Nonlinearities of Monetary Policy across States of Price Rigidity*

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Abstract

This paper shows that monetary policy transmission varies systematically with the frequency of price adjustment. Using CPI microdata from the United Kingdom between 1996 and 2024, I document pronounced cyclical, inflationary, and seasonal variation in price adjustment rates. Embedding this variation in a nonlinear local projection framework, I find that monetary policy has substantially stronger effects on inflation when prices adjust more frequently, whereas real activity responds similarly across states. These results provide microfounded evidence that time-varying price rigidity is a key mechanism behind the state dependence of monetary policy.

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1 Introduction

Nominal rigidities shape the transmission of monetary policy. Standard New Keynesian models, built on the Calvo (1983) framework, assume that firms update prices at random, constant intervals, independent of macroeconomic conditions. A growing body of empirical evidence, however, shows that price rigidity is far from constant. Micro price data reveal that the frequency of price changes co-moves strongly with macroeconomic conditions, rising in economic downturns[1] and increasing in high-inflation[2] periods. It also exhibits pronounced seasonal patterns.[3]

State-dependent pricing models recognize that firms choose when to adjust prices optimally in response to economic conditions (e.g., Dotsey et al., 1999; Golosov and Lucas, 2007; Karadi et al., 2025). These models predict that monetary policy should have stronger and faster effects on inflation when price changes are frequent, and weaker, more sluggish effects when adjustment is infrequent. A flourishing body of empirical research supports this state-dependent view, showing that monetary transmission varies systematically across different states, including business cycles (e.g., Tenreyro and Thwaites, 2016; De Santis and Tornese, 2024), inflation regimes (e.g., Ascari and Haber, 2022), and seasonal patterns in nominal (wage) rigidity (e.g., Olivei and Tenreyro, 2007, 2010; Björklund et al., 2019). However, most of this work defines states using aggregate indicators such as economic growth, inflation, or seasonal dummies, rather than by direct measures of firms' price-setting behavior. What is missing is a microfounded link between observed variation in price rigidity and the state dependence of monetary policy.

This paper provides that link. I show that much of the observed state dependence in monetary policy transmission can be traced back to time variation in the micro-level

[1] CPI microdata from the United States (Vavra, 2014) and the euro area (Gautier et al., 2024) show that price changes become more frequent during recessions. Kryvtsov and Vincent (2021) document countercyclical sales frequencies in the United States and the United Kingdom.

[2] In high-inflation environments, the frequency of price adjustment rises, as documented for Mexico (Gagnon, 2009), the United States (Nakamura et al., 2018), Hungary (Karadi and Reiff, 2019), and Argentina (Alvarez et al., 2019). This pattern reappeared during the post-pandemic inflation surge in CPI microdata (Montag and Villar, 2022; Rudolf and Seiler, 2022), online prices (Cavallo et al., 2024), and survey data (Dedola et al., 2023).

[3] Seasonal adjustment patterns have been documented in CPI microdata for the United States (Nakamura and Steinsson, 2008, 2013), the euro area (Alvarez et al., 2006; Gautier et al., 2024), Belgium (Aucremanne and Dhyne, 2004), and France (Berardi et al., 2015), as well as in PPI data for the United States (Nakamura and Steinsson, 2008) and Denmark (Dedola et al., 2021).

frequency of price adjustment. To do so, I proceed in two steps.

First, to measure price rigidity, I calculate the monthly frequency of price changes for the United Kingdom (UK) from 1996 to 2024 using the micro price quotes collected by the Office for National Statistics (ONS) to construct the Consumer Prices Index (CPI). The dataset contains more than 32 million price observations from over 20,000 outlets across 12 UK regions, covering 1,319 narrowly defined items and over 60 percent of the CPI basket. The measure captures the share of prices that change in a given month and directly reflects firms' price-setting decisions. The data reveal substantial and systematic variation in price rigidity over time. The frequency of price adjustment is strongly countercyclical, rises with inflation, and exhibits a pronounced seasonal pattern: it declines monotonically across the four quarters and within each quarter, generating regular peaks in January, April, July, and October. These patterns are robust to excluding price changes due to temporary sales and product substitution and hold across most CPI components.

Second, I use this microfounded measure of price rigidity to estimate state dependence in the transmission of monetary policy. Specifically, I estimate nonlinear impulse responses to monetary policy shocks using local projections (Jordà, 2005), interacting the monetary policy shocks with the frequency of price changes so that the effects of the shocks vary smoothly with the degree of price rigidity. The remaining structure mirrors a parsimonious monetary VAR and employs external-instrument identification with high-frequency monetary policy shocks. These shocks are identified from asset-price movements around Bank of England monetary policy announcements using the medium-term "Path" factor from the UK Monetary Policy Event-Study Database (Braun et al., 2025). This empirical strategy links the microeconomic timing of price adjustment to the macroeconomic propagation of monetary shocks in a unified framework.

The results provide microfounded evidence of state-dependent monetary transmission. When prices adjust more frequently, the disinflationary effect of a contractionary monetary policy shock becomes substantially stronger, while the timing of the response remains unchanged. By contrast, the output response is largely insensitive to the degree of rigidity, and the path of the policy rate is similar across states. These results imply that variation in micro-level price flexibility primarily amplifies the nominal effects of monetary policy, rather than affecting its real effects.

Quantitatively, moving from low- to high-frequency environments increases the cumulative inflation response by roughly 30 percent—an amplification comparable to differences observed between low- and high-inflation regimes observed over the 1996–2024 sample period, between expansions and downturns, and across the seasonal cycle. By directly linking observed price-setting behavior to the strength of monetary transmission, the paper shows that variation in the frequency of price adjustment is a central mechanism underlying the nonlinearities documented in the broader empirical literature.

Beyond documenting that monetary transmission varies with the aggregate frequency of price adjustment, I show that this state dependence is highly granular. The amplification of monetary tightening arises almost entirely from contractionary shocks, with little corresponding asymmetry for easing episodes. It is driven disproportionately by the downward margin of price flexibility: periods in which firms cut prices more frequently exhibit markedly stronger disinflation in response to tightening, while the frequency of price increases plays a limited role. Sectoral analysis reveals similarly uneven patterns—nonlinear real effects originate mainly in the goods sector, where prices are more flexible, whereas nonlinear price responses are concentrated in the services sector, where rigidity is highest. Other price-setting moments also matter. The size, dispersion, and kurtosis of micro price changes each generate additional nonlinearities consistent with theories linking the cross-sectional distribution of price adjustments to monetary non-neutrality (Alvarez et al., 2016, 2025).

A comprehensive set of robustness exercises confirms the stability of these results. The nonlinear patterns persist when using alternative monetary policy shocks, alternative measures of prices and economic activity, different specifications of the nonlinear local projection model, alternative sample periods, and quarterly rather than monthly data. This indicates that the amplification of inflation responses in high-frequency states does not stem from any particular identification strategy or modeling choice.

Because the frequency of price changes is countercyclical and rises in high-inflation periods, the microfounded regimes map naturally into aggregate macroeconomic states often studied in the empirical literature: flexible-price states coincide with downturns and high-inflation episodes, while rigid-price states align with expansions and low-inflation periods. Seasonal variation provides an additional source of state dependence, generating faster price responses to shocks early in the year than later.

These results provide a microfounded explanation for why monetary policy appears more effective in some environments than others, grounding nonlinear transmission in observed price-setting behavior.

This helps reconcile mixed empirical findings. Some studies conclude that monetary policy is more powerful in recessions than expansions (e.g., Weise, 1999; Garcia and Schaller, 2002; Lo and Piger, 2005; Santoro et al., 2014; Burgard et al., 2019; Bruns and Piffer, 2021; De Santis and Tornese, 2024), while others find the opposite (e.g., Thoma, 1994; Tenreyro and Thwaites, 2016; Jordà et al., 2020; Alpanda et al., 2021). My results support the former by showing that recessions coincide with periods of elevated repricing frequencies, which strengthen disinflationary responses. Inflation provides another dimension of nonlinearity: monetary policy is often found to be less effective when inflation is low (Jordà et al., 2020) and more potent in high-trend-inflation regimes (Ascari and Haber, 2022). These patterns are consistent with state-dependent pricing models in which higher inflation raises the frequency of price adjustment, but existing empirical work typically proxies regimes with aggregate indicators rather than direct evidence on micro rigidities. My results fill this gap by showing that variation in micro price flexibility itself explains these regime differences. Finally, the findings complement research exploiting seasonal or institutional variation in wage rigidity (Olivei and Tenreyro, 2007, 2010; Björklund et al., 2019): whereas that literature focuses on wages, this paper provides a parallel mechanism based on prices.

Overall, the contribution of this paper is to show that many forms of state dependence can be rationalized through a single, unifying channel: variation in the micro-level frequency of price adjustment. By measuring this channel directly, the paper brings the empirical state-dependence literature closer to its microfoundations and clarifies when and why monetary policy has stronger or weaker effects on prices.

The remainder of the paper is organized as follows. Section 2 introduces the micro price data and uses the frequency of price changes to document variation in price rigidity over time. Section 3 presents the econometric method based on a nonlinear local projection framework and introduces the estimation data and monetary policy shocks. Section 4 provides the main results on frequency-dependent transmission of monetary policy. Section 5 examines sources of heterogeneity in price-setting behavior that account for the observed state dependence in monetary policy transmission,

including shock asymmetry, asymmetry in price increases and decreases, heterogeneity across sectors, and additional price-setting moments beyond the frequency of price changes. Section 6 provides an extensive set of robustness checks. Section 7 concludes.

2 Empirical evidence on the non-uniformity of price rigidity over time

This section introduces the micro price data (Section 2.1) and uses the frequency of consumer price changes (Section 2.2) to document the non-uniformity of price rigidity in the United Kingdom from 1996 to 2024 (Section 2.3).

The dataset[4] is uniquely suited for this purpose for several reasons. First, it has broad coverage of the consumption basket, which is essential for analyzing the effects of monetary policy on the broader economy. In particular, the coverage is much broader than with (supermarket or household) scanner data or web-scraped prices. Second, the data contain granular information on price observations, including information on temporary sales and product substitutions, which are important factors to consider when estimating price rigidity. Third, the Office for National Statistics (ONS) publishes some of the microdata underlying the UK CPI on its website, making the data available to the public and for research purposes. Furthermore, the United Kingdom represents a geography for which key data are available at a monthly frequency (e.g., monthly estimates of real growth domestic product) and for which monetary policy shocks have been identified in multiple ways using high-frequency identification schemes, all of which are essential to the empirical strategy and contribute to the quality of the analysis.

2.1 UK CPI microdata

The data I use to estimate price rigidity are the monthly price quotes collected by the Office for National Statistics (ONS) to construct the Consumer Prices Index (CPI) for the United Kingdom. The data span the period from February 1996 to

[4]Several studies have previously analyzed these data. In particular, Bunn and Ellis (2012) document stylized facts about price-setting behavior in consumer prices in the United Kingdom from 1996 to 2006. More recently, Davies (2021a) examines price setting during the COVID-19 pandemic. Other studies include Hahn and Marenčák (2020), Kryvtsov and Vincent (2021), Blanco (2021), Davies (2021b), or Adam and Weber (2023).

December 2024 and cover the economic territory of the United Kingdom, excluding the offshore islands. The raw data consist of over 42 million individual price quotes[5], all sampled on a monthly basis from over 20,000 outlets and 12 regions (plus central collection) across the UK. For any given sales outlet[6], data collectors find the most popular and regularly available goods and services and record the prices as well as the information to identify the products uniquely. At the most granular level of product disaggregation, there are 1,396 so-called expenditure items[7] categorized in the Classification of Individual Consumption by Purpose (COICOP). Using the available product and outlet characteristics, I identify 2.6 million unique price spells.[8]

The publicly[9] available micro price data do not contain all the price information underlying the UK CPI. In particular, they do not include administered prices and prices of central items[10]. In addition, the imputed prices of seasonal items[11] are excluded in their off-season.

To improve the informativeness of the data, I exclude several types of observations a priori. First, I discard duplicate price quotes for a given item, shop, and region

[5]Price quotes reflect transaction prices. The CPI measures consumption expenditure mostly by the acquisition of goods and services, i.e., "the total value of all goods and services delivered during a given period is considered, whether or not they are wholly paid for during the period" (Office for National Statistics, 2019).

[6]The UK is divided into twelve regions, and in each region, a number of locations is determined from which outlets are selected for the local price collection.

[7]While the ONS tracks the prices of products at the barcode level, the true identity of the products is masked by narrowly defined product categories called "expenditure items" (e.g., women's cardigan, flat panel TV 26"–42", or restaurant main course). These items come from the basket of goods and services, which contains all the products that represent household consumption expenditure in the United Kingdom. Items represent the lowest level of the survey scheme and determine the categories for which the ONS collects prices of goods and services.

[8]I define a "price spell" as the sequence of price quotes corresponding to a specific product (identified at item level) collected in a specific outlet (identified at shop code, shop type, region, and stratum levels) over time. In addition, I split price spells if there are missing price quotes for more than one month to ensure that products are not inadvertently grouped because no prices were recorded in the month prior to the price collection month.

[9]The ONS publishes some of the microdata underlying the UK CPI on its website (Office for National Statistics, 2024). Price quote data and item indices are published monthly for the most recent month.

[10]There are around 150 items for which prices are centrally collected, and the index calculation is separate from the main index construction method (Office for National Statistics, 2019). Examples include items related to housing, travel fares, or computer games.

[11]In a few areas covered by the CPI, there are distinct seasonal patterns of purchasing or consumption, for example, for some clothing or garden products. In the past, some fresh fruits and vegetables were also seasonal, although this has become less apparent with the importation of products from around the world. The prices of seasonal items are usually imputed during their off-season.

for the same month. For reasons of confidentiality, the ONS does not always publish all available local information, resulting in product identifiers in the data containing duplicate price quotes. Second, I remove "invalid" price quotes that do not pass ONS cross-checking procedures and do not enter the calculation of the official CPI. Third, whenever possible[12], I exclude imputed prices from the sample.

Because measurement errors raise concerns beyond these restrictions, I further remove outliers by excluding monthly price changes larger than the 99th percentile of absolute log price changes and smaller than the 1st percentile of absolute log price changes for each expenditure item.[13] I also exclude price changes that coincide with changes in the weight or volume of the product, ensuring that price changes due solely to changes in the package size or volume are not counted as price changes. Finally, price changes in months when the UK value-added tax (VAT) rate has changed are excluded[14] when measuring price rigidity.[15] VAT rate changes often occur at the beginning of the year, which could bias the estimation of price rigidity. Table A.1 in the appendix illustrates the effect of all sample restrictions on the sample size. The resulting baseline sample consists of 32 million price quotes and 790,000 products across 1,319 expenditure items and averages more than 60 percent of the CPI basket by expenditure weights.

In measuring price rigidity, the treatment of temporary sales and product substitutions poses two challenges. On the one hand, temporary sales involve price discounts for a limited period, after which the original price is usually restored, leading to high-frequency price fluctuations. On the other hand, product substitutions replace (temporarily) unavailable or discontinued products, which can lead to price changes that result from comparing close substitutes rather than identical products. The baseline sample includes price changes due to both temporary sales and product substitution to capture all price changes relevant to the CPI when analyzing the differential effects of monetary policy across periods of flexible and rigid price adjustment. However,

[12]The ONS flags products that are missing or temporarily out of stock in a store and imputes the prices of these products.

[13]Eichenbaum et al. (2014) raise various concerns regarding mismeasurement in the context of CPI microdata. The treatment chosen to deal with these measurement errors is in line with several contributions in the literature, for example, Gautier et al. (2024) or Rudolf and Seiler (2022).

[14]In regression analyses, I control for VAT rate changes by including dummies in the corresponding months instead of excluding these months altogether.

[15]Over the sample period, the standard VAT rate decreased from 17.5% to 15.0% on 1 December 2008, increased to 17.5% on 1 January 2010, and to 20.0% on 4 January 2011.

since both temporary sales and product substitution can exhibit cyclical or seasonal patterns, I complement the analysis of the non-uniformity of price rigidity over time (in Appendix A.4 in Appendix Appendix A) and their influence on the frequency-dependent monetary policy transmission (in Appendix D) net of these two effects. To do this, I identify temporary sales and product substitution using the flags provided by the ONS.[16] To exclude temporary sales, I replace each sales price with the last observed non-sales price. To exclude product substitutions, I start a new price spell with each product substitution.

2.2 Frequency of price changes as a measure of price rigidity

To measure price rigidity, I use the monthly frequency of price changes, which captures the share of prices that change in a given month. Formally, the frequency of price changes for item i in month t is calculated as

$$f_{i,t} = \frac{\sum_{p,s} \omega_{p,s,t} I_{p,s,t}}{\sum_{p,s} \omega_{p,s,t}}, \tag{1}$$

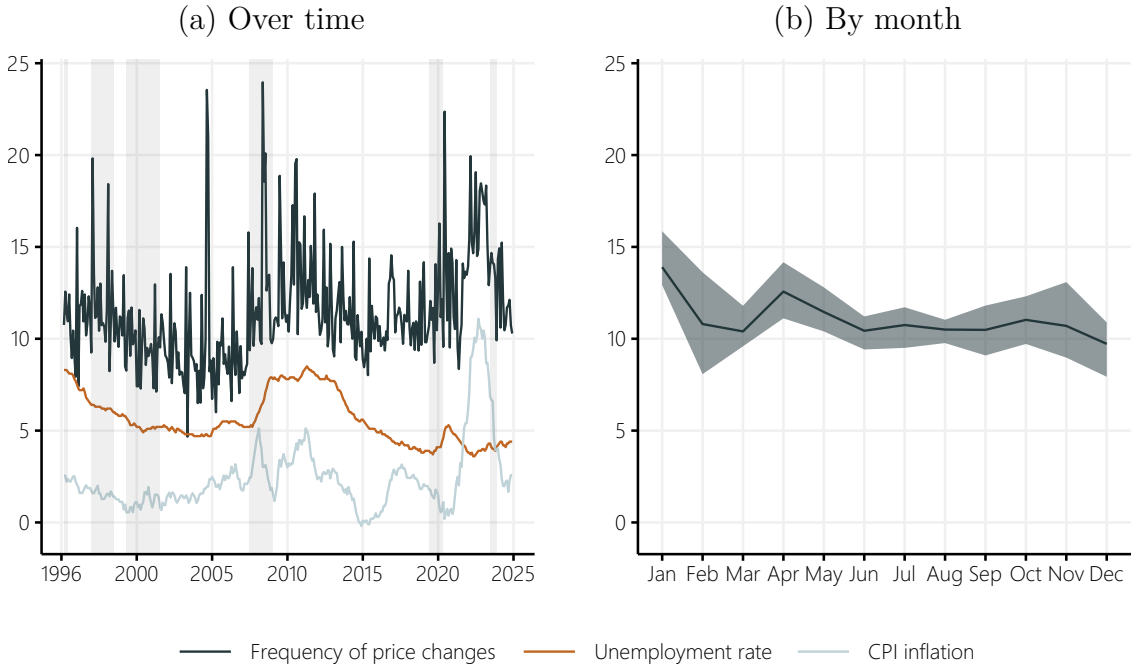
where $I_{p,s,t}$ is an indicator that takes the value one if the price of product p of item i in outlet s has changed from the previous month $t-1$, otherwise it takes the value zero.[17]

The weight $\omega_{p,s,t}$ is calculated as the CPI share of item i divided by the number of prices collected for that item in month t. To compute aggregate statistics, I use average weights over the sample period to minimize the impact of compositional changes in the basket of goods and services and aggregate item-level moments using weighted medians. Hence, the aggregate frequency of price changes, f_t, is the weighted median frequency across items using average CPI expenditure weights.

[16] For temporary sales, the ONS price collectors flag a price observation as a sales price if the discount is temporary and granted to all consumers without restriction. For product substitution, the ONS has established rules to account for quality adjustments when products are substituted (Office for National Statistics, 2019).

[17] By analogy, I calculate the frequency of price increases, $f^{+}_{i,t}$, by considering only price increases in the price change indicator, $I^{+}_{p,s,t}$, and the frequency of price decreases, $f^{-}_{i,t}$, by considering only price decreases in the price change indicator, $I^{-}_{p,s,t}$.

Figure 1. Frequency of price changes



Notes: The figure shows the frequency of consumer price changes in the United Kingdom, including price changes due to temporary sales and product substitution. Moments are computed at the item level and aggregated to weighted medians using CPI expenditure weights. The left panel shows the median frequency of price changes from February 1996 to December 2024, together with the unemployment rate and CPI inflation. Recessions (gray shaded areas) are dated by the Organisation for Economic Co-operation and Development (OECD). The right panel shows the median frequency of price changes by calendar month across the sample period, with the interquartile range indicated by the shaded area. Outliers related to VAT rate changes are excluded. All values are expressed in percent.

2.3 Non-uniformity of price rigidity over time

This section documents the non-uniformity of price rigidity over time using the frequency of consumer price changes in the United Kingdom.[18] It considers two complementary dimensions of time variation: the evolution of the frequency of price changes over the sample period and its seasonal profile across calendar months.

The left panel of Figure 1 shows the monthly frequency of price changes[19] from

[18] Appendix A.5 provides evidence on the time variation of the size of price changes, complementing the evidence on their frequency.

[19] The baseline sample includes price changes due to temporary sales and product substitution

February 1996 to December 2024, together with the unemployment rate and CPI inflation. Far from being constant, price rigidity exhibits pronounced low-frequency fluctuations. The frequency was roughly 12 percent in the mid-1990s, fell to a trough of 8.6 percent in 2003, and rose to above 14 percent around 2009–2011. It declined again in the mid-2010s before increasing sharply—to more than 17 percent—during the 2021–2023 inflation surge.

Two broad patterns stand out. First, the frequency of price changes is strongly countercyclical: it tends to rise when the unemployment rate increases and during recessions, such as the Great Recession. Second, it co-moves positively with inflation, with noticeable increases during high-inflation episodes such as the post-pandemic inflation surge. These patterns align with a growing body of evidence documenting cyclical variation in the frequency of price adjustments. For instance, while the frequency of consumer price changes in the euro area remained broadly flat from 2005 to 2019, it rose markedly during the Great Recession, when inflation reached 4.1 percent (Gautier et al., 2024). Using Swiss CPI microdata, Rudolf and Seiler (2022) find substantial increases in the frequency of price changes following the abandonment of the Swiss franc's minimum exchange rate in early 2015 and again during the inflation surge of 2021–2022. Additional evidence comes from online prices (Cavallo et al., 2024) and survey-based measures (Dedola et al., 2023), both of which show sharp increases in the frequency of price adjustment during the post-pandemic inflation surge.

To validate these visual patterns, I estimate time-series regressions of the monthly frequency of price changes on business-cycle indicators and CPI inflation, controlling for seasonality via month fixed effects and VAT changes. The coefficients (reported in Table A.2 in Appendix A.2) on both inflation and the business-cycle indicators are positive and statistically significant at the 1 percent level. The elasticities are economically meaningful: a 1-percentage-point increase in year-on-year inflation is associated with roughly a 0.7-percentage-point rise in the frequency of price changes, and a 1-percentage-point increase in unemployment is associated with a 0.2-percentage-point rise.

to capture all price changes relevant to the CPI. Appendix A.4 documents how the inclusion or exclusion of these adjustments affects the measured frequency of price changes. Their treatment shifts the level and, to a lesser extent, alters the seasonal profile, while leaving cyclical patterns unchanged.

