*$ that maximizes:

$$E_{t-1} \left( \sum_{s=0}^{\infty} \alpha_{ip}^s Q_{t,t+s} \Pi_{it+s}(i') \right) \tag{3.8}$$

where $Q_{t,t+s}$ is the stochastic discount factor between the period $t$ and $t+s$ and is given by:

$\Pi_{it+s}(i')$, on the other hand, shows the profit of the firm $i'$ in the sector $i$ and is given by:[6]

$$\Pi_{it+s}(i') = p^*_{it+s,t}(i) y_{it+s}(i') - TC_{it+s}(i') \tag{3.9}$$

where $p^*_{it+s,t}(i')$ shows the price set in period $t+s$ by the firm that received a price-change signal at the period $t$ and does not have an opportunity to set an optimal price between $t$ and $t+s$. Due to the backward-indexation rule, one can write $p^*_{it+s,t}(i')$ as:

$$p^*_{it+s,t}(i') = p^*_{it}(i') \chi^p_{t,t+s} \tag{3.10}$$

where

$$\chi^p_{t,t+s} = \begin{cases} \prod_{k=1}^{s} \left( \frac{P_{t+k-1}}{P_{t+k-2}} \right)^{\gamma_p} & \text{if } s \geq 1 \\ 1 & \text{if } s = 0 \end{cases} \tag{3.11}$$

$TC_{it}(i')$ in (3.9), on the other hand, denotes the total cost of the firm. In the *one-sector model* and the *multi-sector model with symmetric cost structure*, all firms have the same total cost structure. Since

[6] It is notable that since the firms are assumed to respond the monetary shocks with a one period delay, they have to condition the optimum price based on the information till the period $t-1$ rather than the period $t$. Correspondingly, $E_{t-1}$ appears in the objective function, rather than $E_t$ in (3.8).

firms are assumed to pay the wage bill in advance, $TC_{it}(i')$ for these models can be written as:

$$TC_{it}(i') = R_t W_t L_{it}(i') \tag{3.12}$$

where $W_t$ and $L_{it}(i')$ denote the nominal aggregate wage and the labor demanded by the firm $i'$ in the sector $i$, respectively.

The optimality condition in (3.8) for $p^*_{it}(i')$ can be expressed as:

$$E_{t-1}\left(\sum_{s=0}^{\infty} \alpha^s_{pi} Q_{t,t+s} \frac{d\Pi_{it+s}(i')}{dp^*_{it}(i')}\right) = 0 \tag{3.13}$$

Using 3.13, one can show the sectoral price inflation ($\pi_{it}$) in the *multi-sector model with symmetric cost structure* evolves according to the following equation:

$$\begin{aligned} \pi_{it} - \gamma_p \pi_{t-1} &= -\xi_{ip}(1+\omega_p\theta_p)E_{t-1}\left(\hat{P}_{it} - \hat{P}_t\right) \\ &+ \xi_{ip}E_{t-1}\left(\hat{R}_t + \hat{w}_t + \omega_p \hat{y}_t\right) + \beta E_{t-1}\left(\pi_{it+1} - \gamma_p \pi_t\right) \end{aligned} \tag{3.14}$$

where $\xi_{ip}$ is the stickiness parameter in each sector and is defined as:

$$\xi_{ip} = \frac{1-\alpha_{ip}}{\alpha_{ip}} \frac{1-\beta\alpha_{ip}}{1+\omega_p\theta_p}, \ \omega_p = \frac{1-\kappa}{\kappa} \tag{3.15}$$

In (3.15), $\omega_p$ and $\kappa$ denote the elasticity of prices with respect to the supply of goods when interest rate and wages paid for composite hours of work stay the same and the reciprocal of the output elasticity of labor demand, respectively. The aggregate inflation equation in the *multi-sector model* is a weighted average of sectoral inflation in the economy:

$$\pi_t = \sum_{i=1}^{M} f_i \pi_{it} \tag{3.16}$$

where $f_i$ is the sector's share of aggregate consumption expenditure at the steady state.

The equation for $\pi_t$ in the *one-sector model*, on the other hand, is given by:

$$\pi_t - \gamma_p \pi_{t-1} = \xi_p E_{t-1}\left(\hat{R}_t + w_t + \omega_p \hat{y}_t\right) + \beta E_{t-1}\left(\pi_{t+1} - \gamma_p \pi_t\right) \tag{3.17}$$

$$\xi_p = \frac{1-\alpha_p}{\alpha_p}\frac{1-\beta\alpha_p}{1+\omega_p\theta_p} \tag{3.18}$$

where $1-\alpha_p$ denotes the frequency of price changes in the model economy. It is measured as the median frequency of price changes in the United States.

Here, I aim to explain the strong positive correlations of the frequency of price changes with the sectoral price responses following the shock ($\rho_{\mathcal{C}^{lnp_{it}}_{1,20},\lambda_i}$) in earlier periods and strong negative correlations in later periods. I explain this by studying how firms in each sector set prices when they have a chance to optimize.

First, one can show from (3.13) that the following equation holds:

$$E_{t-1}\left(\sum_{s=0}^{\infty}\alpha_{ip}^{s}Q_{t,t+s}\chi^{p}_{t,t+s}\left(\frac{P_{t+s}}{p^{*}_{it+s,t}(i')}\right)^{1+\theta_p} y_{t+s}\left(\frac{p^{*}_{it+s,t}(i')}{P_{it+s}}\frac{P_{it+s}}{P_{t+s}} - \mu_p\frac{S_{it+s}(i')}{P_{it+s}}\frac{P_{it+s}}{P_{t+s}}\right)\right) = 0 \tag{3.19}$$

where $\mu_p \geq 1$ (i.e. $\theta_p \geq 1$) shows the steady-state markup. $S_{jt+s}(j')$ denotes the marginal cost of the firm. Letting $y_{it}(i')$ be the output of the firms, $S_{jt+s}(j')$ is defined as:

$$S_{it+s}(i') = \frac{\partial TC_{it}(i')}{\partial y_{it}(i')} \tag{3.20}$$

Log-linearizing (3.19) yields:

$$E_{t-1}\sum_{s=0}^{\infty}(\beta\alpha_{ip})^{s}\left[\hat{p}^{*}_{it}(i') - \hat{P}_{it+s} + \hat{P}_{it+s} - \hat{P}_{t+s} + \hat{\chi}^{p}_{t,t+s} - \left(\hat{R}_{t+s} + \hat{w}_{t+s} + \omega_p\hat{y}_{it+s}(i')\right)\right] = 0 \tag{3.21}$$

where $\hat{\chi}^{p}_{t,t+s}$ is the log-deviation of $\chi^{p}_{t,t+s}$ from its steady state and is given by:

$$\hat{\chi}^{p}_{t,t+s} = \begin{cases} \gamma_p\pi_t + \gamma_p\pi_{t+1} + \cdots + \gamma_p\pi_{t+s-1} & \text{if } s \geq 1 \\ 0 & \text{if } s = 0 \end{cases} \tag{3.22}$$

