MONETARY POLICY, MARKUP DISPERSION, AND AGGREGATE TFP

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Abstract—We study the role of markup dispersion and aggregate TFP for monetary transmission. Empirically, we show that the response of markup dispersion to monetary policy shocks can account for a significant fraction of the aggregate TFP response in the first two years after the shock. Analytically, we show that heterogeneous price rigidity can explain the response of markup dispersion if firms have a precautionary price-setting motive, which is present in common New Keynesian environments. We provide empirical evidence in support of this explanation. Finally, we study the mechanism and its implications in a quantitative model.

I. Introduction

WE revisit one of the long-standing questions in macroeconomics: What are the channels through which monetary policy affects real economic outcomes? Our paper is motivated by empirical evidence that monetary policy shocks have sizable effects on measured aggregate productivity.[1] A potential explanation for fluctuations in measured aggregate TFP is changing resource misallocation across firms. The TFP-misallocation link has been widely studied in the macro-development literature (e.g., Hsieh & Klenow, 2009) and is well understood in the New Keynesian literature. Although in New Keynesian models misallocation is commonly captured by price dispersion, our preferred empirical measure of misallocation is dispersion in markups across firms. Markup dispersion is price dispersion when controlling for differences in marginal costs across firms.

We study the role of markup dispersion for monetary transmission by asking two questions: First, does markup dispersion respond to monetary policy shocks? Using U.S. data, we document a significant response of markup dispersion, which can account for a significant fraction of the aggregate TFP response up to two years after the shock. Second, what explains the response of markup dispersion? We show analytically that heterogeneity in price-setting frictions—in an otherwise standard New Keynesian framework—can explain the response of markup dispersion. The fundamental reason is that firms with stickier prices have a stronger precautionary price setting motive. This channel has testable implications, which, as we show, are supported empirically. Finally, we study the mechanism and its implications in a quantitative model.

We estimate the response of markup dispersion to monetary policy shocks based on quarterly balance-sheet data and high-frequency identified monetary policy shocks. A central contribution of this paper is to show that the dispersion of markups across firms (within industries) significantly increases after contractionary monetary policy shocks and decreases after expansionary monetary policy shocks. The response is persistent and peaks about two years after the shock. We establish this empirical pattern for a host of markup measures, following, among others, De Loecker and Warzynski (2012) and Gutiérrez and Philippon (2017). To translate the estimated response of markup dispersion into an aggregate TFP response, we follow Hsieh and Klenow (2009) and Baqaee and Farhi (2020). The response of markup dispersion implies a response in aggregate TFP between -0.2% and -0.4% two years after a one standard deviation contractionary monetary policy shock. For comparison, the directly estimated empirical response of utilization-adjusted aggregate TFP is -0.4% at a two-year horizon. At more distant horizons, markup dispersion accounts for a decreasing fraction of the aggregate TFP response.

Our evidence sheds new light on the TFP effects of monetary policy. Strikingly, the estimated response of markup dispersion cannot be explained by a large class of New Keynesian models, at least when solved with standard perturbation methods. In many New Keynesian models, including medium-scale models (e.g., Christiano et al., 2005) and models with heterogeneous price rigidity (e.g., Carvalho, 2006), markup dispersion does not respond to monetary policy shocks up to a first-order approximation around the deterministic steady state. In the second-order approximation, markup dispersion responds, but counterfactually increases in response to both positive and negative shocks. In models with trend inflation (e.g., Ascari & Sbordone, 2014), markup dispersion decreases after contractionary and increases after expansionary monetary shocks, which contradicts our empirical evidence.

What can explain the response of markup dispersion to monetary policy shocks instead? We propose a novel mechanism that arises from heterogeneity in the severity of price-setting frictions across firms. A sufficient condition for higher markup dispersion after a monetary tightening is that firms with higher markups have lower pass-through from marginal costs to prices, that is, relatively strong price-setting frictions. A contractionary monetary shock that lowers marginal costs increases the relative markup of

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[1]Using U.S. data, we document that monetary policy shocks lower measured aggregate productivity, which reconfirms the evidence in Evans and Santos (2002), Christiano et al. (2005), Moran and Queralto (2018), Jordà et al. (2020), and Garga and Singh (2021).

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low pass-through firms, which increases markup dispersion. Analogously, expansionary monetary shocks that raise marginal costs will lower markup dispersion. We show that a negative correlation between firm-level markup and pass-through can arise endogenously from heterogeneity in price-setting frictions. The types of price-setting frictions we consider are a Calvo (1983) friction, Taylor (1979) staggered price setting, Rotemberg (1982) convex adjustment costs, and Barro (1972) menu costs. The intuition for this negative correlation is a precautionary price setting motive. The firm profit function in the common New Keynesian environment is asymmetric; that is, it penalizes markups below more than markups above the statically-optimal one. A higher reset markup provides insurance against low profits before the next price adjustment opportunity (Calvo/Taylor) or lowers the expected costs of future price readjustments (Rotemberg/Barro).[2] To summarize, heterogeneous price-setting frictions imply markup dispersion and hence TFP effects of monetary policy. Importantly, precautionary price setting is absent in the deterministic steady state. By extension, our transmission mechanism is absent in models with heterogeneous price-setting frictions when solved around the deterministic steady state.

We empirically test two implications of this transmission mechanism. First, precautionary price setting implies that firms with stickier prices charge higher markups. Second, the markups of firms with stickier prices should increase by relatively more. A caveat is that we do not observe firm-specific price adjustment frequencies. Instead, we capture variation in price adjustment frequencies across firms using price adjustment frequencies in five-digit industries together with the firm-specific sales composition across industries. We find that firms with stickier prices indeed have higher markups on average and increase their markups by more after monetary policy shocks. These two results hold when controlling for two-digit sector fixed effects, firm size, leverage, and liquidity.

Finally, we study the mechanism and its implications in a quantitative New Keynesian model with heterogeneous price rigidity. To capture precautionary price setting, we use nonlinear solution methods to solve the model dynamics around the stochastic steady state, to which the economy converges in the presence of uncertainty but absence of shocks. We find that indeed firms with stickier prices set higher markups on average, and monetary policy shocks raise markup dispersion. Quantitatively, a one standard deviation contractionary monetary policy shock lowers aggregate TFP by 0.34%. We use the model to study two implications of our mechanism. Whereas a contractionary monetary shock increases aggregate markups in many New Keynesian models, the empirically estimated responses of aggregate markups in Nekarda and Ramey (2020) have the opposite sign. In our model, the aggregate markup falls if contractionary monetary shocks lower aggregate TFP sufficiently strongly. This argument extends to sector or firm-level markups if price rigidities are heterogeneous within sectors or firms such that sector or firm-level TFP responds to monetary policy. We further analyze the effectiveness of monetary policy when the endogenous TFP effects are ignored by the monetary authority. If the monetary authority attributes all TFP fluctuations to technology shocks, interest rates are adjusted less aggressively and monetary policy shocks lead to larger GDP fluctuations.

This paper is closely related to four branches of the literature. First, a growing literature studies the positive and normative implications of heterogeneous price rigidity; see, for example, Aoki (2001), Carvalho (2006), Nakamura and Steinsson (2010), Eusepi et al. (2011), Carvalho and Schwartzman (2015), Castro Cienfuegos and Loria (2017), Pasten et al. (2020), and Rubbo (2020). We show that such heterogeneity gives rise to productivity effects of monetary policy. Similarly, Baqaee and Farhi (2017) show that negative money supply shocks lower aggregate TFP if sticky-price firms have exogenously higher ex ante markups than flexible-price firms. We provide empirical evidence which supports this transmission channel and show that the rigidity-markup correlation can arise endogenously from differences in price rigidity.

Second, this paper relates to a literature that studies the productivity effects of monetary policy, for example, Evans and Santos (2002), Christiano et al. (2005), Comin and Gertler (2006), Moran and Queralto (2018), Jordà et al. (2020), and Garga and Singh (2021). We confirm the empirical finding that monetary policy shocks lower aggregate productivity, but provide a novel explanation based on markup dispersion. In terms of alternative explanations, Christiano et al. (2005) show that variable utilization and fixed costs explain a relatively small fraction of the aggregate productivity response. Moran and Queralto (2018) and Garga and Singh (2021) show that R&D investment falls after monetary policy shocks, which may ultimately lower productivity. However, it is unclear whether the R&D response can explain a large response of aggregate productivity at short horizons. For example, Comin and Mestieri (2018) show that recent technologies are adopted with an average lag of five years. Conversely, price rigidities are a more natural candidate for the effects at shorter horizons.

Third, our paper relates to a literature on the relation between inflation and price dispersion. Whereas we show that contractionary monetary policy shocks raise markup dispersion, Nakamura et al. (2018) document flat price dispersion across periods of high and low inflation since the 1970s. This suggests that long-lived changes in inflation have different effects than short-lived monetary policy shocks. For example, when trend inflation increases, managers may schedule more frequent meetings to discuss price changes (Levin & Yun, 2007), whereas monetary policy shocks are less likely to trigger such responses.

[2] Relatedly, in a setup with homogeneous price-setting frictions Fernandez-Villaverde et al. (2015) study precautionary price setting as a channel through which higher uncertainty leads to higher markups.

Fourth, this paper relates to a growing literature that studies allocative efficiency over the business cycle. Eisfeldt and Rampini (2006) show that capital misallocation is countercyclical. Fluctuations in allocative efficiency may be driven by various business cycle shocks, for example, aggregate productivity shocks (Khan & Thomas, 2008), uncertainty shocks (Bloom, 2009), financial shocks (Khan & Thomas, 2013), or supply chain disruptions (Meier, 2020). We relate to this literature by studying the transmission of monetary policy shocks through allocative efficiency. Interestingly, the effects of short- versus long-run changes in interest rates on allocative efficiency seem to differ in sign. Whereas we show that short-run expansionary monetary policy decreases misallocation, Gopinath et al. (2017) show that, in the case of southern Europe, persistently lower interest rates have increased misallocation. Relatedly, Oikawa and Ueda (2018) study the long-run effects of nominal growth through reallocation across heterogeneous firms.