Beyond low-frequency movements, the frequency of price changes displays pronounced seasonality. The right panel of Figure 1 shows the median frequency of price changes by calendar month, along with its interquartile range across years.

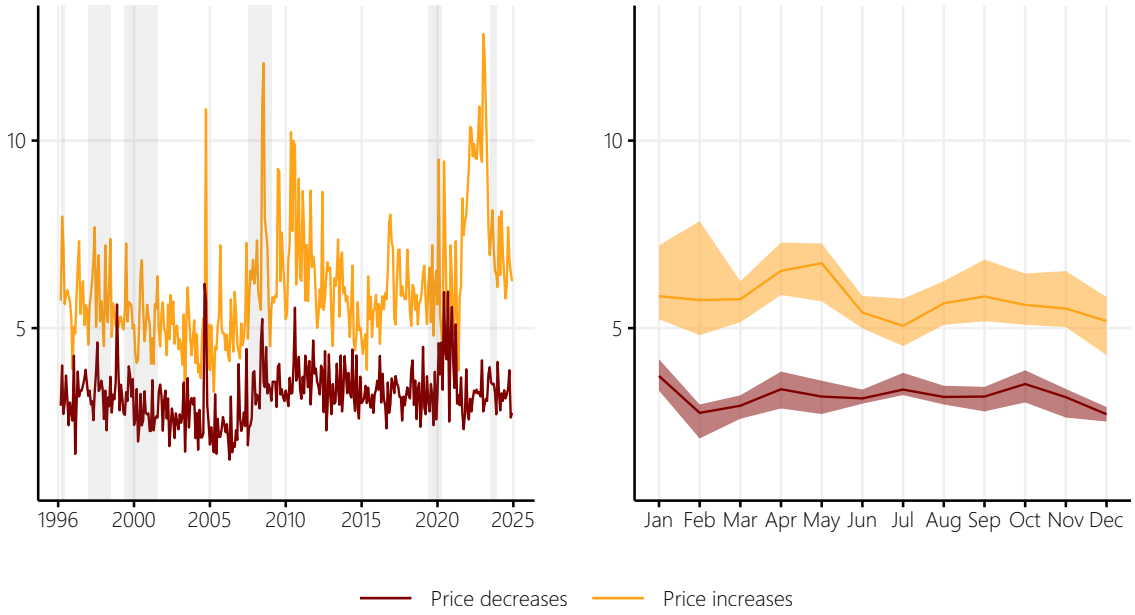
Four observations emerge: First, the frequency of price changes declines monotonically across the four quarters, with the highest levels in the first quarter and the lowest in the fourth. The difference between the peak month (January) and the trough (December) is about 4 percentage points—roughly one third of the average frequency. Second, within each quarter, the frequency is highest in the first month, declining monotonically thereafter. This gives rise to a pattern of local peaks in January, April, July, and October. Third, the intra-quarter amplitude between the months with the highest and lowest frequencies declines throughout the year: it is largest in Q1 (3.5 pp) and smallest in Q3 (0.2 pp). Fourth, January stands out as an exceptional month, with far more frequent price adjustments than any other month. Table A.3 in Appendix A.2 reports regressions of January indicators on item-level frequencies, showing significant "January effects" for 48% of items and an average weighted effect of 11 percentage points.

These seasonal patterns are consistent with evidence from a wide range of countries and datasets, including CPI microdata for the United States (Nakamura and Steinsson, 2008, 2013), the euro area (Alvarez et al., 2006; Gautier et al., 2024), Belgium (Aucremanne and Dhyne, 2004), and France (Berardi et al., 2015), as well as producer-price microdata for the United States (Nakamura and Steinsson, 2008) and Denmark (Dedola et al., 2021). The literature attributes these patterns to several mechanisms, including seasonality in costs (such as wages), product life cycles, product turnover, or seasonal sales.

To identify which types of price adjustments drive aggregate variation in price flexibility, Figure 2 decomposes the frequency of price changes into price increases (in yellow) and price decreases (in red), both over time (in the left panel) and across calendar months (in the right panel).

Throughout the sample, price increases occur almost twice as frequently as price decreases, averaging roughly 6.1 and 3.2 percent, respectively. Cyclical variation is also more pronounced for increases: the standard deviation of the frequency of increases is 1.6 percent, compared with 0.7 percent for decreases. This suggests that the business-cycle sensitivity of the overall frequency reflects primarily firms raising prices rather

Figure 2. Frequency of price increases and decreases



Notes: The figure shows the frequency of price increases (in yellow) and the frequency of price decreases (in red) in the United Kingdom, including price changes due to temporary sales and product substitution. Moments are computed at the item level and aggregated to weighted medians using CPI expenditure weights. The left panel shows the median frequencies of price increases and decreases from February 1996 to December 2024. The right panel shows the median frequencies of price increases and decreases by calendar month across the sample period, with the interquartile ranges indicated by the shaded area. Outliers related to VAT rate changes are excluded. Recessions (gray shaded areas) are dated by the Organisation for Economic Co-operation and Development (OECD). All values are expressed in percent.

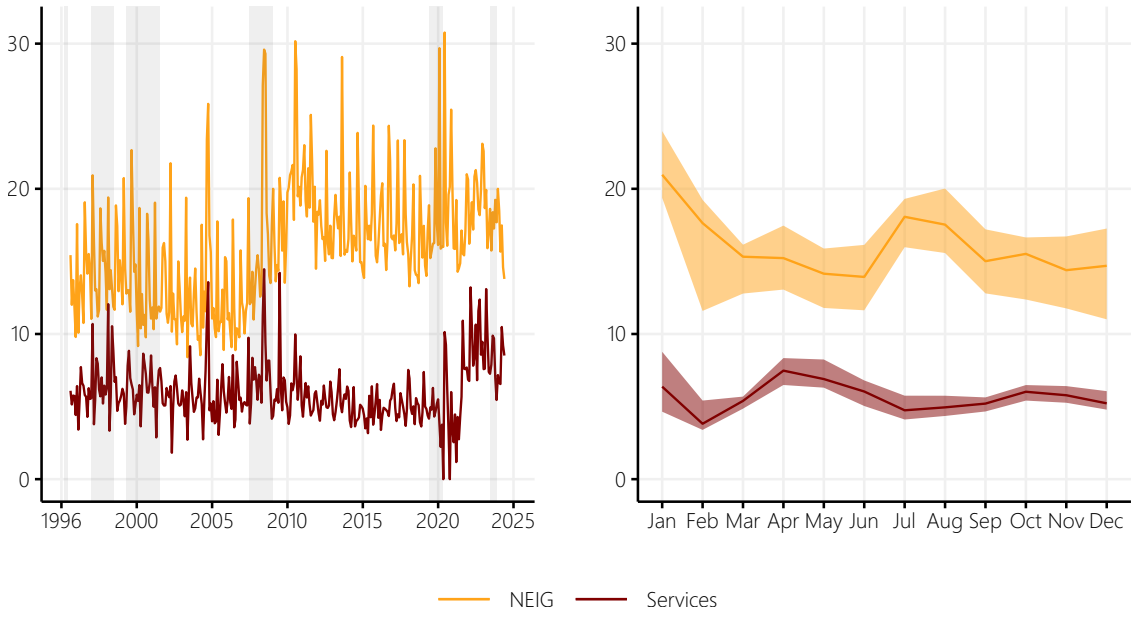
than cutting them. Regression results in Table A.4 in Appendix A.3 confirm that both increases and decreases are countercyclical, but only price increases co-move positively with inflation. Thus, the inflation–frequency relationship documented above is driven primarily by upward price adjustments.

Seasonal patterns also differ sharply across the two margins. The frequency of price changes is strongly front-loaded: it rises over the first two quarters, peaks in May at 6.7 percent, and then declines toward a flat profile of roughly 5.4 percent for the remainder of the year. By contrast, price decreases display limited seasonality, except for spikes in the first months of the quarter—at roughly 3.5 percent—with

an otherwise flat profile at around 3 percent. Figure A.1 in Appendix A.3 reports month-effect regressions confirming that seasonality stems almost entirely from price increases. When temporary sales are excluded, the seasonal component of decreases virtually disappears, while the seasonal pattern for increases remains.

Finally, Figure 3 documents substantial sectoral heterogeneity in price rigidity by comparing non-energy industrial goods (NEIG) and services.

Figure 3. Frequency of price changes in non-energy industrial goods and services



Notes: The figure shows the frequency of price changes for non-energy industrial goods (in yellow) and services (in red) in the United Kingdom, including price changes due to temporary sales and product substitution. Moments are computed at the item level and aggregated to weighted medians using CPI expenditure weights. The left panel shows the median frequencies of NEIG and services price changes from February 1996 to December 2024. The right panel shows the median frequencies of NEIG and services price changes by calendar month across the sample period, with the interquartile ranges indicated by the shaded area. Outliers related to VAT rate changes are excluded. Recessions (gray shaded areas) are dated by the Organisation for Economic Co-operation and Development (OECD). All values are expressed in percent.

Price changes are far more frequent in goods than in services. Over the sample, the median frequency in NEIG averages 16.1%, compared with 6.0% in services, consistent with the well-known sectoral gradient in nominal rigidity: goods prices adjust roughly

three times as often as service prices.

Goods prices also exhibit much stronger high-frequency volatility. The standard deviation of NEIG frequencies is 4.2 percent—considerably larger than for services (2.1 percent)—reflecting more pronounced seasonal dynamics. NEIG displays two clear seasonal peaks, a large spike in January (21.0 percent) and a second one in July (18.1 percent). By contrast, seasonality in services is weak. The frequency of service-sector price changes shows only a mild rise in April–May (peaking at 7.5 percent), and otherwise follows a nearly flat profile.

In sum, price rigidity displays substantial time variation at both low and high frequencies. The frequency of price changes moves strongly with macroeconomic conditions—rising in downturns and during high-inflation periods—and follows a pronounced seasonal cycle, with far more adjustments early in the year than late. These patterns are driven primarily by price increases and amplified by temporary sales. These temporal variations point to meaningful state-dependent elements in firms' price-setting behavior: firms adjust prices more often in some states of the economy than in others. This raises the possibility that the transmission of monetary policy itself may vary across states of the economy. The next section develops an econometric framework to test this hypothesis.

3 Econometric method

To examine whether time variation in price-setting behavior shapes the transmission of monetary policy, this section introduces an econometric framework that allows the effects of monetary policy shocks to depend on the observed frequency of price adjustment. In particular, I estimate nonlinear[20] impulse responses using local projections (Jordà, 2005) in which monetary policy shocks are interacted with the frequency of price changes. This approach provides a flexible, microfounded way of testing whether monetary policy shocks propagate differently during periods of relatively rigid versus flexible price adjustment.

[20] A growing body of literature uses nonlinear local projections to analyze state-dependent effects in empirical impulse responses, e.g., Auerbach and Gorodnichenko (2012); Caggiano et al. (2014); Tenreyro and Thwaites (2016); Ramey and Zubairy (2018); Ascari and Haber (2022).

3.1 Nonlinear local projection model

To test whether impulse responses to monetary policy shocks depend on the frequency of price adjustment, I estimate the following nonlinear local projection model:

$$y_{t+h} = \alpha_h + \beta_h e_t + \gamma_h (e_t \cdot f_{t-1}) + \sum_{k=1}^{K} \delta_{h,k} w_{t,k} + \nu_{t+h}, \tag{2}$$

where y_{t+h} denotes the variable of interest at horizon h, e_t represents the monetary policy shocks identified using an external instruments approach (see Section 3.2), f_{t-1} is the lagged frequency of price changes, $w_{t,k}$ are the K control variables (each included with up to L lags), and ν_{t+h} captures the possibly heteroskedastic and serially correlated estimation error.

Equation (2) extends the standard linear local projection framework of Jordà (2005) by allowing the effects of monetary policy shocks to vary continuously with the degree of price rigidity. The interaction term between the monetary policy shock and the lagged frequency of price changes captures the potential state dependence of impulse responses with respect to price rigidity.

The coefficient β_h measures the average, frequency-independent response of y_{t+h} at horizon h to a monetary policy shock e_t at time t, while γ_h quantifies how this response changes with the frequency of price adjustment. If γ_h has the same sign as β_h, more frequent price changes amplify the linear response implied by β_h. By contrast, if γ_h has the opposite sign, higher price flexibility dampens—or potentially reverses—the linear impulse response for sufficiently large values of the interaction variable. Whenever $\gamma_h = 0$, the impulse response to a monetary policy shock is linear and independent of the degree of price rigidity. Thus, γ_h is the main coefficient of interest for testing whether and how the propagation of monetary policy shocks depends on the frequency of price adjustment.

Under a contractionary monetary policy shock, i.e., a positive value of e_t, we expect prices and output to decline. Hence, β_h is expected to be negative for both variables. In the presence of more flexible pricing, firms adjust prices more quickly, which should lead to a larger (in absolute value) decline in prices and a smaller decline in output. Consequently, one would expect γ_h to be negative when the dependent variable is a price measure (amplifying the price response) and positive when the dependent

variable is output (attenuating the output response).

In the interaction term, the frequency of price changes enters lagged by one period to mitigate contemporaneous feedback from monetary policy shocks to price-setting behavior. A potential concern, however, is that the frequency of price changes may nonetheless respond to monetary policy, violating the exogeneity assumption required for consistent estimation of γ_h. Two arguments mitigate this concern.

First, existing empirical evidence suggests that firms adjust primarily the size rather than the frequency of price changes in response to aggregate shocks, including monetary policy shocks (e.g., Alvarez et al., 2019; Gautier et al., 2024). These studies emphasize that aggregate shocks are small relative to firm-specific shocks, which dominate pricing decisions. As a result, variation in the frequency of price changes is largely driven by idiosyncratic, rather than aggregate, shocks.

Second, I empirically test the exogeneity of the frequency of price changes within my dataset and econometric framework. Specifically, I estimate the response of the frequency and size of price changes to monetary policy shocks using linear local projections. The results, displayed in Figure B.1 in Appendix B, confirm that the frequency of price changes does not react significantly to monetary policy shocks, whereas the size of price changes decreases following a contractionary shock. These findings alleviate endogeneity concerns and support the validity of using the frequency of price changes as an exogenous interaction variable in the nonlinear local projection model.

3.2 Data and model specification

The baseline specification follows a standard monetary VAR specification, chosen to be as parsimonious as possible[21] while capturing the main transmission channels of monetary policy (e.g., Christiano et al., 2005). The vector of endogenous variables y_t comprises real gross domestic product (GDP), the core consumer price index (CPI), and the policy rate of the Bank of England ("Bank Rate"). To avoid the "price puzzle"—a counter-intuitive rise in inflation following monetary policy tightening (Eichenbaum, 1992; Sims, 1992)—I include a global commodity price index in US dollars as an exogenous forward-looking variable. To express this index in sterling,

[21] In Section 6.2, I report results from extended specifications that include additional variables. The main findings remain robust across specifications.

I further include the nominal exchange rate between the British pound and the US dollar as an endogenous variable. A detailed overview of all variables and data sources is provided in Table B.1 in Appendix B.2.

The frequency of price changes is calculated as described in Section 2.2 from CPI microdata for the United Kingdom. The measure uses the sample that includes price changes due to temporary sales and product substitution to capture all price changes relevant to the CPI.[22]

Monetary policy shocks are obtained from the UK Monetary Policy Event-Study Database (UKMPD) of Braun et al. (2025), which provides high-frequency monetary policy surprises for the United Kingdom.[23] Building on the seminal work of Gürkaynak et al. (2005) and Swanson (2021), the UKMPD identifies monetary policy surprises as intra-day asset price revisions—across interest rate futures, gilt yields, overnight index swaps, stock prices, and exchange rates—around monetary policy events[24]. The UKMPD further provides three distinct monetary policy factors that capture market reactions at different points of the yield curve: a short-term factor reflecting policy rate surprises ("Target"), a medium-term factor reflecting expected policy path revisions ("Path"), and a long-term factor reflecting quantitative easing announcements ("QE").

In the baseline model, I use the "Path" factor[25] as my preferred measure of monetary policy shocks, for several reasons. First, because it is extracted from a broad set of contracts at different maturities, it captures monetary policy actions more comprehensively than any single-instrument surprise. Second, during much of the sample period, the Bank Rate was constrained by the effective lower bound following the 2008 financial crisis, and policy operated mainly through unconventional measures such as forward guidance and quantitative easing. The "Path" factor captures these

[22]In Section 6.1, I report sensitivity checks excluding price changes due to temporary sales and product substitution. The main results prove robust to excluding these types of price changes from estimates of the frequency of price changes.

[23]In Section 6.1, I perform extensive robustness analyses using alternative monetary policy shock measures. These include variations based on different underlying contracts and maturities within the UKMPD, as well as shocks obtained from alternative identification schemes in the broader monetary policy literature. Across all specifications, the results remain qualitatively and quantitatively robust.

[24]These events include announcements by the Monetary Policy Committee (MPC) and press conferences accompanying the publication of the quarterly Monetary Policy Report. As shown in Section 6.1, the main results are robust to restricting the shocks to MPC announcement days only.

[25]Figure B.2 in Appendix B.2 shows the shock series over the sample period 1997–2024.

medium-term policy shifts. Third, it provides meaningful variation throughout the entire sample, in contrast to the "Target" factor (which is largely inert during the lower-bound period) and the "QE" factor (which captures only discrete announcement episodes). Braun et al. (2025) show that macroeconomic responses are stronger and statistically significant for the "Path" factor, confirming its empirical relevance.

To construct monthly shocks for estimation, I follow the procedure of Gertler and Karadi (2015). Specifically, I cumulate all intra-day shocks occurring within the previous 31 days and then take the monthly average. Finally, I re-scale the shocks such that a one-unit innovation corresponds to a positive surprise of 100 basis points in the Bank Rate.

All data are monthly and cover the period from 1997:07 to 2024:12.[26] The beginning of this period is determined by the date on which the Bank of England was granted operational independence over monetary policy. To control for extreme outliers during the COVID-19 pandemic, I include a series of dummy variables from March 2020 to August 2022.[27] The policy rate is expressed in percent, and all other variables are expressed in natural logarithms multiplied by 100, allowing the coefficients to be interpreted as approximate percentage-point changes. Figure B.2 in Appendix B.2 displays the series included in the baseline model.

I estimate the local projection model for horizons $h = 0, 1, \ldots H$, where I set $H = 48$, corresponding to an impulse-response horizon of four years. The lag order is set to $L = 12$. In terms of deterministics, only a constant term is included (α_h). However, the results are robust to alternative choices of H, L, and deterministic terms, as discussed in Section 6.2.

For each endogenous variable, the $H + 1$ equations corresponding to the impulse response function are estimated jointly as a system of seemingly unrelated regressions. To ensure valid inference in the presence of potential serial correlation and heteroskedasticity, I compute heteroskedasticity- and autocorrelation-consistent (HAC) standard errors using the estimator proposed by Newey and West (1987).

[26] In Section 6.2, I present the results of sensitivity checks for different subsamples. For example, I exclude the period before the Great Recession or after the onset of the COVID-19 pandemic. The main results are robust across subsamples.

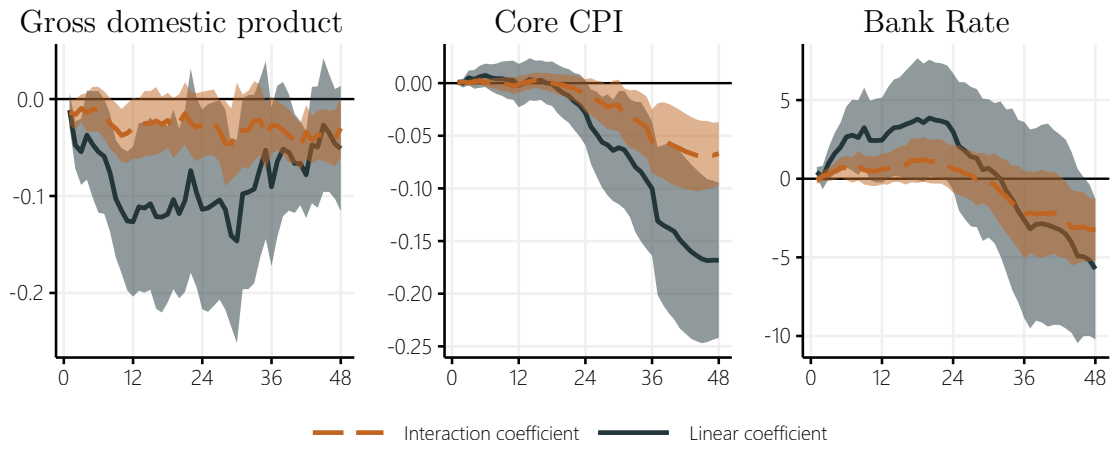
[27] Although most pandemic-related volatility fades by mid-2021, I extend the dummies for one additional year to prevent large residuals affecting lagged coefficients in the local projection estimation.

4 Frequency-Dependent Monetary Policy Transmission

This section presents the main results on the differential effects of monetary policy in periods of flexible and rigid price adjustment.

Figure 4 displays the estimated coefficients from the nonlinear local projection model in Equation (2) for each response variable: real GDP (in the left panel), core CPI inflation (in the middle panel), and the Bank Rate (in the right panel).[28] The solid black lines plot the coefficients of the linear term, $\hat{\beta}_h$, while the dashed orange lines plot the coefficients of the nonlinear interaction term, $\hat{\gamma}_h$, together with their 95% confidence intervals.

Figure 4. Nonlinear local projection coefficients



Notes: The figure shows the coefficients from estimating the nonlinear local projection model in Equation (2) for each response variable: real GDP (in the left panel), core CPI inflation (in the middle panel), and the Bank Rate (in the right panel). The solid black lines plot the coefficients of the linear term, $\hat{\beta}_h$, and the dashed orange lines plot the ones of the nonlinear interaction term, $\hat{\gamma}_h$, together with their respective 95% confidence interval bands. The data used for estimation span from 1997:07 to 2024:12. The coefficients are shown over a four-year horizon (48 months).

The linear coefficients $\hat{\beta}_h$ reproduce the well-documented[29] dynamic responses of

[28] Table C.2 in Appendix C reports the corresponding coefficients from the baseline nonlinear local projection model in Equation (2). For each response variable—real GDP, consumer prices, and the Bank Rate—the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months.

[29] Table C.1 in Appendix C provides an overview of the responses of economic activity and prices to monetary policy shocks from previous studies for the United Kingdom. For example, using the

macroeconomic aggregates to monetary policy shocks (e.g., Christiano et al., 1999). The shock has no effect on the Bank Rate on impact, but it gradually raises the rate thereafter, consistent with the interpretation of the shock as capturing unexpected changes in the anticipated policy path at medium-term horizons. The hump-shaped response that peaks after about one and a half years generates a statistically significant contraction in real activity. GDP begins to decline within six months and reaches its trough after about two and a half years. Inflation shows the familiar delayed and persistent response to a contractionary monetary policy shock. The initial effect is muted and statistically insignificant, but consumer prices start to fall around eight quarters after the shock, with the decline remaining significant throughout the estimation horizon.

