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Contents

Abstract

Nominal rigidities play a central role in monetary policy transmission, shaping how inflation and real activity respond to various shocks. Using a large longitudinal dataset of granular price data for Poland covering 2000–2024, we contribute to the literature on price stickiness across several dimensions. First, we document the price-setting behaviour in Poland. Second, we show how the process was affected by the turbulent post-COVID-19 period and Russia's full-scale invasion of Ukraine, evaluating the importance of the intensive and extensive margins of price adjustments. Third, we distinguish between sticky and flexible sectors, revealing heterogeneity in prices response to economic variables and shocks. We also compare the price-setting process in Poland with those in the euro area and the US. Finally, we complement these analyses with simulations performed on a two-sector DSGE model. Overall, we provide new, comprehensive evidence on price rigidity and discuss its implications for monetary policy transmission.

JEL classification: D40; E31.

Keywords: price rigidity; stickiness; consumer prices; micro data; inflation.

1 Introduction

We study the nature and evolution of price rigidities in Poland over the period 2000–2024 using a unique, large-scale dataset of granular consumer prices. We document key patterns in the frequency, size, and duration of price changes across sectors and compare them with evidence for the euro area and the United States. The analysis also covers the turbulent post-COVID-19 period and the aftermath of Russia's full-scale invasion of Ukraine, which profoundly affected inflation dynamics and the price-setting behaviour. By harmonizing our approach with the methodology of Gautier et al. (2024, 2026), we ensure broad comparability of our results with findings for other advanced economies.

Understanding price rigidity is crucial for assessing how monetary policy and macroeconomic shocks are transmitted to prices and real activity. A widely accepted view in macroeconomics is that monetary policy affects real economic activity in the short run precisely because prices and wages do not adjust instantly (for discussion see Nakamura and Steinsson, 2013). Differences in the price-setting behaviour across countries, sectors, and time periods determine the speed and strength of policy effects and the persistence of inflation. Investigating how these mechanisms operate in different institutional and structural environments helps to better understand the nature of inflation dynamics and the effectiveness of monetary policy.

These questions are especially interesting in the context of Poland, for at least two reasons. First, Poland has undergone a remarkable economic transformation over the past three decades—from an emerging economy characterized by high inflation and multiple macroeconomic imbalances to the sixth-largest economy in the European Union, with robust growth, low pre-pandemic inflation, and an independent monetary policy framework. Second, it demonstrated strong resilience during the COVID-19 pandemic in terms of growth, but also high sensitivity to the inflationary effects of Russia's full-scale invasion of Ukraine, largely due to its geographic proximity to the war. As such, examining the price-setting in Poland offers valuable insights into how an open, fast-growing economy outside the euro area responds to both global and regional shocks.

We contribute to the literature across several dimensions. We start by providing a thorough documentation of the price-setting behaviour in Poland using a comprehensive micro-level dataset covering the period 2000–2024. We describe in detail the frequency, size, and duration of price changes across sectors and product categories, offering new evidence on the structure and dynamics of price rigidity in a large European economy outside the euro area. This rich descriptive analysis helps bridge a significant empirical gap in the existing literature, which has so far focused predominantly on the euro area and the United States. This information is essential not only for the central bank—seeking to understand the degree of price stickiness and thus the potential effectiveness of monetary policy—but also for researchers calibrating theoretical and empirical models, such as dynamic stochastic general equilibrium (DSGE) models, where the price stickiness parameter plays a crucial role.

Our main contribution, however, lies elsewhere. Specifically, we scrutinize how the price-setting behaviour evolved during the turbulent post-COVID-19 period and in the aftermath of Russia's full-scale invasion of Ukraine. We show that while the pandemic had only a limited impact on the price formation process, the invasion triggered a sharp and quite persistent increase in the frequency of price changes, reflecting the intensity and persistence of inflationary pressures. To further investigate these mechanisms, we build a DSGE model, which allows us to assess how changes in the price-setting behaviour influence inflation dynamics and monetary policy transmission. This analysis provides new insights into how firms adjust prices in response to large, overlapping shocks that simultaneously affect costs, demand, and uncertainty.

We also investigate whether price adjustments in Poland occur primarily along the intensive margin (changes in the size of price adjustments) or the extensive margin (changes in the frequency of price adjustments). We discuss the implications of these findings, contributing to the broader debate on the nature of nominal rigidities and their relevance for monetary policy transmission.

Furthermore, recognizing the substantial heterogeneity in the price-setting behaviour across sectors, we divide the sample into "sticky" and "flexible" sectors and document meaningful differences between them. We find that prices in flexible sectors tend to adjust much more frequently but in smaller magnitudes, and are responsive to both cyclical and external factors. In contrast, prices in the sticky sectors exhibit pronounced rigidity, influenced by the downward nominal wage rigidity, menu costs, and long-term contractual arrangements. To further investigate these mechanisms, we build a two-sector DSGE model with sticky and flexible sectors, which allows us to assess how sectoral heterogeneity in price rigidities shapes inflation persistence and the transmission of monetary policy. These findings shed light on the sectoral origins of inflation persistence and the uneven transmission of monetary policy across the economy.

We also provide a carefully designed international comparison of price rigidity between Poland, the euro area, and the United States, using a harmonized methodology and a common basket of goods with consistent weighting schemes. We contrast our findings with recent results for euro area countries presented by Gautier et al. (2024) and Gautier et al. (2026), enabling a coherent assessment of cross-country differences in price-setting behaviour. Our results show that Poland's price-setting patterns resemble those of the euro area more closely than those of the United States, although notable distinctions emerge in the services sector and in the response to recent geopolitical disruptions.

Finally, our research shows how shifts in price-setting behaviour can strengthen the effectiveness of monetary policy in bringing inflation back to target. Our evidence shows that the post-invasion surge in the frequency of price adjustments made prices in Poland more responsive to macroeconomic conditions, thereby increasing the sensitivity of inflation to interest rate changes. In such an environment, monetary policy gains traction because higher adjustment frequencies accelerate the pass-through of policy impulses to consumer prices. As inflationary pressures gradually receded, the elevated frequency of price adjustments allowed a given increase in policy rates to generate a stronger disinflationary effect at a given cost in terms of output than would have been possible under pre-invasion levels of nominal rigidity. This mechanism—confirmed in our DSGE framework—helps explain why inflation quickly and markedly declined despite persistent sectoral heterogeneity and why the recent tightening cycle in Poland proved more effective than historical patterns of price stickiness would have suggested.

1.1 Literature review

There are two major strands of literature on price rigidity. The first relies on extensive datasets of micro-level prices used in the compilation of consumer price indices (such as the CPI or HICP) to examine key features of the price-setting process and draw implications for monetary policy. The second focuses on price formation in online markets, comparing the behaviour of online prices with that observed in traditional micro price data underlying official consumer price indices.

Our paper fits into the first strand of the literature, which uses CPI micro data to study price rigidity. This line of research began with the seminal work of Bils and Klenow (2004). Since then, a growing number of empirical studies have examined price behaviour using micro-level data. For the U.S., Klenow and Kryvtsov (2008) and Nakamura and Steinsson (2008) provide in-depth analyses of price stickiness. In the European context, Álvarez et al. (2006) and Dhyne et al. (2005, 2006) offer comprehensive evidence based on harmonized price data.[1]

Klenow and Malin (2010) provide a comprehensive summary of the empirical literature on price rigidity that developed in the early 2000s. Several robust findings have emerged from this body of work:

i. Overall price rigidity is substantial, with average price durations of approximately one year. This rigidity tends to be more pronounced in the euro area than in the United States.
ii. There is considerable heterogeneity and asymmetry at the sectoral level, leading to substantial differences between mean and median price durations.
iii. The average size of price changes is large relative to the inflation rate, underscoring the role of idiosyncratic shocks. Nonetheless, many small (non-zero) price changes are also observed.
iv. There is limited evidence of downward nominal price rigidity in most sectors, with the notable exception of services.
v. Properly accounting for price discounts is essential for accurately measuring price rigidity. Discounts explain a large share of the cross-country variation in results and have significant implications for the calibration of theoretical models.
vi. Price-setting behaviour is affected by a variety of sector-specific factors but price rigidity tends to be higher in labour-intensive sectors.[2]
vii. The synchronization of price changes is low.

More recent research extends this strand of literature to different inflationary environments. Alvarez et al. (2019) analyse micro price data for Argentina covering periods of low inflation, high inflation, and hyperinflation.[3] Wulfsberg (2016) investigates price behaviour in Norway during episodes of low and high inflation. Similarly, Berardi et al. (2015) provide evidence

[1] The works by Dhyne et al. (2005, 2006) and Álvarez et al. (2006) gather results from the coordinated exercise carried out within the Inflation Persistence Network established by the ECB.

[2] Among others, these include: market structure (i.e., concentration and competition), business cycle, seasonality, the formation of producer prices or wage setting mechanisms.

[3] Previous study of this kind was conducted by Gagnon (2009) for Mexico.

on price-setting dynamics in France during a period of deflation. In parallel with these studies, additional evidence on price stickiness has emerged for various other economies.[4]

Most recently, Gautier et al. (2024) made a significant contribution to the study of price rigidity by analysing a comprehensive dataset of granular price data as part of a large-scale, harmonized exercise covering eleven euro area countries. This study documents price-setting behaviour during a period of low inflation. It was subsequently updated by Gautier et al. (2026) to incorporate data up to 2024, including the high-inflation period of 2022 and 2023. These studies provide the most directly comparable evidence and we will use them as a consistent reference throughout our analysis.