The remainder of this paper is organized as follows. Section II presents the main empirical evidence. Section III studies monetary transmission with heterogeneous price rigidity. Section IV presents a quantitative model. Section V concludes.

II. Evidence on Markup Dispersion and TFP

In this section, we present novel empirical evidence that monetary policy shocks increase the markup dispersion across firms. We further show that aggregate TFP falls after monetary policy shocks and that a sizable share of this response can be accounted for by the response of markup dispersion.

A. Data

Firm-level markups. We use quarterly balance sheet data of publicly-listed U.S. firms from Compustat. We estimate markups through a variety of methods. Our baseline method is the ratio estimator pioneered by Hall (1986) and more recently used in De Loecker and Warzynski (2012), Flynn et al. (2019), De Loecker et al. (2020), and Traina (2020). We further consider markups using the accounting profits and user cost approaches in Gutiérrez and Philippon (2017), Basu (2019), and Baqaee and Farhi (2020).

The ratio estimator of the markup can be obtained from the cost minimization problem. With a flexible input V_{it}, the markup μ_{it} of firm i in quarter t can be computed as

$$\mu_{it} = \frac{\text{output elasticity of } V_{it}}{\text{revenue share of } V_{it}}. \tag{1}$$

We assume that firms in the same two-digit industry and quarter have a common output elasticity. All our subsequent empirical analysis focuses on differences of firm-level log markups from their industry-quarter average. Under our assumption, these markup differences do not depend on the output elasticities. Hence, our empirical results are not affected by challenges to identify output elasticities from revenue data, as recently emphasized by Bond et al. (2021).[3] By controlling for industry-quarter fixed effects in log markups, we also difference out industry and time-specific characteristics such as differences in competitiveness and production technology.

Formally, we define differences of firm-level log markups from their industry-time average as $\hat{\mu}_{it} \equiv \log \mu_{it} - \frac{1}{N_{st}} \sum_{j \in \mathcal{J}_{st}} \log \mu_{jt}$, where $\mathcal{J}_{st}$ is the set of firms j in industry s and quarter t, and N_{st} is the cardinality of $\mathcal{J}_{st}$. Following De Loecker et al. (2020) we assume firms produce output using capital and a composite input of labor and materials, with the latter the flexible factor. We estimate the revenue share as the firm-quarter-specific ratio of costs of goods sold (`cogsq` in Compustat) to sales (`saleq`).

We further consider a host of alternative markup estimation methods in section IID below. First, we construct (nonratio estimator) markups through an accounting profit approach and a user cost approach, following Gutiérrez and Philippon (2017) and Baqaee and Farhi (2020). Second, following Traina (2020), we add selling, general, and administrative expenses (SGA) to the costs of goods sold in the baseline markup measure. Third, we estimate a four-digit industry-specific translog production technology, which implies variation in output elasticities within industry and time. Fourth, we estimate four-digit industry-quarter-specific output elasticities through cost shares.

We consider all industries except public administration, finance, insurance, real estate, and utilities. We drop firm-quarter observations if sales, costs of goods sold, or fixed assets are reported only once in the associated year. We further drop observations if quarterly sales growth is above 100% or below −67% or if real sales are below 1 million USD. We finally drop the bottom and top 5% of the estimated markups. Appendix A.1 provides more details and summary statistics in table A1. Our results are robust to alternative data treatments as we discuss toward the end of this section.

Monetary policy shocks. Using high-frequency data of federal fund future prices, we identify monetary policy shocks through changes of the future price in a narrow time window around FOMC announcements. The identifying restrictions are that the risk premium does not change and that no other macroeconomic shock materializes within the time window. We denote the price of a future by f, and by τ the time of a monetary announcement.[4] We use a thirty-minute window around FOMC announcements, as in

[3]Our baseline approach assumes the ratio estimator to be valid in principle. This excludes the case when the input is not perfectly flexible or when its choice affects demand; see Bond et al. (2021). We also consider nonratio estimators of markups; see section IID.

[4]We obtain time and classification of FOMC meetings from Nakamura and Steinsson (2018) and the FRB. We obtain time stamps of the press release from Lucca and Moench (2015) and Gorodnichenko and Weber (2016).

Gorodnichenko and Weber (2016). Let $\Delta\tau^- = 10$ minutes and $\Delta\tau^+ = 20$ minutes; then monetary policy shocks are

$$\varepsilon_\tau^{\mathrm{MP}} = \mathrm{f}_{\tau+\Delta\tau^+} - \mathrm{f}_{\tau-\Delta\tau^-}. \tag{2}$$

To aggregate the shocks to quarterly frequency, we follow Ottonello and Winberry (2020). We assign daily shocks fully to the current quarter if they occur on the first day of the quarter. If they occur within the quarter, we partially assign the shock to the subsequent quarter. This procedure weights shocks across quarters corresponding to the amount of time agents have to respond. Formally, we compute quarterly shocks as

$$\varepsilon_t^{\mathrm{MP}} = \sum_{\tau\in\mathcal{D}(t)} \phi(\tau)\varepsilon_\tau^{\mathrm{MP}} + \sum_{\tau\in\mathcal{D}(t-1)} (1-\phi(\tau))\varepsilon_\tau^{\mathrm{MP}}, \tag{3}$$

where $\mathcal{D}(t)$ is the set of days in quarter t and $\phi(\tau) =$ (remaining number of days in quarter t after announcement in τ) / (total number of days in quarter t).

As a baseline, we construct monetary policy shocks from the three-months-ahead federal funds future, as in Gertler and Karadi (2015). Our baseline excludes unscheduled meetings and conference calls.[5] Following Nakamura and Steinsson (2018), our baseline further excludes the apex of the financial crisis from 2008Q3 to 2009Q2.[6] The monetary policy shock series covers 1995Q2 through 2017Q3. We discuss alternative monetary policy shocks in section IID. Table A2 in appendix A reports summary statistics and figures A1a and A1b shows the shock series.

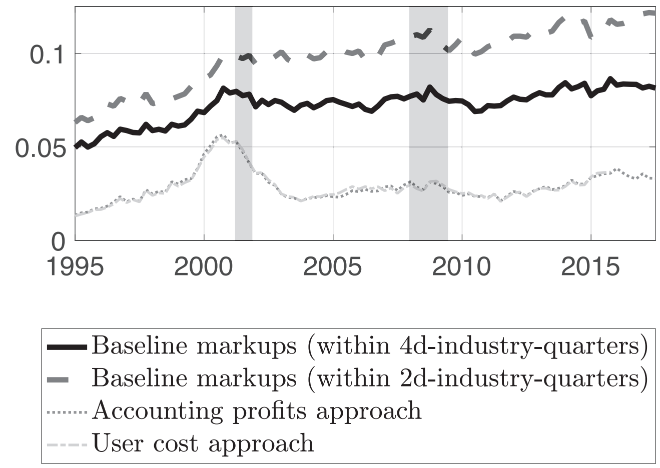
B. *Markup Dispersion*

We estimate the response of markup dispersion to monetary policy shocks. Our baseline measure of markup dispersion is the cross-sectional variance $\mathbb{V}_t(\hat{\mu}_{it})$, where $\hat{\mu}_{it}$ denotes firm-level log markups in deviation from their respective industry-quarter mean. Recall that our baseline estimator of $\hat{\mu}_{it}$ does not depend on an estimator of the output elasticity under our assumption that firms within a two-digit industry-quarter have a common output elasticity. Figure 1 shows time series of markup dispersion for our baseline ratio estimator within four-digit industry-quarters, the same estimator but within two-digit industry-quarters, and for markups based on account profits and user costs. Figure A1c in appendix A shows time series for further alternative estimators, notably the ratio estimator when including SGA, the translog-based markups, and markups based on cost shares.

[5]Unscheduled meetings and conference calls often occur after adverse economic developments. Price changes around such meetings may reflect these developments, invalidating the identifying restriction. Our results remain broadly robust when including these meetings.

[6]We discard shocks during 2008Q3 to 2009Q2, and we do not regress post-2009Q2 outcomes on pre-2008Q3 shocks. Our results are robust to including this period.

FIGURE 1.—EVOLUTION OF MARKUP DISPERSION



Evolution of markup dispersion for different markup measures from 1995Q1 to 2017Q3. Markup dispersion is the variance of log markups across firms, $\mathbb{V}_t(\hat{\mu}_{it})$, where $\hat{\mu}_{it}$ is the difference of a firm's log markup from the mean log markup across firms in the same industry-quarter. Baseline markups are constructed according to equation (1) assuming a common output elasticity for firms in the same second-industry-quarter. Further details on the accounting profits and user cost approaches are provided in section IID.