Turning to the interaction coefficients, the nonlinear effects captured by $\hat{\gamma}_h$ provide evidence that the transmission of monetary policy depends on the degree of price rigidity. For consumer prices, the $\hat{\gamma}_h$ coefficients are close to zero and statistically insignificant during the first eight quarters after the shock. Thereafter, they turn negative and gradually decline, indicating that inflation responds more strongly to a monetary policy shock when the frequency of price changes is higher. Because the timing of significance for both the linear ($\hat{\beta}_h$) and nonlinear ($\hat{\gamma}_h$) coefficients coincides—both becoming negative around eight quarters after the shock—these results suggest that the frequency of price adjustment neither accelerates nor delays the timing of the price response. Instead, it amplifies its magnitude: the disinflationary effect of a contractionary monetary policy shock is stronger when prices adjust more frequently.

By contrast, the nonlinear interaction coefficients for economic activity are negative but remain statistically indistinguishable from zero throughout most of the estimation horizon. This implies that the response of real GDP to monetary policy shocks is largely independent of the degree of price rigidity. In other words, variation in the frequency of price changes primarily affects the pricing margin rather than the output margin of monetary transmission. Finally, it is important to note that these nonlinearities do not stem from differences in the monetary policy reaction itself. The coefficients on the nonlinear term for the Bank Rate are close to zero and statistically

same monetary policy shocks as in this paper, Braun et al. (2025) find a peak decline of 1.5 percent in activity and 0.45 percent in prices to a 100 basis point increase in the policy rate.

insignificant at all horizons, indicating that the policy rate response to shocks does not vary systematically with the degree of price rigidity.

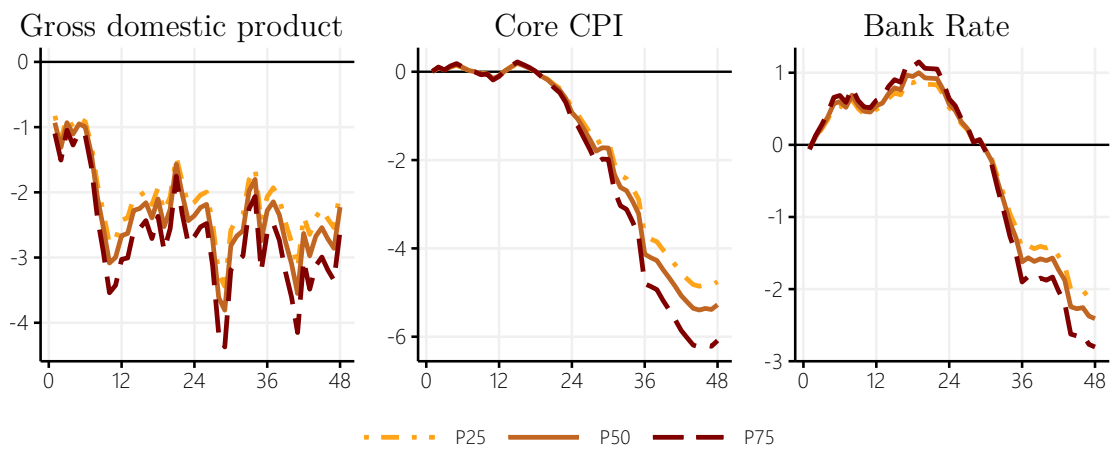
Taken together, the results indicate that the transmission of monetary policy to prices—and to a lesser extent to real activity—systematically varies with the degree of price rigidity. Inflation responds more strongly to a contractionary monetary policy shock when price changes are more frequent, while the output response remains largely invariant across states. This pattern aligns with previous evidence that items characterized by more frequent price adjustment are more responsive to monetary policy (Hong et al., 2023; Gautier et al., 2024; Alvarez et al., 2025), and supports the view that state-dependent pricing steepens the Phillips curve when firms adjust more often, thereby amplifying the nominal rather than the real effects of policy interventions.

At the same time, the findings refine and unify several strands of the empirical literature that document nonlinear monetary transmission across different macroeconomic environments. The empirical patterns documented in Section 2 show that episodes of high inflation, economic downturns, and the early-year seasonal window are the states in which price changes occur more frequently. The stronger price effects in states of more frequent adjustment are consistent with the strand of work documenting stronger monetary transmission in low-growth or recessionary states (e.g., De Santis and Tornese, 2024) and echo the empirical work showing larger monetary policy effects in high-inflation environments (Ascari and Haber, 2022). They also mirror the seasonal evidence on wage rigidity in Olivei and Tenreyro (2007, 2010), but extend the mechanism to price setting itself using microdata rather than relying on assumed timing conventions in wage bargaining. Relatedly, Cravino et al. (2020) highlight that heterogeneity in price stickiness across households generates distributional consequences of monetary shocks: prices of goods consumed by high-income households are stickier and less volatile, leading to impulse responses of their consumer price indices that are smaller than those of middle-income households. In contrast to these studies, which infer state dependence indirectly from macroeconomic indicators such as recessions or high- versus low-inflation regimes (e.g., Tenreyro and Thwaites, 2016; Jordà et al., 2020; Ascari and Haber, 2022; De Santis and Tornese, 2024), my framework ties these nonlinearities directly to the underlying microeconomic timing of actual price adjustments. By doing so, it provides empirical validation for the

central mechanism emphasized in state-dependent pricing models (e.g., Vavra, 2014), namely that changes in the frequency of adjustment alter the slope of the Phillips curve and thereby the strength of monetary policy transmission.

To further illustrate the nonlinear effects implied by the estimated coefficients, Figure 5 plots the impulse response functions (IRFs) of the three main variables—real GDP, core CPI inflation, and the Bank Rate—for three representative values of the frequency of price changes corresponding to the 25th (9.6 percent), 50th (10.9 percent), and 75th (13.0 percent) percentiles of its empirical distribution. The responses are computed over a four-year horizon following a contractionary monetary policy shock equivalent to a 100 basis point increase in the policy rate.

Figure 5. Frequency-dependent impulse response functions



Notes: The figure shows impulse response functions following a contractionary monetary policy shock estimated from the nonlinear local projection model in Equation (2) for each response variable: real GDP (in the left panel), core CPI inflation (in the middle panel), and the Bank Rate (in the right panel). The dashed-dotted yellow, solid orange, and dashed red lines plot the responses for the 25th (9.6 percent), 50th (10.9 percent), and 75th (13.0 percent) percentiles of the frequency of price changes, respectively. The data used for estimation span from 1997:07 to 2024:12. The coefficients are shown over a four-year horizon (48 months).

At the median frequency of price adjustment (10.9 percent), the impulse responses display the conventional transmission pattern following a 100-basis-point monetary tightening. Real GDP declines gradually and reaches a trough of 3.6 percent after roughly two and a half years years. The response of core CPI inflation is initially close to zero, turns negative after one and a half years, and then falls steadily, reaching a

cumulative decline of about 5.3 percentage points by the end of the four-year window. The policy rate path shows the expected mechanical increase to 1 percentage point and remains positive for two and a half years. Taken together, the median-state IRFs provide a benchmark against which the role of time-varying price rigidity can be assessed.

Varying the frequency of price adjustment significantly alters the magnitude of the inflation response. When the frequency is low (9.6 percent), the cumulative decline in core CPI inflation is considerably weaker, stabilizing at around 4.8 percentage points by the end of the horizon. When prices adjust more frequently (13.0 percent), the disinflationary effect rises to about 6.1 percentage points. This corresponds to roughly a 30 percent amplification when moving from the 25th to the 75th percentile of the empirical distribution—a sizable difference given the modest variation in the underlying microeconomic frequency. For context, these frequencies historically correspond to periods with average core CPI inflation of 1.9 percent versus 2.8 percent (Table A.2 in Appendix A.2), and they align closely with the typical seasonal swing from the beginning to the end of the year (Figure 1). Thus, the estimated state dependence is quantitatively meaningful in light of realistic time-series variation in the frequency of adjustment.

Despite the substantial change in magnitude, the timing of the inflation response—the delay before turning negative and the horizon of the trough—is nearly identical across all three states. This suggests that greater price flexibility strengthens the transmission of monetary policy without altering its speed. Output responses, by contrast, vary only modestly across states, and policy rate paths overlap almost perfectly. Taken together, these results reinforce the core finding: fluctuations in the frequency of price adjustment primarily amplify nominal effects of monetary policy while leaving real effects relatively stable.

5 Dimensions of Heterogeneity in Frequency-Dependent Monetary Transmission

Having established the baseline nonlinear effects of monetary policy as a function of the aggregate frequency of price adjustment, this section examines which specific features of price-setting behavior account for the observed state dependence in monetary

policy transmission. I explore four sources of heterogeneity. First, I test whether the nonlinearities differ between contractionary and expansionary monetary policy shocks (Section 5.1). Second, I assess whether state dependence is driven more strongly by the frequency of price increases or price decreases (Section 5.2). Third, I examine whether the results vary across sectors (Section 5.3). Finally, I consider whether alternative price-setting moments—such as the average size or dispersion of price changes—contain additional explanatory power beyond the frequency of adjustment (Section 5.4). Together, these exercises help clarify which granular margins of price flexibility shape the nonlinear transmission of monetary policy.

5.1 Nonlinear Effects Across the Direction of Monetary Policy Shocks

A natural question is whether the role of price rigidity depends on the direction of the monetary policy shock itself. To assess whether the frequency of price adjustment amplifies positive versus negative monetary policy shocks differently, I re-estimate the nonlinear local projection specification separately for contractionary (e_t^+) and expansionary (e_t^-) monetary policy shocks, allowing both the linear effect and its interaction with the frequency of price adjustment to vary across regimes.

Table 1 reports the resulting coefficients for each response variable—real GDP (in the top panel), core CPI inflation (in the middle panel), and the Bank Rate (in the bottom panel)—at horizons from one month to four years. For each shock type, the table lists the linear response coefficient $\hat{\beta}_h$ and the nonlinear interaction coefficient $\hat{\gamma}_h$, which captures how the effect of the shock varies with the frequency of price adjustment.

The linear coefficients reveal the expected qualitative asymmetry between positive and negative monetary shocks. Contractionary shocks generate declines in GDP at short horizons, while expansionary shocks produce small but positive effects on impact that strengthen at one-year horizons. For inflation, contractionary shocks reduce core CPI inflation at medium and long horizons, whereas expansionary shocks generate weaker and more muted price responses, consistent with the well-known difficulty of stimulating inflation during easing episodes. The policy rate panels confirm the mechanical increase (decrease) in response to expansionary (contractionary) shocks and show that the persistence of the shock differs across directions, especially at

Table 1. Effects Across the Direction of Monetary Policy Shocks

		Horizon				
		$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
		Gross domestic product				
Contractionary shocks	$\hat{\beta}_h^{e_t^+}$	-0.155***	-0.187	0.031	-0.146	-0.482
	$\hat{\gamma}_h^{e_t^+}$	-0.014**	-0.096*	-0.175**	-0.101**	-0.141
Expansionary shocks	$\hat{\beta}_h^{e_t^-}$	0.047*	0.452***	0.239	-0.232	0.159**
	$\hat{\gamma}_h^{e_t^-}$	0.039***	0.051	0.163*	-0.006	-0.076***
		Core consumer price index				
Contractionary shocks	$\hat{\beta}_h^{e_t^+}$	0.016	0.043	-0.027	0.127	-0.392**
	$\hat{\gamma}_h^{e_t^+}$	-0.001	-0.013	0.058*	0.066***	-0.072*
Expansionary shocks	$\hat{\beta}_h^{e_t^-}$	0.010	-0.005	-0.190	0.029	0.037
	$\hat{\gamma}_h^{e_t^-}$	0.003	-0.021	-0.107**	-0.030	0.042
		Bank Rate				
Contractionary shocks	$\hat{\beta}_h^{e_t^+}$	-1.076	0.610	-1.349	-1.121	-6.726***
	$\hat{\gamma}_h^{e_t^+}$	-0.376	-0.286	0.852	1.149	-5.075***
Expansionary shocks	$\hat{\beta}_h^{e_t^-}$	1.212**	5.391*	-5.762*	4.325*	-2.980
	$\hat{\gamma}_h^{e_t^-}$	0.089	0.726	-4.431**	-0.105	-1.592

Notes: The table reports the estimated coefficients from the nonlinear local projection model in Equation (2), estimated separately for contractionary (e_t^+) and expansionary (e_t^-) monetary policy shocks. For each response variable—real GDP, core CPI inflation, and the Bank Rate—the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequency of price adjustment. The data used for estimation span from 1997:07 to 2024:12. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.

longer horizons. Together, the linear estimates indicate that tightening and loosening episodes do not simply mirror one another, aligning with evidence of asymmetric monetary transmission in earlier work.

Turning to the nonlinear coefficients, the interaction terms indicate that the amplifying effect of the frequency of price adjustment is present but differs in magnitude across shock types. For contractionary shocks, $\hat{\gamma}_h$ is generally negative for inflation at medium and long horizons, implying that more frequent price adjustment strengthens the disinflationary impact of a monetary tightening—mirroring the baseline results in Section 4. In contrast, the interaction terms for expansionary shocks are smaller in absolute value and often statistically insignificant, suggesting that variation in price rigidity plays a more limited role in shaping the response to monetary easing. For real GDP, the interaction coefficients remain modest in both regimes, indicating that the degree of price flexibility primarily affects nominal rather than real transmission. Taken together, the results show that the state dependence documented earlier is predominantly a feature of contractionary monetary policy shocks: price flexibility amplifies the effects of monetary tightening, but plays a weaker role in easing episodes.

5.2 Nonlinear Effects Across the Direction of Price Changes

To assess whether the state dependence in monetary policy transmission is driven more by upward or downward price adjustments, I re-estimate the nonlinear local projection model in Equation (2) using separately the frequency of price increases (f_t^+) and the frequency of price decreases (f_t^-) as the interaction variable. Table 2 reports the estimated linear and nonlinear coefficients for real GDP, core CPI inflation, and the Bank Rate at yearly horizons.

Across all specifications, the linear coefficients $\hat{\beta}_h$ remain very similar regardless of whether the interaction is based on f_t^+ or f_t^-. For both real GDP and core CPI inflation, the expected signs and magnitudes broadly resemble the baseline estimates. GDP shows significant negative responses, and core CPI inflation displays increasingly negative linear coefficients at longer horizons. Thus, decomposing the overall frequency of price adjustment does not alter the average transmission of monetary policy.

More substantial differences appear in the nonlinear interaction coefficients $\hat{\gamma}_h$, which

Table 2. Effects Across the Direction of Price Changes

		Horizon				
		$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
		Gross domestic product				
Frequency of price increases	$\hat{\beta}_h^{f^+}$	-0.012**	-0.159***	-0.088**	-0.045	-0.090**
	$\hat{\gamma}_h^{f^+}$	-0.020**	-0.051**	-0.006	-0.018	-0.034
Frequency of price decreases	$\hat{\beta}_h^{f^-}$	-0.013**	-0.162***	-0.114**	-0.014	-0.096**
	$\hat{\gamma}_h^{f^-}$	-0.035**	-0.086**	0.001	-0.058*	-0.067**
		Core consumer price index				
Frequency of price increases	$\hat{\beta}_h^{f^+}$	0.001	-0.012	-0.050***	-0.123***	-0.165***
	$\hat{\gamma}_h^{f^+}$	-0.001	-0.003	-0.019**	-0.086***	-0.093***
Frequency of price decreases	$\hat{\beta}_h^{f^-}$	0.001	-0.011	-0.032*	-0.096***	-0.125***
	$\hat{\gamma}_h^{f^-}$	-0.002	-0.003	-0.037*	-0.122***	-0.138***
		Bank Rate				
Frequency of price increases	$\hat{\beta}_h^{f^+}$	0.053	-0.838	-3.525***	-4.387**	-1.171
	$\hat{\gamma}_h^{f^+}$	0.373***	-1.700**	-0.121	-1.658**	-0.493
Frequency of price decreases	$\hat{\beta}_h^{f^-}$	0.047	-1.142	-4.338***	-3.554**	-0.575
	$\hat{\gamma}_h^{f^-}$	0.463***	-2.209**	0.626	-2.748*	-1.068

Notes: The table reports the estimated coefficients from the nonlinear local projection model in Equation (2), estimated separately using the frequency of price increases (f_t^+) and the frequency of price decreases (f_t^-) as the interaction variable. For each response variable—real GDP, core CPI inflation, and the Bank Rate—the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequencies of price increases and decreases, respectively. The data used for estimation span from 1997:07 to 2024:12. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.

capture the frequency dependence of the shock. For core CPI inflation, both interactions yield negative nonlinear effects at medium and long horizons, but their magnitudes differ. When using the frequency of price increases, the interaction coefficients are negative but moderate. When using the frequency of price decreases, the coefficients are more strongly negative at the same horizons. The gap is sizeable across all horizons where coefficients are significant, indicating that states with more frequent price decreases exhibit substantially stronger disinflationary responses to contractionary monetary policy shocks. The interaction terms under f_t^- are also more consistently statistically significant. These results suggest that the state dependence documented in the baseline model is driven more by the downward margin of price adjustment rather than by a greater incidence of price increases.

For real GDP, the nonlinear effects remain small and largely insignificant under both specifications. At longer horizons, the coefficients on both f_t^+ and f_t^- are close to zero or only marginally significant. This mirrors the baseline result that the real effects of monetary policy do not vary systematically with the frequency of either price increases or price decreases. The evidence therefore continues to point to nominal rather than real amplification of monetary policy transmission. Finally, the estimates for the Bank Rate show some significant nonlinearities at shorter horizons, but these patterns do not persist beyond the first year. At longer horizons, the interaction coefficients become inconsistent in sign and significance, indicating that differences in the CPI responses are not mechanically driven by distinct adjustment paths of the policy rate itself.

Taken together, the results show that the frequency dependence of monetary transmission operates predominantly through the extensive margin of price decreases. Inflation falls more strongly in states where firms reduce prices more frequently, while upward adjustment frequencies play a secondary role and real activity remains largely unaffected. This underscores that downward price flexibility is the key microeconomic channel behind the nonlinearities observed in the transmission of monetary policy to inflation.

5.3 Nonlinear Effects Across Sectoral Heterogeneity

The degree of price rigidity varies significantly across sectors. In the goods sector, the frequency of price changes is relatively high, averaging 16.1 percent over the sample period, and exhibits pronounced cyclical and seasonal fluctuations. In contrast, the services sector is characterized by a much lower frequency of price changes, averaging only 6.0 percent, with weaker cyclical and seasonal patterns.

How do these sectoral differences in price setting influence the transmission of monetary policy? To address this question, I examine whether economic activity and prices respond differently to monetary policy shocks in the goods versus services sector. Specifically, I re-estimate the nonlinear local projection model in Equation (2) separately for each sector, distinguishing between flexible and rigid price adjustment regimes based on sector-specific frequencies of price changes as the interaction variable. For economic activity, I use GDP data for the production and services sectors, while for prices, I analyze the goods (non-energy industrial goods, NEIG) and services components of the CPI. Table 3 reports the resulting linear and nonlinear coefficients for the sectoral measures of real activity and consumer prices.

The estimated linear coefficients $\hat{\beta}_h$ reveal meaningful differences across sectors. For real GDP in the goods-producing sector, the response to a contractionary monetary policy shock is strongly negative at medium and long horizons, indicating a sizable and persistent decline in output. In contrast, services GDP exhibits a much weaker response that vanishes by the end of the estimation horizon. This suggests that, on average, monetary policy has more pronounced real effects in the goods sector. The linear price responses display the opposite pattern. Goods prices show limited and only intermittently significant declines, whereas services prices fall more persistently and with larger magnitudes. These differences are consistent with slower and less frequent price adjustment in services, leading to a delayed but ultimately stronger disinflationary response.

Turning to the interaction terms, the sectoral patterns differ sharply. For real GDP, the nonlinear coefficients are small across both sectors, but only the goods sector shows systematic significance. Output in the goods sector becomes more responsive to monetary tightening when prices adjust more frequently. In services, by contrast, the estimates of $\hat{\gamma}_h$ are generally modest and do not display a consistently significant

Table 3. Effects Across Sectoral Heterogeneity

		Horizon				
		$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
		Gross domestic product				
NEIG	$\hat{\beta}^g_h$	-0.010	-0.133***	-0.117*	-0.062	-0.175***
	$\hat{\gamma}^g_h$	-0.010**	-0.019*	-0.029**	-0.024	-0.047***
Services	$\hat{\beta}^g_h$	0.020**	-0.019	-0.048	-0.126***	0.007
	$\hat{\gamma}^s_h$	-0.006	-0.042**	-0.006	-0.068**	-0.032
		Core consumer price index				
NEIG	$\hat{\beta}^g_h$	0.001	-0.050**	-0.010	-0.045	-0.070**
	$\hat{\gamma}^g_h$	0.007**	-0.001	-0.006	-0.018	0.007
Services	$\hat{\beta}^g_h$	0.006	-0.020*	-0.019	-0.072***	-0.077***
	$\hat{\gamma}^s_h$	-0.005	-0.004	-0.015	-0.066**	-0.027***
		Bank Rate				
NEIG	$\hat{\beta}^g_h$	0.496***	2.313***	5.667***	1.944	1.104
	$\hat{\gamma}^g_h$	-0.051	0.004	2.144***	0.525	0.327
Services	$\hat{\beta}^g_h$	0.294	2.089**	5.024***	-3.009	2.421
	$\hat{\gamma}^s_h$	-0.221	0.197	3.298***	-3.320	2.088**

Notes: The table reports the estimated coefficients from the nonlinear local projection model in Equation (2), estimated separately for non-energy industrial good (NEIG) and services using sector-specific frequencies of price adjustment as the interaction variable. For each response variable—sector-specific GDP, sector-specific CPI inflation, and the Bank Rate—the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequencies of price increases and decreases, respectively. The data used for estimation span from 1997:07 to 2024:12. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.

pattern. Overall, state dependence in the real effects of monetary policy appears to originate primarily in the goods sector. For prices, the nonlinearities differ not only in magnitude but also in sign. In the goods sector, the interaction coefficients are small and mostly insignificant, implying that more frequent price adjustment does not meaningfully alter the response of goods prices. In services, however, the nonlinear effects are larger and consistently negative at medium and long horizons. This indicates that when services prices adjust more frequently, the disinflationary effect of monetary policy is substantially amplified.

Taken together, these results show that frequency-dependent nonlinearities in price responses are concentrated in the services sector—where prices are typically most rigid—while nonlinear real effects arise mainly in the goods sector. The aggregate state dependence documented earlier therefore reflects distinct sectoral mechanisms rather than a uniform response across the economy.

5.4 Nonlinear Effects Across Alternative Price-Setting Moments

To assess whether state dependence in the transmission of monetary policy reflects other dimensions of price-setting beyond the frequency of adjustment, I re-estimate the nonlinear local projection model in Equation (2) using alternative interaction variables that summarize additional micro-level moments of price changes. Specifically, I consider (i) the average absolute size of price changes, (ii) the dispersion (standard deviation) of price changes, and (iii) their kurtosis. Table 4 reports the estimated coefficients at yearly horizons.

A common feature across all specifications is that the linear coefficients, $\hat{\beta}_h$, remain virtually unchanged relative to one another and to the baseline using the frequency of price changes as interaction variable. A contractionary monetary policy shock lowers economic activity and produces a delayed but persistent decline in core CPI inflation, regardless of which moment is used. This confirms that the average, frequency-independent transmission of monetary policy is not sensitive to the choice of interaction variable. The relevant variation therefore lies in the nonlinear coefficients, $\hat{\gamma}_h$, which capture whether and how different price-setting moments shape the amplification or attenuation of monetary policy shocks.