The second strand of the literature examines price formation in online markets. This line of research was employed by Lünnemann and Wintr (2011), Gorodnichenko and Talavera (2017), Gorodnichenko et al. (2018), and Cavallo (2018), with more recent contributions by Henkel et al. (2023), who analyse the price-setting behaviour during the COVID-19 period, and Strasser et al. (2023), who compare price dynamics in online and offline markets, or Stelmasiak et al. (2026), who examine price formation around VAT rate changes. While e-commerce has become increasingly pervasive across economies—as documented by Mućk et al. (2025)—price rigidity statistics in online markets may provide a different picture on the price-setting, particularly due to the higher sampling frequency of online data, as shown by Cavallo (2018). As such, this strand of literature remains only loosely connected to our work.[5]

[4] Contributions on price stickiness for selected economies include: Baudry et al. (2007) for France, Hoffman et al. (2006) for Germany, Macias and Makarski (2013) for Poland, Dixon et al. (2017) for the UK or Rudolf and Seiler (2022) for Switzerland. Dhyne et al. (2005, 2006) and Klenow and Malin (2010) provide further references to early, country-level works on large datasets on micro data. However, as the datasets, time span and methodological aspects may vary across these studies, we do not review them here. Álvarez et al. (2006) and Dhyne et al. (2006) summarize country-level results for a harmonized sample of 50 goods and services based on a unifying methodology discussed in Dhyne et al. (2005).

[5] Lünnemann and Wintr (2011) find ambiguous outcomes on the flexibility of online prices, but large heterogeneity in the frequency of price changes. Gorodnichenko and Talavera (2017) show that prices are more flexible in online markets than in regular stores, dispersed, changing weakly, but more sensitive to the exchange rate. Gorodnichenko et al. (2018) share similar findings but argue that despite small physical costs of price adjustment price setting in online markets is characterized by considerable frictions and by many metrics online prices are as imperfect as offline prices. Cavallo (2018) demonstrates that time-averaging and imputing prices can account for the differences in price rigidity in online and offline markets. Hillen and Fedoseeva (2021) argue that high flexibility in online prices is also sector-dependent.

The remainder of this paper is structured as follows. In Section 2 we characterize our dataset and outline the methodology for calculating price rigidity (exact formulas are presented in Appendix A). Section 3 provides the discussion of the main results. In Section 4 we study the importance of intensive and extensive margins for the price-setting process in Poland. Section 5 is devoted to the division of the CPI basket into flexible and sticky components and assessing their sensitivity to economic developments. In Section 6 we compare our results to the international evidence, whereas in Section 7 we simulate a two-sector DSGE model to illustrate how key economic variables respond to a monetary policy shock depends depending on the level of price stickiness. Section 8 concludes and summarizes our findings.

2 Data and methodology

In this section we characterize our micro price dataset. Next, we outline the empirical approach towards calculating price rigidity statistics.

2.1 Data

We conduct the analysis using annual datasets of prices quoted in outlets across Poland. These prices are collected by Statistics Poland and subsequently used to compile the Consumer Price Index (CPI).

In essence, granular data on prices of consumer goods and services are recorded continuously throughout each year. However, every year the composition of the consumption basket is updated. To make the analysis of price rigidities feasible, we first meticulously merge price series for representative products across the entire sample period and manually verify the accuracy of this process. Following this step, we obtain a panel dataset spanning the period from January 2000 to December 2024. Overall, our database contains over 81 million price quotes for up to 2000 representative consumer goods and services, recorded monthly in up to 300 locations, effectively covering the entire CPI basket.

Regarding data cleaning, we first exclude goods and services for which only average prices or price indices are reported, as these do not correspond to individual price quotes.[6] Second, we remove outliers, defined as price decreases exceeding 80% or price increases above 400%.[7] Despite these exclusions, coverage across all elementary CPI groups is preserved.

Our database contains no information on sales, imputations, product substitutions or quantity and quality adjustments. Therefore, we identify promotional prices using filters proposed by

[6] Such price aggregates are reported for certain products (e.g., for the most popular book sold in a specific retail chain). Moreover, Statistics Poland also incorporates scanner data and online prices into the CPI calculation. We exclude these observations as well, since they enter the database in the form of price indices rather than individual price quotes. Importantly, both Macias et al. (2023) and Szafranek et al. (2025) demonstrate that the dynamics of online and traditional retail price indices in Poland are closely aligned, suggesting that the exclusion of these data is unlikely to have a meaningful impact on our results.

[7] This is defined as the monthly rate of change in prices. A total of 5307 price quotes were excluded from the database, representing 0.04 percent of all recorded prices.

Gautier et al. (2024) and Nakamura and Steinsson (2008).[8] As for quantity adjustments, Statistics Poland already controls for the changes in the product package size to capture the "true" price change. Meanwhile, the share of quality-adjusted prices or imputations is relatively small, according to international evidence, and thus should not materially affect our results (Menz et al., 2022; Gautier et al., 2024).

2.2 Method

We study the price-setting process at the level of individual (unit) prices using the frequency approach, widely adopted in the literature (e.g., Bils and Klenow, 2004; Dhyne et al., 2006; Klenow and Kryvtsov, 2008; Nakamura and Steinsson; 2008; Gautier et al., 2024, Gautier et al., 2026).

In our analysis, an individual price quote refers to the price of a specific good or service recorded by Statistics Poland at a particular point of sale in a given month. Based on these observations, we compute key indicators of price stickiness—the frequency of price changes, the size of price changes, and the implied price duration. For the first two measures, we also report separate statistics for price increases and price decreases.

All measures are calculated for all the elementary groups, COICOP[9] classes (at the four-digit level) and divisions (at the two-digit level), as well as for several custom groups that we construct (i.e., main disaggregates of the CPI basket), alongside headline and core inflation baskets.[10] When aggregating the results for Poland, we apply the CPI weighting system. Since this approach is well established in the literature, we provide only a brief overview here; the detailed procedure for calculating the price stickiness measures and all related formulas are presented in Appendix A.

To account for changing macroeconomic conditions, we compute price-rigidity estimates for the entire sample, as well as separately for the pre-COVID-19 period (January 2000–December 2019) and on a yearly basis since 2020 (unless stated otherwise).

[8] The filter by Gautier et al. (2024) refines the procedure originally developed by Nakamura and Steinsson (2008). Therefore, we adopt this approach as our baseline, while also reporting results obtained using the alternative framework in Appendix C. Previewing ahead, the choice of filter has no material effect on our results.

[9] COICOP stands for *Classification of Individual Consumption by Purpose*. In Poland, consumer expenditures are categorized according to this classification system for the purpose of compiling the Consumer Price Index (CPI).

[10] Further results for specific groups are available upon reasonable request.

3 Results

The behavior of individual prices exhibits substantial heterogeneity across product categories. We begin with a simple illustration of two extreme cases of price dynamics. On one end of the spectrum are fuel prices, which are highly flexible and adjust frequently, with price reductions occurring regularly (Figure 1, left panel). At the opposite end are prices for men's haircuts, which are notably sticky and may remain unchanged for several years, with price decreases being virtually absent (Figure 1, right panel). Moreover, for certain products, price changes follow seasonal patterns, whereas for others, sales play a significant role in shaping the observed price patterns.

In the following part of this section, we present summary statistics describing the price-setting behavior in Poland, accounting for heterogeneity across product categories as well as changes over time. The analysis focuses on the frequency and size of price changes, the degree of synchronization across products, and the importance of sales in driving observed price adjustments.

3.1 Frequency of price changes

During the period 2000–2024, an average of 22.1% of prices were adjusted each month, implying that a typical price remained unchanged for approximately 4.0 months (Table 1). The median frequency of price adjustments was lower, at 14.5%, indicating a higher degree of price rigidity and corresponding to an implied price duration of 6.4 months.

At the disaggregated level, substantial heterogeneity in price rigidity is observed. Prices of goods are notably more flexible than prices of services. As expected, energy and food prices are very flexible (with the average frequency of price adjustment at 36.5% and 30.5%, respectively). Non-energy industrial goods are still quite flexible and change on average after roughly 5 months. In turn, services prices are sticky, with the mean and the median frequency of price changes at only 8.8% and 3.0%, respectively.

For the overall CPI basket, the average frequency of price increases amounts to 12.8%, while price decreases occur at a rate of 9.3%, suggesting that price cuts are nearly as common as price hikes. A similar pattern—both in mean and median terms—is found across most major product categories (Table 1, Figure 2-3). The services sector stands out as an exception: prices there increase roughly twice as often as they decrease (6.0% vs. 2.8%), likely reflecting the importance of wages in setting prices and their downward nominal rigidity.

Overall, price changes exhibit a relatively low degree of synchronization, as indicated by the Fisher–Konieczny ratio of 0.27, with the notable exception of energy prices, which tend to adjust more synchronously.

The frequency of price changes remained broadly stable between 2000 and 2019, averaging 20.3% (Table 2). During this period, movements in the frequency of price adjustments were largely seasonal in nature.[11] After accounting for seasonal variation (Figure 4), it becomes evident that the most pronounced shifts in the frequency of price adjustments were associated with one-off shocks, such as Poland's accession to the European Union in 2004 and the VAT rate increase in 2011.