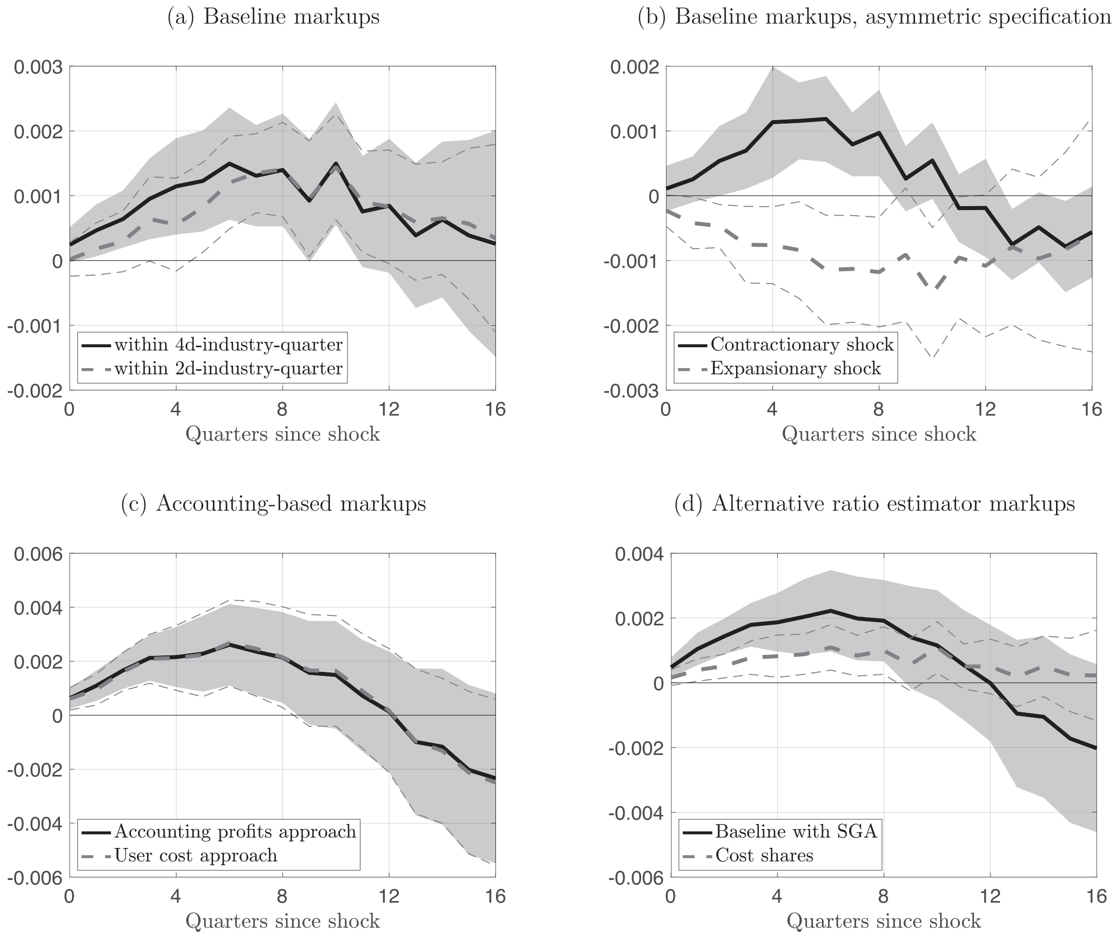
To estimate the effects of monetary policy shocks on markup dispersion, we use the following local projection for $h = 0, \dots, 16$ quarters and where y_t is markup dispersion:

$$y_{t+h} - y_{t-1} = \alpha^h + \beta^h\varepsilon_t^{\mathrm{MP}} + \gamma_0^h\varepsilon_{t-1}^{\mathrm{MP}} + \gamma_1^h(y_{t-1} - y_{t-2}) + u_t^h. \tag{4}$$

The central empirical finding of this paper is shown in panel a of figure 2, which plots the response of markup dispersion, captured by the estimates of coefficients β^h. The key finding is that markup dispersion increases significantly and persistently. The response of markup dispersion peaks at about two years after the shock and reverts back to zero afterwards. Whether we compute markup dispersion within two-digit or four-digit industry-quarters changes this result by little.

The specification of equation (4) implicitly assumes that the effects of monetary policy shocks are symmetric in the sign of the shock. However, in a large class of New Keynesian models, solved via a second-order approximation, markup dispersion increases in response to both positive and negative shocks; cf. figure H5 in appendix H. To investigate whether markup dispersion responds asymmetrically to shocks of different sign, we separately estimate the separate effects of contractionary and expansionary monetary policy shocks. To be precise, we replace $\varepsilon_t^{\mathrm{MP}}$ by the two sign-dependent shocks in specification equation (4). Panel b of figure 2 shows the sign-dependent responses of (within four-digit industry-quarter) markup dispersion. The evidence suggests that the responses are indeed symmetric in shock sign. Although contractionary monetary policy shocks significantly increase markup dispersion, expansionary shocks significantly lower markup dispersion. In addition, the estimated magnitudes are comparable across shock signs. The results in panels a and b prove robust in a large number of

FIGURE 2.—RESPONSES OF MARKUP DISPERSION TO MONETARY POLICY SHOCKS



Panel a shows the responses of markup dispersion to a one standard deviation monetary policy shock, coefficients β^h in equation (4). Panel b shows the sign-dependent responses of markup dispersion to a one standard deviation contractionary and expansionary monetary policy shock, respectively. Panel c shows the response of markup dispersion using the *accounting profits* and *user cost approach*, respectively. Panel d shows the response of markup dispersion using the *baseline with SGA* approach that adds SGA to the costs of goods sold, and the *cost shares* approach that uses a ratio estimator of four-digit industry-quarter-specific cost shares as output elasticities, and constructs markup dispersion within two-digit industry-quarters. The shaded and bordered areas indicate one standard error bands based on the Newey-West estimator.

dimensions, including alternative measures of markups, as we discuss in section IID.

C. *Aggregate Productivity*

Fluctuations in markup dispersion lead to changes in allocative efficiency of inputs across firms and thereby to fluctuations in aggregate TFP. To characterize this link, we build on Hsieh and Klenow (2009) and Baqaee and Farhi (2020). In a model with monopolistic competition and Dixit–Stiglitz aggregation, aggregate TFP approximately follows

$$\Delta \log \mathrm{TFP}_t = -\frac{\eta}{2} \Delta \mathbb{V}_t(\log \mu_{it}) + \left[\Delta \text{ exogenous productivity}\right], \quad (5)$$

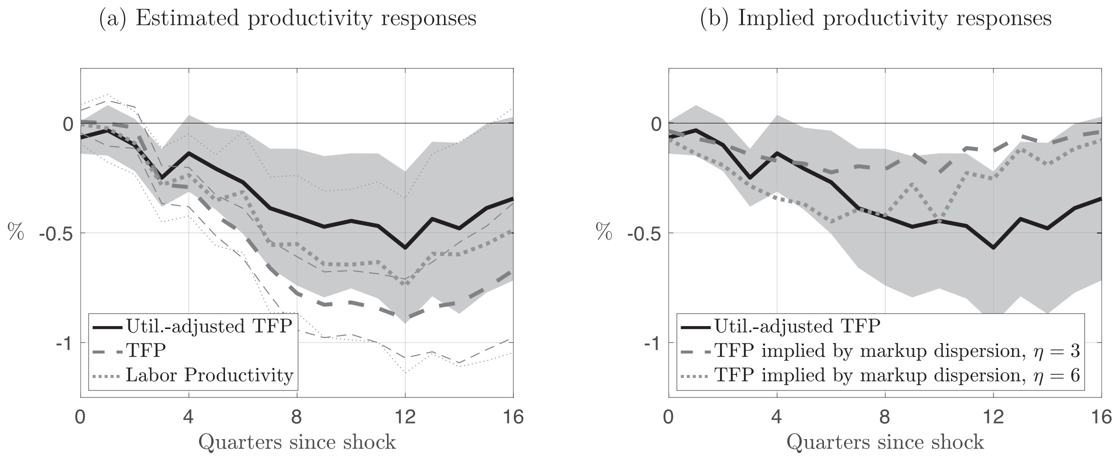
where η is the substitution elasticity between variety goods. The details of the derivation are provided in appendix E.1.[7] An increase in the variance of log markups by 0.01 lowers aggregate TFP by $\frac{\eta}{2}$%. To provide some intuition for this link, first, suppose firms are homogeneous. Aggregate output is maximal for given aggregate inputs if all firms produce the same quantity, which implies equal markups across firms. If instead firms have heterogeneous productivity and demand shifts, the efficient allocation of inputs is not homogeneous across firms, but still implies equal markups. Conversely, markup dispersion is associated with an allocation of inputs across firms that implies aggregate TFP losses.

We empirically estimate the aggregate productivity response to monetary policy shocks and compare it with the implied productivity response according to equation (5) and the estimated response of markup dispersion in figure 2a. We consider aggregate TFP and utilization-adjusted aggregate TFP from Fernald (2014), as well as labor productivity, and estimate their responses to monetary policy shocks through equation (4).[8] Panel a of figure 3 shows that the

[7] In the calibrated New Keynesian model of section IV, equation (5) closely matches the comovement of aggregate TFP and markup dispersion; cf. figures 5b and 5f.

[8] Aggregate TFP is $\Delta \log \mathrm{TFP} = \Delta y - w_k \Delta k - (1 - w_k)\Delta \ell$, with Δy real business output growth, w_k capital income share, Δk real capital

FIGURE 3.—AGGREGATE PRODUCTIVITY RESPONSE TO MONETARY POLICY SHOCKS



Panel a shows the responses of aggregate productivity measures to a one standard deviation contractionary monetary policy shock. Panel b shows the imputed response of TFP, implied by the response of markup dispersion within four-digit industry-quarters, according to $\Delta \log \mathrm{TFP}_t = -\frac{\eta}{2} \Delta \mathbb{V}_t (\log \mu_{it})$ [see equation (5)] and using $\eta = 3$ and $\eta = 6$, respectively. Alongside, it shows the empirical response of utilization-adjusted TFP from panel a. The shaded and bordered areas indicate one standard error bands based on the Newey-West estimator.

responses of all three aggregate productivity measures are significantly and persistently negative. At a two-year horizon, a one standard deviation monetary policy shock lowers aggregate TFP by 0.8%, labor productivity by 0.6%, and utilization-adjusted aggregate TFP by 0.4%. For comparison, a monetary policy shock of the same magnitude raises the federal funds rate by up to 30 basis points and lowers aggregate output by about 1% at a two-year horizon; see figure B2 in appendix B. However, aggregate factor inputs respond little and thus aggregate TFP accounts for 50%–80% of the output response at a two-year horizon.