First, considering the size of price changes, the interaction terms based on the aver-

Table 4. Effects Across Alternative Price-Setting Moments

		Horizon				
		$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
		Gross domestic product				
Size of price changes	$\hat{\beta}_h^{\lvert\Delta p\rvert}$	-0.011	-0.157***	-0.150***	-0.068*	0.012
	$\hat{\gamma}_h^{\lvert\Delta p\rvert}$	-0.003**	-0.008***	-0.009***	-0.004*	-0.001
Dispersion of price changes	$\hat{\beta}_h^{sd(\Delta p)}$	-0.016*	-0.161***	-0.145***	-0.076*	-0.007
	$\hat{\gamma}_h^{sd(\Delta p)}$	-0.012*	-0.041***	-0.042***	-0.021	-0.008
Kurtosis of price changes	$\hat{\beta}_h^{Kur(\Delta p)}$	-0.006	-0.167***	-0.142***	-0.093**	-0.029
	$\hat{\gamma}_h^{Kur(\Delta p)}$	-0.008**	-0.028***	-0.029***	-0.009	-0.002
		Core consumer price index				
Size of price changes	$\hat{\beta}_h^{\lvert\Delta p\rvert}$	0.001	-0.002	0.000	-0.049*	-0.070***
	$\hat{\gamma}_h^{\lvert\Delta p\rvert}$	0.000	0.000	0.000	-0.006***	-0.005***
Dispersion of price changes	$\hat{\beta}_h^{sd(\Delta p)}$	0.001	0.000	-0.010	-0.085***	-0.123***
	$\hat{\gamma}_h^{sd(\Delta p)}$	0.000	0.000	-0.006	-0.041***	-0.043***
Kurtosis of price changes	$\hat{\beta}_h^{Kur(\Delta p)}$	0.000	-0.004	-0.025	-0.089***	-0.134***
	$\hat{\gamma}_h^{Kur(\Delta p)}$	0.000	-0.001	-0.007	-0.029***	-0.028***
		Bank Rate				
Size of price changes	$\hat{\beta}_h^{\lvert\Delta p\rvert}$	0.616***	3.110*	4.102**	-1.890	-4.778**
	$\hat{\gamma}_h^{\lvert\Delta p\rvert}$	-0.018	0.168	0.231*	-0.305	-0.502***
Dispersion of price changes	$\hat{\beta}_h^{sd(\Delta p)}$	0.615***	2.947**	2.925*	-2.959	-7.385***
	$\hat{\gamma}_h^{sd(\Delta p)}$	-0.065	1.014**	1.091	-1.308	-2.992***
Kurtosis of price changes	$\hat{\beta}_h^{Kur(\Delta p)}$	0.654***	5.694***	5.419**	-1.607	-6.265***
	$\hat{\gamma}_h^{Kur(\Delta p)}$	-0.052	0.936**	1.029**	-0.840	-2.089***

Notes: The table reports the estimated coefficients from the nonlinear local projection model in Equation (2), estimated separately using alternative moments of micro price adjustments (absolute size, dispersion, and kurtosis of price changes) as interaction variables. For each response variable—real GDP, core CPI inflation, and the Bank Rate—the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequencies of price increases and decreases, respectively. The data used for estimation span from 1997:07 to 2024:12. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.

age absolute size of adjustments generate small but statistically significant nonlinear effects for real activity at short horizons, though these are economically modest and dissipate quickly. For inflation, the coefficients become negative and significant at longer horizons, indicating that in periods when individual price adjustments are larger, contractionary monetary policy produces a stronger disinflationary response.

These findings are consistent with the selection-effect mechanism emphasized in standard menu cost models (Golosov and Lucas, 2007; Karadi and Reiff, 2019; Bonomo et al., 2022). Periods in which individual price adjustments are larger reflect stronger selection, in the sense that only firms with sufficiently large desired price changes choose to adjust. In such environments, a contractionary monetary policy shock induces larger disinflation because the firms that do adjust respond disproportionately, even though the number of adjusting firms remains limited. As a result, nominal adjustment absorbs much of the shock, leaving real activity largely unaffected. The concentration of nonlinear effects in inflation rather than output in the estimates aligns closely with this prediction and provides empirical support for selection as a key channel shaping monetary transmission.

Second, for the higher-order moments of price changes, both the dispersion and kurtosis of price adjustments yield nonlinear effects on real activity: the coefficients are negative and statistically significant at shorter horizons. For inflation as well, the interaction terms are negative and statistically significant at horizons beyond three years, implying that greater cross-sectional variation in price changes strengthens the disinflationary response to a monetary tightening.

The role of kurtosis is consistent with the interpretation that periods of lower kurtosis—characterized by more price selection and hence greater flexibility in the aggregate price level—correspond to lower monetary non-neutrality Alvarez et al. (2016, 2025). In such states, the pass-through of monetary shocks to consumer prices is faster but smaller overall and less persistent, and the Phillips curve is steeper. The flexibility in the aggregate price level decouples the nominal and real sides of the economy, thereby reducing the impact of a monetary shock on output. By contrast, higher kurtosis signals less price selection and greater rigidity, which reduces output and slows but prolongs the disinflationary response and flattens the Phillips curve.

Overall, alternative price-setting moments reveal further nonlinearities in monetary transmission. In particular, the impact of monetary policy on prices and output

is larger in periods when our measures indicate high monetary non-neutrality and smaller when they indicate low monetary non-neutrality. The nonlinear estimates therefore provide time-series evidence that the distributional shape of price changes, not only their frequency, is a key determinant of monetary non-neutrality.

6 Robustness

In this section, I conduct a comprehensive set of robustness exercises. In particular, I perform additional tests regarding the identification of both price rigidity and the monetary policy shocks used in the nonlinear local projection estimates (Section 6.1). I also analyze how sensitive the results are to the model specification and data choices (Section 6.2). All tables corresponding to these sensitivity checks can be found in Appendix D.

6.1 Identification

Measurement of price rigidity The baseline specification uses a measure of the frequency of price changes that includes price changes due to temporary sales and product substitution, consistent with the idea that all price changes relevant for the CPI should enter the state variable. As a robustness check, I re-estimate Equation (2) using interaction variables constructed from all combinations of alternative treatments that include or exclude sales and substitution-related price changes. The results, summarized in Table D.1, show that the main conclusions regarding the state dependence of monetary transmission are qualitatively robust to these alternative measures of price rigidity.

Overall, the findings are more sensitive to the treatment of temporary sales than to the treatment of product substitutions. Excluding product substitutions leaves the results essentially unchanged relative to the baseline, whereas excluding sales leads to somewhat smaller and more muted responses of both output and prices. This pattern suggests that temporary sales constitute an important margin of price flexibility and may play a non-negligible role in how firms adjust to aggregate shocks, consistent with evidence on the macroeconomic relevance of sales behavior (Kryvtsov and Vincent, 2021; Rudolf and Seiler, 2022; Henkel et al., 2023).

Alternative monetary policy shocks The robustness of the results to the choice of monetary policy instrument is an important concern, given the variety of available approaches for identifying policy surprises. In addition to the baseline medium-term "Path" shocks from the UK Monetary Policy Event-Study Database (Braun et al., 2025), I re-estimate the nonlinear local projection model using a comprehensive set of alternative shocks drawn both from within the UKMPD and from external sources. These include alternative event-study measures based on different definitions of policy events, individual-asset surprises at various maturities, high-frequency shocks from Cesa-Bianchi et al. (2020), and narrative monetary policy shocks following Romer and Romer (2004) from Cloyne and Hürtgen (2016).

Across these specifications, summarized in Table D.2, the qualitative results remain stable. Using alternative event constructions within the UKMPD yields responses very similar to the baseline, confirming that the key findings do not depend on the precise definition of monetary policy announcements. Shocks derived from individual assets display weaker and less precise responses, but they nonetheless preserve the overall pattern of stronger inflation responses in states of more frequent price adjustment. The external high-frequency and narrative shocks also produce dynamics aligned with the main results. Taken together, the evidence shows that the state-dependent effects of monetary policy documented in the paper are robust to a wide range of identification strategies and are not an artifact of any particular shock series.

6.2 Specification and data choices

Variable selection and model specification The baseline specification is intentionally parsimonious, including only the variables needed to capture the core transmission of monetary policy. To assess whether omitted channels materially affect the results, I estimate an extended version of the model that augments the baseline model with the 1-year government gilt yield, the producer price index (PPI), the US BAA corporate spread, and the FTSE 100 index (Table D.3). The coefficients on the main variables remain highly stable across horizons, confirming that the baseline dynamics are not sensitive to a richer information set. The estimated responses of the additional variables are both economically plausible (such as a decline in PPI following a contractionary shock) and broadly consistent with previous findings using the same shocks (Braun et al., 2025).

I then evaluate the robustness of the results to the choice of activity and price indicators. Specifically, I replace real GDP with the industrial production index, and I estimate versions of the model using alternative inflation measures: the headline CPI, the CPI including owner-occupiers' housing costs (CPIH), the Retail Price Index (RPI), and the producer price index (PPI). Across all alternatives, the core patterns in both the linear and nonlinear coefficients remain very similar to the baseline (Table D.4). This confirms that the documented state dependence does not hinge on a specific choice of output or price series.

Finally, I examine the sensitivity of the results to the lag structure and the deterministic components of the model. Varying the lag order from two to twenty-four lags (Table D.5) produces highly consistent results, with only minor changes in precision. Likewise, estimating specifications without a constant, with a constant only, and with both a constant and linear trend (Table D.6) leaves the main qualitative conclusions unchanged. Overall, these robustness checks indicate that the findings are not driven by particular modeling choices and that the nonlinear responses are a stable feature of the data.

Sample and data frequency To ensure that the estimated state dependence is not driven by specific historical episodes, I re-estimate the model on a range of subsamples. Excluding either the Great Recession or the COVID-19 pandemic yields responses that closely mirror the baseline patterns, both for output and prices, indicating that the main findings are not dominated by these large macroeconomic disruptions. While some subsample estimates are naturally less precise—reflecting shorter samples and calmer interest rate environments—the underlying nonlinear relationship between monetary policy and the frequency of price adjustment remains intact (Table D.7).

I also examine robustness to the use of quarterly data by re-estimating the model with GDP and the GDP deflator, aggregating all remaining monthly variables (including the state variable) to the quarterly frequency. The resulting dynamic responses are broadly aligned with those obtained from monthly data (Table D.8). As expected, the estimation is less precise when moving to the lower-frequency specification, but the core patterns in both the linear and nonlinear coefficients remain qualitatively similar.

7 Conclusion

This paper studied how time variation in price rigidity shapes the transmission of monetary policy. Using CPI microdata for the United Kingdom from 1996 to 2024, I constructed a direct, microfounded measure of price rigidity based on the monthly frequency of price changes. The data reveal substantial and systematic variation in price adjustment: prices change more frequently during economic downturns, in high-inflation periods, and early in the calendar year. I then used this measure to estimate nonlinear local projection models in which the effects of monetary policy shocks vary continuously with the degree of price rigidity, linking the microeconomic timing of price setting directly to macroeconomic outcomes.

The results show that monetary policy has significantly stronger effects on inflation when prices adjust more frequently, while the timing of the inflation response remains unchanged. By contrast, the response of real activity is largely invariant to the degree of price rigidity, and the policy rate itself reacts similarly across states. Quantitatively, moving from low- to high-frequency environments amplifies the cumulative inflation response by about 30 percent, an effect comparable to differences observed across business cycle phases, inflation regimes, and the seasonal calendar. Taken together, these findings provide microfounded evidence that variation in the frequency of price adjustment is a central mechanism underlying state-dependent monetary transmission, primarily affecting nominal rather than real outcomes.

From a policy perspective, the results imply that the same monetary policy action can have markedly different inflationary consequences depending on the prevailing degree of price flexibility, even when real effects remain stable (Karadi et al., 2025). Models and empirical analyses that abstract from time variation in price rigidity may therefore mischaracterize the strength of monetary transmission and the trade-offs faced by central banks. Incorporating observable, micro-level measures of price-setting behavior offers a promising avenue for improving both positive and normative analyses of monetary policy.

References

Adam, Klaus and Henning Weber (2023) "Estimating the optimal inflation target from trends in relative prices," American Economic Journal: Macroeconomics, Vol. 15, pp. 1–42.

Alpanda, Sami, Eleonora Granziera, and Sarah Zubairy (2021) "State dependence of monetary policy across business, credit and interest rate cycles," European Economic Review, Vol. 140, p. 103936.

Alvarez, Fernando, Martin Beraja, Martin Gonzalez-Rozada, and Pablo Andrés Neumeyer (2019) "From hyperinflation to stable prices: Argentina's evidence on menu cost models," The Quarterly Journal of Economics, Vol. 134, pp. 451–505.

Alvarez, Fernando, Andrea Ferrara, Erwan Gautier, Hervé Le Bihan, and Francesco Lippi (2025) "Empirical investigation of a sufficient statistic for monetary shocks," Review of Economic Studies, Vol. 92, pp. 2165–2196.

Alvarez, Fernando, Herve Le Bihan, and Francesco Lippi (2016) "The real effects of monetary shocks in sticky price models: A sufficient statistic approach," American Economic Review, Vol. 106, pp. 2817–51.

Alvarez, Luis J, Emmanuel Dhyne, Marco Hoeberichts, Claudia Kwapil, Hervé Le Bihan, Patrick Lünnemann, Fernando Martins, Roberto Sabbatini, Harald Stahl, Philip Vermeulen et al. (2006) "Sticky prices in the Euro Area: A summary of new micro-evidence," Journal of the European Economic Association, Vol. 4, pp. 575–584.

Ascari, Guido and Timo Haber (2022) "Non-linearities, state-dependent prices and the transmission mechanism of monetary policy," The Economic Journal, Vol. 132, pp. 37–57.

Aucremanne, Luc and Emmanuel Dhyne (2004) "How frequently do prices change? Evidence based on the micro data underlying the Belgian CPI," Working Paper 331, European Central Bank.

Auerbach, Alan J and Yuriy Gorodnichenko (2012) "Measuring the output responses to fiscal policy," American Economic Journal: Economic Policy, Vol. 4, pp. 1–27.

Bachmann, Rüdiger, Benjamin Born, Steffen Elstner, and Christian Grimme (2019) "Time-varying business volatility and the price setting of firms," Journal of Monetary Economics, Vol. 101, pp. 82–99.

Berardi, Nicoletta, Erwan Gautier, and Herve Le Bihan (2015) "More facts about prices: France before and during the Great Recession," Journal of Money, Credit and Banking, Vol. 47, pp. 1465–1502.

Björklund, Maria, Mikael Carlsson, and Oskar Nordström Skans (2019) "Fixed-wage contracts and monetary non-neutrality," American Economic Journal: Macroeconomics, Vol. 11, pp. 171–192.

Blanco, Andres (2021) "Optimal inflation target in an economy with menu costs and a zero lower bound," American Economic Journal: Macroeconomics, Vol. 13, pp. 108–141.

Bonomo, Marco, Carlos Carvalho, Oleksiy Kryvtsov, Sigal Ribon, and Rodolfo Rigato (2022) "Multi-Product Pricing: Theory and Evidence from Large Retailers," The Economic Journal.

Braun, Robin, Silvia Miranda-Agrippino, and Tuli Saha (2025) "Measuring monetary policy in the UK: The UK monetary policy event-study database," Journal of Monetary Economics, Vol. 149, p. 103645.

Bruns, Martin and Michele Piffer (2021) "Monetary policy shocks over the business cycle: Extending the Smooth Transition framework," Working Paper 2021-07, University of East Anglia.

Bunn, Philip and Colin Ellis (2012) "Examining the behaviour of individual UK consumer prices," The Economic Journal, Vol. 122, pp. F35–F55.

Burgard, Jan Pablo, Matthias Neuenkirch, and Matthias Nöckel (2019) "State-Dependent Transmission of Monetary Policy in the Euro Area," Journal of Money, Credit and Banking, Vol. 51, pp. 2053–2070.

Caggiano, Giovanni, Efrem Castelnuovo, and Nicolas Groshenny (2014) "Uncertainty shocks and unemployment dynamics in US recessions," Journal of Monetary Economics, Vol. 67, pp. 78–92.

Calvo, Guillermo (1983) "Staggered prices in a utility-maximizing framework," Journal of Monetary Economics, Vol. 12, pp. 383–398.

Cavallo, Alberto, Francesco Lippi, and Ken Miyahara (2024) "Large shocks travel fast," American Economic Review: Insights, Vol. 6, pp. 558–574.

Cesa-Bianchi, Ambrogio, Gregory Thwaites, and Alejandro Vicondoa (2020) "Monetary policy transmission in the United Kingdom: A high frequency identification approach," European Economic Review, Vol. 123, p. 103375.

Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans (1999) "Monetary Policy Shocks: What Have We Learned and to What End?" in John B. Taylor and Michael Woodford eds. Handbook of Macroeconomics, Vol. 1, Part A: Elsevier, Chap. 2, pp. 65–148.

Christiano, Lawrence J, Martin Eichenbaum, and Charles L Evans (2005) "Nominal rigidities and the dynamic effects of a shock to monetary policy," Journal of Political Economy, Vol. 113, pp. 1–45.

Cloyne, James and Patrick Hürtgen (2016) "The macroeconomic effects of monetary policy: A new measure for the United Kingdom," American Economic Journal: Macroeconomics, Vol. 8, pp. 75–102.

Cravino, Javier, Ting Lan, and Andrei A Levchenko (2020) "Price stickiness along the income distribution and the effects of monetary policy," Journal of Monetary Economics, Vol. 110, pp. 19–32.

Davies, Richard (2021a) "Prices and Inflation in a Pandemic - A Micro Data Approach," Covid-19 Analysis 017, Centre for Economic Performance, London School of Economics and Political Science.

——— (2021b) "Prices and Inflation in the UK - A New Dataset," Occasional Paper 055, Centre for Economic Performance, London School of Economics and Political Science.

De Santis, Roberto A and Tommaso Tornese (2024) "US monetary policy is more powerful in low economic growth regimes," Working Paper 2919, European Central Bank.

Dedola, Luca, Erwan Gautier, Anton Nakov, Sergio Santoro, Emmanuel de Veirman, Lukas Henkel, and Bruno Fagandini (2023) "Some implications of micro price-setting evidence for inflation dynamics and monetary transmission," Occasional Paper 321, European Central Bank.

Dedola, Luca, Mark Strøm Kristoffersen, and Gabriel Züllig (2021) "The extensive and intensive margin of price adjustment to cost shocks: Evidence from Danish multiproduct firms," Mimeo, European Central Bank.

Dedola, Luca and Francesco Lippi (2005) "The monetary transmission mechanism: Evidence from the industries of five OECD countries," European Economic Review, Vol. 49, pp. 1543–1569.

Dotsey, Michael, Robert G King, and Alexander L Wolman (1999) "State-Dependent Pricing and the General Equilibrium Dynamics of Money and Output," The Quarterly Journal of Economics, Vol. 114, pp. 655–690.

Eichenbaum, Martin (1992) "Comments on "Interpreting the Time Series Facts: The Effects of Monetary Policy" by Christopher Sims," European Economic Review, Vol. 36, pp. 1001–1011.

Eichenbaum, Martin, Nir Jaimovich, Sergio Rebelo, and Josephine Smith (2014) "How frequent are small price changes?" American Economic Journal: Macroeconomics, Vol. 6, pp. 137–55.

Ellis, Colin, Haroon Mumtaz, and Pawel Zabczyk (2014) "What lies beneath? A time-varying FAVAR model for the UK transmission mechanism," The Economic Journal, Vol. 124, pp. 668–699.

Gagnon, Etienne (2009) "Price setting under low and high inflation: Evidence from Mexico," The Quarterly Journal of Economics, Vol. 124, pp. 1221–63.

Garcia, Rene and Huntley Schaller (2002) "Are the effects of monetary policy asymmetric?" Economic Inquiry, Vol. 40, pp. 102–119.

Gautier, Erwan, Cristina Conflitti, Riemer P Faber, Brian Fabo, Ludmila Fadejeva, Valentin Jouvanceau, Jan-Oliver Menz, Teresa Messner, Pavlos Petroulas, Pau Roldan-Blanco et al. (2024) "New facts on consumer price rigidity in the euro area," American Economic Journal: Macroeconomics, Vol. 16, pp. 386–431.

Gautier, Erwan and Hervé Le Bihan (2022) "Shocks versus menu costs: Patterns of price rigidity in an estimated multisector menu-cost model," Review of Economics and Statistics, Vol. 104, pp. 668–685.

Gerko, Elena and Hélene Rey (2017) "Monetary policy in the capitals of capital," Journal of the European Economic Association, Vol. 15, pp. 721–745.

Gertler, Mark and Peter Karadi (2015) "Monetary policy surprises, credit costs, and economic activity," American Economic Journal: Macroeconomics, Vol. 7, pp. 44–76.

Golosov, Mikhail and Robert E. Lucas (2007) "Menu Costs and Phillips Curves," Journal of Political Economy, Vol. 115, pp. 171–199.

Gürkaynak, Refet S, Brian Sack, and Eric Swanson (2005) "The sensitivity of long-term interest rates to economic news: Evidence and implications for macroeconomic models," American Economic Review, Vol. 95, pp. 425–436.

Hahn, Volker and Michal Marenčák (2020) "Price points and price dynamics," Journal of Monetary Economics, Vol. 115, pp. 127–144.

Henkel, Lukas, Elisabeth Wieland, Aneta Błażejowska, Cristina Conflitti, Brian Fabo, Ludmila Fadejeva, Jana Jonckheere, Peter Karadi, Paweł Macias, Jan-Oliver Menz, Pascal Seiler, and Karol Szafranek (2023) "Price setting during the coronavirus (COVID-19) pandemic," Occasional Paper 324, European Central Bank.

Hong, Gee Hee, Matthew Klepacz, Ernesto Pasten, and Raphael Schoenle (2023) "The real effects of monetary shocks: Evidence from micro pricing moments," Journal of Monetary Economics, Vol. 139, pp. 1–20.

Jordà, Òscar (2005) "Estimation and inference of impulse responses by local projections," American Economic Review, Vol. 95, pp. 161–182.

Jordà, Òscar, Moritz Schularick, and Alan M Taylor (2020) "The effects of quasi-random monetary experiments," Journal of Monetary Economics, Vol. 112, pp. 22–40.

Karadi, Peter, Anton Nakov, Galo Nuño, Ernesto Pastén, and Dominik Thaler (2025) "Strike while the iron is hot – Optimal monetary policy under state-dependent pricing," Working Paper 3068, European Central Bank.

Karadi, Peter and Adam Reiff (2019) "Menu costs, aggregate fluctuations, and large shocks," American Economic Journal: Macroeconomics, Vol. 11, pp. 111–46.

Kryvtsov, Oleksiy and Nicolas Vincent (2021) "The cyclicality of sales and aggregate price flexibility," Review of Economic Studies, Vol. 88, pp. 334–377.

Lo, Ming Chien and Jeremy Piger (2005) "Is the response of output to monetary policy asymmetric? Evidence from a regime-switching coefficients model," Journal of Money, Credit and Banking, Vol. 37, pp. 865–886.