In contrast, the recent period of 2020–2024 presents a markedly different picture. The frequency of price changes rose significantly, averaging around 30%, largely in response to external shocks—most notably the COVID-19 pandemic and the Russian invasion of Ukraine.[12] The frequency of price changes peaked in 2022, before gradually declining in subsequent years, though it has remained elevated relative to the pre-pandemic level. This increase is somewhat more pronounced for the median frequency, indicating that it was broad-based rather than driven by large price movements within a few specific product groups (Table 2-3, Figure C5).

Figure 4 further illustrates several distinct spikes in the frequency of price adjustments over this period—corresponding to the reopening of the economy after COVID-19 lockdowns[13], the shock following the Russian full-scale invasion of Ukraine and, more recently, the phasing out of government shielding mechanisms for households introduced during the pandemic.

[11] Strong seasonality is visible both for the whole basket and main subcomponents (Figure C5 in Appendix C).

[12] Szafranek et al. (2024) document the sources of the inflationary episode in Poland using a structural Bayesian VAR model identified through a combination of zero and sign restrictions. They are mostly external.

[13] In the initial phase of the COVID-19 pandemic, many firms pledged not to adjust their prices. Once the lockdown was lifted, however, the frequency of price changes spiked, indicating that most firms reset their prices in response to various factors, such as higher costs of complying with sanitary restrictions.

Throughout this time, the overall rise in the frequency of price changes stemmed mainly from more frequent price increases.[14]

3.2 Size of price changes

Typical upward and downward price adjustments tend to be relatively large (Table 1). Both the mean and median size of price increases (13.3% and 6.7%, respectively) and decreases (11.3% and 7.5%, respectively) exceed the observed inflation rate. This finding aligns with Gautier et al. (2024), who interpret such a pattern as evidence that idiosyncratic shocks at the firm level play a more significant role in price resetting than aggregate shocks. This interpretation reflects the fact that, at any given time, price increases and decreases occur simultaneously across products, while a large share of prices remains unchanged.

The substantial gap between the mean and the median results from the prevalence of numerous small price adjustments—half of all price changes fall within the range of -6.3% to 8.7 %—as well as from the presence of fat tails in the distribution. Specifically, 5% of prices increased by more than 29.1% or fell by more than 30.6% (Table 4). Moreover, the density of round-number price adjustments rises across the distribution, indicating the importance of sales and psychological pricing behavior (Figure 5).

We find comparable magnitudes of price changes across most product categories. The only notable exception is energy prices, for which both the mean and median of upward (7.2% and 3.4%, respectively) and downward (6.2% and 3.5%, respectively) adjustments are substantially smaller (Table 1). This difference largely reflects the smaller average changes observed in fuel prices, which are also among the most flexible components of the CPI basket.

Over the period 2000–2024, the average size of both upward and downward price adjustments increased. These increases were driven primarily by larger price changes in food and non-alcoholic beverages, as well as in non-energy industrial goods. By contrast, the magnitude of

[14] Figure 4 also shows that when the frequency of price changes spiked, the size of price changes generally remained consistent with its previous trend. This suggests that, following a shock, firms tend to adjust prices more frequently by similar—or even smaller—margins, rather than implementing large, one-off price revisions.

price adjustments for energy and services remained broadly stable throughout the sample period.[15]

3.3 The role of promotions

Promotions play an important role in the price-setting process in Poland. Over the period 2000–2024, promotional prices accounted for 3.4% of all recorded prices (Table 1).[16] Their share is even more pronounced in the context of price adjustments. On average, 4.7% of prices change each month as a direct result of sales. Consequently, as much as 21.3% of all price changes can be attributed to sales[17], of which 24.1% represent all downward and 19.2% all upward adjustments, related to the end of sales.

The sizable contribution of promotions implies that excluding them from the analysis would substantially alter key measures of price rigidity. The average frequency of price changes would decline from 22.1% to 18.0% (median: from 14.5% to 11.1%), while the implied average price duration would extend from 4 to 5 months (or from 6.4 to 8.5 months when considering medians). Similarly, the frequency of both upward and downward price adjustments would fall to 10.7% and 7.3%, respectively.

The average size of promotional price changes is comparable to that of regular price adjustments. As a result, when promotions are excluded, the average upward adjustment equals 12.5%, compared with 13.3% when promotions are included, while the average downward adjustment amounts to 11.0%, relative to 11.3% with promotions. The distribution of regular price changes also becomes slightly more concentrated: roughly 34% of prices change by no more than 3.5%, compared with 28% when promotions are included (Table 4). Furthermore, the distribution exhibits thinner tails (Figure 5). This pattern is particularly pronounced for food and non-energy goods, reflecting the high prevalence of promotions within these categories and—by extent—the competitive nature of the market.

[15] These results can be found in Appendix C (Figure C9 and Figure C10).

[16] In the paper, we present results obtained using the filter proposed by Gautier et al. (2024). For comparison, results based on the Nakamura and Steinsson (2008) filter are very similar and are reported in Appendix C (Table C1).

[17] For services and energy, sales account for less than 10% of all price changes, whereas for food and non-alcoholic beverages, their share is approximately 25%.

4 Intensive and extensive margins of price adjustment

We now turn to examining whether price adjustments in Poland occur primarily along the *intensive* margin—i.e., through changes in the size of price adjustments—or along the *extensive* margin—i.e., through changes in their frequency.

Recomposed and counterfactual inflation. First, we follow Klenow and Kryvtsov (2008) as well as later contributions by decomposing observed inflation into its extensive and intensive components. The key insight is that the monthly rate of price change for any product group l at time t can be approximated by the product of the frequency and the size of the non-zero price changes for that group[18], denoted by $\tilde{\pi}_{lt}$:

$$\tilde{\pi}_{lt} = f_{lt} \times \mathrm{dp}_{lt}, \tag{1}$$

or it can be expressed as the sum of contributions from price increases and decreases:

$$\tilde{\pi}_{lt} = f_{lt}^{+} \times \mathrm{dp}_{lt}^{+} - f_{lt}^{-} \times \mathrm{dp}_{lt}^{-}. \tag{2}$$

This latter formulation recognizes that, at any point in time, substantial variation may occur in the frequency or size of price increases and decreases, potentially offsetting each other. As highlighted by Gagnon (2009), such offsetting dynamics can result in a relatively stable overall inflation, even amid pronounced heterogeneity in underlying price adjustments.

This framework allows us to perform simple counterfactual simulations, following the approach of Gautier et al. (2024). Specifically, we fix one of the two statistics—the frequency or the size of price changes—at its average level and examine whether the resulting counterfactual inflation replicates the observed dynamics of recomposed monthly inflation, as defined in equations (1) and (2). These counterfactual exercises are conducted for the overall CPI.[19]

Let $\pi_t^{\bar{f}}$ denote the counterfactual inflation rate assuming a constant frequency of price changes—consistent with the model of Calvo (1983), which postulates a constant probability of price adjustment. Similarly, let $\pi_t^{\overline{\mathrm{dp}}}$ represent the counterfactual inflation rate assuming a

[18] One clarification is in order. In Equation (1) dp_{lt} is in fact a price index of (all) non-zero price changes, whereas dp_{lt}^{+} and dp_{lt}^{-} denote the absolute sizes of price increases and decreases, respectively. Consequently, recomposed inflation calculated using Equations (5) and (6) will not be identical, as the two formulations capture different aggregation mechanisms for underlying price adjustments.

[19] Thus, from now on we omit subscript l for simplicity.

constant size of price changes, as in state-dependent models, where firms adjust mainly by changing the likelihood of a price change. These measures are computed as follows:

$$\pi_t^{\bar{f}} = \bar{f} \times \mathrm{dp}_t, \tag{3}$$

$$\pi_t^{\overline{\mathrm{dp}}} = f_t \times \overline{\mathrm{dp}}, \tag{4}$$

where $\bar{f}$ and $\overline{\mathrm{dp}}$ denote the average frequency and average size of price changes over the sample period.

For a more detailed decomposition, we define $\bar{f}^+, \overline{\mathrm{dp}}^+, \bar{f}^-, \overline{\mathrm{dp}}^-$ analogously, corresponding to the frequency and the size of price increases and price decreases, respectively. This allows us to construct additional counterfactual inflation measures: one, where the frequencies of price increases and decreases are fixed at their average values:

$$\pi_t^{\bar{f}^+, \bar{f}^-} = \bar{f}^+ \times \mathrm{dp}_t^+ - \bar{f}^- \times \mathrm{dp}_t^-, \tag{5}$$

and another, where the sizes of price increases and decreases remain constant:

$$\pi_t^{\overline{\mathrm{dp}}^+, \overline{\mathrm{dp}}^-} = f_t^+ \times \overline{\mathrm{dp}}^+ - f_t^- \times \overline{\mathrm{dp}}^-. \tag{6}$$

Finally, we account for the fact that the average size of price changes can itself be decomposed as:

$$\mathrm{dp}_t = \frac{f_t^+}{f_t} \times \mathrm{dp}_t^+ - \left(1 - \frac{f_t^+}{f_t}\right) \times \mathrm{dp}_t^-. \tag{7}$$

This implies that variations in dp_t may arise either from shifts in the relative frequency of price increases to total price changes, $\alpha_t = \frac{f_t^+}{f_t}$, or from changes in the size of price increases and decreases.