Using the approximation that $\hat{P}_{t+s} = \hat{P}_t + \pi_{t+s} + \sum_{k=0}^{s-1} \pi_{t+k} - \pi_t$ for $s \geq 1$, (3.21) can be restated as:

$$\begin{aligned} \hat{p}^*_{it}(i') &= E_{t-1}\hat{P}_t + (1-\beta\alpha_{ip})E_{t-1}\sum_{s=1}^{\infty}(\beta\alpha_{ip})^s \left(\pi_{t+s} + (1-\gamma_p)\sum_{k=0}^{s-1}\pi_{t+k} - \pi_t\right) \\ &+ (1-\beta\alpha_{ip})E_{t-1}\hat{s}_{it} + (1-\beta\alpha_{ip})E_{t-1}\sum_{s=1}^{\infty}(\beta\alpha_{ip})^s \hat{s}_{it+s} \end{aligned} \tag{3.23}$$

where $\hat{s}_{it+s}$ shows the log-deviation of the real marginal cost of the firm from its steady state and is given by:

$$\hat{s}_{it+s} = \hat{R}_{t+s} + \hat{w}_{t+s} + \omega_p \hat{y}_{it+s}(i') \tag{3.24}$$

(3.23) can be rewritten as:

$$\begin{aligned} \hat{p}^*_{it}(i') &= E_{t-1}\hat{P}_t - \gamma_p\beta\alpha_{ip}E_{t-1}\pi_t + E_{t-1}\sum_{s=1}^{\infty}(\beta\alpha_{ip})^s((1-\gamma_p\beta\alpha_{ip})\pi_{t+s}) \\ &+ E_{t-1}\hat{s}_{it} + E_{t-1}\sum_{s=1}^{\infty}(\beta\alpha_{ip})^s(\hat{s}_{it+s} - \hat{s}_{it+s-1}) \end{aligned} \tag{3.25}$$

Two points must be emphasized regarding (3.25). First, as $\gamma_p$ increases, firms give less importance to inflation in subsequent periods. When $\gamma_p = 1$, they set prices such that importance given to inflation in subsequent periods is minimized. An intuitive explanation can be given for this: The fact that the prices are optimized only if the Calvo signal is received leads firms to take preemptive measures against expected inflation in subsequent periods. When the degree of backward-looking indexation is high in an economy, firms are able to change prices by taking into account inflation realized in the previous period even if prices are not optimized. This results in a decrease in the degree of firms' preemptive measures against expected inflation in the subsequent periods.

Second, when expected real costs are higher in subsequent periods than today, the percentage increase in prices is higher than the percentage change in the current real marginal cost, holding fixed expected inflation in subsequent periods.[7] Christiano, Eichenbaum & Evans (2005) refer to this as firms "front-load" for the expected real

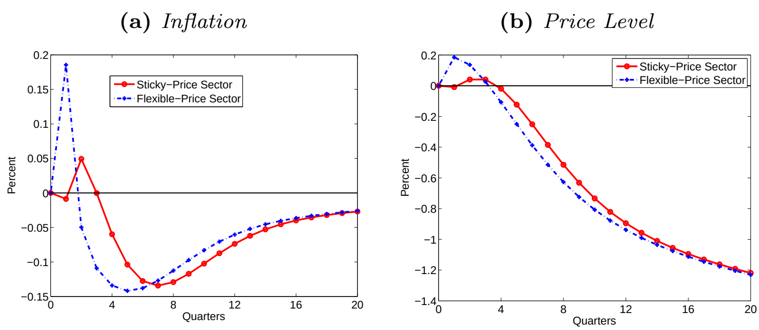
[7] It is notable that since the effect of $\hat{p}^*_{it}(i')$ on $E_{t-1}\hat{P}_t$ is negligible, when setting prices, the firms treats $E_{t-1}\hat{P}_t$ as constant. Hence, whether the percentage change in the optimized prices outweighs that of the real marginal cost today depends entirely on the statement above.

cost increases in subsequent periods in which the chance to optimize their prices is uncertain. The "front-loading" is most relevant for firms in a sector where the frequency of price changes is lower ($\alpha_{ip}$ is higher) since a higher frequency of price change discounts the importance of real marginal costs in subsequent periods on prices set by a firm when a Calvo signal is received. For example, consider an unanticipated rise in interest rate. A rise in today's marginal costs is likely because of the working capital channel in the model. Yet, as interest rate returns to its undistorted level and output and wages decrease, marginal costs are bound to fall in subsequent periods. In the flexible-price sector where firms can optimize prices often, marginal costs today have a decisive effect on prices set. For firms in the sector where price flexibility is low, on the other hand, the extent that marginal costs in subsequent periods are taken into account in price setting is much larger. I illustrate the "front loading" argument in Figure 7. In this figure, the sticky- and flexible-price sectors are defined as the percentile groups with the lowest and highest frequency of price changes among 10 groups in the model, respectively. It is evident from this figure that a contractionary monetary shock results in an initial fall in prices in the sticky-price sector and an initial rise in prices in the flexible-price sector. This results in a positive correlation between the frequency of price changes and sectoral price responses in the early periods following the contractionary monetary shock.

In subsequent periods, firms' marginal costs fall markedly due to a persistent fall in the real wage, output and the interest rate. Because of a higher price flexibility in the flexible-price sector, prices in this sector fall more pronouncedly than those in the sticky-price sector during these periods. This explains the finding in Figure 6 that a higher frequency of price change is associated with a lower price response in the *multi-sector model with symmetric cost structure* in the third period and onwards.