We compute the implied TFP response by multiplying the estimated response of markup dispersion with $-\frac{\eta}{2}\%$. Panel b of figure 3 shows the implied response for $\eta = 6$, which corresponds to the estimate in Christiano et al. (2005), and $\eta = 3$, the assumption in Hsieh and Klenow (2009). The imputed TFP responses closely match the estimated TFP response within the first two years of the shock. This suggests that the response of markup dispersion is quantitatively important to understand the productivity effects of monetary policy.

An alternative explanation why aggregate productivity declines after monetary policy shocks is a reduction in R&D investment. In fact, figure B1 in appendix B shows that aggregate R&D expenditures fall after contractionary monetary policy shocks, which reconfirms the findings in Moran and Queralto (2018) and Garga and Singh (2021). Hence, some scope exists for R&D to explain part of the aggregate TFP response. However, it is less clear how much of the short-run productivity response can be explained by R&D investment. The evidence on technology adoption suggests that R&D has rather medium-run than short-run productivity effects. For example, Comin and Mestieri (2018) estimate an average adoption lag of five years for recent technologies. A sluggish effect of R&D investment on aggregate productivity is consistent with the finding in figure 3b that markup dispersion accounts for a relatively small fraction of the TFP response three to four years after a monetary policy shock.

D. Robustness

Markup estimation. We investigate the robustness of our empirical findings by considering a host of alternative markup measures. Our baseline results are robust to using these alternative markups. First, we construct (nonratio estimator) markups through an accounting profit approach and a user cost approach, following Gutiérrez and Philippon (2017) and Baqaee and Farhi (2020). The accounting profit approach uses operating income after depreciation, which is sales (`saleq`) minus costs of goods sold (`cogsq`), selling, general and administrative expenses (`xsgaq`), and depreciation and amortization (`dpq`). We compute markups from these accounting profits via $(\text{accounting profit})_{it} = (1 - \mu_{it}^{-1})\mathtt{saleq}_{it}$.

For the user cost approach, we additionally subtract the firm's capital costs (excluding depreciation) from accounting profits as in Baqaee and Farhi (2020). We construct firm-level capital stocks k_{it} via a perpetual inventory method to property, plant, and equipment; see appendix A.1. The user cost of capital is $r_t = r_t^f + RP_{jt} - (1 - \delta_{jt})\Pi_{jt+1}^K$, where r^f is the risk-free real rate, RP_j the industry-specific risk premium, δ_j the industry-specific BEA depreciation rate, and Π_j^K the industry-specific growth in the relative price of capital, based on data in Gutiérrez and Philippon (2017).[9] In

growth, and $\Delta\ell$ growth of hours worked plus growth in labor quality. Utilization adjustment follows Basu et al. (2006). Labor productivity is real output per hour in the nonfarm business sector. Figure A1d in appendix A shows the different aggregate productivity time series.

[9]The Gutiérrez and Philippon (2017) user cost is at annual frequency; we divide through by four to arrive at a quarterly rate. The data from Gutiérrez and Philippon (2017) end in 2015, so that the time sample of user cost approach markups is shorter.

general, the size of capital costs relative to total costs is modest with an average of 3.2%. This may explain the small differences between the accounting profit and user cost approaches.

Second, we construct a ratio estimator that adds selling, general, and administrative expenses (SGA) to the costs of goods sold, following Traina (2020). Third, we estimate a four-digit industry-specific translog production technology, which implies firm-quarter-specific output elasticities as in De Loecker et al. (2020). We then compute markups by combining output elasticities with revenue shares according to equation (1). Fourth, we compute four-digit industry-quarter-specific cost shares to estimate output elasticities. Specifically, we follow De Loecker et al. (2020) and compute the industry-quarter median of costs of goods sold plus 3% of the capital stock (which approximates the user cost of capital by an annual rate of 12% that includes risk premium and depreciation) divided by sales. This is a valid estimator of the output elasticity if all factors are flexible.

Our results are robust to computing markups based on these alternative measures.[10] Figure 2c shows the response of markup dispersion within four-digit industry-quarters to monetary policy shocks when using the accounting profits and user cost approach. Figure 2d shows the markup dispersion response when including SGA as well as for the cost share approach. In appendix C, we show additional results. Figure C1 shows the responses of all alternative markup dispersion measures within two-digit and four-digit industry-quarters. Figure C2 shows the responses of all markup dispersion measures conditional on the sign of the monetary policy shock.

Firm-level data treatment. We show the robustness of our results under alternative data treatments. First, we keep firms with real sales growth above 100% or below −67%. Second, we keep small firms with real quarterly sales below 1 million 2012 USD. Third, instead of dropping the top/bottom 5% of the markup distribution per quarter, we drop the top/bottom 1%. Fourth, we condition on firms with at least sixteen quarters of consecutive observations. Figure C3 in appendix C shows that markup dispersion robustly increases after contractionary monetary policy shocks. Figure C4 shows the responses of markup dispersion remain symmetric in the sign of the monetary policy shock. A well-known recent trend is the delisting of public firms. We address the concern that this may affect our results in two ways. First, when only considering firms that are in the sample for at least sixteen consecutive quarters, we find our results to be robust, as discussed above. Second, we estimate whether the number of firms in the sample responds to monetary policy shocks. Figure C5 shows that the response is insignificant and small.

[10]For the comparability of our results across markup measures, we include only firms in the robustness checks for which the baseline markup is nonmissing after the data treatment steps. Additionally we trim the alternative markups at the 1% and 99% quantiles of the quarterly markup distributions.

Monetary policy shocks. We show that our results are robust to a variety of alternative monetary policy shock series. Similar to Nakamura and Steinsson (2018), we consider the first principal component of the current/three-month federal funds futures and the 2/3/4-quarters-ahead Eurodollar futures. High-frequency future price changes may release private central bank information about the state of the economy. To control for such information effects we employ two alternative strategies. First, following Miranda-Agrippino and Ricco (2021), we regress daily monetary policy shocks on internal Greenbook forecasts and revisions for output growth, inflation, and unemployment. Second, following Jarocinski and Karadi (2020), we discard daily monetary policy shocks if the associated high-frequency change in the S&P500 moves in the same direction. To address the concern that unconventional monetary policy may drive our result, we set daily monetary policy shocks at Quantitative Easing (QE) announcements to zero. Figure C6 in appendix C shows the response of markup dispersion for all monetary policy shock series. Figure C7 shows the sign-dependent responses of markup dispersion to monetary policy shock. Figure C8 in appendix C shows the responses of aggregate productivity for all monetary policy shock series.

Great Recession We exclude the apex of the Great Recession from 2008Q3 to 2009Q2 in our baseline estimations. However, our results are robust to using the pre–Great Recession period until 2008Q2; see panels d and e of figures C3 and C4 in appendix C.

LP-IV. To revisit our main results with the LP-IV method, we replace the shocks ε_t^{MP} by the quarterly change in the one-year Treasury rate and use ε_t^{MP} as an instrument. Figures C9a and C9b in appendix C shows that our results are robust to the LP-IV method.

Proxy SVAR. Additionally, we revisit our main results through a proxy SVAR model following Gertler and Karadi (2015).[11] Figure C10 in appendix C shows the responses to monetary policy shocks in a VAR, including the one-year rate, (log) industrial production, (log) CPI, the excess bond premium of Gilchrist and Zakrajsek (2012), (log) TFP, and the baseline measure of markup dispersion (within four-digit industry-quarters). At a horizon between one and five quarters after the shock, the responses of TFP and markup dispersion are similar to our local projection results.

TFP measurement. Hall (1986) shows that the Solow residual is misspecified in the presence of market power. Hall shows that the correct Solow weights are not the income share for capital w_{kt} and labor $1 - w_{kt}$, but instead $\mu_t w_{kt}$ and $1 - \mu_t w_{kt}$, where μ_t is the aggregate markup. We examine

[11]In contrast to the proxy SVAR model, both our baseline LP approach in equation (4) and the LP-IV approach are robust to noninvertibility; see Plagborg-Møller and Wolf (2021).

the response of markup-corrected (utilization-adjusted) aggregate TFP to monetary policy shocks. We use the average markup series from De Loecker et al. (2020) to compute Hall's weights. Figure C11a in appendix C shows that the TFP response is barely different from figure 3a. In response to expansionary monetary shocks, figure C12 shows a significant increase of TFP, while the response to contractionary shocks is insignificant. We further investigate whether measurement error in quarterly TFP data is responsible for the effects of monetary policy. This problem was flagged for defense spending shocks by Zeev and Pappa (2015). We follow them in recomputing TFP using measurement-error-corrected quarterly GDP from Aruoba et al. (2016). Figure C11b shows that measurement error corrected TFP also falls after monetary policy shocks. We further show that Fernald's (2014) investment-specific and consumption-specific TFP significantly fall after contractionary monetary policy shocks; see figures C11c and C11d.

III. Heterogeneous Price-Setting Frictions

In this section, we characterize a novel mechanism through which firm heterogeneity in price-setting frictions may explain why markup dispersion increases in response to contractionary monetary policy shocks, and decreases after expansionary ones. In addition, we provide empirical evidence in support of this mechanism, and discuss alternative mechanisms.