Following Gautier et al. (2024), we therefore define an additional counterfactual inflation measure that holds the size of price changes constant while allowing the relative frequency to vary:

$$\pi_t^{\bar{f}, \overline{\mathrm{dp}}^+, \overline{\mathrm{dp}}^-} = \alpha_t \times \bar{f} \times \overline{\mathrm{dp}}^+ - (1 - \alpha_{jt}) \times \bar{f} \times \overline{\mathrm{dp}}^-. \tag{8}$$

We evaluate the tracking performance of all counterfactual inflation measures using the Pearson correlation coefficient, comparing them against the observed recomposed inflation series.

Our results align closely with the findings of Gautier et al. (2024). We show that recomposed inflation—calculated as the product of the frequency and size of price changes, as in eq. (1)—is more strongly correlated with the counterfactual assuming a constant frequency of price changes (eq. 3) than with the counterfactual assuming a constant size of price changes (eq. 4). The Pearson correlation coefficient for the former pair equals 0.92, while for the latter it falls slightly to 0.81. Despite the relatively small difference in correlation values, the tracking performance of the counterfactual with constant size deteriorates markedly (see the upper panel of Figure 6). This suggests that the short-term variation in inflation primarily reflects changes in the size of price adjustments, consistent with Calvo-type or menu cost models operating in a low-inflation environment (e.g., Alvarez et al., 2019; Nakamura et al., 2018).

The conclusion changes once we separately account for the variation in the frequency and size of price increases and decreases. In this case, recomposed inflation—calculated as in eq. (2) – is highly correlated with the counterfactual measure that captures price adjustment through the extensive margin (eq. 6), with a Pearson correlation coefficient of 0.93. By contrast, the counterfactual measure based on a constant frequency of price adjustment (eq. 5) shows only limited correlation with the recomposed measure, as indicated by a Pearson coefficient of 0.35. Moreover, the counterfactual measure that allows for price adjustments via the extensive margin tracks recomposed inflation very closely throughout the entire sample period, whereas the measure that captures changes via the intensive margin mimics recomposed inflation poorly, with large discrepancies between the two series almost at every point in time. The lower row of Figure 6 illustrates these differences. Overall, these results indicate that inflation dynamics are primarily driven by movements in the frequency of price increases and decreases (i.e., adjustments along the extensive margin) rather than by changes in their average size (i.e., adjustments along the intensive margin).[20]

[20] We also examine the role of the share of price increases in the overall frequency of price changes. In this case, the correlation remains high, at 0.85, although the absolute differences between the counterfactual and recomposed measures widen in the more recent period. Importantly, the way the extensive margin operates varies systematically across inflation regimes: during periods of higher inflation, price adjustments are predominantly driven by a larger share of price increases, whereas during low-inflation periods, price changes more often take the form of price decreases. Over most of the sample, however, the overall frequency of price changes remains relatively stable, indicating that changes in the composition of adjustments—rather than in their total incidence—drive the dynamics of price setting. Only in the most recent period of elevated inflation do we observe a noticeable increase in the overall frequency of price changes.

Quantile regression. We also check whether the margin of price adjustment depends on the prevailing level of inflation. To this end, we estimate the following quantile regression equation:

$$Q_\tau(\pi_t|X_t) = X_t\beta_\tau, \tag{9}$$

where $Q_\tau(\pi_t|X_t)$ denotes the τth quantile of the monthly growth rate in consumer prices, π_t, conditional on the set of explanatory variables, $X_t = [1, \pi_t^{\bar{f}^+,\bar{f}^-}, \pi_t^{\overline{\mathrm{dp}}^+,\overline{\mathrm{dp}}^-}, I(\pi_t^{yy} > d)\pi_t^{\bar{f}^+,\bar{f}^-}, I(\pi_t^{yy} > d)\pi_t^{\overline{\mathrm{dp}}^+,\overline{\mathrm{dp}}^-}]$. This set includes a constant term, counterfactual inflation under the assumption of constant frequency of price increases and decreases, $\pi_t^{\bar{f}^+,\bar{f}^-}$, counterfactual inflation under the assumption of constant size of price increases and decreases, $\pi_t^{\overline{\mathrm{dp}}^+,\overline{\mathrm{dp}}^-}$, as well as their interactions with an indicator function $I(\cdot)$ that equals one when the underlying condition is satisfied—specifically, when the annual headline inflation rate π_t^{yy} exceeds a threshold $d = 3.5\%$. We estimate parameters β_τ by minimizing the following loss function:

$$\mathcal{L}(\beta) = \sum_{t=1}^{T} Q_\tau(\pi_t - X_t\beta), \tag{10}$$

where $Q_\tau(u) = u\big(\tau - I(u < 0)\big)$. In our estimation the quantile parameter τ changes from 0.05 to 0.95, increasing in increments of 0.025.

We plot the outcomes of this exercise on Figure 7, which contrasts the outcomes from our quantile regression with those obtained from a simple linear model estimated using OLS (represented by red solid and dashed lines). The figure shows that the intensive margin of price adjustment plays a limited role in explaining the monthly rate of change in prices, regardless of the inflation environment. In most cases, the confidence intervals from both the quantile regression and OLS estimations overlap and include zero, indicating a lack of statistical significance of both linear and non-linear effects of the intensive margin.

In contrast, the extensive margin is statistically significantly associated with monthly inflation across all quantiles, with its impact increasing slightly toward the right tail of the inflation distribution. Notably, for the highest quantiles, the strength of the adjustment becomes more pronounced when the annual inflation rate exceeds the upper bound of the inflation target.

This suggests that in periods of high inflation—both on a monthly and annual basis—price adjustment through the extensive margin becomes more dominant compared to low-inflation regimes.

Overall, the results from both exercises discussed in this section align closely with evidence from other countries (e.g., Gautier et al., 2024; Gautier et al., 2026; Nakamura and Steinsson, 2008; Nakamura et al., 2018; Gagnon, 2009; Alvarez et al., 2019). Consistent with these studies, we find that the size of individual price changes (the intensive margin) has only a weak relationship with inflation, whereas the frequency of price adjustments (the extensive margin) plays the central role in explaining inflation dynamics—especially during periods of elevated inflation.

5 Flexible and sticky inflation

Having established the degree of price rigidity at the elementary group level, we divide goods and services into two baskets: *flexible* and *sticky*. We classify each elementary group $j \in J$ as either flexible or sticky based on the median frequency of price changes across all groups over the sample period (as in Bryan and Meyer, 2010). The rationale is straightforward: if the median frequency of price changes for group j exceeds the overall median, it is categorized as flexible; if it falls below the overall median, it is categorized as sticky. Once all elementary groups are classified, we compute the price indices for these two baskets following the official methodology used by Statistics Poland.

The results show that the frequency of price changes for the flexible basket is substantially higher, at 33.3%, compared with 8.6% for the sticky basket (Table 5). This difference implies a pronounced gap in implied average price duration—approximately 2.5 months for flexible prices versus 11 months for sticky ones. We also observe that the average size of both upward and downward price adjustments is considerably larger for sticky prices. Thus, while flexible prices change frequently and by relatively small amounts, sticky prices adjust less often but by a greater magnitude. Finally, the groups with flexible prices make up a slight majority of the overall CPI basket, accounting for roughly 55% on average.

Figure 8 illustrates the evolution of these inflation measures. Two key observations emerge. First, the monthly rate of change in flexible inflation (left panel) exhibits greater volatility than its sticky counterpart. This heightened variability primarily reflects seasonal variation, with the index rising and falling more frequently throughout the year. Moreover, during the COVID-19 period, flexible inflation responded more sharply and more rapidly, whereas adjustments in sticky inflation occurred later but were of similar magnitude.

Second, when examining annual rates of change, flexible inflation displays more pronounced cyclicality and variability. After the initial disinflation triggered by the COVID-19 outbreak, this measure quickly exceeded 20% following Russia's full-scale invasion of Ukraine. The recent disinflation episode has also been swift, mirroring previous cycles (such as the early 2000s and other low-inflation periods).

In contrast, sticky inflation is more persistent over time—it builds up and declines gradually, peaks later, and only rarely turns negative. Interestingly, sticky inflation began to rise already at the onset of the pandemic, suggesting that the shock induced notable price adjustments in the sticky-price sectors, while inflation in the flexible sectors initially declined sharply.

Having constructed consumer price indices based on flexible and sticky price baskets, we next estimate a set of regressions, which can be perceived as a reduced-form, stylized Phillip curve:

$$\pi_t^{**} = \mu + \sum_{j=1}^{P} \rho_j \pi_{t-j}^{**} + \beta \hat{y}_t + \lambda_1 e_t + \lambda_2 \pi_t^{en} + \lambda_3 gvc_t + \gamma E_t[\pi_{t+1}^{**}] + \eta_t, \qquad (11)$$

where $\pi_t^{**} \in \{\pi_t, \pi_t^{FL}, \pi_t^{ST}\}$ denotes the quarterly log growth rate in headline inflation (π_t), flexible inflation (π_t^{FL}), or sticky inflation (π_t^{ST}), μ is the constant term, $\hat{y}_t$ measures economic slack (output gap approximated using the production function methodology), e_t is the quarterly log growth rate in nominal effective exchange rate for Poland (taken from BIS), π_t^{en} is the quarterly log growth rate in a wide energy price index (from World Bank), and gvc_t denotes country-specific supply chain disruptions (measured as in Mućk and Postek, 2025).[21] Finally, $E_t[\pi_{t+1}^{**}]$ is the expected quarterly log growth rate in the dependent variable, and η_t is the error term.[22]

To capture persistence and distributed effects of the explanatory variables, we include two lags of the dependent variable, $P = 2$. To address potential endogeneity, the model is estimated using two-step GMM with HAC standard errors.[23] The Sargan test of overidentifying restrictions is employed to verify the statistical validity of the results.[24]

[21] We follow the approach by Mućk and Postek (2025) and measure supply shortages as the share of enterprises in Poland reporting actual shortages in materials and equipment. The data are sourced from Eurostat. By that we acknowledge the fact that the GSCPI proxy (an alternative measure for supply disruptions, see Benigno et al., 2022) is less suitable, as it combines information from several main economies and is more related to the overall cost pressure.