As evident in Figure 6, while the *multi-sector model with symmetric cost structure* explains the correlations successfully in *qualitative terms*, an undesirable feature of this model is that the correlations predicted by the model are too high compared to those found in the data. This is a natural consequence of the fact that sectors in the model are assumed to be identical apart from their frequency of price changes. With this assumption, price responses in sectors in any period are or-

**Figure 7:** *The Front-Loading Argument*
*(The Multi-Sector Model with Symmetric Cost Structure)*

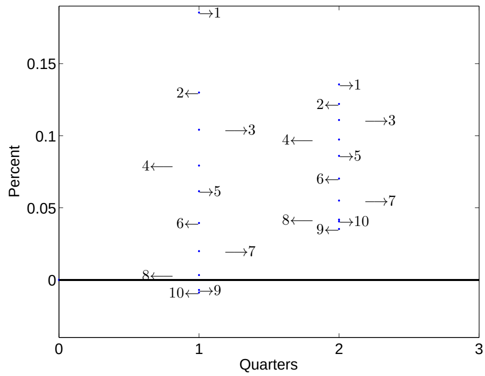



**Note**: The dot-dashed lines marked with a plus sign and the solid lines marked with circles show the model-based impulse responses of inflation and the price level in the flexible- and sticky-price sectors to a 1% contractionary shock to the federal funds rate, respectively.

dered to a large extent according to sectoral frequency of price changes after a contractionary interest rate shock. Figure 8 illustrates this. In Panel (a) of the figure, the model-based sectoral price responses in the first and second periods are shown for the *multi-sector model with symmetric cost structure*. Sectors are indicated by numbers and are ordered according to their frequency of price changes from highest to lowest. For example, *1* in the figure denotes the price response of the sector that has the highest frequency of price changes among sectors. As evident from Panel (a) of the figure, price responses in sectors in the first and second periods are largely ordered according to sectoral frequency of price changes. This causes the predicted correlations between the frequency of price changes and the sectoral price responses in the model to be exceedingly high compared to those in the data.

Briefly, the discussion in this section notes that the strong positive

**Figure 8:** *Model-Based* $\mathcal{C}_{2,3}^{lnp_{it}}$
*(The Multi-Sector Model with* ***Symmetric*** *Cost Structure)*




**Note**: The points with numbers inside the figures show the model-based price responses in each sector to a 1% contractionary shock in the federal funds rate in the first and second quarters. Sectors in the figure are ordered according to the frequency of price changes from highest to lowest. For example, *1* in the figure denotes the price response of the sector with the highest frequency of price changes in the first and second periods.

correlations between the frequency of price changes and sectoral price responses in the initial periods and the strong negative correlations in subsequent periods are a direct consequence of the front-loading argument. However, such strong correlations conflict with the low VAR-based correlations shown in Figure 3.

## 3.4 The Multi-Sector Model with *Asymmetric* Cost Structure

In this section, I show that when there is asymmetry in the cost structure of firms in different sectors, not only is it possible to account for the low correlations in the data, but one can also explain the wide distribution of sectoral price responses to the shock which is evident in Figure 2.

I consider a *multi-sector model with asymmetric output elasticity of labor demand* where sectors differ not only in terms of price flexibility but also in terms of their cost structure since production functions used by firms differ among sectors in this model. This contrasts sharply with the *multi-sector model with symmetric cost structure* where firms use the same production function. To explain this model, it is useful to first write the production function that firms use to produce their output ($y_{it}(i')$) in the *multi-sector model with symmetric cost structure*,

$$y_{it}(i') = Z_t H_{it}(i')^{\kappa} \quad 0 < \kappa < 1 \tag{3.26}$$

where $Z_t$ and $H_{it}(i')$ denote the technology level and the demand of the firm for the composite labor, respectively. Lastly, $\kappa$ denotes the reciprocal of the output elasticity of labor demand in sectors which is assumed to be identical among all sectors. In the *multi-sector model with asymmetric cost structure*, a differential $\kappa_i$ for each sector is considered:

$$y_{it}(i') = Z_t H_{it}(i')^{\kappa_i} \tag{3.27}$$

Except the sectoral inflation equation ($\pi_{it}$) and the nominal wage inflation equation ($\pi_t^w$), the structural equations are the same as in Section 3.1. The only change in the sectoral inflation equation is that $\kappa$ in (3.15) should be replaced by $\kappa_i$. $\pi_t^w$, on the other hand, is now given as:

$$\pi_t^w - \gamma_w \pi_{t-1} = \xi_w E_{t-1}\left(\sigma_H^{-1}\hat{H}_t + \varphi^{-1}\hat{x}_t - \hat{w}_t\right) + \beta E_{t-1}\left(\pi_{t+1}^w - \gamma_w \pi_t\right) \tag{3.28}$$

where $\hat{H}_t$ denotes the total composite labor demand. Let $\hat{Y}_{it}$ and $n_i$ stand for sectoral output and the sectoral weight in total output, respectively. Then, $\hat{H}_t$ can be written as:

$$\hat{H}_t = \sum_{i=1}^{10} \frac{n_i}{\kappa_i} \hat{Y}_{it} \tag{3.29}$$

where sectoral output is a function of sectoral relative price and total output:

$$\hat{Y}_{it} = -\theta_p \left( \hat{P}_{it} - \hat{P}_t \right) + \hat{Y}_t \tag{3.30}$$

### 3.4.1 Calibration for the Multi-Sector Model with *Asymmetric* Cost Structure

It is notable that in the *multi-sector model with symmetric cost structure*, the number of sectors is reduced to ten since it is impractical to solve the model if all disaggregated sectors, for which the frequency of price changes is available, are included. When grouping disaggregated sectors into ten groups, sectors are ordered by their frequency of price changes, and they are included in one of the ten groups. The frequency of price changes in a sector is then approximated by the median frequency of price changes in its group. Since sectors in the *multi-sector model with symmetric cost structure* only differ in their frequency of price changes and the frequency of price changes in all sectors contained in a group is approximated by the median frequency of price changes in that group, sectors within the same group must have the same sectoral inflation equation.

However, in the *multi-sector model with asymmetric cost structure*, grouping disaggregated sectors based only on the frequency of price changes may not be justified. This results from the fact that even when such sectors have a similar frequency of price changes, sectoral price dynamics following a monetary shock may be markedly dissimilar if their output elasticity of labor demand largely differs. Consequently, in the *multi-sector model with asymmetric cost structure*, both the frequency of price changes and labor shares in sectors are needed to solve the model. To calibrate these parameters, we first match 125 PCE categories, for which the frequency of price changes is available and whose price responses are shown in Figure 2, with the industries reported by Close & Shulenburger (1971).[8] If an industry is matched with only one PCE category, the frequency of price changes in that industry is taken as the one in the PCE category. If there are multiple PCE categories that match with a single industry, the frequency of

[8] It may be useful here to exemplify our matching. For example, the PCE categories "Tires" and "Accessories and parts" are matched with the industry of "Motor vehicles and equipment" in Close & Shulenburger (1971).