Montag, Hugh and Daniel Villar (2022) "Price-Setting During the Covid Era," Working Paper 547, U.S. Bureau of Labor Statistics.

Mountford, Andrew (2005) "Leaning into the wind: A structural VAR investigation of UK monetary policy," Oxford Bulletin of Economics and Statistics, Vol. 67, pp. 597–621.

Nakamura, Emi and Jón Steinsson (2008) "Five facts about prices: A reevaluation of menu cost models," The Quarterly Journal of Economics, Vol. 123, pp. 1415–1464.

——— (2013) "Price Rigidity: Microeconomic Evidence and Macroeconomic Implications," Annual Review of Economics, Vol. 5, pp. 133–163.

Nakamura, Emi, Jón Steinsson, Patrick Sun, and Daniel Villar (2018) "The elusive costs of inflation: Price dispersion during the US great inflation," The Quarterly Journal of Economics, Vol. 133, pp. 1933–1980.

Newey, Whitney K and Kenneth D West (1987) "Hypothesis testing with efficient method of moments estimation," International Economic Review, Vol. 28, pp. 777–787.

Office for National Statistics (2019) "Consumer Prices Indices Technical Manual, 2019," https://www.ons.gov.uk/economy/inflationandpriceindices/methodologies/consumerpricesindicestechnicalmanual2019 [Accessed: 2025 12 15].

——— (2024) "Dataset: Consumer price inflation item indices and price quotes," https://www.ons.gov.uk/economy/inflationandpriceindices/datasets/

consumerpriceindicescpiandretailpricesindexrpiitemindicesandpricequotes [Accessed: 2025 12 15].

Olivei, Giovanni and Silvana Tenreyro (2007) "The timing of monetary policy shocks," American Economic Review, Vol. 97, pp. 636–663.

——— (2010) "Wage-setting patterns and monetary policy: International evidence," Journal of Monetary Economics, Vol. 57, pp. 785–802.

Ramey, Valerie A and Sarah Zubairy (2018) "Government spending multipliers in good times and in bad: Evidence from US historical data," Journal of Political Economy, Vol. 126, pp. 850–901.

Romer, Christina D and David H Romer (2004) "A new measure of monetary shocks: Derivation and implications," American Economic Review, Vol. 94, pp. 1055–1084.

Rudolf, Barbara and Pascal Seiler (2022) "Price Setting Before and During the Pandemic: Evidence from Swiss Consumer Prices," Working Paper 2748, European Central Bank.

Santoro, Emiliano, Ivan Petrella, Damjan Pfajfar, and Edoardo Gaffeo (2014) "Loss aversion and the asymmetric transmission of monetary policy," Journal of Monetary Economics, Vol. 68, pp. 19–36.

Sims, Christopher A (1992) "Interpreting the macroeconomic time series facts: The effects of monetary policy," European Economic Review, Vol. 36, pp. 975–1000.

Swanson, Eric T (2021) "Measuring the effects of federal reserve forward guidance and asset purchases on financial markets," Journal of Monetary Economics, Vol. 118, pp. 32–53.

Tenreyro, Silvana and Gregory Thwaites (2016) "Pushing on a string: US monetary policy is less powerful in recessions," American Economic Journal: Macroeconomics, Vol. 8, pp. 43–74.

Thoma, Mark A (1994) "Subsample instability and asymmetries in money-income causality," Journal of Econometrics, Vol. 64, pp. 279–306.

Vavra, Joseph (2014) "Inflation dynamics and time-varying volatility: New evidence and an Ss interpretation," The Quarterly Journal of Economics, Vol. 129, pp. 215–258.

Weise, Charles L (1999) "The asymmetric effects of monetary policy: A nonlinear vector autoregression approach," Journal of Money, Credit and Banking, Vol. 31, pp. 85–108.

A Empirical evidence on the non-uniformity of price rigidity over time

This appendix refers to Section 2 and provides additional information on the empirical evidence of the non-uniformity of price rigidity over time. It provides figures and tables that are not included in the main body of the paper.

A.1 UK CPI microdata

Table A.1 illustrates the effect of all sample restrictions on the sample size. I exclude duplicate price quotes (because of reasons of confidentiality, the ONS does not always publish all available local information, resulting in product identifiers containing duplicate price quotes), "invalid" price quotes (which do not pass the ONS cross-checking procedures and are not included in the calculation of the official CPI), imputed prices, outliers (by excluding monthly price changes larger than the 99th percentile of absolute log price changes and smaller than the 1st percentile of absolute log price changes for each expenditure item), price changes that coincide with changes in product units, and price changes that coincide with changes in VAT rates.

Table A.1. Sample restrictions and sample size

	# price quotes	# products	# price changes
Raw data	42,018,324	906,866	9,846,656
– exclude invalid price quotes	3,685,454	571,693	1,752,741
– exclude duplicate price quotes	6,747,419	65,585	2,054,872
– exclude imputed prices	1,992,487	326,848	1,094,871
– exclude $\lvert\Delta p\rvert < \text{P1}$, $\lvert\Delta p\rvert > \text{P99}$	100,786	65,160	100,786
– exclude product unit changes	10,335	8,510	4,527
– exclude/dummy VAT price changes	348,408	158,476	151,606
Baseline sample	32,009,659	789,382	5,499,732

Notes: The table illustrates the effect of all sample restrictions on the sample size. I exclude duplicate price quotes (because of reasons of confidentiality, the ONS does not always publish all available local information, resulting in product identifiers containing duplicate price quotes), "invalid" price quotes (which do not pass the ONS cross-checking procedures and are not included in the calculation of the official CPI), imputed prices, outliers (by excluding monthly price changes larger than the 99th percentile of absolute log price changes and smaller than the 1st percentile of absolute log price changes for each expenditure item), price changes that coincide with changes in product units, and price changes that coincide with changes in VAT rates.

A.2 Further evidence on the aggregate frequency of price changes

The left panel of Figure 1 illustrates cyclical patterns of price rigidity: the frequency of price changes is countercyclical and positively associated with inflation. To validate these visual patterns, I formally examine the relationship between the frequency of price changes, the unemployment rate, and CPI inflation in regression analyses. Specifically, I estimate OLS time-series regressions of the monthly frequency of price changes on business-cycle indicators (the unemployment rate and recession dummies) and CPI inflation, controlling for seasonality via month fixed effects and VAT changes. The regressions are run for all sample definitions, including and excluding price changes due to temporary sales and product substitution.

Table A.2 reports the estimated coefficients. Across specifications, the coefficients on CPI inflation and business-cycle indicators are statistically significant at the 1% level, even when a linear time trend or a lagged dependent variable is included. The elasticity of the frequency of price changes with respect to inflation is consistently larger than with respect to the unemployment rate, though both effects are economically meaningful. A 1-percentage-point increase in inflation is associated with a 0.7 percentage point (pp) higher frequency of price changes, while a 1-percentage-point increase in the unemployment rate is associated with a 0.2 pp higher frequency.

The right panel of Figure 1 presents the seasonal profile of price rigidity, highlighting a pronounced "January effect" whereby the frequency of price changes is substantially higher in January than in any other month. To validate this visual pattern, I regress January dummies on item-level frequencies of price changes, increases, and decreases for UK consumer prices from February 1996 to December 2024, using the samples with and without price changes due to temporary sales. The regressions include month and year fixed effects as well as VAT dummies.

Table A.2. Cyclical properties of the frequency of price changes

	Frequency of price changes						
	incl. sales, incl. subst.				incl. sales excl. subst.	excl. sales incl. subst.	excl. sales excl. subst.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
CPI inflation	0.719***	0.707***	0.485***	0.587***	0.559***	0.594***	0.468***
	(0.052)	(0.053)	(0.063)	(0.053)	(0.041)	(0.029)	(0.024)
Unemployment rate	0.201***		0.134*	0.419***	0.235***	0.279***	0.266***
	(0.074)		(0.071)	(0.078)	(0.059)	(0.042)	(0.033)
Recession dummy		0.094					
		(0.252)					
Frequency (lagged)			0.325***				
			(0.053)				
Constant	8.324***	9.456***	5.608***	6.181***	6.386***	4.340***	2.889***
	(0.455)	(0.176)	(0.613)	(0.548)	(0.364)	(0.256)	(0.205)
Linear time trend	N	N	N	Y	N	N	N
Observations	335	335	327	335	335	335	335
R^2	0.499	0.487	0.561	0.554	0.494	0.627	0.613

Notes: Linear regressions of the frequency of price changes on CPI inflation and aggregate business cycle indicators (the unemployment rate and recession dummies). All regressions include calendar month dummies and control for VAT rate changes. The frequency of price changes is calculated at the item level and aggregated to weighted medians using CPI expenditure weights using the samples of UK CPI microdata from February 1996 to December 2024, including and excluding price changes due to temporary sales and product substitution. The unemployment rate is the monthly unemployment rate for civilians aged 15 and over. CPI inflation is the year-on-year change in the consumer price index excluding owner-occupier housing costs. Recessions are dated by the Organisation for Economic Co-operation and Development (OECD). *p<0.1; **p<0.05; ***p<0.01.

Table A.3. "January effect" on the frequency of price changes

		Including sales			Excluding sales		
Sector	Variable	Effect	N	% CPI	Effect	N	% CPI
Aggregate	Price changes	11.02	462	47.94	6.80	330	37.56
	Price decreases	6.15	484	50.28	2.59	242	27.19
	Price increases	5.62	370	40.90	5.14	328	38.43
Food	Price changes	8.76	36	3.48	7.95	33	2.98
	Price decreases	-1.80	82	6.69	1.80	29	2.34
	Price increases	9.92	71	6.85	6.94	53	5.07
NEIG	Price changes	11.45	334	27.82	3.30	204	17.61
	Price decreases	10.03	331	28.16	3.20	149	11.84
	Price increases	1.22	204	17.67	1.94	179	16.02
Services	Price changes	10.77	92	16.64	10.22	93	16.97
	Price decreases	2.52	71	15.43	2.18	64	13.01
	Price increases	8.65	93	16.27	7.62	95	17.27

Notes: "January effects" from item-level regressions of January dummies on the frequency of price changes, increases, and decreases for consumer prices in the United Kingdom from February 1996 to December 2024 using the samples including and excluding price changes due to temporary sales. Price changes due to product substitution are excluded. Regressions include month and year fixed effects and dummies for VAT changes. "Effect" indicates the weighted average size of significant January-dummy coefficients. "N" and "% CPI" indicate the absolute number and weighted share of items with significant January-dummy coefficients. Significance is evaluated at the 5% significance level. The total number of items used in the estimation is 1,260. The CPI weights refer to the average CPI expenditure weight over the sample period.

Table A.3 summarizes the estimation results. The "January effect" is statistically significant for 48% of items, with an average weighted effect of 11.0 pp for the frequency of price changes when sales are included. The effect for price decreases is 6.2 pp, the effect for price increases is 5.6 pp.

A.3 Further evidence on the frequency of price increases and decreases

The left panel of Figure 2 illustrates cyclical patterns in the frequencies of price increases and decreases: both types of adjustments are countercyclical, rising when economic activity slows, but only price increases co-move positively with inflation. To validate these visual patterns, I formally examine the relationship between the frequency of price increases and decreases, the unemployment rate, and CPI inflation in regression analyses. Specifically, I estimate OLS time-series regressions of the monthly frequency of price increases and decreases on business-cycle indicators

(the unemployment rate and recession dummies) and CPI inflation, controlling for seasonality via month fixed effects and VAT changes.

Table A.4 reports the estimated coefficients. The results show that both the frequency of price increases and decreases are equally countercyclical: a 1-percentage-point increase in the unemployment rate is associated with a 0.17 pp higher frequency of either increases or decreases. By contrast, the positive co-movement with inflation is driven solely by price increases. A 1-percentage-point increase in inflation is associated with a 0.50 pp higher frequency of price increases, while the frequency of price decreases shows no significant response to inflation.

The right panel of Figure 2 illustrates seasonal patterns in the frequencies of price increases and decreases: both types of adjustment are more frequent early in the year than later, with the seasonal pattern being more pronounced for price increases than for price decreases. To further investigate the role of increases and decreases in driving these seasonal patterns, I adopt the approach of Gautier et al. (2024) and estimate weighted panel regressions relating item-level frequencies of price increases and decreases to month and year fixed effects. Specifically, I estimate

$$y_{i,t} = \mu_i + \mu_m + \mu_y + \beta \mathrm{VAT}_{i,t} + \varepsilon_{i,t}, \tag{3}$$

where μ_i is an item-level fixed effect, μ_m is a calendar-month fixed effect, μ_y is a year fixed effect, and $\mathrm{VAT}_{i,t}$ is a dummy variable controlling for VAT changes in December 2008, January 2010, and January 2011. This regression framework is independent of both the choice of the measure of central tendency[30] used for cross-sectional aggregation (since month fixed effects are estimated at the most disaggregated product level) and of any trend in the frequency of price changes (since year fixed effects are included). The dependent variables are the frequencies of price increases and decreases, calculated using samples both including and excluding temporary sales.

Figure A.1 plots the estimated month-effects. The left panel shows results using the sample including temporary sales, while the right panel shows results excluding them.

The evidence points to strong seasonal patterns in both frequencies. Price increases and decreases are roughly 2 percentage points more frequent in January than in other months. Temporary sales play a key role in shaping these patterns: when sales are excluded, the seasonal pattern for price increases persists, while the seasonality of price decreases largely disappears. Thus, most of the seasonality in the overall frequency of price changes is driven by the frequency of price increases.

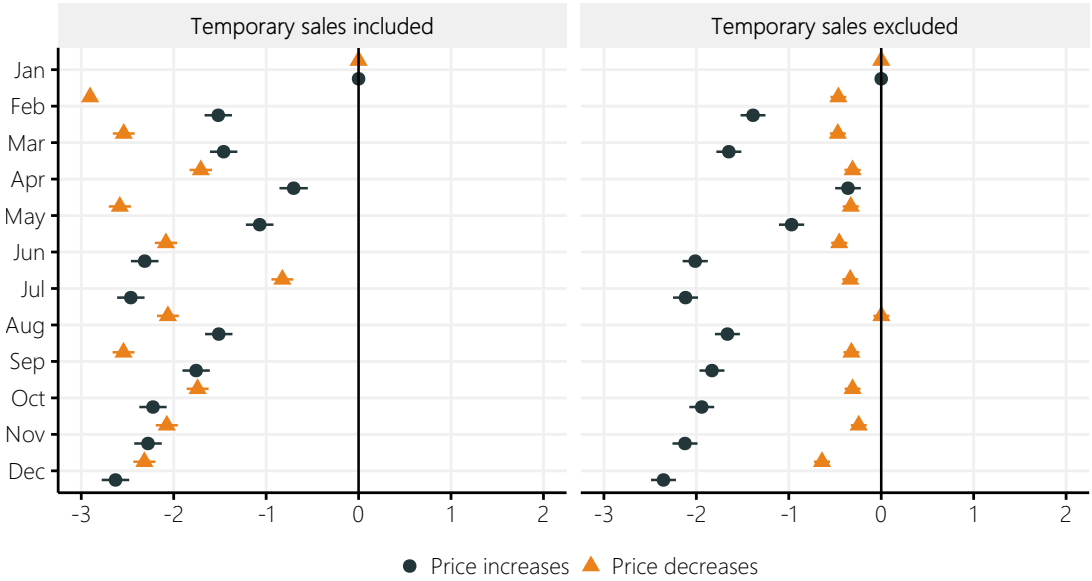
[30]Estimates aggregated using mean and median moments can vary widely depending on cross-sectoral heterogeneity. In the euro area, the median frequency of consumer price changes is only about 2.5 pp lower than the mean frequency (Gautier et al., 2024). By contrast, the difference is larger than 6 pp in the United States (Nakamura and Steinsson, 2008), France (Gautier and Le Bihan, 2022), or Switzerland (Rudolf and Seiler, 2022).

Table A.4. Cyclical properties of the frequency of price increases and decreases

	Frequency of price increases				Frequency of price decreases			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
CPI inflation	0.495***	0.499***	0.320***	0.422***	0.031**	0.022	0.022*	−0.021
	(0.030)	(0.031)	(0.040)	(0.031)	(0.015)	(0.015)	(0.013)	(0.015)
Unemployment rate	0.056		0.004	0.173***	0.095***		0.056***	0.179***
	(0.039)		(0.037)	(0.041)	(0.022)		(0.020)	(0.022)
Recession dummy		0.129				−0.057		
		(0.129)				(0.076)		
Frequency (lagged)			0.356***				0.422***	
			(0.052)				(0.051)	
Constant	4.472***	4.752***	3.042***	3.337***	2.514***	3.083***	1.424***	1.679***
	(0.235)	(0.094)	(0.320)	(0.285)	(0.134)	(0.053)	(0.174)	(0.154)
Time trend	No	No	No	Yes	No	No	No	Yes
Observations	335	335	326	335	335	335	325	335
R^2	0.514	0.513	0.534	0.569	0.300	0.259	0.451	0.434

Notes: Linear regressions of the frequency of price increases and decreases on CPI inflation and aggregate business cycle indicators (the unemployment rate and recession dummies). All regressions include calendar month dummies and control for VAT rate changes. The frequency of price increases and decreases is calculated at the item level and aggregated to weighted medians using CPI expenditure weights using the samples of UK CPI microdata from February 1996 to December 2024, including price changes due to temporary sales and product substitution. The unemployment rate is the monthly unemployment rate for civilians aged 15 and over. CPI inflation is the year-on-year change in the consumer price index excluding owner-occupier housing costs. Recessions are dated by the Organisation for Economic Co-operation and Development (OECD) *p<0.1; **p<0.05; ***p<0.01.

Figure A.1. Month-effects of the frequency of price increases and decreases



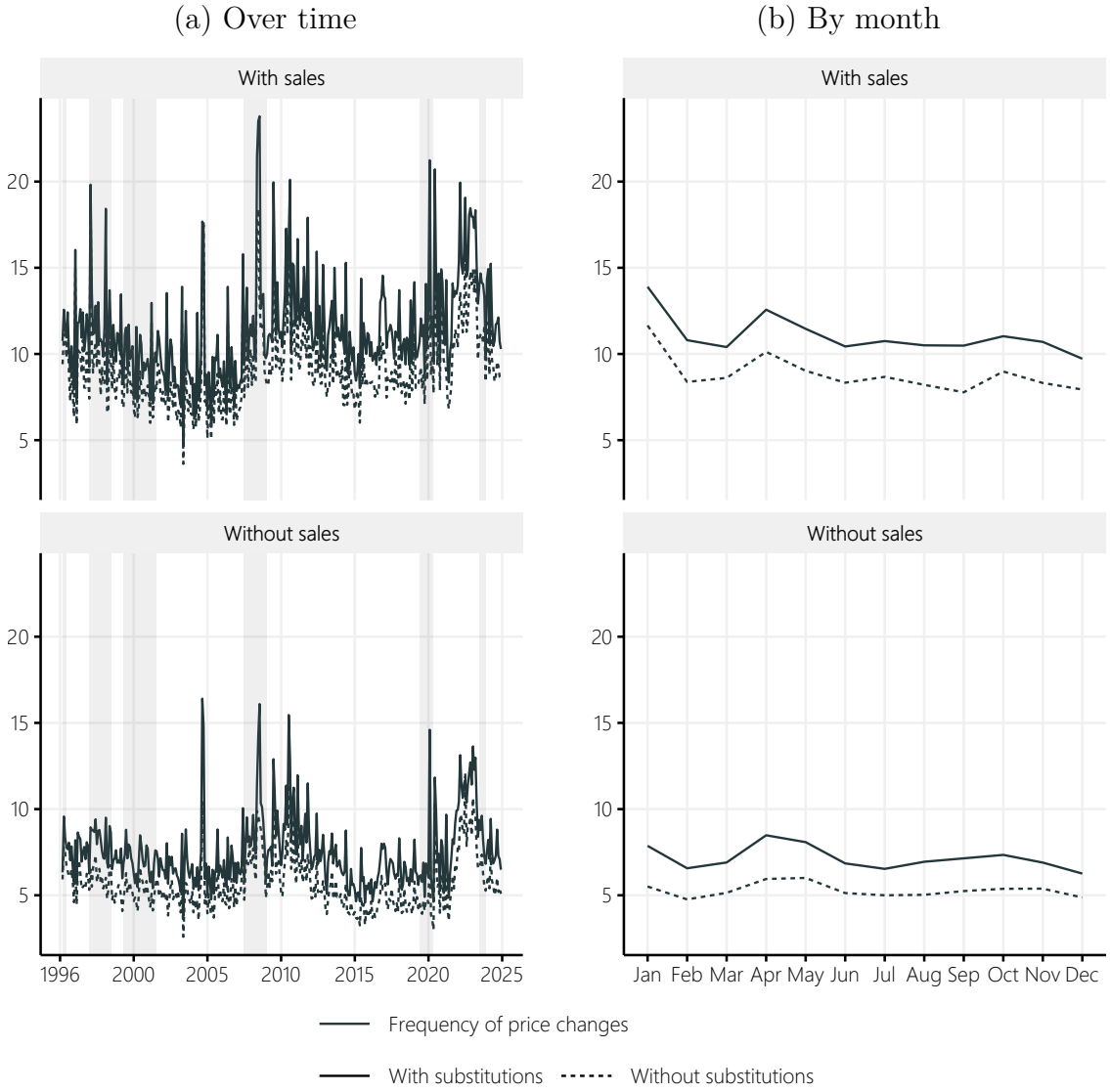
Notes: Coefficient plots of month-effects from weighted item-level panel regressions as specified in Equation (3) with robust standard errors and 95% confidence intervals. The dependent variables are the frequencies of price increases and decreases for consumer prices in the United Kingdom from February 1996 to December 2024. The base month is January. Price changes due to product substitution are included. Outliers identified as values smaller (larger) than the 1st (99th) percentiles of the dependent variables are excluded beforehand.

A.4 The role of temporary sales and product substitutions

This appendix documents how the treatment of price changes due to temporary sales and product substitution affects the measured frequency of price changes. Figure A.2 illustrates the evolution of the frequency of price changes over time (left column) and its seasonal profile (right column) across four sample definitions: including versus excluding temporary sales (upper versus lower panels) and including versus excluding product substitution (solid versus dashed lines).

The time-series patterns in the left column of Figure A.2 show that the treatment of temporary sales and product substitution primarily affects the level of the series rather than its cyclical co-movement with macroeconomic conditions. Excluding product substitutions lowers the average frequency by around two percentage points, reflecting the small but non-negligible contribution of replacement items—particularly during periods of product shortages such as the COVID-19 pandemic, when substitution-related price changes briefly reached 5.5 percent. Excluding temporary sales has a much larger effect: the frequency measured without sales is roughly three

Figure A.2. Frequency of price changes



Notes: The figure shows the frequency of consumer price changes in the United Kingdom, including price changes due to temporary sales and product substitution. Moments are computed at the item level and aggregated to weighted medians using CPI expenditure weights. The left panel shows the median frequency of price changes from February 1996 to December 2024, together with the unemployment rate and CPI inflation. Recessions (gray shaded areas) are dated by the Organisation for Economic Co-operation and Development (OECD). The right panel shows the median frequency of price changes by calendar month across the sample period, with the interquartile range indicated by the shaded area. Outliers related to VAT rate changes are excluded. All values are expressed in percent.