A. *Sufficient Condition*

We first propose a sufficient condition for monetary policy shocks, which lower real marginal costs, to increase the dispersion of markups across firms. Let i index a firm and t time. A firm's markup is $\mu_{it} \equiv P_{it}/(P_t X_t)$, where P_{it} is the firm's price, P_t the aggregate price, and X_t real marginal cost. Let pass-through from marginal cost to price be defined as

$$\rho_{it} \equiv \frac{\partial \log P_{it}}{\partial \log X_t}. \tag{6}$$

This is the percentage price change in response to a percentage change in real marginal cost (without conditioning on price adjustment). The correlation between firm-level markup and firm-level pass-through is a key moment for the response of markup dispersion to shocks.

Proposition 1. *If* $\mathrm{Corr}_t(\rho_{it}, \log \mu_{it}) < 0$, *markup dispersion decreases in real marginal costs,*

$$\frac{\partial \mathbb{V}_t(\log \mu_{it})}{\partial \log X_t} < 0,$$

and markup dispersion increases if $\mathrm{Corr}_t(\rho_{it}, \log \mu_{it}) > 0$.

Proof. See appendix E.2.

Contractionary monetary policy shocks that lower real marginal costs increase the dispersion of markups if firms with higher markups have lower pass-through. Although we focus on monetary policy shocks in this paper, in principle any shock that lowers real marginal costs will raise markup dispersion as long as markups and pass-through are negatively correlated across firms.

B. *Precautionary Price Setting*

We next show that firm-level heterogeneity in the severity of various price-setting frictions may explain a negative correlation between firm-level pass-through and markup. It follows from proposition 1 that heterogeneous price-setting frictions can explain why contractionary monetary policy shocks raise markup dispersion.

Consider a risk-neutral investor that sets prices in a monopolistically competitive environment with an isoelastic demand curve and subject to adjustment costs:

$$\max_{\{P_{it+j}\}_{j=0}^{\infty}} \mathbb{E}_t \sum_{j=0}^{\infty} \beta^t \left[\left(\frac{P_{it+j}}{P_{t+j}} - X_{t+j} \right) \left(\frac{P_{it+j}}{P_{t+j}} \right)^{-\eta} Y_{t+j} - \text{adjustment cost}_{it+j} \right]. \tag{7}$$

Adjustment costs differ across firms and may be deterministic or stochastic. This formulation nests the Calvo (1983) random adjustment, Taylor (1979) staggered price setting, Rotemberg (1982) convex adjustment costs, and Barro (1972) menu costs.

Importantly, the period profit (net of adjustment costs) is asymmetric in the price P_{it} and hence in the markup μ_{it}. Profits fall more rapidly for low markups than for high markups. This gives rise to a precautionary price-setting motive: when price adjustment is frictional, firms have an incentive to set a markup above the frictionless optimal markup. Setting a higher markup today provides some insurance against low profits before the next price adjustment opportunity (Calvo/Taylor) or lowers the expected costs of future price readjustments (Rotemberg/Barro).

To characterize precautionary price setting, we study the problem in partial equilibrium. Analytically solving the nonlinear price-setting problem with adjustment costs and aggregate uncertainty in general equilibrium is not feasible. We assume that aggregate price, real marginal costs, and aggregate demand, denoted by (P_t, X_t, Y_t), follow an i.i.d. joint log-normal process around the unconditional means $\bar{P}$, $\bar{X}$, and $\bar{Y}$. The (co-)variances of innovations are σ_k^2 and σ_{kl} for $k, l \in \{p, x, y\}$.

Calvo friction. Consider a Calvo (1983) friction, parametrized by a *firm-specific* price adjustment probability $1 - \theta_i \in (0, 1)$. The profit-maximizing reset price is

$$P_{it}^* = \frac{\eta}{\eta - 1} P_t X_t \frac{\mathbb{E}_t \left[\sum_{j=0}^{\infty} \beta^j \theta_i^j \frac{X_{t+j}}{X_t} \left(\frac{P_{t+j}}{P_t} \right)^{\eta} \frac{Y_{t+j}}{Y_t} \right]}{\mathbb{E}_t \left[\sum_{j=0}^{\infty} \beta^j \theta_i^j \left(\frac{P_{t+j}}{P_t} \right)^{\eta - 1} \frac{Y_{t+j}}{Y_t} \right]}, \tag{8}$$

and we denote the associated markup by μ_{it}^*. To isolate the role of uncertainty in price setting, we focus on the dynamics around the stochastic steady state, which is described by the unconditional means $(\bar{P}, \bar{X}, \bar{Y})$. The following proposition characterizes the precautionary upward price-setting bias—relative to the frictionless environment—as a function of θ_i, and establishes a condition under which firms with lower pass-through set higher markups.

Proposition 2. *If $P_t = \bar{P}$, $X_t = \bar{X}$, $Y_t = \bar{Y}$, and $(\eta - 1)\sigma_p^2 + \sigma_{py} + \eta\sigma_{px} + \sigma_{xy} > 0$, the firm sets a markup above the frictionless optimal one and the markup further increases the less likely price readjustment is,*

$$\mu_{it}^* > \frac{\eta}{\eta - 1} \quad \text{and} \quad \frac{\partial \mu_{it}^*}{\partial \theta_i} > 0.$$

Pass-through ρ_{it} is zero with probability θ_i and positive otherwise. Expected pass-through, denoted by $\bar{\rho}_{it}$, of either a transitory or permanent change in X_t, falls monotonically in θ_i,

$$\frac{\partial \bar{\rho}_{it}}{\partial \theta_i} < 0.$$

If the above conditions are satisfied, then $\mathrm{Corr}_t(\rho_{it}, \log \mu_{it}^*) < 0$.

Proof. See appendix E.3.

A permanent decrease in real marginal costs leads to an permanent increase in the optimal reset price by the same factor. The pass-through is hence one for adjusting firms and zero for nonadjusting firms. A transitory decrease in real marginal costs increases the optimal reset price by less than the marginal cost change if the future reset probability is below one. The pass-through of adjusting firms is hence less than one and falling in price stickiness.

Staggered price setting. Consider Taylor (1979) staggered price setting and assume that firms adjust asynchronously and at different deterministic frequencies. Staggered price setting is a deterministic variant of the Calvo setup and yields very similar results.

Rotemberg friction. Consider the price-setting problem subject to Rotemberg (1982) quadratic price adjustment costs, parametrized by a *firm-specific* cost shifter $\phi_i \geq 0$, that is, adjustment $\text{cost}_{it} = \frac{\phi_i}{2}\left(\frac{P_{it}}{P_{it-1}} - 1\right)^2$. The first-order condition for P_{it} is

$$\left[(1-\eta)\frac{P_{it}}{P_t} + \eta X_t\right]\left(\frac{P_{it}}{P_t}\right)^{-\eta} Y_t = \phi_i\left(\frac{P_{it}}{P_{it-1}} - 1\right)\frac{P_{it}}{P_{it-1}} - \phi_i\beta\mathbb{E}_t\left[\left(\frac{P_{it+1}}{P_{it}} - 1\right)\frac{P_{it+1}}{P_{it}}\right]. \quad (9)$$

The following proposition summarizes our analytical results.

Proposition 3. *If $P_{t-1} = P_t = \bar{P}$, $X_t = \bar{X}$, $Y_t = \bar{Y}$, and $\frac{\sigma_{px}}{\sigma_p\sigma_x} > -1$, then up to a first-order approximation of equation (9) around $\phi_i = 0$, it holds that*

$$\mu_{it} \geq \frac{\eta}{\eta - 1} \quad \text{and} \quad \frac{\partial \mu_{it}}{\partial \phi_i} \geq 0, \quad \text{with strict inequality if } \phi_i > 0.$$

If in addition $\eta \in (1, \tilde{\eta})$, where $\tilde{\eta} = 1 + (\exp\{\frac{3}{2}\sigma_p^2 + \frac{3}{2}\sigma_x^2 + 4\sigma_{px}\} - \exp\{\sigma_{px}\})^{-1}$, the pass-through, of either a transitory or permanent change in X_t, falls monotonically in ϕ_i,

$$\frac{\partial \rho_{it}}{\partial \phi_i} < 0.$$

If the above conditions are satisfied, then $\mathrm{Corr}_t(\rho_{it}, \log \mu_{it}) < 0$.

Proof. See appendix E.4.

Menu costs. Consider the price-setting problem subject to firm-specific menu costs. Because of the asymmetry of the profit function, price adjustment is more rapidly triggered for markups below the frictionless optimal markup than above. Thus, a higher reset markup may be optimal to economize on adjustment costs. Analytical results, however, are not available for the fully nonlinear menu cost problem. Instead, we investigate this problem quantitatively. We find that markups increase in menu costs, consistent with precautionary price setting. Consequently, the correlation between pass-through and markup is negative. More details on calibration, solution, and results are provided in appendix F.

C. *Empirical Evidence for the Mechanism*

We corroborate the mechanism by considering two testable implications. First, firms with higher markups adjust prices less frequently. Second, monetary policy shocks increase the relative markup of firms that adjust prices less frequently. We show that both implications are supported empirically.

For the subsequent empirical analysis, we use data on price adjustment frequencies together with the data described in section II. We observe average price adjustment frequencies over 2005–2011 for five-digit industries, computed in Pasten et al. (2020) from PPI microdata.[12] We further use the Compustat segment files, which provide sales and industry codes of business segments within firms. The firm-specific sales composition across industries allows us to compute firm-specific price adjustment frequencies as sales-weighted average of industry-specific price adjustment frequencies. We expect this procedure to underestimate the true extent of heterogeneity across firms, which we expect will bias our subsequent regression coefficients toward zero because of attenuation bias.[13] For some firms, Compustat segment files are not available, and for others, they report only one segment per firm. We can construct firm-specific price

[12] We thank Michael Weber for generously sharing the data with us.