price changes in this industry is measured as the weighted average of the frequency of price changes in these PCE categories, with the weights given as the sum of the expenditure shares of the ELIs in 2000 that are mapped with the PCE categories in Section 2.2. Labor shares in industries are calibrated as those in 1948 reported by Close & Shulenburger (1971). Weights of each industry are calibrated as the sum of the weights of the PCE categories that match with the industry. However, some PCE categories may not be matched with an industry, causing the sum of the industries' weights to be less than one. Consequently, the weights of the industries need to be rescaled so that their sum is equal to one. Table 2 reports the calibrated values for industries' labor share, the frequencies of price changes and weight. It is notable that while the petroleum and air-transportation industries have virtually the same frequency of price changes,[9] they markedly differ in their labor shares.[10]

In Table 3, I report the structural parameter estimates in the *multi-sector model with asymmetric cost structure*. It is notable that the value of the objective function in the *multi-sector model with asymmetric cost structure* is lower compared the ones in the *one-sector model* and *multi-sector model with symmetric cost structure* as reported in Table 1, suggesting that the *multi-sector model with asymmetric cost structure* is the most successful in accounting for the aggregate dynamics and the correlations.

In Figure 9, I display the aggregate dynamics in the *multi-sector model with asymmetric cost structure*, after an unanticipated 1% increase in the federal funds rate, which are largely in conformity with the aggregate dynamics in the previous two models as shown in Figure 4 and Figure 5.

However, as evident in Figure 10, the correlations between the frequency of price changes and sectoral price responses in the *multi-sector model with asymmetric cost structure* differ markedly from those in the *multi-sector model with symmetric cost structure*. As a matter of fact, contrasting with the latter, the former can more successfully explain the low VAR-based correlations in the data.

[9] The frequency of price changes in the petroleum and air-transportation industries are 0.97 and 0.94 in the former and latter, respectively

[10] The labor shares in the petroleum and air-transportation industries are 0.32 and 0.88 in the former and the latter, respectively

**Table 2:** *Calibrated Parameters*
*(The Multi-Sector Model with* Asymmetric *Cost Structure)*

| Industry | Labor Share | Frequency | Weight |
|---|---|---|---|
| Food | 0.74 | 0.66 | 0.143 |
| Tobacco | 0.52 | 0.69 | 0.019 |
| Textiles | 0.80 | 0.64 | 0.002 |
| Apparel | 0.87 | 0.66 | 0.064 |
| Paper | 0.64 | 0.58 | 0.003 |
| Printing | 0.79 | 0.20 | 0.013 |
| Chemicals | 0.56 | 0.39 | 0.030 |
| Petroleum | 0.32 | 0.97 | 0.060 |
| Furniture | 0.82 | 0.51 | 0.023 |
| Fabricated metal | 0.78 | 0.37 | 0.009 |
| Electrical machinery | 0.77 | 0.50 | 0.019 |
| Transportation equipment and ordinance | 0.89 | 0.21 | 0.035 |
| Motor vehicles and equipment | 0.63 | 0.60 | 0.049 |
| Instruments | 0.79 | 0.25 | 0.001 |
| Miscellaneous manufacturing industries | 0.75 | 0.41 | 0.011 |
| Railroad trasportation | 0.82 | 0.56 | 0.001 |
| Local, suburban, highway passanger transportation | 0.87 | 0.14 | 0.005 |
| Water transportation | 0.87 | 0.65 | 0.001 |
| Air transportation | 0.88 | 0.94 | 0.014 |
| Trasporation services | 0.80 | 0.29 | 0.002 |
| Telephone and telegraph | 0.80 | 0.65 | 0.037 |
| Radio broadcasting and television | 0.83 | 0.34 | 0.014 |
| Electric, gas, and sanitary services | 0.55 | 0.71 | 0.067 |
| Wholesale trade | 0.71 | 0.27 | 0.005 |
| Retail trade | 0.6 | 0.38 | 0.008 |
| Hotels and other lodging places | 0.69 | 0.75 | 0.040 |
| Personal services | 0.64 | 0.12 | 0.028 |
| Miscellaneous business services | 0.7 | 0.18 | 0.136 |
| Automobile repair | 0.62 | 0.42 | 0.016 |
| Miscellaneous repair services | 0.48 | 0.22 | 0.001 |
| Motion pictures | 0.77 | 0.44 | 0.001 |
| Amusements | 0.75 | 0.20 | 0.004 |
| Medical and other health services | 0.4 | 0.13 | 0.061 |
| Educational Services | 0.90 | 0.18 | 0.037 |
| Nonprofit membership organizations | 0.98 | 0.25 | 0.025 |
| Miscellaneous professional services | 0.58 | 0.15 | 0.013 |

**Note:** *Labor shares* in industries are calibrated from the labor shares in industries in 1948 as reported by Close & Shulenburger (1971). See text for explanations related with the frequency of price changes in industries and the weight of industries.

This can be attributed to the fact that when the asymmetric cost structure across industries is introduced in the multi-sector model, sectoral price responses may delink from sectoral frequency of price

**Table 3:** *Estimates of Structural Parameters*
*(Multi-Sector Model with* Asymmetric *Cost Structure)*

| $\varphi$ | $\sigma_H$ | $\theta_p$ | $\theta_w$ | $b$ | $\gamma_p$ | $\gamma_w$ | *Obj. Func.* |
|---|---|---|---|---|---|---|---|
| 0.014 | 0.001 | 81.4 | 12.68 | 0.78 | 0.97 | 0.87 | 53.59 |

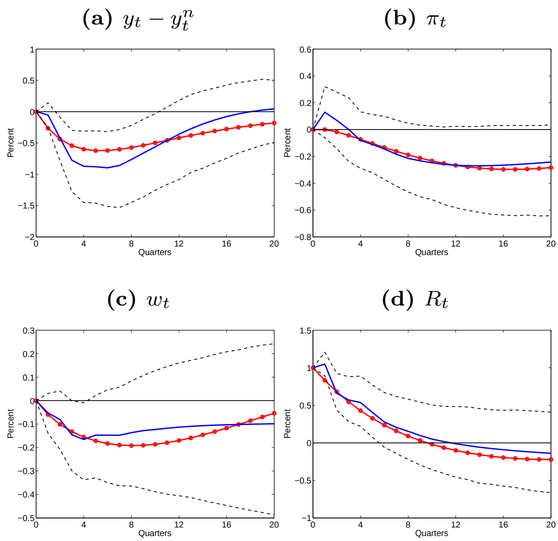
**Note:** *Obj. Func.* indicates the estimated value for the minimization problem discussed in (3.5).