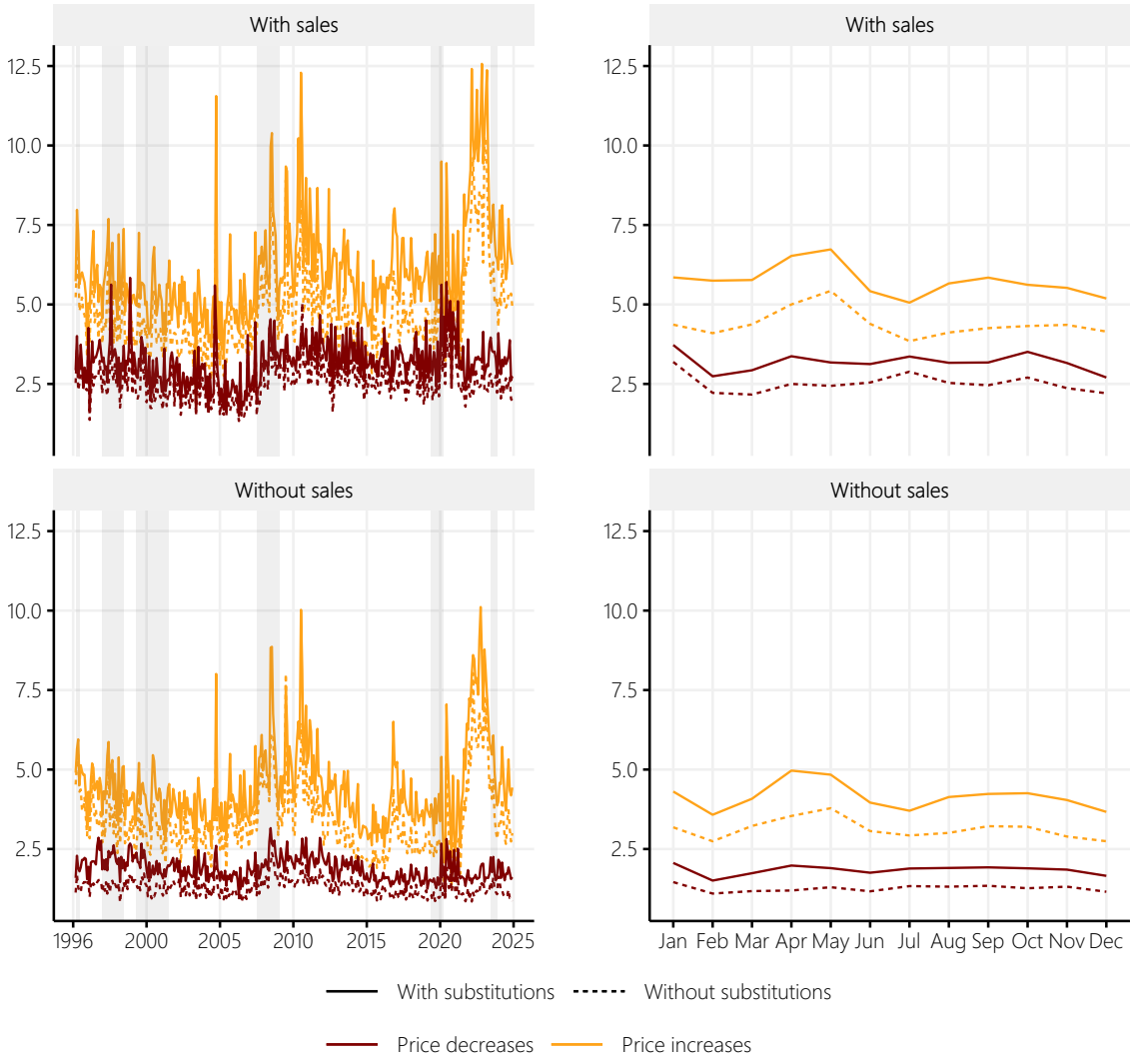
to four percentage points lower, and the contribution of sale-related price changes has increased over the sample, stabilizing around 4 percent in recent years. Despite these level differences, the cyclical properties of the series remain unchanged. Frequencies remain countercyclical and positively correlated with inflation across all sample definitions; the regression coefficients reported in Table A.2 are robust to including or excluding sales and substitution. However, the elasticities of the frequency of price changes to fluctuations in inflation are much smaller when temporary sales are excluded. This is consistent with the view that firms' sales policies are important for adjusting to large aggregate shocks. Using CPI microdata, Kryvtsov and Vincent (2021) show that the frequency of sales is strongly countercyclical in the United Kingdom and the United States, more than doubling during the Great Recession. Evidence for a more recent period suggests that firms responded to the large demand shock caused by the COVID-19 lockdowns in the euro area (Henkel et al., 2023) and Switzerland (Rudolf and Seiler, 2022) by adjusting the frequency of their temporary sales.

The seasonal patterns differ more starkly across sample definitions. As shown in the right column of Figure A.2, product substitution has only a modest effect on seasonality, whereas temporary sales play a central role. Excluding sales substantially attenuates the seasonal amplitude and removes the characteristic within-quarter monotonic decline. The January peak remains visible when sales are excluded, but it is notably smaller: the January effect averages 6.8 percentage points (compared with 10.2 pp in the sample including temporary sales). These results indicate that temporary sales disproportionately shape the seasonal structure of aggregate price flexibility, whereas regular (non-sale) price adjustments show weaker and less systematic seasonal variation. When sales are excluded, the frequency of price changes is still higher in January than in other months for some items (around 30 percent). However, the overall effect is smaller (6.8 pp), driven by a much weaker effect for price decreases (2.1 pp). By contrast, the January effect for price increases remains about the same as when sales are included (5.7 pp). This suggests that the exclusion of temporary sales particularly affects the seasonal profile of price decreases.

To identify which types of price adjustments account for this variation, Figure A.3 decomposes frequencies into price increases and price decreases under different sample definitions. Consistent with the aggregate results, the treatment of temporary sales primarily affects the level and seasonal profile of price decreases, while price increases are less sensitive to sample construction.

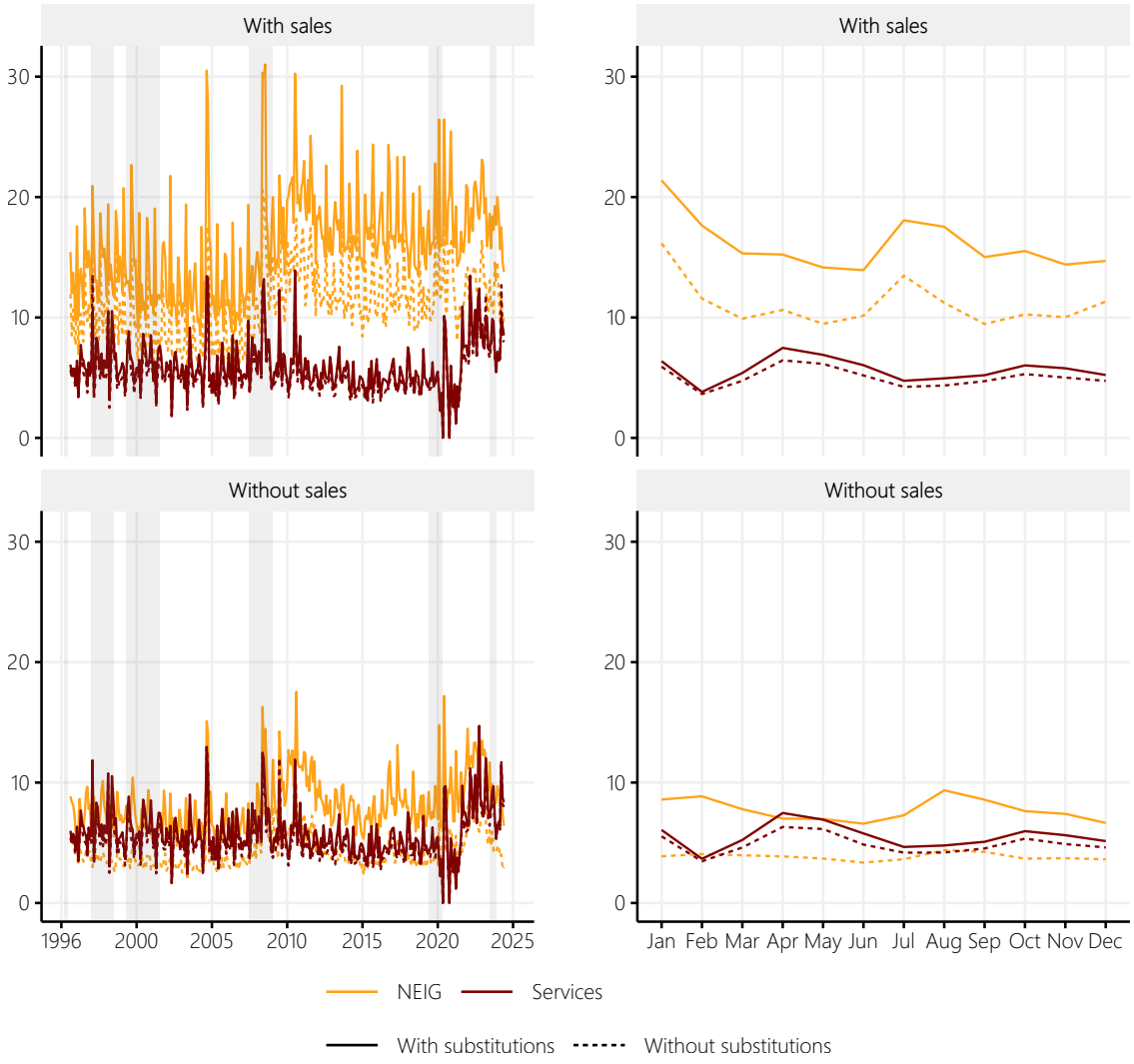
Finally, Figure A.4 reports the same set of sample definitions separately for non-energy industrial goods and services. The influence of temporary sales is concentrated in the goods sector, where discounts and promotions significantly elevate both the level and seasonality of price adjustment frequencies. By contrast, services exhibit low frequencies and limited seasonal variation regardless of sample definition, reflecting the near absence of temporary sales in that sector.

Figure A.3. Frequency of price increases and decreases



Notes: The figure shows the frequency of price increases (in yellow) and the frequency of price decreases (in red) in the United Kingdom, including price changes due to temporary sales and product substitution. Moments are computed at the item level and aggregated to weighted medians using CPI expenditure weights. The left panel shows the median frequencies of price increases and decreases from February 1996 to December 2024. The right panel shows the median frequencies of price increases and decreases by calendar month across the sample period, with the interquartile ranges indicated by the shaded area. Outliers related to VAT rate changes are excluded. Recessions (gray shaded areas) are dated by the Organisation for Economic Co-operation and Development (OECD). All values are expressed in percent.

Figure A.4. Frequency of price changes in non-energy industrial goods and services



Notes: The figure shows the frequency of price changes for non-energy industrial goods (in yellow) and services (in red) in the United Kingdom, including price changes due to temporary sales and product substitution. Moments are computed at the item level and aggregated to weighted medians using CPI expenditure weights. The left panel shows the median frequencies of NEIG and services price changes from February 1996 to December 2024. The right panel shows the median frequencies of NEIG and services price changes by calendar month across the sample period, with the interquartile ranges indicated by the shaded area. Outliers related to VAT rate changes are excluded. Recessions (gray shaded areas) are dated by the Organisation for Economic Co-operation and Development (OECD). All values are expressed in percent.

A.5 Size of price changes

The absolute size of price changes for item i in month t captures the average of all non-zero log price changes of product p of item i that change from one month to another in absolute value and is calculated as

$$|\Delta p_{i,t}| = \frac{\sum_{p,s} \omega_{p,s,t} I_{p,s,t} (|\log P_{p,s,t} - \log P_{p,s,t-1}|)}{\sum_{p,s} \omega_{p,s,t} I_{p,s,t}}. \tag{4}$$

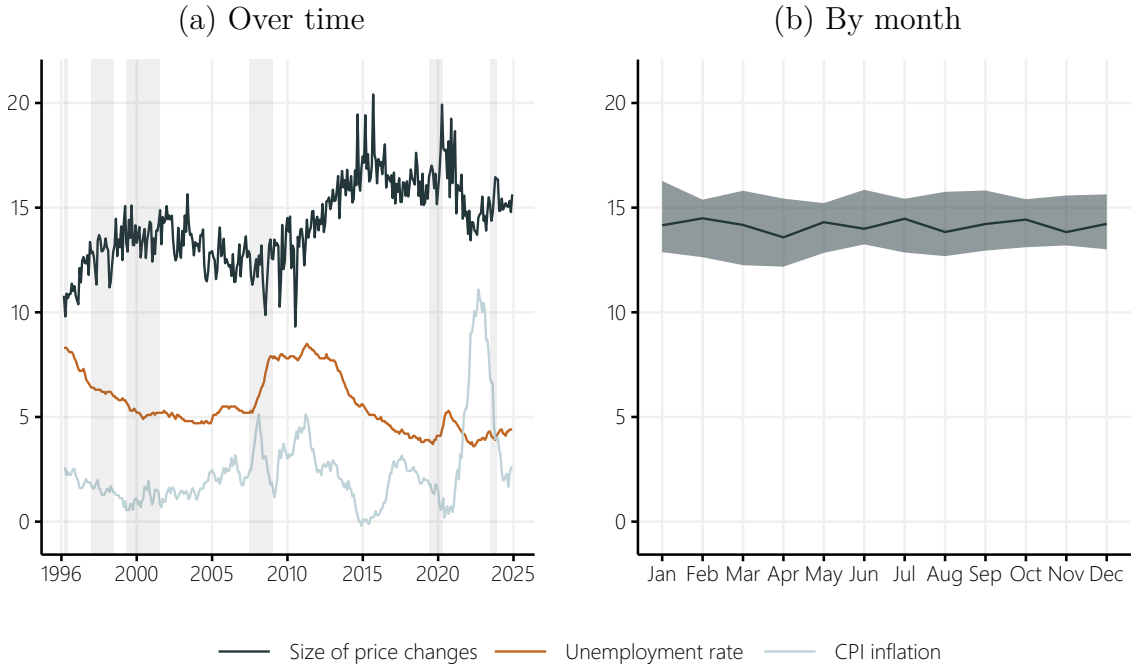
When calculating the size of price increases, $\Delta p^{+}_{i,t}$, and the size of price decreases, $\Delta p^{-}_{i,t}$, the price change indicators, $I^{+}_{p,s,t}$ and $I^{-}_{p,s,t}$, only consider price increases or price decreases, respectively. The aggregate absolute size of price changes, $|\Delta p_t|$, is the weighted median absolute size across items using average CPI expenditure weights.

The left panel of Figure A.5 plots the absolute size of price changes from February 1996 to December 2024, alongside the unemployment rate and CPI inflation. The size of price changes exhibits a distinct hump-shaped pattern in the 2000s: it rises from roughly 10 percent at the start of the sample to about 14 percent in the mid-2000s and declines to around 12 percent by the end of that decade. After 2010, the average size increases again, stabilizing above 16 percent for several years. During the post-pandemic inflation surge, the average size temporarily falls below 15 percent. Visually, these fluctuations appear to move opposite to those of the frequency of price changes. The size of price changes is clearly cyclical and negatively associated with inflation.

Time-series regressions confirm these impressions. Table A.5 reports OLS estimates of monthly absolute price-change size on business-cycle indicators and CPI inflation. Both inflation and the business-cycle variables (unemployment rate and a recession dummy) enter with negative and statistically significant coefficients at the 1 percent level. These results hold across alternative sample definitions, including and excluding price changes due to temporary sales, and remain robust to the inclusion of a linear time trend or a lagged dependent variable.

The right panel of Figure A.5 displays the seasonal profile of the absolute size of price changes. In contrast to the seasonal patterns documented for the frequency of price changes, the size shows little systematic intra-annual variation: monthly averages hover around 14 percent with no clear "January effect." This is corroborated by item-level regressions of January dummies on the size of price changes—estimated separately for all price changes, increases, and decreases, and for samples with and without temporary sales. The regressions include month and year fixed effects as well as VAT dummies. The results, summarized in Table A.6, show that fewer than one-fifth of items exhibit a statistically significant January effect. When present, the

Figure A.5. Size of price changes



Notes: The figure shows the absolute size of consumer price changes in the United Kingdom, including price changes due to temporary sales and product substitution. Moments are computed at the item level and aggregated to weighted medians using CPI expenditure weights. The left panel shows the median absolute size of price changes from February 1996 to December 2024, together with the unemployment rate and CPI inflation. Recessions (gray shaded areas) are dated by the Organisation for Economic Co-operation and Development (OECD). The right panel shows the median absolute size of price changes by calendar month across the sample period, with the interquartile range indicated by the shaded area. Outliers related to VAT rate changes are excluded. All values are expressed in percent.

Table A.5. Cyclical properties of the size of price changes

	Size of price changes						
	incl. sales, incl. subst.				incl. sales excl. subst.	excl. sales incl. subst.	excl. sales excl. subst.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
CPI inflation	-0.164^{***}	-0.166^{***}	-0.027	-0.438^{***}	-0.159^{***}	-0.175^{***}	-0.097^{**}
	(0.047)	(0.051)	(0.028)	(0.034)	(0.055)	(0.039)	(0.039)
Unemployment rate	-0.623^{***}		-0.093^{**}	-0.164^{***}	-0.696^{***}	-0.431^{***}	-0.421^{***}
	(0.068)		(0.045)	(0.050)	(0.079)	(0.056)	(0.056)
Recession dummy		-1.329^{***}					
		(0.245)					
Size (lagged)			0.817^{***}				
			(0.032)				
Constant	18.093^{***}	14.887^{***}	3.194^{***}	13.536^{***}	17.991^{***}	14.038^{***}	12.821^{***}
	(0.416)	(0.172)	(0.633)	(0.352)	(0.489)	(0.343)	(0.345)
Linear time trend	N	N	N	Y	N	N	N
Observations	346	346	345	346	346	346	346
R^2	0.227	0.107	0.737	0.663	0.206	0.193	0.162

Notes: Linear regressions of the absolute size of price changes on CPI inflation and aggregate business cycle indicators (the unemployment rate and recession dummies). All regressions include calendar month dummies and control for VAT rate changes. The absolute size of price changes is calculated at the item level and aggregated to weighted medians using CPI expenditure weights using the samples of UK CPI microdata from February 1996 to December 2024, including and excluding price changes due to temporary sales and product substitution. The unemployment rate is the monthly unemployment rate for civilians aged 15 and over. CPI inflation is the year-on-year change in the consumer price index excluding owner-occupier housing costs. Recessions are dated by the Organisation for Economic Co-operation and Development (OECD). $^{*}p<0.1$; $^{**}p<0.05$; $^{***}p<0.01$.

effect is modest: on average, 3.5 percentage points for the absolute size of all price changes in the sample including sales, 6.1 percentage points for decreases, and -4.7 percentage points for increases.

Table A.6. "January effect" on the size of price changes

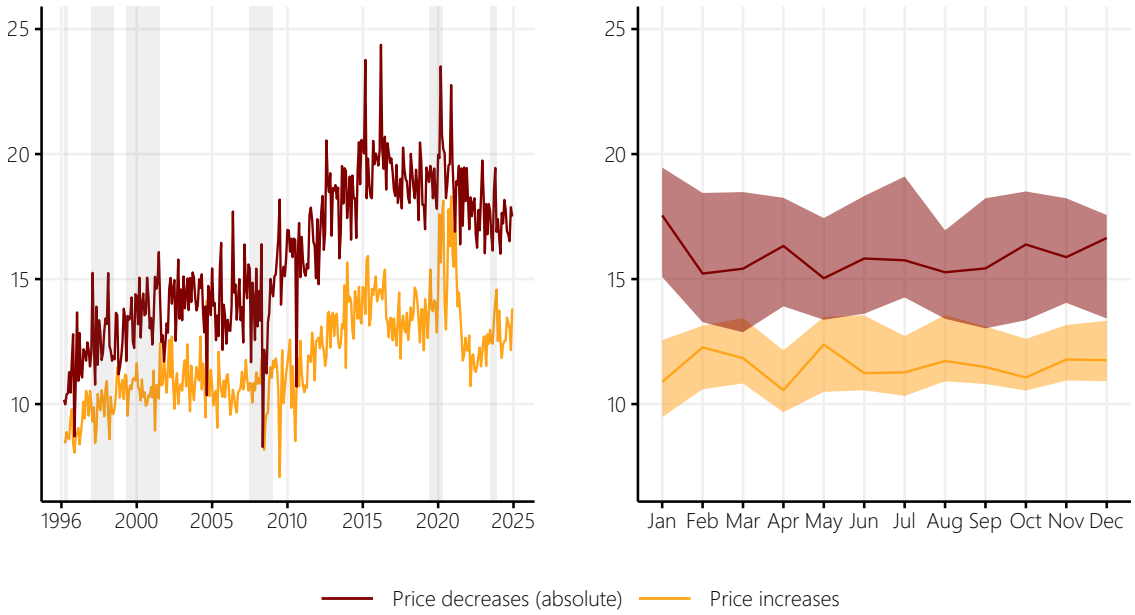
		Including sales			Excluding sales		
Sector	Variable	Effect	N	% CPI	Effect	N	% CPI
Aggregate	Price changes	3.52	194	19.01	-2.86	97	11.40
	Price decreases	6.06	245	21.18	2.72	93	8.33
	Price increases	-4.71	147	15.45	-5.41	132	13.29
Food	Price changes	2.64	37	2.67	1.20	23	1.79
	Price decreases	1.96	74	5.73	4.17	36	2.51
	Price increases	1.11	47	3.73	0.87	35	3.05
NEIG	Price changes	4.76	136	11.07	-5.19	53	4.77
	Price decreases	7.94	164	14.25	3.39	52	4.82
	Price increases	-8.25	83	8.09	-8.57	83	7.42
Services	Price changes	1.35	20	5.26	-2.06	20	4.82
	Price decreases	3.36	7	1.21	-4.10	5	1.00
	Price increases	-3.31	14	2.67	-4.17	13	2.70

Notes: "January effects" from item-level regressions of January dummies on the absolute size of price changes, increases, and decreases for consumer prices in the United Kingdom from February 1996 to December 2024 using the samples including and excluding price changes due to temporary sales. Price changes due to product substitution are excluded. Regressions include month and year fixed effects and dummies for VAT changes. "Effect" indicates the weighted average size of significant January-dummy coefficients. "N" and "% CPI" indicate the absolute number and weighted share of items with significant January-dummy coefficients. Significance is evaluated at the 5% significance level. The total number of items used in the estimation is 1,260. The CPI weights refer to the average CPI expenditure weight over the sample period.

Figure A.6 decomposes the absolute size of adjustments into price increases (in yellow) and price decreases (in red). Throughout the sample, price decreases are almost consistently larger in absolute terms than price increases (16 percent vs. about 12 percent on average). Both series trend upward over the sample, with increases rising from roughly 10.5 percent before 2010 to 13.1 percent afterward, and decreases rising from around 13.4 percent to 18.2 percent over the same periods. Despite these long-run changes, neither series exhibits pronounced seasonal variation. Price decreases tend to be slightly larger in December and January, while the size of price increases is marginally smaller in January and April, but these differences are small and not systematic.