[22] The choice of the variables is guided by previous studies on the Polish economy (e.g., Szafranek, 2017). To avoid understating inflationary pressures from the energy sector, we use a global, broad measure of energy commodity prices, as suggested by Szafranek et al. (2024).

[23] The default bandwidth parameter in the HAC weighting matrix is used (Andrews, 1991). To compute the covariance matrix of the vector of sample moments conditions Bartlett kernel is used. We have also run the iterative GMM (Hansen et al., 1996). Most results remain qualitatively robust.

[24] As instruments we use four lags of the dependent variable, two lags of output gap and unemployment gap, and up to two lags of the exogenous variables included in the specification.

We have several expectations regarding the estimated parameters. First, we anticipate that the persistence of the inflation series, calculated as $\hat{\rho} = \sum_{j=1}^{P} \hat{\rho}_j$, should be highest for the sticky price index and lowest for the flexible one. Moreover, we expect that flexible inflation should respond more strongly to fluctuations in the output gap and external factors than sticky inflation.

The estimation results presented in Table 6 confirm these expectations. The sensitivity of flexible inflation to variations in the output gap is higher than for headline inflation and substantially higher than for sticky inflation. By comparing 95% confidence intervals for these estimates we also note that the response of flexible and sticky inflation is statistically different. A similar pattern emerges for the effects of the exchange rate and energy prices. For both these cases the magnitude of the prices response is statistically different. As regards the impact of supply chain disruptions we observe that this factor affects flexible inflation but the impact on sticky inflation is insignificant. The effect on headline inflation is significant only at α =10%. This again is in line with economic intuition as the chosen measure reflects mostly shortages originating in the manufacturing sector, thus should affect mostly flexible inflation (mostly goods), while sticky inflation (mostly services) should not be influenced. Finally, we observe that the forward-looking component is statistically significant only in the equation for sticky inflation. This can be interpreted that expectations are more important when prices change infrequently.

Consistent with our expectations, sticky inflation displays greater persistence than flexible inflation. In all estimations, we fail to reject the null hypothesis of the Sargan test of overidentifying restrictions, confirming the validity of the moment conditions. The autocorrelation functions for model residuals do not point to issues with error autocorrelation.

Overall, these results suggest that inflation in sticky-price sectors adjusts only weakly to current economic conditions and exhibits markedly higher persistence. By contrast, flexible inflation tends to revert more quickly toward the inflation target in the absence of shocks, consistent with Dedola et al. (2025). This implies that central banks should pay close attention to price developments in the sticky sectors, as distortions arising from nominal rigidities may exert a prolonged influence on inflation dynamics.

6 International comparison

6.1 Data and methodology

In this section, we provide a comprehensive comparison of our results with those obtained for euro area countries and the United States by Gautier et al. (2024) in a harmonized exercise. To ensure consistency and comparability, we use the same basket of goods and services defined at the COICOP5 level and apply the same item weights in the aggregation of results. We report both average and median frequencies and sizes of price changes, following the methodology of Gautier et al. (2024). We expand the analysis by reporting also implied price durations based on those frequencies.[25]

Since the dataset provided by Statistics Poland does not include flags for product replacements, our results are compared with the "including replacements" (substitution) statistics reported in Appendix Section B.2, Table A5, of Gautier et al. (2024), and not the baseline. Regarding product sales, for Poland the statistics labelled "excluding sales" are based solely on a sales filter, whereas for the euro area sales flags are available in the datasets for a fair number of countries in the sample.[26] Similarly, the data for the U.S. also include sales flags.

These issues highlight that full harmonization of the frameworks used in our study and that of Gautier et al. (2024) is infeasible for several reasons (for further discussion of cross-country discrepancies see the Appendix of Gautier et al., 2024). First, due to different data availability at the country level, datasets span slightly different periods, although almost all country

[25] Some of the price rigidity metrics are reported in Appendix C. To align with the methodology of Gautier et al. (2024), the median frequency is calculated as the weighted median across COICOP groups, while the median size is obtained as the weighted average of median sizes at the COICOP level. The reported size of price changes is based on log-price changes. Price durations are harmonized across countries and calculated using the aggregated frequency of price changes. This approach prevents outliers, which could arise when calculating durations for highly sticky prices (COICOP groups). We also make the usual assumption that prices can change more than once per month. The average time between price changes is then defined as in equation (A.25).

[26] In the euro area sample, only Greece, Luxembourg, Slovakia, and Spain do not provide sales flags. For these countries, a sales filter is used instead. Sales filters—even sophisticated ones such as that of Nakamura and Steinsson (2008), with clearance sale adjustments introduced by Gautier et al. (2024)—may underreport sales in sticky price groups and overreport them in highly flexible groups.

samples end in 2019.[27] Second, as indicated, treatment of sales, imputed prices, product replacements, and quality and quantity adjustments may slightly vary between countries. Third, data preprocessing procedures (such as the treatment of outliers) can also differ slightly from one country to another. Finally, frequency and size of price change are directly linked and dependent on the inflation level; therefore, these measures should not be interpreted in isolation when making international comparisons. Overall, while exact comparability of price rigidity statistics is inherently limited, we have made every effort to align our framework as closely as possible with that of Gautier et al. (2024).

The amount of information used for studying price rigidities varies substantially across countries used in this comparison. For Poland, the coverage of the consumption basket is higher than in any of the euro area countries or the United States, with 98.0% of the common basket represented, spanning 276 COICOP5 groups. For the common subsample of goods and services, the share of the basket covered stands at 58.2%, which is very similar to the euro area (58.9%), and the United States (58.5%). The number of COICOP5 groups in this common basket is also comparable across Poland (165), the euro area (166), and the United States (163).[28]

The share of sales for Poland amounts to 3.6%, very close to the euro area (3.8%) and notably below the United States (7.4%). The sales effect is much more pronounced in the U.S. than in other economies, which significantly impacts all the statistics, including the frequency and size changes as well as price duration.

6.2 Frequency and size of price changes

Results presented in Table 7 show that the overall frequency of price changes in Poland is 17.0%, higher than in the euro area (13.6%) but lower than in the United States (20.2%). This pattern also holds for the frequencies of price increases and decreases. The share of price increases in Poland (61.2%) is slightly below that of the euro area (63.8%) and the United States (62.0%). When sales are excluded, the frequency of price changes falls to 12.3% in Poland, 9.8% in the euro area, and 11.6% in the United States. Thus, the sales effect is

[27] For Poland, price rigidity statistics are calculated using the 2014–2019 sample, which is limited by the continuity of COICOP5 level data provided by Statistics Poland. The results for the U.S. are also taken from Gautier et al. (2024), based on the original estimates of Nakamura and Steinsson (2008).

[28] We report detailed information on data coverage by country in Appendix C, Table C4.

moderate in Poland and the euro area but much stronger in the United States, where it reduces the frequency of price changes by almost half.

Across sectors, the frequency of price changes is higher in Poland than in the euro area for most goods categories but slightly lower for services, while in the United States the highest figures are generally reported due to the strong influence of sales. For unprocessed food, the frequency of price changes in Poland (45.0%) exceeds that in the euro area (32.0%) and is close to the estimates for the United States (43.0%), reflecting the inherent volatility of this category. For processed food, prices in Poland (22.6%) are more flexible than in the euro area (16.1%) but marginally more sticky than in the United States. Again, these differences become smaller once sales are excluded. For non-energy industrial goods, the frequency of price changes in Poland (19.4%) is above the euro area (15.9%) but below the United States (23.5%), where it drops strongly when sales are excluded (8.9%), highlighting their predominance. Finally, prices of services are very sticky in Poland (6.5%), similarly as in the euro area (6.8%), and notably stickier than in the United States (10.0%). Overall, the services sector is most sticky across countries, and the sales effect there is negligible.

We also present counterfactual "equalised" figures that exclude the impact of different inflation levels[29] between Poland and the euro area. These approximate estimates serve as additional tool for comparison, particularly to assess whether the relationship between the frequency and size of price changes differs across economies, which may indicate distinct price adjustment mechanisms. When the effects of varying inflation levels are excluded, the frequency of price changes for processed food and non-energy industrial goods is still higher in Poland (Figure 9, Panel B). However, for unprocessed food, the frequency drops to the euro area level. Prices in Poland are relatively more flexible in sectors with a higher degree of tradability (openness). In services, typically considered as non-tradable, the frequency of price changes is lower in Poland than in the euro area.

The median size of price increases and decreases is very similar in Poland and the euro area (around 11% for increases and 13% for decreases, Table 8). However, the average figures reveal that the size of price increases in Poland is larger than in the euro area (15.0% vs

[29] Discrepancies in price rigidities across countries may arise from two factors: (i) differences in relation between the frequency and size of prices changes, and (ii) variations in the inflation level over a given period. Ultimately, frequency and size are related and jointly form inflation (via multiplication). Therefore, statistics on price rigidities should not be interpreted in isolation.