changes following the contractionary monetary shock. This point is illustrated in Figure 11 where the sectoral inflation dynamics in the petroleum and air-transportation industries. As suggested by the frequency of price changes in Table 2, almost all firms in both industries optimize their prices each month. However, the former has a much lower labor share than the latter. For this reason, in Figure 11, the former and the latter are labeled as the low and high labor-share industries, respectively. As evident in the figure, inflation dynamics initially differ largely in the former and the latter despite having virtually the same frequency of price changes. Indeed, while inflation is almost unchanged in the former one period after the shock, inflation in the latter shows a strong increase one period after the shock, suggesting sectoral price responses may substantially differ across industries with a similar frequency of price changes in the *multi-sector model with asymmetric cost structure*, causing the association between the frequency of price changes and sectoral price responses in the sectors following a contractionary shock to be low compared to that in the *multi-sector model with symmetric cost structure.* [11]

[11] This can be explained as follows: Since the frequency of price changes in both of the industries is almost one, it is reasonable to assume firms optimize their prices each period. Under this assumption, it can be shown that firms set prices relative to the aggregate price by imposing some constant mark up over real marginal costs ($s_{it+s}(i')$) which can be written in its log-deviation as:

$$\hat{s}_{it+s}(i') = \frac{1-\kappa_i}{\kappa_i}\hat{y}_{it}(i') + \hat{R}_t + \hat{w}_t$$

**Figure 9:** *Impulse Responses to an Unanticipated 1% Rise in* $R_t$
*(Multi-Sector Model with* Asymmetric *Cost Structure)*




**Note**: The solid lines show the VAR-based impulse responses and the area between dotted lines indicate the 95% confidence intervals estimated with the method suggested by Sims & Zha (1999). The solid lines marked with circles represent the dynamic responses of the variables as predicted by the model.

In addition to bringing the correlations closer to those found in

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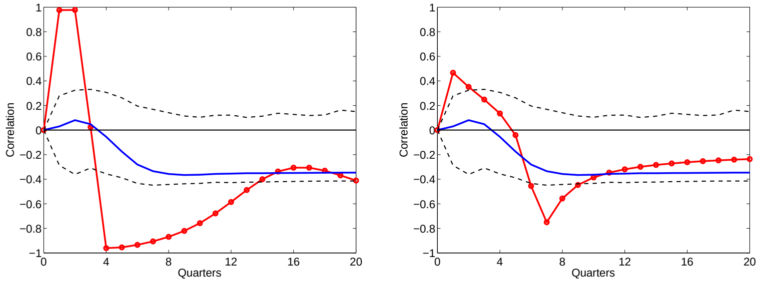
A contractionary shock has two effects on prices which work in opposite directions. The first is that marginal costs increase due to the working-capital channel in the model and an increase in $\hat{R}_t$. The second is that a fall in output results in a fall in marginal costs. The second effect is more decisive in the low labor-share industry since the real marginal costs faced by firms in the low labor-share industry would fall much more markedly compared to those in the high labor-share industry for a given fall in their output as the former has much lower $\kappa_i$. This explains why prices in the low labor-share industry fall, while they increase strongly in the high labor-share industry one period after the shock.

**Figure 10:** *Model- and VAR-Based Correlations of* $\lambda_i$ *with* $\mathcal{C}^{lnp_{it}}_{1,20}$

**(a)** *The Multi-Sector Model with Symmetric Cost Structure*

**(b)** *The Multi-Sector Model with* Asymmetric *Cost Structure*



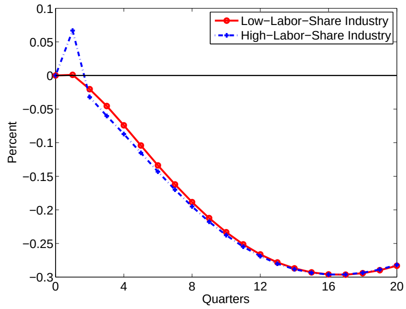

**Note**: The solid lines show the VAR-based $\rho_{\mathcal{C}^{lnp_{it}}_{1,20},\lambda_i}$ and the area between dotted lines indicates the 95% confidence interval for $\rho_{\mathcal{C}^{lnp_{it}}_{1,20},\lambda_i}$ that is estimated using the block-bootstrap method described in Appendix B. The solid lines marked with circles represent $\rho_{\mathcal{C}^{lnp_{it}}_{1,20},\lambda_i}$ predicted by the model.

the data, the *multi-sector model with asymmetric cost structure* can also explain the wide distrubution of sectoral price responses to a contractionary interest rate shock displayed in Figure 2. My findings in this section suggest adding asymmetries in the cost structure is crucial in explaining the low correlations between the frequency of price changes and the sectoral price responses to an interest rate shock.

# 4 Conclusion

In this paper, the implications of heterogeneity in price flexibility at the disaggregated level are studied. I have found that price responses

**Figure 11:** *Inflation Dynamics in the Low and High Labor-Share Industries*
*(The Multi-Sector Model with* Asymmetric *Cost Structure)*




to an unanticipated change in the interest rate differ substantially among sectors. Based on this finding, it is safe to claim that interest rate shocks have strong relative price effects at the disaggregate level. Next, I have investigated whether this differential price response across sectors can be associated with the wide distribution of the frequency of price changes in the United States. The findings in this paper indicate that the association is weak. Lastly, the performances of three DSGE models are evaluated in explaining the aforementioned findings in the empirical section. It has been shown that the *one-sector model* may not explain the wide distribution of sectoral price responses following the shock. It is possible to account for this finding by using a multi-sector model where sectors differ only in their frequency of price changes. However, contrary to the weak association of the frequency of price changes with sectoral price responses in the data, this model predicts a strong correlation between these variables. For this reason, an alternative multi-sector model has been considered. In this model, sectors differ not only in the frequency of price changes but also in their cost structure. Such sectoral asymmetries in the cost structure and the frequency of price changes have proved important in successfully explaining the strong relative price effects of the interest rate shock and the weak association of the frequency of price changes with sectoral price responses in the data.

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# A The Bils, Klenow & Kryvtsov (2003) Model Reconsidered

In this section, I aim to explain why the empirical strategy for studying the relative price effects of monetary shocks in the United States in Bils, Klenow & Kryvtsov (2003) produces different outcomes than those in Section 2.2. Before such an analysis, it is useful to review the Bils, Klenow & Kryvtsov (2003) model.