Figure A.7 separates the size of price changes for non-energy industrial goods (in yellow) and services (in red). Service-sector price changes are remarkably stable:

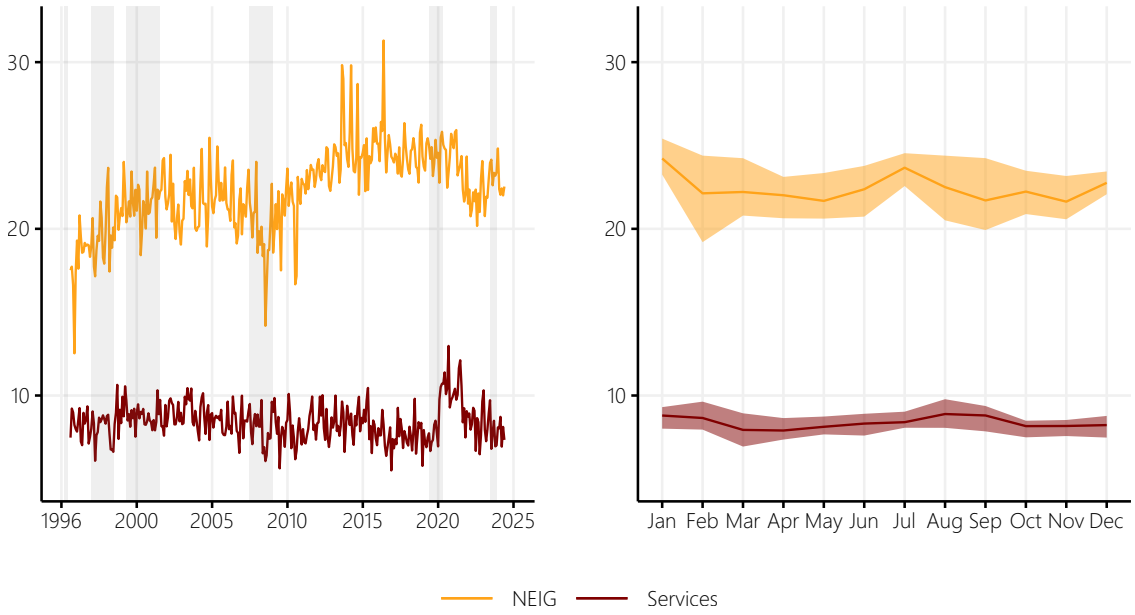
Figure A.6. Size of price increases and decreases



Notes: The figure shows the frequency of price increases (in yellow) and the frequency of price decreases (in red) in the United Kingdom, including price changes due to temporary sales and product substitution. Moments are computed at the item level and aggregated to weighted medians using CPI expenditure weights. The left panel shows the median frequencies of price increases and decreases from February 1996 to December 2024. The right panel shows the median frequencies of price increases and decreases by calendar month across the sample period, with the interquartile ranges indicated by the shaded area. Outliers related to VAT rate changes are excluded. Recessions (gray shaded areas) are dated by the Organisation for Economic Co-operation and Development (OECD). All values are expressed in percent.

they average 8.4 percent over the full sample and remain below 10 percent for nearly the entire period, with the post-pandemic inflation surge as the only notable deviation (when the size briefly exceeds 13 percent). In contrast, the absolute size of price changes for non-energy industrial goods is substantially higher—averaging 22.3 percent—and displays more pronounced cyclical variation. The seasonal profiles on the right-hand side remain flat for both sectors, consistent with the overall finding that the size of price changes exhibits little systematic seasonal pattern.

Figure A.7. Size of price changes in non-energy industrial goods and services



Notes: The figure shows the frequency of price changes for non-energy industrial goods (in yellow) and services (in red) in the United Kingdom, including price changes due to temporary sales and product substitution. Moments are computed at the item level and aggregated to weighted medians using CPI expenditure weights. The left panel shows the median frequencies of NEIG and services price changes from February 1996 to December 2024. The right panel shows the median frequencies of NEIG and services price changes by calendar month across the sample period, with the interquartile ranges indicated by the shaded area. Outliers related to VAT rate changes are excluded. Recessions (gray shaded areas) are dated by the Organisation for Economic Co-operation and Development (OECD). All values are expressed in percent.

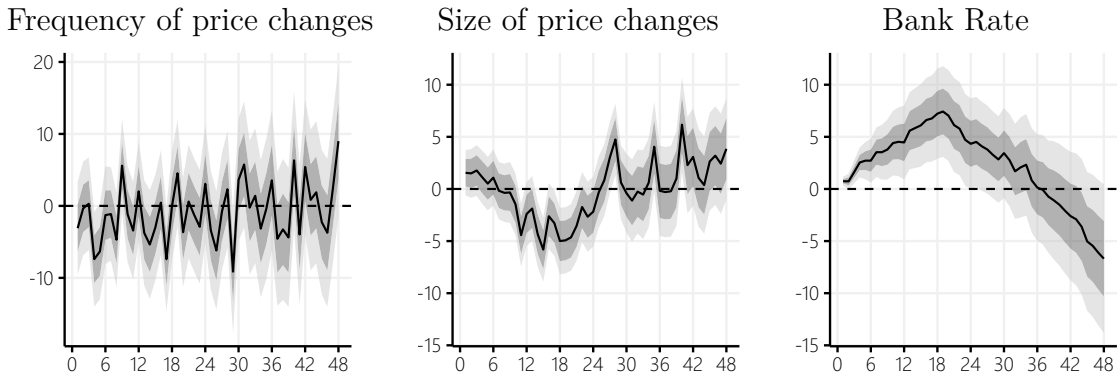
B Econometric method

This appendix refers to Section 3 and provides additional information not included in the main body of the paper.

B.1 Exogeneity of the frequency of price changes

To test the exogeneity of the frequency of price changes empirically within my dataset and econometric framework, I estimate the response of the frequency and (absolute) size of price changes to monetary policy shocks. Figure B.1 shows the estimated coefficients together with their 68% and 95% confidence intervals using linear local projections. The left panel reports the results for the frequency of price changes, the middle panel for the absolute size of price changes, and the right panel for the Bank Rate. The data used for estimation span from 1997:07 to 2024:12. The coefficients are shown over a four-year horizon (48 months). The results confirm that the frequency of price changes does not react significantly to monetary policy shocks, whereas the size of price changes decreases following a contractionary shock.

Figure B.1. Exogeneity of the frequency of price changes



Notes: The figure shows the coefficients together with their 68% and 95% confidence intervals estimated using linear local projections. The left panel reports the results for the frequency of price changes, the middle panel for the absolute size of price changes, and the right panel for the Bank Rate. The data used for estimation span from 1997:07 to 2024:12. The coefficients are shown over a four-year horizon (48 months).

B.2 Data sources

Table B.1 gives details on the data used in the paper, including information on the coverage and data sources.

Table B.1. Data description, sources, and coverage

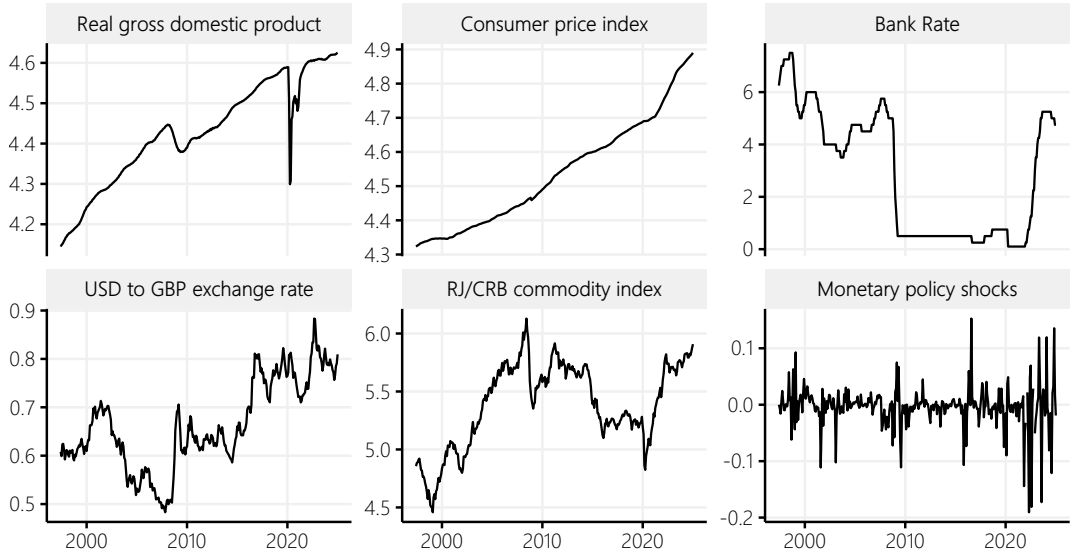
Variable	Description	Source	Sample
Microdata			
CPI price quotes	Price quote data that underpin consumer price inflation statistics in the United Kingdom	ONS	1996:01–2024:12
CPI item indices	Item index data that underpin consumer price inflation statistics in the United Kingdom	ONS	1996:01–2024:12
Monetary policy shocks			
Path	Monetary MPC factors estimated in multiple sterling futures and gilt yields (Baseline)	UKMPD	1997-06-06–2024-12-19
Events	Monetary event factors estimated in multiple sterling futures and gilt yields (Baseline)	UKMPD	1997-06-06–2024-12-19
FSScm1	Monetary MPC surprises in the 3M Libor 1st quarterly contract	UKMPD	1997-06-06–2024-12-19
SON3c1	Monetary MPC surprises in the 3M Sonia 1st quarterly contract	UKMPD	2021-05-06–2024-12-19
GBP1YT=RR	Monetary MPC surprises in 1-year gilt yields	UKMPD	2021-05-06–2023-06-22
GBP10YT=RR	Monetary MPC surprises in 10-year gilt yields	UKMPD	2021-05-06–2023-06-22
CTV	High-frequency monetary policy surprises	Cesa-Bianchi et al. (2020)	1997:06–2015:01
CH	Narrative monetary policy surprises of Cloyne and Hürtgen (2016), extended by Cesa-Bianchi et al. (2020)	Cesa-Bianchi et al. (2020)	1997:06–2009:02
Baseline variables			
GDP	Real gross domestic product (GDP) monthly estimate, seasonally adjusted	ONS	1997:01–2024:12
DKC6	UK consumer price index (CPI), all items excluding energy, food, alcoholic beverages & tobacco (2015=100), seasonally adjusted	ONS	1988:01–2024:12
RATE	Bank Rate	BoE	1972-10-16–2024-11-07
DEXUSUK	US Dollars to UK Pound Sterling Spot Exchange Rate	FRED	1971-01-04–2024-12-31
BAA10Y	Moody's Seasoned Baa Corporate Bond Yield Relative to Yield on 10-Year Treasury Constant Maturity	Federal Reserve Bank of St. Louis	1986-01-02–2024-12-31
Additional variables			
Economic activity			
GDP Production	Real gross domestic product (GDP) monthly estimate, Production (B-E), seasonally adjusted	ONS	1997:01–2024:12
GDP Services	Real gross domestic product (GDP) monthly estimate, Services (G-T), seasonally adjusted	ONS	1997:01–2024:12
GBRRECDM	Recession Indicators for the United Kingdom from the Peak through the Trough	OECD	1952:02–2022:09
UKNGDP	Real gross domestic product (GDP), quarterly	ONS	1955:I–2024:IV

gbprod00571	Industrial production index, Total, Constant prices, SA	ONS	1948:01–2024:12
LRHUTTTTGBM156S	Infra-Annual Labor Statistics: Monthly unemployment rate total: 15 years or over	OECD	1983:01–2024:12
Prices			
D7BT	UK consumer price index (CPI), all items (2015=100), seasonally adjusted	ONS	1988:01–2024:12
D7F4	UK consumer price index (CPI), Goods (2015=100), seasonally adjusted	ONS	1988:01–2024:12
D7F5	UK consumer price index (CPI), Services (2015=100), seasonally adjusted	ONS	1988:01–2024:12
CHAW	UK retail price index (RPI), all items (Jan 1987=100)	ONS	1987:01–2024:12
L522	UK consumer price index (CPIH) including owner occupiers' housing costs, all items (2015=100)	ONS	1988:01–2024:12
gbpric27001	UK producer price index (PPI), output, manufactured products for the domestic market, index (2015=100)	ONS	1957:01–2024:12
GBRGDPDEFQISMEI	GDP deflator, SA, index (2015=100)	ONS	1960:I–2024:IV
Financial variables			
gbpric27001	UK government benchmarks, 1-year gilt yield	Macrobond	1979-01-02–2024-12-31
BAMLC0A0CM	Investment-grade non-financial corporate bond spreads: ICE BofA US Corporate Index Option-Adjusted Spread	Ice Data Indices	1996-12-31–2024-12-31

Notes: The table provides details of the data used in the study, including information on coverage and data sources. Unless otherwise stated, the geographical scope of the variables refers to the United Kingdom.

Figure B.2 shows the series included in the baseline local projection model over the sample period 1997–2024.

Figure B.2. Transformed data series in the baseline local projection model



Notes: The figure shows the series included in the baseline local projection model over the sample period 1997–2024. Real GDP, the core consumer price index, and the commodity price index are in log levels multiplied by 100. The Bank Rate is in percent.

C Results

This appendix refers to Section 4 and provides additional information not included in the main body of the paper.

C.1 Linear impulse responses

Table C.1 is an update and extension of the tables reported in Cloyne and Hürtgen (2016) and Cesa-Bianchi et al. (2020) and provides an overview of the responses of economic activity and prices to monetary policy shocks from previous studies.

Table C.1. The effects of monetary policy shocks in the United Kingdom in previous studies

Study	Method	Period	Peak Effects		
			Quarter	Activity (in %)	Prices (in %)
Dedola and Lippi (2005)	VAR	1975:1-1997:3	6–8	-0.5 (IP)	0.2 (CPI)
Mountford (2005)	Sign Rest.	1974:I-2001:II	6	-0.6 (GDP)	-0.15 (GDP Defl)
Ellis et al. (2014)	FAVAR	1992:I-2005:IV	3	-2.0 (IP)	-2 (CPI)
Cloyne and Hürtgen (2016)	Narrative	1975:3-2007:12	3-4 (IP), 11 (CPI)	-0.5 (IP)	-1.0 (CPI Infl)
Gerko and Rey (2017)	Proxy-SVAR	1982:1-2015:1	10	-1.8 (IP)	1.0 (RPIX)
Cesa-Bianchi et al. (2020)	Proxy-SVAR	1992:1-2015:1	8 (GDP), 3-4 (CPI)	-1.6 (GDP)	-0.3 (CPI)
Braun et al. (2025)	Proxy-SVAR	1997:1-2019:12	8	ca. -1.5 (GDP)	ca. -0.45 (CPI)

Notes: The table is an update and extension of the tables reported in Cloyne and Hürtgen (2016) and Cesa-Bianchi et al. (2020). The results are from impulse responses displayed in previous studies for the United Kingdom and are the response of prices and indicators of economic activity to a one percentage point increase in the short-term nominal rate (Exceptions: In Gerko and Rey (2017), the shock increases the UK 5-year-rate by 20 basis points on impact; in Cesa-Bianchi et al. (2020), the shock increases the 1-year gilt rate by 25 basis points on impact). In brackets, I include the specific measures of economic activity and prices (IP denotes industrial production, GDP denotes gross domestic product, CPI denotes consumer price index, and RPIX denotes the Retail Price Index excluding mortgage interest payments).

C.2 Baseline results

Table C.2. Nonlinear local projection coefficients from the baseline model

	Horizon				
	$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
	Gross domestic product				
$\hat{\beta}_h$	-0.011	-0.127***	-0.114**	-0.090**	-0.051
$\hat{\gamma}_h$	-0.014**	-0.030*	-0.026	-0.027*	-0.030*
	Core consumer price index				
$\hat{\beta}_h$	0.001	-0.002	-0.028*	-0.100***	-0.168***
$\hat{\gamma}_h$	0.000	-0.001	-0.012	-0.055***	-0.067***
	Bank rate				
$\hat{\beta}_h$	0.429***	2.452*	2.990	-2.626	-5.741**
$\hat{\gamma}_h$	-0.116	0.619	0.604	-2.291	-3.233***

Notes: The table reports the estimated coefficients from the baseline nonlinear local projection model in Equation (2). For each response variable—real GDP, consumer prices, and the Bank Rate—the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequency of price adjustment. The data used for estimation span from 1997:07 to 2024:12. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.

D Sensitivity analyses

This appendix refers to Section 6 and provides the tables of the sensitivity checks that are not included in the main body of the paper.

D.1 Measurement of price rigidity

The baseline specification uses a measure of the frequency of price changes that includes price changes due to temporary sales and product substitution, consistent with the idea that all price changes relevant for the CPI should enter the state variable. As a robustness check, I re-estimate Equation (2) using interaction variables constructed from all combinations of alternative treatments that include or exclude price changes due to temporary sales and product substitution. I identify temporary sales and product substitution using flags provided by the ONS. To exclude temporary sales, I replace each sales price with the last observed non-sales price. To exclude product substitutions, I start a new price spell with each product substitution. Table D.1 shows that the main conclusions regarding the state dependence of monetary transmission are qualitatively robust to these alternative measures of price rigidity.

D.2 Alternative monetary policy shocks

A key concern in empirical monetary policy analysis is the sensitivity of estimated transmission effects to the choice of shock measure. To assess the robustness of the state-dependent responses documented in the baseline specification, I re-estimate the nonlinear local projection model using a broad set of alternative monetary policy shocks drawn from both the UK Monetary Policy Event-Study Database (UKMPD) and external sources.

Within the UKMPD, I consider several variants of high-frequency event-study shocks. First, I employ a broader event definition that includes the press conferences accompanying the release of the Bank of England's quarterly Monetary Policy Report, in addition to the baseline shocks around monetary policy events defined by market surprises following Monetary Policy Committee (MPC) announcements. In addition, rather than relying on the factor structure underpinning the Path shocks, I re-estimate the model using individual-asset surprises at different maturities: the first quarterly contracts of the 3-month interest rate futures (based on Libor until 2021 and on Sonia thereafter) as well as the 1-year and 10-year benchmark gilt yields. These shocks are informative because they isolate movements on specific parts of the yield curve and allow me to examine whether the state dependence identified in the baseline is an artifact of the factor extraction procedure.

Table D.1. Robustness to the treatment of price changes due to temporary sales and product substitution in estimating the frequency of price changes

		Horizon				
		$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
		Gross domestic product				
With sales, with substitutions	$\hat{\beta}_h$	-0.011	-0.127***	-0.114**	-0.090**	-0.051
	$\hat{\gamma}_h$	-0.014**	-0.030*	-0.026	-0.027*	-0.030*
With sales, without substitutions	$\hat{\beta}_h$	-0.011	-0.127***	-0.112**	-0.081*	-0.044
	$\hat{\gamma}_h$	-0.015**	-0.035**	-0.031	-0.021	-0.029*
Without sales, with substitutions	$\hat{\beta}_h$	-0.011	-0.137***	-0.123**	-0.078*	-0.055*
	$\hat{\gamma}_h$	-0.015*	-0.040**	-0.040**	-0.034*	-0.034**
Without sales, without substitutions	$\hat{\beta}_h$	-0.011	-0.128***	-0.114**	-0.082**	-0.053*
	$\hat{\gamma}_h$	-0.017*	-0.044**	-0.041*	-0.035*	-0.036*
		Core consumer price index				
With sales, with substitutions	$\hat{\beta}_h$	0.001	-0.002	-0.028*	-0.100***	-0.168***
	$\hat{\gamma}_h$	0.000	-0.001	-0.012	-0.055***	-0.067***
With sales, without substitutions	$\hat{\beta}_h$	0.001	-0.002	-0.022	-0.089***	-0.149***
	$\hat{\gamma}_h$	0.000	-0.002	-0.012	-0.054***	-0.068***
Without sales, with substitutions	$\hat{\beta}_h$	0.001	-0.002	-0.019	-0.081**	-0.146***
	$\hat{\gamma}_h$	0.000	-0.002	-0.007	-0.055***	-0.069***
Without sales, without substitutions	$\hat{\beta}_h$	0.001	-0.003	-0.020	-0.077**	-0.132***
	$\hat{\gamma}_h$	0.000	-0.002	-0.009	-0.059***	-0.074***
		Bank rate				
With sales, with substitutions	$\hat{\beta}_h$	0.429***	2.452*	2.990	-2.626	-5.741**
	$\hat{\gamma}_h$	-0.116	0.619	0.604	-2.291	-3.233***
With sales, without substitutions	$\hat{\beta}_h$	0.398***	2.361*	2.410	-3.242	-5.402**
	$\hat{\gamma}_h$	-0.119	0.746	0.550	-2.520*	-3.400***
Without sales, with substitutions	$\hat{\beta}_h$	0.409**	1.533	2.477	-3.081	-5.851***
	$\hat{\gamma}_h$	-0.125	0.461	0.480	-2.611*	-3.533***
Without sales, without substitutions	$\hat{\beta}_h$	0.399**	1.576	2.373	-3.046	-5.078**
	$\hat{\gamma}_h$	-0.141	0.687	0.619	-2.970*	-3.912***

Notes: The table reports the estimated coefficients from the nonlinear local projection model in Equation (2), using interaction terms constructed from alternative measures of the frequency of price changes that either include or exclude price changes due to temporary sales and product substitution. For each response variable—real GDP, core CPI inflation, and the Bank Rate—the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequency of price adjustment. The data used for estimation span from 1997:07 to 2024:12. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.

From outside the UKMPD, I use two widely employed shock measures. The first is the series of high-frequency monetary policy surprises constructed by Cesa-Bianchi et al. (2020), available for the period 1997:06–2015:01. The second consists of narrative monetary policy shocks inspired by Romer and Romer (2004); I rely on the extended series compiled by Cloyne and Hürtgen (2016), which extends the original dataset through 2009:02.

Table D.2 shows that the qualitative results remain stable across specifications. Using alternative event constructions within the UKMPD yields responses very similar to the baseline, confirming that the key findings do not depend on the precise definition of monetary policy announcements. Shocks derived from individual assets display weaker and less precise responses, but they nonetheless preserve the overall pattern of stronger inflation responses in states of more frequent price adjustment. The external high-frequency and narrative shocks also produce dynamics aligned with the main results. Taken together, the evidence shows that the state-dependent effects of monetary policy documented in the paper are robust to a wide range of identification strategies and are not an artifact of any particular shock series.

D.3 Extended model specification

The baseline specification is intentionally parsimonious, including only the variables needed to capture the core transmission of monetary policy. To assess whether omitted channels materially affect the results, I estimate an extended version of the model that augments the baseline model with the 1-year government gilt yield, the producer price index (PPI), the US BAA corporate spread (which is the difference between the Moody's BAA corporate yield and the yield on the 10-year US Treasury constant maturity), and the FTSE 100 index.

The coefficients on the main variables in Table D.3 remain highly stable across horizons, confirming that the baseline dynamics are not sensitive to a richer information set. The estimated responses of the additional variables are both economically plausible (such as a decline in PPI following a contractionary shock) and broadly consistent with previous findings using the same shocks (Braun et al., 2025).