13.9%). This difference is driven primarily by the services sector, where the average size of price increases is roughly twice as large—16.5% compared to 8.5% in the euro area—while the size of price decreases remains similar. This pattern may be associated with the higher level of inflation in services in Poland. Once we control for the inflation level, the equalised size of price increases declines to around 8%, the same as in the euro area (Figure 9, Panel B). This observation may be linked to the Balassa-Samuelson effect—as the role of non-tradables (services) in economy increases, it tends to generate higher inflation in services, which materialises through the larger price increases rather than more frequent price adjustments. In contrast, the goods sectors in Poland exhibit an average size of price increases notably smaller than in both the euro area and the United States.

Taking into account both the frequency and size of price changes, firms in Poland tend to adjust prices of tradable goods more often but by smaller magnitudes than firms in the euro area (Figure 9). Several factors may explain this pricing behaviour. First, firms in Poland face greater volatility in inputs prices (e.g., due to exchange rate fluctuations). Second, the domestic market is more competitive, making it preferable to adjust prices in smaller steps but more frequently to maintain competitiveness (Vermeulen et al., 2012). Third, and related, consumers may be more sensitive to large price changes, which could trigger "consumer anger" during periods of economic turbulence (Stelmasiak et al., 2024). In contrast, the opposite pattern is observed in the services sector. The frequency of price changes, particularly of price increases, is slightly lower in Poland. This difference becomes even more pronounced after accounting for inflation differentials. It points to an inherent high level of stickiness of services prices, given which firms in Poland—compared to those in the euro area—prefer to change their prices rather via the intensive margin.

6.3 Price duration

At the aggregate level, the average price duration in Poland stands at 5.4 months, while the median duration is slightly higher at 6.3 months. Thus, price spells are somewhat shorter than in the euro area (6.8 and 8.5 months, respectively), but longer than in the United States (4.4 and 4.9 months). This indicates that prices in Poland change more often than in the euro area but less often than in the U.S. (Table 9 and Figure 10).

For the unprocessed food, prices change very frequently across all economies. The average implied price duration in Poland stands at 1.7 months—similar to the United States (1.8 months) but shorter than in euro area (2.6 months). In all cases, the short duration reflects a high degree of price flexibility, due to, *inter alia,* fast turnover of perishable products such as fresh fruits, vegetables, and meat. For processed food, prices in Poland remain fixed on average for 3.9 months, shorter than in the euro area (5.7 months) but slightly longer than in the United States (3.2 months), pointing to greater —partly due to product characteristics (longer shelf life) and cost structure.

Prices in the non-energy industrial goods sector in Poland are more flexible (4.6 months) than in the euro area (5.8 months), yet still more sticky than in the U.S. (3.7 months). The services sector stands out as the stickiest across all economies. In Poland, services prices remain unchanged for an average of nearly 15 months, similar to the euro area. In contrast, services prices in the United States while still very sticky, change roughly every 10 months. Median durations in services are notably higher than mean durations, reaching 28 months in Poland. This gap between the mean and median implied durations indicates that a large share of prices in services is extremely rigid.

Sales also shorten price durations in Poland, as the average implied duration decreases from 7.6 months (excluding sales) to 5.4 months (including sales). The impact of sales is therefore similar to that observed in the euro area but notably smaller than in the United States. However, drawing firm conclusions in this regard is somewhat limited, as Gautier et al. (2024) note that sales flags may be reported differently in the U.S. compared with Europe. Moreover, as previously mentioned, for Poland the figures “excluding sales” are based solely on a sales filter, while for most euro area countries official sales flags are available.

Overall, in Poland, prices of goods exhibit relatively high flexibility, while services display substantial stickiness—comparable to the euro area and much higher than in the U.S. This result is likely connected to relatively greater volatility in input costs for goods (including exchange rate fluctuations), whereas in case of services, wage contracts and menu cost play a key role in both Poland and the euro. In contrast, the U.S. labour market is more flexible, allowing for more frequent price adjustments. As noted earlier, higher inflation in services in Poland compared with the euro area stems from larger size of price increases rather than higher frequency. Consequently, services prices in Poland remain relatively more sticky despite the higher inflation level.

6.4 Post-COVID-19 period

Finally, we compare changes in price behaviour in Poland and the euro area during the post-COVID-19 period (2020–2024). A thorough comparison between the two economies—particularly in terms of levels—is limited by methodological differences relative to Gautier et al. (2026).[30] The results are, however, still based on a common basket and weights.[31]

Examining developments in the frequency of price changes after 2019, it is evident that the primary driver of the substantial increase was Russia's full-scale invasion of Ukraine in 2022 (Figure 11). In contrast, the onset of COVID-19 pandemic did not have any notable impact on the price formation process. The Russia full-scale invasion of Ukraine was not only a key factor behind the inflation surge in 2022, but its effects have also proven persistent, underscoring the deep and lasting influence of wartime conditions on all sectors of the economy in both Poland and the euro area. The subsequent recovery is reflected in both a decline in the frequency of price increases and a rise in the frequency of price decreases in 2023–2024. Despite many similarities across the two economies, the Polish economy experienced a more pronounced and widespread impact of the invasion, reflecting its geopolitical proximity and greater vulnerability to regional shocks.

To illustrate the extent to which Russia's full-scale invasion of Ukraine affected the price-setting process across economies, we examine a suitable measure of overall inflation pressure—the share of price increases. In Poland, this indicator rose from 63.2% in 2020 to 74.2% in 2022, before easing slightly to 68.6% in 2023 and further to 66.5% in 2024. In the euro area, this share followed a similar pattern: starting at 60.9% in 2020, peaking at 74.8% in 2022, and then declining to 69.3% in 2023 and 62.0% in 2024. The pace of normalisation following the Russia full-scale invasion of Ukraine is very similar across European economies in 2023 but appears to have slowed down in 2024 for Poland. However, this slowdown was to some extent driven by regulatory changes —specifically, the VAT rate adjustment for food products in April 2024 (a deferred reversal of earlier rate cuts). As noted, the results for euro area, taken from Gautier et al. (2026), exclude tax rate or regulatory changes beforehand, which results in smaller variation in the compared measures.

[30] Methodological differences include the treatment of replacements, which considerably affects the measured frequency of price changes (based on the pre-COVID-19 period, the bias exceeds 1 percentage point for the euro area), as well as handling imputations and tax rate changes in Gautier et al. (2026). In the euro area these effects are removed through data filtering procedures.

[31] Euro area weights are averaged over 2017–2024 for the post-COVID-19 comparison.

7 The implications of changing price rigidity

This section examines the implications of the observed price stickiness in Poland for the performance of monetary policy through the lens of a DSGE model. Full details of the model are provided in Appendix B. Below we summarize the key building blocks of the model.

7.1 The model

We employ a standard new Keynesian framework with a competitive final-goods sector and two monopolistically competitive intermediate-goods sectors that differ in their degree of price rigidity. Final output is produced using a constant elasticity of substitution (CES) aggregator that combines the two sectors. The intermediate sectors, referred to as the flexible and sticky sectors, follow Calvo price setting, where only a fraction of firms can adjust prices in each period. By construction, prices change more frequently in the flexible sector than in the sticky one.

Firms in each sector use a Cobb-Douglas technology that combines capital and labour inputs and are subject to technology shocks that follow a persistent autoregressive process. Total capital is fixed, while households decide how much labour to supply and how much to consume and save.

Aggregate labour supply is modelled as a CES composite of sectoral labour inputs, which allows for imperfect mobility of workers across sectors. The degree of labour mobility is governed by a single parameter: perfect mobility represents a frictionless case, while low mobility corresponds to limited substitution between sectors. This formulation ensures that labour frictions influence only the adjustment dynamics of the model, not its steady state.

Monetary policy is conducted according to a standard Taylor rule with interest rate smoothing. The central bank adjusts the policy rate in response to deviations of inflation and output from their steady-state levels.

This setup nests the conventional one-sector New Keynesian model as a special case in which both sectors exhibit the same degree of price stickiness. As it is standard in new Keynesian model, in our economy each period corresponds to a quarter.

7.2 One- and Two-Sectors Frameworks

In new Keynesian models, it is common practice to use the overall frequency of price changes. However, the data show substantial heterogeneity in the frequency of price adjustments, especially between services and the rest of the economy. We begin by examining the implications of this heterogeneity for monetary policy analysis.

To this end, we compare the behaviour of a one-sector economy with that of a two-sector economy. The two-sector setup provides an instructive environment for studying inflation and monetary policy effects because it allows us to trace the dynamics of relative prices in response to shocks such as monetary policy disturbances. Since movements in relative prices are a key channel through which inflation generates welfare costs, this framework is particularly useful for illustrating how inflation distorts relative prices and, in turn, alters the transmission of monetary policy. We consider two versions of the two-sector model: (i) with mobile labour across sectors, (ii) with quasi-immobile labour. The second setup allows us to assess additionally how the degree of labour mobility influences the transmission and persistence of monetary policy shocks.

We compute quarterly Calvo probabilities using median implied price durations rather than means. The median provides a robust measure of central tendency that better captures the pricing behaviour of the representative firm within each sector. In the context of Calvo pricing, where the probability of price adjustment is intended to summarize the behaviour of a typical price setter, the median implied duration offers a more economically meaningful calibration target. Using the median also aligns with the objective of distinguishing between flexible- and sticky-price sectors without allowing a small number of highly rigid prices to dominate the calibration. As shown in Table 5, the median implied duration equals 5.3 months in the flexible-price sector (denoted by f) and 25.5 months in the sticky-price sector (denoted by s), thereby capturing a clear and economically intuitive contrast in price rigidity across sectors.