## A.1 The Bils, Klenow & Kryvtsov (2003) Model

Bils, Klenow & Kryvtsov (2003) investigate sectoral price responses to a monetary policy shock using the following empirical method:

$$lnp_{it} = \lambda_i \sum_{k=k_{min}}^{k_{max}} \beta_k \epsilon_{t-k} + \mu_i + \tau_i t + \eta_t + \nu_{it} \qquad \text{(A.1)}$$

where $lnp_{it}$ denotes the logged price of *sector i*. $\lambda_i$ and $\varepsilon_{t-k}$ show the frequency of price changes in *sector i* and innovations in the monetary policy instrument, respectively. The error component in the panel data estimation of (A.1) is composed of sector- and time-specific terms. Sector-specific terms, which are included to allow each disaggregated serial to have a different intercept and a different trend, are denoted by $\mu_i$ and $\tau_i$, respectively. Time-specific terms, on the other hand, are meant to capture factors unobservable to the researcher and are assumed to affect all prices by the same magnitude in *period t* and are denoted by $\eta_t$. The maximum number of periods that a monetary policy innovation may have an impact on $p_{it}$ is denoted by $k_{max}$.

I assume some delays may occur for prices in a sector when responding to monetary shocks. $k_{min}$ denotes the number of lags in sectoral price responses. In estimating (A.1), I maintain there is a quarter delay in sectoral price responses to monetary shocks ($k_{min} = 1$). Differently, Bils, Klenow & Kryvtsov (2003) assume monetary shocks have a contemporaneous impact on sectoral prices ($k_{min} = 0$). However, such an assumption is at odds with the empirical strategy for isolating monetary shocks in Bils, Klenow & Kryvtsov (2003), which requires aggregate price level to respond to monetary shocks with a lag.

Lastly, it is notable that sector-specific errors ($\nu_{it}$) in the Bils, Klenow & Kryvtsov (2003) model are assumed to follow an AR(2) process and is given by:

$$\nu_{it} = \rho_1 \nu_{it-1} + \varrho_1 \lambda_i \nu_{it-1} + \rho_2 \nu_{it-2} + \varrho_2 \lambda_i \nu_{it-2} + u_{it} \tag{A.2}$$

The specification for sector-specific shocks in (A.2) implies that persistence in sector-specific shocks depends on the frequency of price changes in sectors. To explain why such an assumption is made in the Bils, Klenow & Kryvtsov (2003) model, consider first that a sector-specific shock emerges in the fully flexible-price sector. As all prices in this sector can adjust instantaneously to any type of shock, it is expected to have transitory effects on this sector's price, $p_{it}$. Next, consider a sector-specific shock hits a sticky-price sector where only a small fraction of firms can reset prices each period. Since it may take quite a while for firms in this sector to adjust fully to the shock, the shock is likely to have more persistent effects on this sector's price. Lower persistence of sector-specific shocks in flexible-price categories are reflected in the Bils, Klenow & Kryvtsov (2003) model in the conjecture that $\varrho_1$ and $\varrho_2$ have negative signs in (A.2).

The first step in Bils, Klenow & Kryvtsov (2003) is to obtain structural monetary shocks ($\varepsilon_t$). To do so, I assume that the Federal Reserve uses the federal funds rate as its policy instrument and uses the following interest rate rule:

$$\begin{aligned} R_t &= \theta_0 + \textstyle\sum_{k=0}^{4} \theta_{y-y^n,k}(y_{t-k} - y^n_{t-k}) + \sum_{k=0}^{4} \theta_{\pi,k}\pi_{t-k} \\ &+ \textstyle\sum_{k=0}^{4} \theta_{w,k} w_{t-k} + \sum_{k=1}^{4} \theta_{i,k} R_{t-k} + \epsilon_t \end{aligned} \tag{A.3}$$

where $R_t$, $y_t - y^n_t$, $\pi_t$ and $w_t$ are defined in (2.2). After obtaining monetary shocks by using (A.3), the Cochrane-Orcutt procedure is used to estimate the coefficients in the Bils, Klenow & Kryvtsov (2003) model.[12] In this model, the relative price effects of a contractionary monetary policy shock are given as:

[12]Since $\nu_{it}$ in (A.1) is autoregressive, the OLS estimates of the coefficients in (A.1) are inefficient. In addition, OLS standard errors of those coefficients are incorrect. If the true values of the autoregressive coefficients were known, (A.1) could easily be estimated by multiplying each side with the following term:

$$\beta_k(\lambda_{90} - \lambda_{10}) \tag{A.5}$$

where $\lambda_{90}$ and $\lambda_{10}$ show the frequencies of the sectors which lie on the 90th and 10th percentile of price flexibility, respectively. I call the categories which lie at these percentiles of price flexibility as the flexible- and sticky-price categories, respectively. The intuition for (A.5) is as follows. Note that when the Federal Reserve introduces an unanticipated 1% increase in the federal funds rate, the percentage change in $p_{it}$ of the flexible-price sector (the sticky-price sector) is given by $\lambda_{90}\beta_k$ ($\lambda_{10}\beta_k$). Accordingly, (A.5) measures the percentage change in $p_{it}$ of the flexible-price sector relative to that of the sticky-price sector in the $k^{th}$ period following the shock. Since the frequencies of price changes in the sticky- and flexible-price categories differ significantly, the dynamic behavior of $p_{it}$ in both sectors may show substantial differences unless the estimate of $\beta_k$s are too small.

---

$$1 - (\rho_1 + \varrho_1\lambda_i)L - (\rho_2 + \varrho_2\lambda_i)L^2 \tag{A.4}$$

where $L$ is the lag operator. It needs to be emphasized that the transformation required for each sector is different as $\lambda_i$ varies across sectors. It is intuitive to transform (A.1) this way since multiplying $\nu_{it}$ with (A.4) yields errors of the transformed model($\zeta_{it}$) which are uncorrelated, and thus, the OLS with the transformed model is efficient. Yet, autoregressive coefficients are unknown and need to be estimated. In estimating these coefficients, Bils, Klenow & Kryvtsov (2003) employed the well-known Cochrane-Orkutt iterative procedure. In this procedure, parameters in (A.1) are first estimated with OLS. Then, estimated OLS residuals in (A.1) are used to obtain the first round estimate of $\varrho_1$,$\rho_1$, $\varrho_2$ and $\rho_2$ in (A.2). Then, both dependent and independent variables are transformed using these first round autoregressive estimates instead of the true autoregressive coefficients in (A.4). After both the dependent and independent variables are transformed this way, the second round coefficient estimates are obtained as well as the second round OLS residuals in (A.1). Using the second round OLS residuals, the second round autoregressive coefficients in (A.2) are obtained and the variables are transformed using these second round autoregressive coefficients once again. This iteration continues until the autoregressive coefficient estimates in two consecutive rounds differ no more than some threshold. As a convergence criteria, I chose estimates of $\rho_1$ from two consecutive rounds with a change of less than 0.01.