D.4 Alternative variables for economic activity and prices

The baseline specification employs monthly estimates of real GDP as the measure of economic activity and core CPI inflation as the measure of prices. Since the choice

Table D.2. Robustness to alternative monetary policy shocks

		Horizon				
		$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
		Gross domestic product				
Events	$\hat{\beta}_h$	0.021**	-0.037	-0.092**	-0.131***	-0.031
	$\hat{\gamma}_h$	-0.029**	-0.032*	-0.046***	-0.050**	-0.020
3-month Libor/Sonia futures	$\hat{\beta}_h$	0.023**	-0.054	-0.092**	-0.124***	-0.021
	$\hat{\gamma}_h$	-0.026***	-0.039**	-0.043***	-0.045**	-0.012
1-year gilt yield	$\hat{\beta}_h$	0.035*	0.015	-0.023	-0.113*	-0.036
	$\hat{\gamma}_h$	-0.021**	-0.011	-0.017	-0.047	-0.015
10-years gilt yield	$\hat{\beta}_h$	0.053	0.029	-0.032	-0.240***	-0.119**
	$\hat{\gamma}_h$	-0.025*	-0.007	-0.031	-0.125***	-0.064***
Cesa-Bianchi et al. (2020)	$\hat{\beta}_h$	0.023**	-0.072	-0.081*	-0.089***	-0.019
	$\hat{\gamma}_h$	-0.010***	-0.032	-0.062**	-0.060***	-0.021
Cloyne and Hürtgen (2016)	$\hat{\beta}_h$	0.001	0.017*	0.027***	0.016	0.005
	$\hat{\gamma}_h$	0.000	-0.002	0.003	0.006*	0.002
		Core consumer price index				
Events	$\hat{\beta}_h$	0.001	-0.001	0.006	-0.012	-0.071**
	$\hat{\gamma}_h$	0.000	0.000	0.001	-0.018	-0.045***
3-month Libor/Sonia futures	$\hat{\beta}_h$	0.000	0.012***	0.022**	0.028*	0.026
	$\hat{\gamma}_h$	0.003***	0.007***	0.011***	0.009	0.012
1-year gilt yield	$\hat{\beta}_h$	0.000	0.004	0.014	0.013	0.000
	$\hat{\gamma}_h$	0.002***	0.003	0.010**	0.006	-0.001
10-years gilt yield	$\hat{\beta}_h$	0.000	0.005	0.017	0.044	0.072**
	$\hat{\gamma}_h$	0.002**	0.003	0.015**	0.029*	0.039**
Cesa-Bianchi et al. (2020)	$\hat{\beta}_h$	0.000	0.005	0.011*	0.014**	0.014***
	$\hat{\gamma}_h$	0.001***	0.004**	0.008**	0.013***	0.009**
Cloyne and Hürtgen (2016)	$\hat{\beta}_h$	0.000	-0.001	0.001	0.002**	0.001
	$\hat{\gamma}_h$	0.000	0.000	0.000	0.001*	0.000
		Bank rate				
Events	$\hat{\beta}_h$	0.484**	0.661	2.192	0.904	-0.496
	$\hat{\gamma}_h$	-0.172*	-0.041	0.508	-0.329	-1.670
3-month Libor/Sonia futures	$\hat{\beta}_h$	0.021	-2.768**	-1.586	-2.002	-0.530
	$\hat{\gamma}_h$	0.453***	-1.164*	-0.760*	-0.861	-0.435
1-year gilt yield	$\hat{\beta}_h$	0.054	0.105	-1.751	-3.484*	0.053
	$\hat{\gamma}_h$	0.288***	-0.258	-0.894*	-1.230*	-0.332
10-years gilt yield	$\hat{\beta}_h$	-0.217	-2.774**	-3.864**	-3.322	3.874
	$\hat{\gamma}_h$	0.165	-1.069**	-1.930***	-1.084	1.328
Cesa-Bianchi et al. (2020)	$\hat{\beta}_h$	0.037	-0.641	-0.858	-1.579*	-1.131
	$\hat{\gamma}_h$	0.569***	-0.641	-1.170	-1.317***	-1.000
Cloyne and Hürtgen (2016)	$\hat{\beta}_h$	-0.282***	-0.718***	0.199	0.406*	0.490**
	$\hat{\gamma}_h$	0.228***	-0.338***	0.136*	0.252**	0.072

Notes: The table reports the estimated coefficients from the nonlinear local projection model in Equation (2), estimated separately using a range of alternative monetary policy shock measures, including surprises from the UKMPD based on a broader set of monetary policy events ("Events"), individual-asset shock series (3-month Libor/Sonia interest rate futures; 1-year and 10-years gilt yield), and external high-frequency (Cesa-Bianchi et al., 2020) and narrative shocks (Cloyne and Hürtgen, 2016). For each response variable—real GDP, core CPI inflation, and the Bank Rate—the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequency of price adjustment. The data used for estimation span from 1997:07 to 2024:12. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.

Table D.3. Robustness to an extended model specification

	Horizon				
	$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
Gross domestic product					
$\hat{\beta}_h$	-0.010	-0.190***	-0.092***	0.018	-0.007
$\hat{\gamma}_h$	-0.015***	-0.064***	-0.003	0.002	-0.005
Core consumer price index					
$\hat{\beta}_h$	0.001	-0.009	-0.047***	-0.120***	-0.140***
$\hat{\gamma}_h$	0.000	-0.005**	-0.020***	-0.055***	-0.054***
Bank rate					
$\hat{\beta}_h$	0.639***	0.707	1.938**	-6.508***	-5.030**
$\hat{\gamma}_h$	-0.113**	0.374	1.202***	-3.520***	-2.331***
1-year government gilt yield					
$\hat{\beta}_h$	0.649*	1.900**	1.486*	-5.537***	-1.977
$\hat{\gamma}_h$	-0.213	0.904**	0.692	-2.534***	-1.300**
Producer price index					
$\hat{\beta}_h$	0.022***	-0.024	-0.069	-0.295***	-0.239***
$\hat{\gamma}_h$	-0.007**	-0.019	-0.037	-0.132***	-0.098***
US BAA corporate spread					
$\hat{\beta}_h$	-0.197*	-0.804***	0.628**	0.118	0.137
$\hat{\gamma}_h$	-0.136*	-0.455***	0.131	-0.106	0.084
FTSE 100					
$\hat{\beta}_h$	-0.059	0.338***	-0.255**	-0.079	0.176
$\hat{\gamma}_h$	0.078*	0.204***	0.000	-0.114*	0.025
CRB					
$\hat{\beta}_h$	-0.052	-0.156	-0.486**	-0.965***	-0.887***
$\hat{\gamma}_h$	0.056	-0.027	-0.165*	-0.433***	-0.252**
USD-GBP exchange rate					
$\hat{\beta}_h$	-0.014	0.348***	-0.103	-0.128	-0.099
$\hat{\gamma}_h$	-0.001	0.183***	0.005	-0.111	-0.019

Notes: The table reports the estimated coefficients from the nonlinear local projection model in Equation (2), estimated for a model that extends the baseline with the 1-year government gilt yield, the producer price index, the US BAA corporate spread and the FTSE 100 index. For each response variable, the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequency of price adjustment. The data used for estimation span from 1997:07 to 2024:12. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.

of variables can be influential in parsimonious local projection frameworks, I assess the robustness of the main findings to alternative indicators for both activity and inflation. For economic activity, I re-estimate the nonlinear local projection model replacing real GDP with the industrial production index (IPI), while keeping the remaining variables consistent with the baseline specification. To evaluate robustness with respect to the price measure, I estimate versions of the model using several alternatives one at a time: the headline CPI, the CPI including owner-occupiers' housing costs (CPIH), the Retail Price Index (RPI), and the producer price index (PPI).

Across all alternative indicators, the results remain highly consistent with the baseline. As shown in Table D.4, both the linear coefficients $\hat{\beta}_h$ and the interaction coefficients $\hat{\gamma}_h$ retain their qualitative patterns, including the stronger price responses in periods of more frequent price adjustment. The economic activity responses also remain stable when substituting IPI for GDP. These findings confirm that the state-dependent effects of monetary policy documented in the paper do not depend on the specific choice of output or price series.

D.5 Model specification: Selection of lag order and model deterministics

A further set of robustness exercises examines whether the main results depend on specific modeling choices within the local projection framework, in particular the selection of lag length and the inclusion of deterministic terms. To this end, I re-estimate the model using a wide range of lag orders (Table D.5) and alternative deterministic components (Table D.6), assessing both the linear coefficients and the interaction terms for stability across specifications. Across all these variants, the core patterns remain intact.

D.6 Samples excluding the Great Recession and the COVID-19 pandemic

To verify that the detected state dependence is not driven by specific large macroeconomic disruptions, I re-estimate the model across a range of subsamples. In particular, I exclude the Great Recession by estimating the model either on the period before 2008 or on the period after 2009, and I exclude the COVID-19 period by restricting the sample to observations before 2020. Across these alternative samples, the responses

Table D.4. Robustness to alternative variables for economic activity and prices

		Horizon				
		$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
		Economic activity (baseline: monthly GDP)				
Alternative price variable: CPI	$\hat{\beta}_h$	-0.010	-0.123***	-0.129***	-0.107***	-0.053
	$\hat{\gamma}_h$	-0.013*	-0.024	-0.027	-0.025*	-0.033*
Alternative price variable: CPIH	$\hat{\beta}_h$	-0.019**	-0.134***	-0.124**	-0.080*	-0.045
	$\hat{\gamma}_h$	-0.014*	-0.029**	-0.018	-0.020	-0.034*
Alternative price variable: RPI	$\hat{\beta}_h$	-0.022***	-0.130***	-0.112**	-0.061	-0.051
	$\hat{\gamma}_h$	-0.014*	-0.030**	-0.022	-0.022	-0.038**
Alternative price variable: PPI	$\hat{\beta}_h$	-0.018**	-0.145***	-0.131***	-0.073	-0.051
	$\hat{\gamma}_h$	-0.011*	-0.037**	-0.025	-0.034**	-0.037**
Alternative economic activity: IPI	$\hat{\beta}_h$	0.038**	0.023	-0.042	-0.082**	-0.037
	$\hat{\gamma}_h$	-0.008	-0.004	-0.032***	-0.020	-0.017
		Prices (baseline: Core CPI)				
Alternative price variable: CPI	$\hat{\beta}_h$	0.004	-0.003	0.000	-0.156**	-0.242***
	$\hat{\gamma}_h$	0.001	-0.006	-0.008	-0.092***	-0.104***
Alternative price variable: CPIH	$\hat{\beta}_h$	0.000	-0.017	-0.017	-0.123**	-0.192***
	$\hat{\gamma}_h$	0.003**	-0.003	-0.004	-0.072***	-0.083***
Alternative price variable: RPI	$\hat{\beta}_h$	0.000	0.004	-0.025	-0.175**	-0.246***
	$\hat{\gamma}_h$	0.002	0.002	-0.006	-0.103***	-0.109***
Alternative price variable: PPI	$\hat{\beta}_h$	0.025***	-0.006	-0.041	-0.262***	-0.285***
	$\hat{\gamma}_h$	-0.008**	-0.013	-0.033	-0.151***	-0.110***
Alternative economic activity: IPI	$\hat{\beta}_h$	0.001	0.000	-0.047***	-0.130***	-0.168***
	$\hat{\gamma}_h$	0.000	0.002	-0.018**	-0.062***	-0.071***
		Bank rate				
Alternative price variable: CPI	$\hat{\beta}_h$	0.452**	1.599	2.866	-2.847	-4.783**
	$\hat{\gamma}_h$	-0.067	-0.156	0.718	-2.456*	-2.945***
Alternative price variable: CPIH	$\hat{\beta}_h$	0.496***	2.617**	2.061	-1.816	-2.943
	$\hat{\gamma}_h$	-0.033	0.675	0.531	-1.627	-1.530
Alternative price variable: RPI	$\hat{\beta}_h$	0.479**	3.830***	2.669	-2.579	-4.582*
	$\hat{\gamma}_h$	-0.002	1.066**	0.610	-1.746	-2.224**
Alternative price variable: PPI	$\hat{\beta}_h$	0.609***	2.255	0.662	-3.237	-4.621*
	$\hat{\gamma}_h$	-0.145	0.431	0.437	-2.234	-2.201**
Alternative economic activity: IPI	$\hat{\beta}_h$	0.321*	2.884**	2.571	-3.775	-5.895***
	$\hat{\gamma}_h$	-0.272***	0.883**	0.091	-2.761	-3.178***

Notes: The table reports the estimated coefficients from the nonlinear local projection model in Equation (2), estimated using alternative variables of prices (CPI, CPIH, RPI, PPI) and economic activity (IPI). For each response variable, the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequency of price adjustment. The data used for estimation span from 1997:07 to 2024:12. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.

Table D.5. Robustness to different lag orders

		Horizon				
		$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
		Gross domestic product				
Lag 6	$\hat{\beta}_h$	-0.007	-0.132***	-0.125**	-0.096*	-0.039
	$\hat{\gamma}_h$	-0.011**	-0.047***	-0.044***	-0.022	-0.011
Lag 8	$\hat{\beta}_h$	-0.012*	-0.132***	-0.135***	-0.080	-0.017
	$\hat{\gamma}_h$	-0.011**	-0.044***	-0.045***	-0.027	-0.016
Lag 10	$\hat{\beta}_h$	-0.015**	-0.146***	-0.140**	-0.085	-0.010
	$\hat{\gamma}_h$	-0.015**	-0.046***	-0.035**	-0.009	0.005
Lag 12	$\hat{\beta}_h$	-0.013	-0.155***	-0.134**	-0.075*	0.006
	$\hat{\gamma}_h$	-0.014**	-0.043**	-0.047***	-0.024	-0.004
Lag 16	$\hat{\beta}_h$	-0.010	-0.181***	-0.156*	-0.061	0.012
	$\hat{\gamma}_h$	-0.014**	-0.056**	-0.062**	-0.016	0.007
Lag 18	$\hat{\beta}_h$	-0.013	-0.153***	-0.142***	0.023	0.034
	$\hat{\gamma}_h$	-0.016**	-0.042*	-0.075***	0.007	-0.004
		Core consumer price index				
Lag 6	$\hat{\beta}_h$	0.003*	0.012	-0.018	-0.056	-0.082*
	$\hat{\gamma}_h$	0.000	0.002	-0.007	-0.035**	-0.040**
Lag 8	$\hat{\beta}_h$	0.002	0.001	-0.008	-0.043	-0.064
	$\hat{\gamma}_h$	0.000	0.003	-0.006	-0.034**	-0.038**
Lag 10	$\hat{\beta}_h$	0.002	-0.003	-0.014	-0.070**	-0.094**
	$\hat{\gamma}_h$	0.000	-0.001	-0.009	-0.035**	-0.040**
Lag 12	$\hat{\beta}_h$	0.002	0.000	-0.025	-0.086***	-0.118***
	$\hat{\gamma}_h$	0.000	0.000	-0.010	-0.048***	-0.050***
Lag 16	$\hat{\beta}_h$	0.003***	0.014	-0.011	-0.057**	-0.051
	$\hat{\gamma}_h$	0.000	0.003	-0.008	-0.036***	-0.034***
Lag 18	$\hat{\beta}_h$	0.003***	0.015*	-0.010	-0.017	-0.014
	$\hat{\gamma}_h$	0.000	0.008**	-0.002	-0.036***	-0.034**
		Bank rate				
Lag 6	$\hat{\beta}_h$	0.601***	6.901***	5.935**	-0.948	-7.238***
	$\hat{\gamma}_h$	-0.094	1.465*	1.434*	-0.742	-2.824***
Lag 8	$\hat{\beta}_h$	0.689***	5.060***	4.649**	-1.376	-5.637**
	$\hat{\gamma}_h$	-0.089	2.166***	1.934**	-1.017	-3.129***
Lag 10	$\hat{\beta}_h$	0.621***	4.176**	4.606**	-1.591	-5.161**
	$\hat{\gamma}_h$	-0.108	1.808***	1.645*	-1.298	-3.285***
Lag 12	$\hat{\beta}_h$	0.704***	5.315***	5.907***	-1.301	-6.108***
	$\hat{\gamma}_h$	-0.076	1.646***	1.958**	-1.542	-3.488***
Lag 16	$\hat{\beta}_h$	0.647**	4.527**	4.217**	-5.571***	-5.352***
	$\hat{\gamma}_h$	-0.102	1.658**	1.434*	-1.821	-2.652***
Lag 18	$\hat{\beta}_h$	0.606**	4.497***	4.533**	-4.088**	-3.145
	$\hat{\gamma}_h$	-0.053	1.278*	1.352*	-2.603**	-3.749***

Notes: The table reports the estimated coefficients from the nonlinear local projection model in Equation (2), estimated using different numbers of lags included in the model. For each response variable, the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequency of price adjustment. The data used for estimation span from 1997:07 to 2024:12. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.

Table D.6. Robustness to deterministic components of the model

		Horizon				
		$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
		Gross domestic product				
No constant	$\hat{\beta}_h$	-0.017**	-0.209***	-0.238***	-0.236***	-0.204**
	$\hat{\gamma}_h$	-0.015**	-0.061***	-0.085***	-0.086***	-0.083**
Constant	$\hat{\beta}_h$	-0.013	-0.155***	-0.134**	-0.075*	0.006
	$\hat{\gamma}_h$	-0.014**	-0.043**	-0.047***	-0.024	-0.004
Constant and trend	$\hat{\beta}_h$	-0.013	-0.155***	-0.136**	-0.076*	0.001
	$\hat{\gamma}_h$	-0.014**	-0.042**	-0.047***	-0.024	-0.005
		Core consumer price index				
No constant	$\hat{\beta}_h$	0.002	0.000	-0.025	-0.089*	-0.120*
	$\hat{\gamma}_h$	0.000	0.000	-0.011	-0.056***	-0.063**
Constant	$\hat{\beta}_h$	0.002	0.000	-0.025	-0.086***	-0.118***
	$\hat{\gamma}_h$	0.000	0.000	-0.010	-0.048***	-0.050***
Constant and trend	$\hat{\beta}_h$	0.002	-0.001	-0.025***	-0.086***	-0.112***
	$\hat{\gamma}_h$	0.000	-0.002	-0.015***	-0.043***	-0.050***
		Bank rate				
No constant	$\hat{\beta}_h$	0.697***	5.332***	5.792*	-1.015	-6.085***
	$\hat{\gamma}_h$	-0.075	1.756**	2.329	-0.829	-3.217***
Constant	$\hat{\beta}_h$	0.704***	5.315***	5.907***	-1.301	-6.108***
	$\hat{\gamma}_h$	-0.076	1.646***	1.958**	-1.542	-3.488***
Constant and trend	$\hat{\beta}_h$	0.704***	5.406***	5.918***	-1.301	-6.488***
	$\hat{\gamma}_h$	-0.065	1.934***	2.087**	-1.545	-3.474***

Notes: The table reports the estimated coefficients from the nonlinear local projection model in Equation (2), estimated using different deterministic components of the model (no constant, constant, constant and linear trend). For each response variable, the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequency of price adjustment. The data used for estimation span from 1997:07 to 2024:12. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.

of output, prices, and the policy rate closely track the baseline dynamics, indicating that the main findings are not dominated by these extreme episodes. While some estimates become less precise due to shorter sample periods and more tranquil policy environments, the fundamental nonlinear relationship between monetary policy shocks and the frequency of price adjustment remains clearly visible (Table D.7).

Table D.7. Robustness to different subsamples

		Horizon				
		$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
		Gross domestic product				
Before 2008	$\hat{\beta}_h$	-0.007***	-0.064***	-0.029	0.030**	-0.003
	$\hat{\gamma}_h$	-0.002*	-0.015*	-0.016	0.007	0.007
After 2009	$\hat{\beta}_h$	-0.040**	0.013	0.185**	-0.144***	-0.176**
	$\hat{\gamma}_h$	-0.030*	0.008	0.002	0.009	-0.277***
Before 2020	$\hat{\beta}_h$	-0.015***	-0.128***	-0.143***	-0.046	0.008
	$\hat{\gamma}_h$	-0.006***	-0.038**	-0.042**	-0.017	-0.003
		Core consumer price index				
Before 2008	$\hat{\beta}_h$	0.000	0.015*	0.012*	0.013**	0.001
	$\hat{\gamma}_h$	0.001	0.005	0.006	0.009***	0.001
After 2009	$\hat{\beta}_h$	0.002	0.004	0.020**	0.017**	-0.003
	$\hat{\gamma}_h$	0.000	0.002	0.010**	0.030***	0.013***
Before 2020	$\hat{\beta}_h$	0.000	-0.007	-0.023*	-0.034*	-0.059***
	$\hat{\gamma}_h$	0.000	-0.001	-0.006	-0.013**	-0.025***
		Bank rate				
Before 2008	$\hat{\beta}_h$	0.290	-1.284	-0.899	1.086	0.834
	$\hat{\gamma}_h$	0.277	0.004	0.138	-0.248	0.267
After 2009	$\hat{\beta}_h$	0.042	1.294***	0.054	-0.462	-0.672**
	$\hat{\gamma}_h$	-0.230***	0.329*	-0.283*	0.689	0.353**
Before 2020	$\hat{\beta}_h$	0.620**	8.100***	9.154***	3.073	-3.889**
	$\hat{\gamma}_h$	-0.012	3.047***	3.546***	1.014	-2.101***

Notes: The table reports the estimated coefficients from the nonlinear local projection model in Equation (2), estimated on different subsamples: excluding the Great Recession (estimated on the period before 2008; estimated on the period after 2009) and the COVID-19 pandemic (estimated on the period before 2020). For each response variable, the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequency of price adjustment. The data used for estimation span from 1997:07 to 2024:12. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.

D.7 Quarterly data

I further assess robustness to the frequency of the data by re-estimating the model at the quarterly level. Specifically, I use quarterly GDP and the GDP deflator, and aggregate all remaining monthly series—including the frequency of price changes—by taking simple averages within each quarter. The resulting impulse responses are broadly consistent with those obtained from the monthly baseline (Table D.8). Although the lower-frequency specification naturally yields somewhat less precise estimates, the qualitative patterns in both the linear coefficients and the nonlinear interaction terms remain intact, reinforcing the stability of the main conclusions.

Table D.8. Robustness to quarterly data

	Horizon				
	$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
	Gross domestic product				
$\hat{\beta}_h$	-0.046	0.025	0.000	-0.121**	-0.075
$\hat{\gamma}_h$	0.025	0.016	-0.044	-0.044*	-0.009
	GDP deflator				
$\hat{\beta}_h$	0.034**	0.015	0.007	0.028	0.028
$\hat{\gamma}_h$	-0.022*	-0.028	0.002	-0.010	0.002
	Bank rate				
$\hat{\beta}_h$	-0.925	6.322***	10.185***	2.068	-4.624*
$\hat{\gamma}_h$	0.509	1.968**	0.788	-1.057	-2.553***

Notes: The table reports the estimated coefficients from the nonlinear local projection model in Equation (2), estimated using quarterly data. For each response variable, the table lists the linear coefficients $\hat{\beta}_h$ and the nonlinear interaction coefficients $\hat{\gamma}_h$ at horizons $h = 1, 12, 24, 36,$ and 48 months. The linear coefficients $\hat{\beta}_h$ capture the direct effect of the monetary policy shock, while the interaction coefficients $\hat{\gamma}_h$ measure how the effect varies with the frequency of price adjustment. The data used for estimation span from 1997:III to 2024:IV. Statistical significance is indicated by * $p < 0.10$, ** $p < 0.05$, and *** $p < 0.01$.