The Calvo parameter in sector $j \in \{f, s\}$, denoted by θ_j, represents the probability that a firm does not change its price in a given period. It therefore captures the degree of nominal rigidity: higher values of θ_j indicate more persistent prices and less frequent adjustments. Using the formula $\theta_j = 1 - \frac{3}{D_j}$, where D_j denotes the median duration in months, we obtain quarterly

Calvo parameters: $\theta_f = 0.435$ for the flexible-price sector and $\theta_s = 0.882$ for the sticky-price sector. These sector-specific probabilities are then used as inputs in the model.

For the one sector version of the model, we set the Calvo parameter to $\theta = 0.636$, computed as a weighted average of the flexible and sticky sectors, with the weight of the flexible-price sector equal to 0.55, as described in Section 5.

Comparison of inflation and GDP responses to a monetary policy shock. We first compare the responses of the one-sector and two-sector models to a monetary policy shock under the assumption that all economies exhibit the same aggregate degree of price stickiness. In the one-sector economy, this corresponds to a uniform frequency of price adjustment across all firms, equal to the aggregate average. In contrast, in the two-sector economy, the frequency of price adjustment differs across sectors, while the aggregate level of price stickiness matches that of the one-sector model.

We analyse both the impulse response functions and the disinflation costs in terms of foregone output. To quantify the latter, we use a quasi-sacrifice ratio based on impulse response functions (QSR-IRF). This metric captures how costly a disinflation episode is in terms of cumulative output losses implied by the models IRFs. Intuitively, it measures how much cumulative output is lost for each unit of cumulative reduction in inflation. Formally, it is defined as:

$$QSR\text{-}IRF = \frac{\sum_{t=0}^{T} \min\{0, \Delta\tilde{y}_t\}}{\sum_{t=0}^{T} \min\{0, \Delta\pi_t\}}, \tag{12}$$

where $\Delta\tilde{y}_t$ and $\Delta\pi_t$ denote, respectively, the responses of GDP and inflation at horizon t, and T is the length of the horizon considered. We take only the negative parts because the metric is designed to capture the economic cost of disinflation, not benefits or rebound dynamics, ensuring consistency with the classical sacrifice ratio and preventing cancellation of losses by later gains.

A higher quasi-sacrifice ratio indicates a more costly disinflation in terms of cumulative GDP loss. Since the economy typically returns close to its steady state within this period, we set $T = 19$ quarters. To assess the dynamics, we analyse the responses of GDP and inflation to a one-percentage-point monetary policy shock, as shown in Figure 12. All impulse responses are reported as deviations from the steady state, and the numerical values discussed below correspond to these deviations.

In the one-sector model, the monetary policy shock generates a sharp disinflation and a relatively modest decline in output. Inflation drops steeply on impact by 1.45 pp and reaches its maximum decline of about 1.56 pp in the first period, then gradually returning toward zero with only a brief and small positive rebound. Output falls by 0.40 pp on impact and declines slightly further in the first period, after which the deviation from steady state fades quickly and is followed by a modest temporary overshooting.

In the two-sector model with perfectly mobile labour inflation declines by a similar magnitude but returns to its steady state more rapidly. On impact, inflation falls by 1.62 pp, after which it converges back to the steady state. Output reaction on impact is very similar to the one-sector case, as it is declining by 0.38 pp on impact and reaching a maximum decline of 0.42 pp in the subsequent period. However, output remains below its steady state for a longer period, reflecting the amplification of real effects under sectoral heterogeneity despite full labour mobility.

In the two-sector model with quasi-immobile labour, inflation exhibits the smallest initial decline by 1.17 pp and rebounds gradually over an extended period with almost no overshooting. Output falls by 0.45 pp on impact and reaches maximum decline of 0.57 pp, also remaining below its steady state for an extended period, with only a very small late positive deviation. Limited labour mobility prolongs real adjustment and produces the most persistent output decline among the three model specifications.

Taken together, these IRFs demonstrate how sectoral structure and labour mobility shape the transmission of monetary policy. The one-sector model yields a sharp disinflation alongside a short-lived output contraction. With mobile labour, the two-sector model smooths nominal adjustment but amplifies real deviations, resulting in a more persistent fall in output. When labour mobility is restricted, nominal responses become stronger and real responses more prolonged, as reallocation frictions hinder adjustment.

These differences are reflected in the quasi-sacrifice ratio given by equation (12), which rises from 0.27 in the one-sector model to 0.49 in the mobile-labour two-sector economy and further to 0.68 when labour is quasi-immobile. This pattern highlights that sectoral heterogeneity in price stickiness plays a central role in magnifying the output cost of lowering inflation. When sectors adjust prices at different speeds, monetary policy shocks generate misaligned nominal responses across sectors, slowing aggregate price adjustment and increasing the real cost of disinflation. Limited labour mobility strengthens this mechanism but is secondary to the underlying dispersion in price rigidities.

The comparison across models underscores that assuming uniform price rigidity can obscure important adjustment channels. In two-sector economies, differences in price stickiness generate distinct sectoral responses that shape aggregate inflation dynamics, and labour mobility influences both the strength and persistence of disinflation. The effects on GDP are even more marked: while the initial impact is similar across models, the persistence of real deviations is substantially higher in the two-sector settings, especially when labour mobility is limited. In such environments, sectoral misallocation and delayed relative price adjustment prolong the return of output to steady state. Overall, the one-sector framework tends to understate how long output remains away from steady state after a monetary policy shock, suggesting that it may mismeasure the welfare costs of business cycle in economies characterised by heterogeneous price stickiness. We examine the welfare implications of these differences next and discuss their underlying sources in more detail in Appendix B5.

Welfare Cost of Business Cycle. We examine how accounting for heterogeneity in price stickiness affects the welfare cost of business cycles. The welfare cost is measured as the consumption-equivalent loss, expressed in percent of permanent consumption. It represents how much consumption households would be willing to give up to live in a smooth steady-state economy without cyclical fluctuations. In other words, it quantifies the welfare gain from eliminating business cycle volatility.

We consider our economy that is disturbed by two shocks—technology and monetary—with the following properties. The technology shock follows an AR(1) process with persistence $\rho = 0.95$ and a standard deviation of innovations equal to 0.007, while the monetary policy shocks, are assumed to be independently and identically distributed (i.i.d.), with a standard deviation of 0.002.

The results, presented in Table 10, show that introducing sectoral heterogeneity in price stickiness modestly increases the welfare cost of business cycles. In the benchmark one-sector model, the welfare loss amounts to 0.135% of lifetime consumption. Allowing for two sectors with perfect labour mobility raises the cost slightly to 0.149%, reflecting the additional propagation channel through relative prices. When labour becomes quasi-immobile ($\lambda = 0.1$), the welfare cost rises further to 0.155%. This increase—around 15% relative to the one-sector benchmark—indicates that two-sector framework with limited labour mobility amplifies output fluctuations and delays recovery after shocks.

Overall, these findings suggest that ignoring sectoral asymmetries and labour mobility constraints leads to an underestimation of the welfare costs associated with business cycle volatility. Accounting for heterogeneity in price stickiness therefore provides a more accurate assessment of the welfare implications of cyclical fluctuations.

7.3 High- and Low-Price Stickiness Framework

We compare how the effectiveness of monetary policy changes with different degrees of price stickiness. Specifically, we contrast the behaviour of the economy when price rigidity corresponds to its standard level observed during the 2000–2019 period—characterized by relatively stable prices—with the more recent period following the COVID-19 pandemic and the full-scale Russian invasion of Ukraine, which was associated with substantially higher inflation and more frequent price adjustments.

Building on the methodology described in the previous section, we measure price stickiness using median implied price durations, as shown in Table 5, computed separately for the pre-2020 period (2000–2019) and the high-inflation period (2020–2024), which corresponds to the post-COVID-19 phase and the period of the full-scale war. We refer to the latter as the *recent simulation*. This distinction allows us to capture changes in pricing behaviour under markedly different macroeconomic conditions.

For the 2000–2019 sample, the median implied price duration equals 6.0 months in the flexible-price sector and 29.5 months in the sticky-price sector. These durations imply quarterly Calvo parameters of $\theta_f = 0.498$ and $\theta_s = 0.898$, respectively, indicating a clear segmentation between sectors, with prices in the sticky sector exhibiting substantially greater persistence.

For the recent simulation (2020–2024), median implied price durations decline markedly in both sectors, falling to 4.1 months in the flexible sector and 16.0 months in the sticky sector. Applying the same mapping from durations to Calvo probabilities yields $\theta_f = 0.265$ and $\theta_s = 0.813$. The reduction in implied durations and Calvo parameters reflects a pronounced increase in the frequency of price adjustment during the post-COVID and wartime period.
All simulations are conducted using the two-sector New Keynesian model with perfect labour mobility described above.

We next examine how monetary transmission changes when the degree of price stickiness differs across pricing environments. For this purpose, we analyse the responses of GDP and inflation to a one-percentage point increase in the policy interest rate, as shown in Figure 13. All impulse responses are reported as deviations from the steady state, and the numerical values discussed below correspond to these deviations.