[13] A sufficient condition for downward bias is that the error in the measured firm-specific price adjustment frequencies is independent of the true unobserved firm-specific price adjustment frequencies.

TABLE 1.—MARKUPS AND PRICE STICKINESS

(a) Regressions of markups on implied price duration

	log(markup)			
	Baseline		Accounting profits	User cost approach
Implied price duration	0.0537	0.0472	0.00706	0.00882
	(0.0180)	(0.0155)	(0.00300)	(0.00344)
Additional controls	No	Yes	Yes	Yes
Two-digit industry FE	Yes	Yes	Yes	Yes
Observations	3,857	3,857	3,806	3,798
Adjusted R^2	0.145	0.228	0.237	0.184

(b) Regressions of markups on price adjustment frequency

	log(markup)			
	Baseline		Accounting profits	User cost approach
Price adjustment frequency	−0.391	−0.336	−0.0501	−0.0600
	(0.0999)	(0.0860)	(0.0199)	(0.0214)
Additional controls	No	Yes	Yes	Yes
Two-digit industry FE	Yes	Yes	Yes	Yes
Observations	3,857	3,857	3,806	3,798
Adjusted R^2	0.151	0.231	0.237	0.184

Regressions of firm-level markup on firm-level price adjustment frequency and implied price duration, respectively. The regressions with additional controls include firm-level size, liquidity, and leverage as regressors. Standard errors are clustered at the two-digit industry level and shown in parentheses.

adjustment frequencies for 25% of firms. For the remaining firms, we use the price adjustment frequency of the five-digit industry they operate in.[14] More details are provided in appendix A.4. To measure price rigidity, we consider both the price adjustment frequency and the implied price duration, defined as $-1/\log(1 - \text{price adjustment frequency})$.

Testable implication 1: Firms with stickier prices charge higher markups. We provide empirical evidence that firms with stickier prices tend to charge higher markups. To compare markups with average price adjustment frequencies and implied price durations for 2005–2011, we compute average firm-level markups over the same time period. Columns 1 and 3 of table 1 show that firms that have more rigid prices than other firms in the same two-digit industry charge markups significantly above the industry average. The correlation is statistically significant for both implied price duration and price adjustment frequency as measures of price rigidity. Although this correlation is consistent with precautionary price setting, it may reflect omitted factors. In columns 2 and 4 we control for firm-specific size, leverage, and liquidity, all averages over 2005–2011. The conditional correlations remain of the same sign and statistically significant at the 1% level. In table 1, we have excluded firms for which price-setting frictions are practically irrelevant, in particular, firms with a price adjustment frequency above 99% per quarter, which are about 3% of all firms. When including these, the relation between stickiness and markup remains positive, albeit somewhat less significant; see table D1 in appendix D. Note that we have not considered four-digit industry fixed effects, because for many firms our measure of rigidity is based on the five-digit industry average, which limits the variation in rigidity measures within four-digit industries.

Testable implication 2: Monetary policy shocks increase the relative markups of firms with stickier prices. We investigate whether contractionary monetary policy shocks increase the relative markup of firms with stickier prices. This is not necessarily the case if the average stickiness differs from the stickiness after monetary policy shocks, or if the marginal costs of firms with stickier prices respond differently from other firms.

We estimate panel local projections of firm-level log markups on the interaction between monetary policy shocks and firm-level price rigidity. We measure firm-level price rigidity by the price adjustment frequency or the implied price duration. Let Z_{it} denote a vector of firm-specific characteristics. We consider two specifications for Z_{it}: (i) including one of the two rigidity measures and (ii) additionally including lags of firm size (log of total assets), leverage (total debt per total assets), and the ratio of liquid assets to total assets (all in deviation from their firm-level mean). Our selection of controls is motivated by recent work in Ottonello and Winberry (2020), who study the transmission of monetary policy shocks through financial constraints. We use the panel local projection

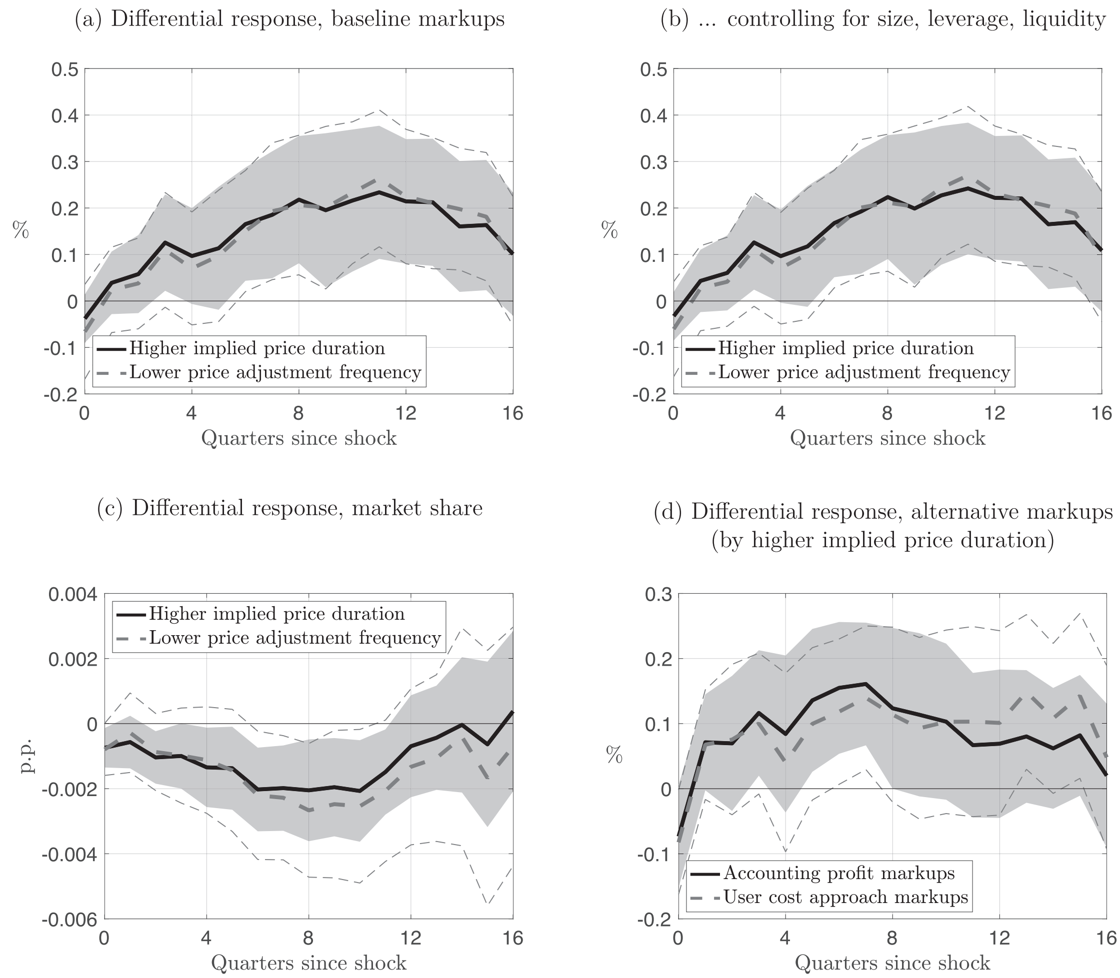
$$y_{it+h} - y_{it-1} = \alpha_i^h + \alpha_{st}^h + B^h Z_{it} \varepsilon_t^{\mathrm{MP}} + \Gamma^h Z_{it} + \gamma^h (y_{it-1} - y_{it-2}) + u_{it}^h \quad (10)$$

for $h = 0, \ldots, 16$ quarters, in which we include two-digit industry-time and firm fixed effects. To focus on the within-industry variation in the interaction between monetary policy shock and price rigidity, we subtract the corresponding two-digit industry mean from the measure of price rigidity. The main coefficients of interest are the coefficients in $\{B^h\}$ associated with price rigidity. These capture the relative markup increase for firms with stickier prices. Figure 4 shows the results. The markups of firms with stickier prices increase by significantly more after monetary policy shocks.[15] Firms with a price adjustment frequency one standard deviation above the associated two-digit industry mean increase their markup by up to 0.2% more. Importantly, the estimates are almost identical when adding controls; see panel b of figure 4. We additionally investigate the relative size response of firms with stickier prices. In particular, we consider firm-level sales market shares at the two-digit industry-quarter level. As a relative increase in markup implies relatively lower demand, we expect that firms with stickier prices become relatively smaller after contractionary

[14] Our results are robust when using only sectoral price adjustment frequencies.

[15] Driscoll–Krayy standard errors yield almost the same confidence bands as in figure 4.

FIGURE 4.—RELATIVE MARKUP AND MARKET SHARE RESPONSES OF FIRMS WITH STICKIER PRICES



Relative responses of firms with a price adjustment frequency one standard deviation below (or with an implied price duration one standard deviation above) the two-digit industry mean to a one standard deviation monetary policy shock. We plot the appropriately scaled coefficients in B^h that are associated with price rigidity in the panel local projections [equation (10)]. In panel a, Z_{it} contains only price stickiness. In panels b–d, Z_{it} also contains lagged log assets, leverage, and liquidity. The shaded and bordered areas indicate 90% error bands two-way clustered by firmand quarter.

monetary policy shocks. Indeed, we find that firms with stickier prices lose market share after contractionary monetary policy shocks, as can be seen in panel c of figure 4.[16]

Robustness. Our findings are robust along various dimensions, similar to section IID. Figure 4d shows the differential markup response of firms with more sticky prices based on the accounting profits and user costs approach. We show further robustness checks in appendix D.