## A.2 Findings from the Bils, Klenow & Kryvtsov (2003) Model

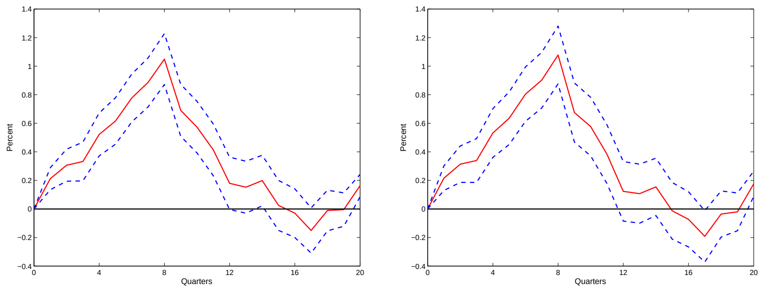
Next, I study the relative price effects of monetary policy shocks in the Bils, Klenow & Kryvtsov (2003) model using my sample, which spans the period of 1959Q1-2013Q1. Figure 12 shows the percentage change in $p_{it}$ of the flexible-price sector relative to that of the sticky-price sector following an unanticipated 1% increase in the federal funds rate in the Bils, Klenow & Kryvtsov (2003) method ( $\beta_k(\lambda_{90} - \lambda_{10})$ ) . In *Panel (a)* and *Panel (b)* of this figure, the dynamics of the relative price after the shock are estimated using *the frequency of non-sales price changes* and *the frequency of price changes including sales*, respectively. Two points are noteworthy regarding Figure 12. First, the inclusion of price changes during sales in measuring the frequency of price changes in sectors has only a small effect on our results since the relative price dynamics after the shock are rather similar when the frequency of price changes includes or excludes price changes during sales. Second, our results are even more striking than the results in Bils, Klenow & Kryvtsov (2003). Indeed, following a contractionary monetary shock, I find the relative price stays above its undistorted path for *about three years* compared to only *three quarters* as found in Bils, Klenow & Kryvtsov (2003).

Barth & Ramey (2001) claim a rise in inflation after a contractionary monetary shock can be explained with the working-capital channel, which they show to be operative in the United States. Indeed, firms' requirement to pay input costs in advance raises factor costs when the interest rate increases. Consequently, when there is an unanticipated rise in the interest rate, firms raise their prices for some periods after the shock despite the downward pressure on prices from a reduction in output following the shock. If the relative price responses in the flexible-price sector were positive *only for a few periods*, such a channel could be invoked to explain the relative price puzzle. Yet, this channel may not explain the positive responses of the relative price for three years as found in this paper.

**Figure 12:** *The Bils, Klenow & Kryvtsov (2003) Model-Based Impulse Responses of the Relative Price to Monetary Shocks*

**(a)** *With the Frequency of Non-Sales Price Changes*

**(b)** *With the Frequency of Price Changes Including Sales*




**Note:** In the figure, the solid line indicates the estimated point-wise impulse responses. The area between the dashed lines shows the two standard deviation confidence intervals for the estimate of $\beta_k(\lambda_{90} - \lambda_{10})$ in (A.1). As in Bils, Klenow & Kryvtsov (2003), in estimating these confidence intervals, the uncertainty in estimating structural monetary shocks in (A.3) is not taken into account.

## A.3 Testing a Critical Assumption in the Bils, Klenow & Kryvtsov (2003) Model

The reliability of the findings in Figure 12 depends on whether monetary shocks are orthogonal to sector-specific errors in period $t$ in (A.1). To see this, note from (A.1) that the unbiasedness of coefficients requires $E(\nu_{it}|\epsilon_t) = 0$. That is, sector-specific shocks should be orthogonal to monetary shocks to have unbiased estimates. If this condition does not hold, the GLS estimates of the parameters in (A.1) will have bias and the results obtained with the Bils, Klenow & Kryvtsov (2003) model may be questionable. Next, I aim to test this hypothesis. Note, when obtaining monetary shocks, Bils, Klenow & Kryvtsov (2003) assume the Federal Reserve only responds to inflation in the general price level and does not take into account movements in any sectoral price. If this really holds, then, the sector specific shocks will be orthogonal to monetary shocks and the coefficients in the Bils, Klenow & Kryvtsov (2003) model can be estimated unbiasedly. One way to test whether the Federal Reserve responds to sectoral prices, apart from the general price level, is to incorporate a single sectoral price index in its policy reaction function in (A.3) and test whether the coefficients pertinent to the Federal Reserve's response to sectoral prices are jointly zero. The following regression is considered for this test:

$$\begin{aligned} R_t &= \theta_0 + \textstyle\sum_{k=0}^4 \theta_{y-y^n,k}(y_{t-k} - y_{t-k}^n) + \sum_{k=0}^4 \theta_{\pi,k}\pi_{t-k} \\ &+ \textstyle\sum_{k=0}^4 \theta_{lnp_i,k} lnp_{it-k} + \sum_{k=0}^4 \theta_{w,k} w_{t-k} + \sum_{k=1}^4 \theta_{i,k} R_{t-k} + \tilde{\epsilon}_t \end{aligned} \tag{A.6}$$

Estimating (A.6) requires the assumption that the Federal Reserve observes sectoral prices and may respond to them if this is desired. Also, it requires sectoral prices to respond to monetary shocks with at least a quarter lag. The only difference between (A.3) and (A.6) is the term $\sum_{k=0}^4 \theta_{lnp_i,k} lnp_{it-k}$. The structural monetary shocks in (A.3) can be associated with those in (A.6) in the following way:

$$\epsilon_t = \tilde{\epsilon}_t + \sum_{k=0}^{4} \theta_{lnp_i,k} lnp_{it-k} \tag{A.7}$$

The maintained assumption in the Bils, Klenow & Kryvtsov (2003) model is that $\theta_{lnp_i,0} = \theta_{lnp_i,1} = \cdots = \theta_{lnp_i,4} = 0$. If this assumption does not hold, $E(\epsilon_t \nu_{it}) \neq 0$ and the GLS estimators in the Bils, Klenow & Kryvtsov (2003) model will be inconsistent. One way to interpret the rejection of the null is that the response of the Federal Reserve to price movements in these sectors is not confined to their marginal effects on the aggregate price level. Apart from this, the Federal Reserve gives a statistically different response to sectoral shocks in these sectors. By using (A.6), I perform an F-test for the null hypothesis for each sector. I find among 125 sectors, the null hypothesis is rejected for 19 sectors. Table 4 lists the sectors for which the null is rejected. For the remaining sectors, the Federal Reserve's response is confined to its response to the change in the general price level inflation caused by sector-specific price shocks in these sectors.