D. *Alternative Mechanisms*

A key condition to explain the response of markup dispersion to monetary policy shocks is a negative correlation between firm-level markups and pass-through (proposition 1). We show that firm heterogeneity in price-setting frictions can explain this correlation, and we provide empirical evidence in support of this explanation. However, this does not preclude other mechanisms. In the following, we discuss three alternative mechanisms.

First, a nonisoelastic demand system as proposed by Kimball (1995) can explain a negative correlation between markup and pass-through and thus the response of markup dispersion.[17] Indeed, recent work by Baqaee et al. (2021) shows that under Kimball preferences (also applied, e.g., by Edmond et al., 2021), firms with a higher market share may have higher markups and lower pass-through. Even in the absence of heterogeneous price-setting frictions, this environment can qualitatively explain our empirically estimated response of markup dispersion to monetary policy shocks. Second, a negative correlation between markup and pass-through can arise in an environment with oligopolistic competition and different elasticities of substitution across and within sectors, as proposed by Atkeson and Burstein (2008). Third, heterogeneity in pass-through across firms

[16]The response of dispersion in firm-level market shares increases after monetary policy shocks, similar to markup dispersion; see figure C1f in appendix C.

[17]The evidence for Kimball-type demand curves is mixed; see Klenow and Willis (2016).

TABLE 2.—CALIBRATION

Parameter		Value	Source/target
Discount factor	β	$1.03^{-1/4}$	Risk-free rate of 3%
Elasticity of intertemporal substitution	γ	2	Standard
Elasticity of substitution between goods	η	6	Christiano et al. (2005)
Interest rate smoothing	ρ_r	0.85	Christiano et al. (2016)
Policy reaction to inflation	ϕ_π	1.5	Christiano et al. (2016)
Policy reaction to output	ϕ_y	0.05	Christiano et al. (2016)
Standard deviation of MP shock	σ_ν	0.00411	30 bp effect on nominal rate
Frisch elasticity of labor supply	φ	0.1135	Relative hours response of 11.7%
Distribution of price adjustment frequencies			
Firm type k		Share	Price adjustment frequency $1-\theta_k$
1		0.2	0.0231
2		0.2	0.0678
3		0.2	0.1396
4		0.2	0.2829
5		0.2	0.8470

The distribution of price adjustment frequencies is chosen to match the within-sector distribution reported in Gorodnichenko and Weber (2016).

can arise from financial frictions. For example, markup dispersion may increase if contractionary monetary policy shocks increase by more the financing costs of firms with lower markups.

IV. Quantitative Example

In this section, we investigate the transmission mechanism and its implications in a New Keynesian model with heterogeneous price rigidity.

A. *Model Setup*

Our model setup builds on Carvalho (2006), Kara (2015), and Gorodnichenko and Weber (2016). We discuss the model only briefly and relegate a formal description to appendix G. An infinitely lived representative household has additively separable preference in consumption and leisure, and discounts future utility by β. The intertemporal elasticity of substitution for consumption is γ, and the Frisch elasticity of labor supply is φ. The consumption good is a Dixit-Stiglitz aggregate of differentiated goods with constant elasticity of substitution η.

The economy is populated by five types of monopolistically competitive intermediate goods firms. There is an equal mass of firms of each type. All firms produce differentiated output goods with the same linear technology in labor. The only ex ante difference across firms is the exogenous price adjustment probability $1-\theta_k$, which is specific to type k. Firms set prices to maximize the value of the firm to the households. In contrast to Carvalho (2006) and the subsequent literature, which consider models with cross-sector differences in price rigidity, our model is a one-sector economy, in which price rigidity differs between firms. This speaks more directly to our empirical within-industry evidence. The monetary authority aims to stabilize inflation and the output gap. The output gap is defined as deviations of aggregate output from its natural level, defined as the flexible-price equilibrium output. Monetary policy follows a Taylor rule with interest rate smoothing and is subject to monetary policy shocks, $\nu_t \sim \mathcal{N}(0, \sigma_\nu^2)$.

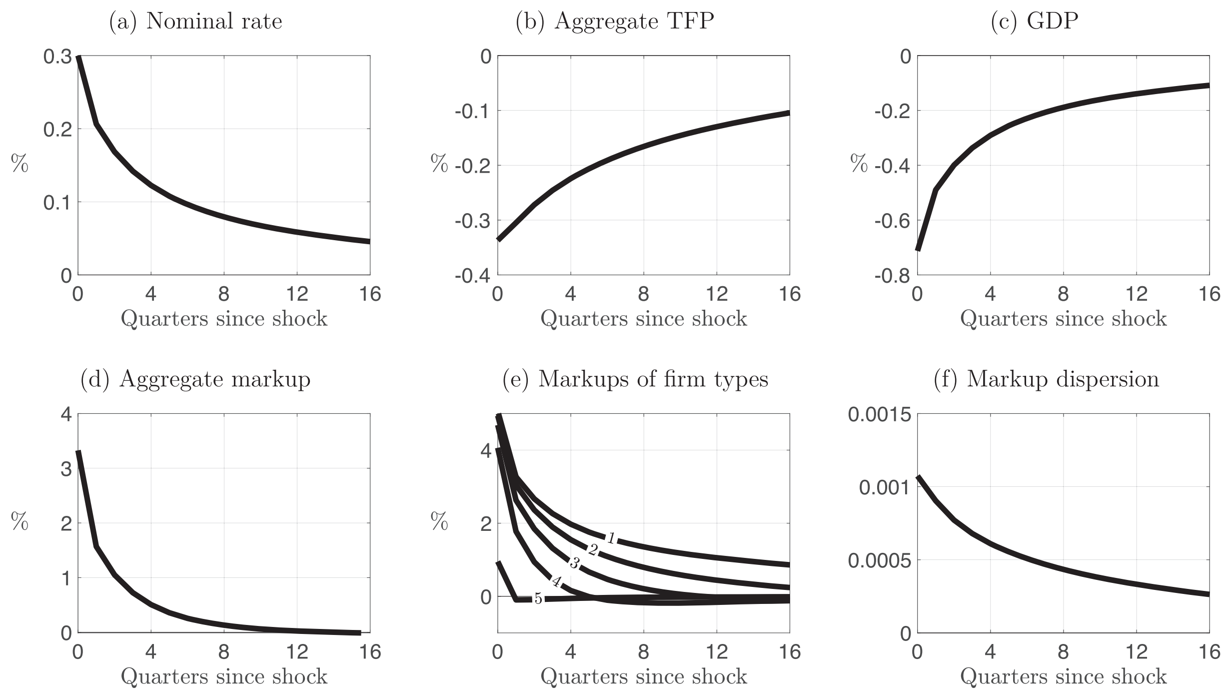
B. *Calibration and Solution*

A model period is a quarter. Table 2 summarizes the model calibration. We set the elasticity of substitution between differentiated goods at $\eta = 6$, as estimated in Christiano et al. (2005). This is conservative when compared to $\eta = 21$ in Fernandez-Villaverde et al. (2015), who study precautionary price setting as transmission of uncertainty shocks. A higher η means more curvature in the profit function, hence more precautionary price setting, and larger TFP losses from markup dispersion. We use standard values for the discount factor β and the intertemporal elasticity of substitution γ. We set the former to match an annual real interest rate of 3%, and the latter to a value of 2. We use the estimates in Christiano et al. (2016) for the Taylor rule and set $\rho_r = 0.85$, $\phi_\pi = 1.5$, and $\phi_y = 0.05$.

The parameters which play a key role in this model are the price adjustment frequencies. For the five types of firms, we calibrate θ_k for $k = 1, \ldots, 5$ to match the empirical distribution of within-industry price adjustment frequencies based on Gorodnichenko and Weber (2016). They document mean and standard deviation of monthly price adjustment frequencies for five sectors. We first compute the value-added-weighted average of the means and variances. The monthly mean price adjustment frequency is 0.1315, and the standard deviation is 0.1131. Second, we fit a log-normal distribution to these moments. Third, we compute the mean frequencies within the five quintile groups of the fitted distribution. Finally, we transform the monthly frequencies into quarterly ones to obtain $\{\theta_k\}$.

We calibrate the Frisch elasticity of labor supply internally. The hours response to monetary policy shocks is small on impact, but larger at longer horizons; see figure B2 in appendix B. The utilization-adjusted TFP response is

FIGURE 5.—MODEL RESPONSES TO MONETARY POLICY SHOCKS



Responses to a one standard deviation contractionary monetary policy shock. In panel e, the responses are the average markup responses of the firm types $k = 1, \ldots, 5$, where $k = 1$ is the stickiest and $k = 5$ the most flexible type of firms.

immediately negative but has a flatter profile at longer horizons. On average, the two responses have similar magnitude. The average difference of the response of utilization-adjusted TFP relative to the hours response, computed as the mean of $\frac{1-\text{response of util-adj. TFP in }\%}{1-\text{response of hours in }\%} - 1$ up to sixteen quarters after the shock, is 11.7%. In the model, we compute the relative hours response in the same way and target 11.7% to calibrate the Frisch elasticity. Importantly, we do not directly target the absolute magnitude of the TFP response, but only a relative quantity. The calibrated Frisch elasticity is $\varphi = 0.1135$, which is low compared to the macroeconomics literature, but is within the range of empirical estimates surveyed by Ashenfelter et al. (2010). The remaining parameter is the standard deviation of monetary policy shocks σ_ν, which we also calibrate internally. The target is the peak nominal interest rate response to a one standard deviation monetary policy shock of 30 bp; see figure B2 in appendix B. This yields $\sigma_\nu = 0.00411$.