**Table 4:** *Sectors with a Significant Response from the Federal Reserve*

| | |
|---|---|
| Tires | Dental Services |
| Watches | Nursing homes |
| Poultry | Other recreational services |
| Fish and Seafood | Musical instruments |
| Vegetables (fresh) | Commercial and vocational schools |
| Children's and infants' clothing | Professional and other services |
| Lubricants and fluid | Clothing repair, rental, and alterations |
| Miscellaneous household products | Carpets and other floor coverings |
| Newspapers and periodicals | Telecommunication services |
| Electric appliances for personal care | |

Next, I study the association between the likelihood that the Federal Reserve provides a significant response to sectoral price shocks and the frequency of price changes. For this purpose, a scatter plot of F-values and the frequency of *non-sales* price changes and a scatter plot of F-values and the frequency of price changes *including sales* are displayed with circles in Panel (a) and Panel (b) of Figure 13, respectively. The 90% critical value for an $F(5, T-25)$ random variable, where T indicates sample size, is graphically represented by the thick solid line. If an F-value is greater than the critical value, the null hypothesis, that the Federal Reserve is only concerned with the general price level inflation and does not respond to sectoral prices, should be rejected.

In Figure 13, I also show the fitted line for the following regression:

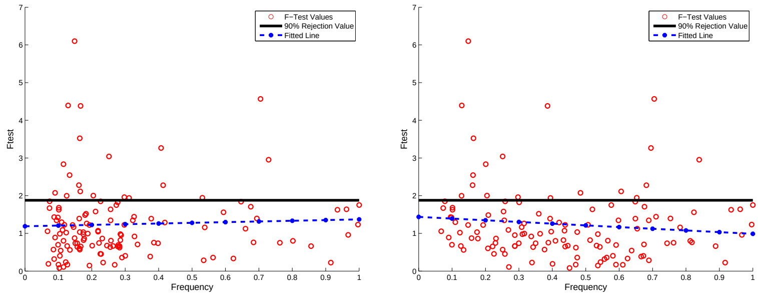
$$F_i = a_0 + a_1 \lambda_i \tag{A.8}$$

A positive (negative) slope of the fitted line suggests the higher (lower) the frequency of price changes in a sector, the more (less) probable is the Federal Reserve to provide a significant response to sectoral price shocks. I estimate (A.8) with both the frequency of price changes including sales ($\lambda_i^{sales}$) and the frequency of non-sales price changes ($\lambda_i^{regular}$) and report the findings in (A.9).

**Figure 13:** *Testing for the Significance of the Federal Reserve's Response for Sectoral Prices*

**(a)** *With the Frequency of Non-Sales Price Changes*

**(b)** *With the Frequency of Price Changes Including Sales*




**Note:** In the figure, the circles show a scatter plot of the F-value for the null hypothesis and the frequency of price changes in sectors. The thick solid line indicates the 90% critical value for the F-test. The dotted lines marked with asterisks show the fitted line for the regression in (A.8).

$$F_i = 1.19 + 0.18\lambda_i^{regular} \qquad\qquad F_i = 1.44 - 0.45\lambda_i^{sales} \tag{A.9}$$

These findings indicate that while the likelihood of the Federal Reserve to provide a significant response to sectoral price shocks is positively associated with the frequency of *non-sales* price changes, it is negatively associated with the frequency of price changes *including sales*. However, both the positive and negative associations are weak. To see this point, note that the fitted lines in both cases are flat and they always remain below the critical value within the admissible region of the frequency of price changes in sectors. Hence, the frequency of price changes in a sector does not seem to be an important factor in the decision of the Federal Reserve once the effect of these shocks on the general price level inflation is controlled.

To sum up, the findings in this section indicate that structural monetary shocks that are needed to estimate the Bils, Klenow & Kryvtsov (2003) model ($\epsilon_t$ in (A.1)) are correlated with sector-specific price shocks ($\nu_{it}$ in (A.1)). Since this assumption is critical in the Bils, Klenow & Kryvtsov (2003) model and is shown to be violated for a non-negligible number of sectors in our sample, it can be argued that the results in the Bils, Klenow & Kryvtsov (2003) model are questionable.

# B Estimation of Confidence Intervals for Figure 3 Using a Block-Bootstrap Method

The 'block-of-blocks' bootstrap method of Politis & Romano (1992) is used to estimate the confidence interval for the correlation between the frequency of price changes in sectors and sectoral price responses to a 1% increase in the federal funds rate. The following steps are followed to construct the confidence intervals.

1. Let $Z_t$ and $\mathcal{Z}_t$ be defined as:

$$
\begin{aligned}
Z_t &= \left|\; y_t - y_t^n, \quad \pi_t, \quad w_t, \quad R_t, \quad \pi_{1t}, \quad \pi_{2t}, \quad \ldots \quad \pi_{it}, \quad \ldots \quad \pi_{125t} \;\right| \\
\mathcal{Z}_t &= \left|\; Z_t \quad Z_{t-1} \quad Z_{t-2} \quad Z_{t-3} \quad Z_{t-4} \;\right|
\end{aligned}
\tag{B.1}
$$

   where $\pi_{it}$ in $Z_t$ represents the annualized percentage change in the price of the sector $i$.

2. Next, $T - b + 1$ overlapping blocks are formed where $T$ is the sample size and $b$ is the fixed length of blocks. The first contains observations $|\; \mathcal{Z}_1 \quad \mathcal{Z}_2 \quad \ldots \quad \mathcal{Z}_b \;|'$. The second contains observations $|\; \mathcal{Z}_2 \quad \mathcal{Z}_3 \quad \ldots \quad \mathcal{Z}_{b+1} \;|'$. The last contains observations $|\; \mathcal{Z}_{T-b+1} \quad \mathcal{Z}_{T-b+2} \quad \ldots \quad \mathcal{Z}_T \;|'$.

3. For the block length, the following values are considered $b \in \{6, 9, 12\}$. Since the results are similar, only the results for $b = 12$ are displayed in Figure 3.

4. A random sample of size T is constructed by resampling from these blocks with replacement. Since $T/b$ is not an integer, the last block has been truncated.

5. Next, for each random sample, the correlation between the frequency of price changes and sectoral price responses to a 1% increase in the federal funds rate is estimated by the method discussed in Section 2.2.

6. These steps are repeated 500 times. The confidence intervals displayed in Figure 3 give the area between the 2.5th and 97.5th percentiles of these randomly generated correlations.