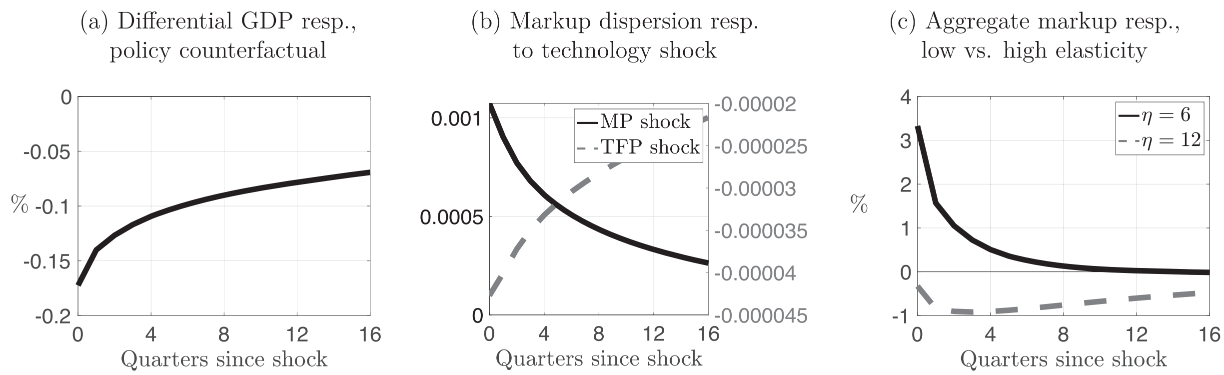
For markup dispersion to arise from precautionary price setting, it is important to use an adequate model solution technique. We rely on local solution techniques, but, importantly, solve the model around its stochastic steady state. Whereas markups are the same across firms in the deterministic steady state, differences across firms may exist in the stochastic steady state. We apply the method developed by Meyer-Gohde (2014), which uses a third-order perturbation around the deterministic steady state to compute the stochastic steady state as well as a first-order approximation of the model dynamics around it. In the stochastic steady state, precautionary price setting has large effects. Firms with the most rigid prices have 11.5% higher markups than firms with the most flexible prices.[18] As follows from proposition 1, the negative correlation between markups and pass-through implies that contractionary monetary policy shocks increase markup dispersion and lower aggregate TFP.

C. *Results*

Figure 5 shows the responses to a one-standard deviation monetary policy shock. The shock depresses aggregate demand and lowers real marginal costs. In response, firms want to lower their prices. For firms with stickier prices, however, pass-through is lower, and on average their markups increase by more. Firms with stickier prices have higher initial markups, and so markup dispersion increases. This worsens the allocation of factors across firms and thereby depresses aggregate TFP. The mechanism is quantitatively important. The increase in markup dispersion is about 75% of the peak empirical response (see figure 2), and the model explains 60% of the peak empirical response in utilization-adjusted TFP (see figure 3). In addition, the responses show the frequency composition effect described by Carvalho (2006).

[18]The only source of uncertainty in the stochastic steady state are monetary policy shocks. In principle, considering multiple shocks may increase or decrease the precautionary price setting motive. As proposition 2 shows, precautionary price setting depends on the comovement of prices, marginal costs, and aggregate demand. A sufficient condition for precautionary price setting is that all covariances between these variables are positive. This is commonly satisfied by monetary policy shocks.

FIGURE 6.—POLICY COUNTERFACTUAL AND ADDITIONAL MODEL RESULTS



Panel a shows the difference between the response to a monetary policy shock in the baseline model and the same model using a Taylor rule in which the output gap is computed by counterfactually assuming the TFP responses are driven by technology shocks. Panel b compares the response of markup dispersion to a monetary policy shock (left *y* axis) with a technology shock (right *y* axis). Panel c compares the response of the aggregate markup to a monetary policy shock for two values of the elasticity of substitution between differentiated goods.

The firms with flexible prices are quick to adjust. Hence, at longer horizons, the distribution of firms with nonadjusted prices is dominated by the stickier type of firms. This generates additional persistence in the responses.

In the model, contractionary monetary policy shocks raise markup dispersion and expansionary shocks lower markup dispersion, consistent with our empirical evidence. This response of markup dispersion critically depends on solving the model around the stochastic steady state, which allows us to capture precautionary price setting. In contrast, the deterministic steady state is characterized by zero markup dispersion. If we solve the model using a second-order approximation around the deterministic steady state, markup dispersion increases in response to both expansionary and contractionary monetary policy shocks, and irrespective of whether price rigidity is heterogeneous or homogeneous; see figure H5 in appendix H.

Even when capturing precautionary price setting, contractionary monetary policy shocks do not necessarily increase markup dispersion outside a local neighborhood around the stochastic steady state. After sufficiently large expansionary monetary policy shocks, markups of stickier firms may fall below the markups of more flexible firms. At this point, contractionary monetary policy may lower markup dispersion. We study the behavior of the model away from the stochastic steady state using a stochastic simulation of the model. The estimated response of markup dispersion on simulated data is similar and only somewhat smaller than the baseline response in figure 5; see appendix H.1 for details.

An important aspect of the monetary transmission channel in our model is the response of aggregate TFP. In contrast, traditional business cycle models assume that fluctuations in aggregate TFP are solely driven by exogenous technology shocks. This motivates us to examine the success of a Taylor rule in stabilizing output if the monetary authority in the model (mis-)perceives the aggregate TFP response to demand shocks as originating from technology shocks. Specifically, we construct a policy counterfactual, in which the only counterfactual element is natural output, and thus the output gap in the Taylor rule. Whereas model-consistent natural output responds to aggregate technology shocks but not to monetary policy shocks, counterfactual natural output responds to all changes in aggregate TFP.

We then compare the effects of a monetary policy shock in the baseline and counterfactual model.[19] Panel a in figure 6 shows the difference between the response of GDP in the counterfactual versus the baseline response.[20] Output drops by up to 0.17 percentage points more if the monetary authority attributes aggregate TFP fluctuations to technology shocks, and the response is markedly more persistent. In the counterfactual, the output gap response is dampened, which implies a less aggressive response of (systematic) monetary policy. This is similar to a lower Taylor coefficient on the output gap, and hence output falls by more. For further details and discussion, see appendix H.2.

Panel b in figure 6 shows the response of markup dispersion to a negative technology shock with the size and persistence that matches the endogenous response of TFP to a monetary policy shock.[21] The behavior of markup dispersion helps to discriminate between productivity and monetary policy shocks. It increases after contractionary monetary policy shocks but decreases after contractionary productivity shocks. So, to avoid the cost of misattributing changes in aggregate TFP to technology shocks, the monetary authority could monitor changes in markup dispersion.

The fact that aggregate TFP responds to monetary policy shocks can change the sign of the (aggregate) markup response to monetary policy shocks. This relates to a recent debate. Although monetary policy shocks raise markups in a large class of New Keynesian models, recent evidence in Nekarda and Ramey (2020) points in the opposite direction.

[19]We ensure the same interest rate response (30 bp) in baseline and counterfactual, by scaling up the size of the shock to 1.147 standard deviations in the counterfactual.

[20]Figure H2 in appendix H provides further responses for this counterfactual scenario.

[21]Figure H3 in appendix H provides further responses for the technology shock.

Following Hall (1986), the aggregate markup in our model is

$$\mu_t = \frac{\mathrm{TFP}_t}{W_t/P_t}, \tag{11}$$

where W_t/P_t denotes the real wage. In standard New Keynesian models, tighter monetary policy reduces aggregate demand, which lowers real marginal costs, and, hence, markups increase. In contrast, equation (11) shows that the aggregate markup falls if aggregate TFP falls sufficiently strongly in response to tighter monetary policy. This argument extends to sectoral and even firm-level markups, if monetary policy shocks affect TFP at more disaggregated levels. In general equilibrium, an endogenous decline in aggregate TFP will feed back into real marginal costs, which also affects markups.

Panel c in figure 6 shows the aggregate markup response to monetary policy shocks. In our baseline calibration with an elasticity of substitution $\eta = 6$ the aggregate markup raises. In some sense, that is because aggregate TFP does not fall strongly enough. We next compare our baseline results with the results when doubling the elasticity to $\eta = 12$. A larger η increases the misallocation costs of markup dispersion and thus the TFP loss after a monetary policy shock. For $\eta = 12$, the aggregate TFP response is almost twice as large; see figure H4 in appendix H. This is sufficient to explain lower aggregate markups after monetary policy shocks. Dynamically, the TFP loss leads to an increase in hours worked, which additionally increases marginal costs and lowers firm-level markups, reinforcing the effect on the aggregate markup.

To investigate the robustness of our quantitative results, we analyze the effects of monetary policy shocks in a number of model variations, including a model with real rigidities, a model with Rotemberg price adjustment, and a model with trend inflation; see appendix I.

V. Conclusion

This paper studies how markup dispersion matters for monetary transmission. Monetary policy shocks increase the dispersion of markups across firms if firms with stickier prices have higher preshock markups. Increased markup dispersion implies a change in the allocation of inputs across firms, which lowers measured aggregate TFP. Using aggregate and firm-level data, we document three new facts, which are consistent with this mechanism. First, firms that adjust prices less frequently have higher markups. Second, monetary policy shocks increase the relative markup of firms with stickier prices. Third, monetary policy shocks increase the markup dispersion across firms, and lower aggregate productivity. The empirically estimated magnitudes suggest that the response in markup dispersion is quantitatively important to understand the response of aggregate productivity. We show that an explanation for the negative correlation between markup and price stickiness is differences in price stickiness across firms. Firms with stickier prices optimally set higher markups for precautionary reasons. We show that our novel mechanism has implications for monetary policy and for the markup response to monetary policy shocks.

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