In the baseline simulation, inflation declines sharply by 1.45 pp on impact and then gradually converges back toward zero. Output falls by 0.40 pp on impact and reaches its maximum decline of about 0.47 pp in the first period, after which it recovers steadily and turns positive after several quarters. Overall, this calibration implies a sizeable and persistent disinflation accompanied by a relatively large and prolonged output contraction.

In the recent simulation—corresponding to the high-inflation environment observed in the post-COVID and full-scale war period—inflation responds more forcefully but stabilises more quickly. Inflation falls by 2.15 pp on impact, representing the strongest immediate response across the two simulations, and returns rapidly toward zero. Output declines by 0.31 pp on impact and reaches a maximum decline of approximately 0.32 pp in the first period, before recovering more quickly than in the baseline simulation. The adjustment pattern is therefore characterised by a sharp disinflation together with a smaller and less persistent output decline.

These differences in adjustment dynamics are summarised by the quasi-sacrifice ratio defined in equation (12). In the baseline simulation, the quasi-sacrifice ratio equals 0.54, indicating a relatively high output cost of disinflation. In the recent simulation, the ratio declines to 0.30, implying that disinflation is achieved at a substantially lower output cost when prices adjust more frequently. Taken together, the comparison across the two simulations highlights that greater price flexibility strengthens the immediate response of inflation to monetary policy while attenuating the persistence and magnitude of the associated output decline.

The underlying mechanism is straightforward. When firms adjust prices more frequently, monetary policy shocks pass through more rapidly to inflation, and less of the adjustment occurs through quantities. Conversely, when price changes are infrequent, inflation responds sluggishly, and a larger share of the adjustment is borne by output, generating more protracted real effects.

These mechanisms can also be viewed through the lens of the Phillips curve. In the baseline simulation, where price stickiness is higher, the Phillips curve is relatively flat, indicating that inflation responds weakly to movements in economic activity. As a result, reducing inflation requires larger and more persistent output losses. In the recent simulation, where prices adjust more often, the Phillips curve becomes steeper, with inflation reacting more strongly to changes in marginal costs and demand conditions. Importantly, the increase in the frequency of price changes observed during the recent high-inflation period may itself be interpreted as evidence of such a steepening. Consequently, this development may have increased the ability of central banks to bring down inflation at a lower real cost.

8 Conclusions

This paper provides a comprehensive assessment of price rigidity in Poland using a unique, large-scale micro dataset of CPI prices. By documenting the frequency, size and duration of price changes across goods and services and by harmonising our methodology with Gautier et al. (2024, 2026), we add new evidence for an open economy outside the euro area that has undergone rapid structural transformation and recently experienced sizable shocks. Our analysis covers both the relatively stable pre-COVID-19 period and the subsequent years marked by the pandemic and Russia's full-scale invasion of Ukraine, offering a detailed view of how price-setting behaviour responds to large and overlapping disturbances.

Our empirical results show that overall price rigidity in Poland is substantial. On average, around one-fifth of prices change each month, implying typical price durations of about four to six months, depending on whether one considers mean or median statistics. Behind these aggregates lies pronounced heterogeneity across sectors: prices of tradable goods—especially energy and food—are highly flexible and adjust frequently, whereas prices in services are very sticky, often remaining unchanged for more than a year and, in many cases, for several years. The large gap between mean and median durations in services indicates that a significant share of prices in this sector is extremely rigid. Promotions play a non-trivial role in shaping observed price patterns, accounting for a sizable fraction of price changes and materially affecting measured frequencies, durations and the distribution of price adjustments.

The post-2020 period reveals a clear regime change in the behaviour of prices. While the COVID-19 pandemic itself had only a limited impact on the price formation process, Russia's full-scale invasion of Ukraine triggered a sharp and persistent increase in the frequency of price changes in both Poland and the euro area, driven mainly by more frequent price increases. In Poland, this shift was particularly pronounced, reflecting its greater exposure to regional shocks. Although the frequency of price changes has declined from its 2022 peak, it remains elevated relative to the pre-pandemic period, suggesting that the recent high-inflation episode was associated with a higher degree of effective price flexibility.

In line with the international evidence, we find that inflation dynamics in Poland are shaped primarily by movements in the extensive margin of price adjustment—that is, by changes in the frequency of price increases and decreases—rather than by systematic changes in their size. Counterfactual decompositions show that recomposed inflation is much better tracked by measures that allow the frequency of price changes to vary than by those that restrict adjustment to the intensive margin. Quantile regression results confirm that this finding holds across the entire distribution of monthly inflation and becomes even more pronounced in high-inflation regimes. Thus, when inflation rises, firms respond mainly by adjusting prices more often rather than by making proportionally larger changes each time.

By separating the CPI basket into flexible and sticky components, we show that these two segments of the economy exhibit markedly different behaviour. Flexible prices respond quickly and strongly to changes in the output gap, the exchange rate, energy prices and global supply chain pressures, and display relatively low persistence. In contrast, inflation in sticky-price sectors is highly persistent, only weakly related to contemporaneous macroeconomic conditions and heavily influenced by backward- and forward-looking components. These findings underscore the importance of developments in sticky sectors—particularly services—since nominal rigidities there are a key source of inflation persistence and can keep inflation away from the target for prolonged periods.

Our international comparison reveals that Poland's price-setting behaviour is closer to that of the euro area than to that of the United States, especially in terms of the degree of stickiness in services. At the aggregate level, prices in Poland change more frequently than in the euro area but less often than in the U.S. For tradable goods, firms in Poland tend to adjust prices more often but by smaller amounts, consistent with stronger competitive pressures, greater input cost volatility and higher consumer sensitivity to large price changes. In services, by contrast, prices are as sticky as in the euro area and considerably more rigid than in the United States, with higher inflation in Poland being driven mainly by larger, not more frequent, price increases. These cross-country differences highlight the role of institutional features, wage-setting practices and market structure in shaping price rigidity.

Building on these empirical findings, we study the macroeconomic implications of price stickiness using one- and two-sector DSGE models. The two-sector framework, calibrated to the observed degree of flexibility in goods and rigidity in services, shows that sectoral heterogeneity in price stickiness materially affects the transmission of monetary policy. Relative price distortions arising from different adjustment speeds across sectors lead to smoother nominal but more persistent real responses to monetary shocks. As a result, the quasi-sacrifice ratio—the output loss required to achieve a given disinflation—increases substantially in the two-sector model relative to a standard one-sector model, especially when labour mobility across sectors is limited. Similarly, we show that when the overall degree of price rigidity declines—as observed in the recent high-inflation period—monetary policy can reduce inflation more quickly and at a lower output cost, consistent with a steeper Phillips curve.

Overall, our results carry several important policy implications. First, sectoral heterogeneity in price rigidity is not a second-order detail but a central determinant of both inflation persistence and the real cost of disinflation. Models that abstract from this heterogeneity are likely to understate the duration and magnitude of output losses following monetary policy tightening. Second, the recent increase in price flexibility implies that, under certain conditions, disinflation can be achieved at a relatively lower real cost than in the past, provided that expectations remain anchored and monetary policy is credible.

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Tables and Figures

Table 1 Price stickiness metrics

			Frequency of price change	Frequency of price increase	Frequency of price decrease	Size of price increase	Size of price decrease	Average duration	Share of sales	Price synch.
CPI	Incl. sales	mean	22.1	12.8	9.3	13.3	11.3	4.0	3.4	0.27
		med.	14.5	7.7	4.3	6.7	7.5	6.4	1.7	0.23
	Excl. sales	mean	18.0	10.7	7.3	12.5	11.0	5.0	.	0.26
		med.	11.1	5.9	3.2	6.1	6.9	8.5	.	0.21
By sector										
Core inflation	Incl. sales	mean	14.3	8.6	5.6	15.2	12.6	6.5	2.5	0.26
		med.	8.9	5.3	2.4	7.7	8.3	10.7	1.0	0.23
	Excl. sales	mean	11.5	7.2	4.3	14.7	12.5	8.2	.	0.24
		med.	7.5	4.4	1.9	7.1	8.0	12.9	.	0.20
Food and non-alcoholic beverages	Incl. sales	mean	30.5	17.3	13.2	12.0	11.1	2.7	6.5	0.18
		med.	25.1	13.0	9.4	6.7	7.6	3.5	5.7	0.20
	Excl. sales	mean	22.5	13.2	9.2	10.6	10.3	3.9	.	0.18
		med.	17.1	9.1	5.8	5.6	6.7	5.3	.	0.21
Energy	Incl. sales	mean	36.5	20.7	15.9	7.2	6.2	2.2	1.7	0.46
		med.	12.6	3.7	1.2	3.4	3.5	7.5	0.0	0.40
	Excl. sales	mean	33.6	19.1	14.5	7.2	6.2	2.4	.	0.47
		med.	10.0	3.1	1.1	3.4	3.4	9.5	.	0.35
Non-energy industrial goods	Incl. sales	mean	19.3	11.1	8.2	12.7	10.9	4.7	4.0	0.22
		med.	15.3	8.5	5.3	6.2	7.1	6.0	2.4	0.18
	Excl. sales	mean	14.7	8.7	5.9	11.7	10.5	6.3	.	0.19
		med.	11.5	6.4	3.9	5.7	6.8	8.2	.	0.14
Services	Incl. sales	mean	8.8	6.0	2.8	18.4	14.9	10.9	0.8	0.31
		med.	3.0	2.2	0.5	11.1	10.7	32.8	0.0	0.33
	Excl. sales	mean	8.0	5.5	2.5	18.2	15.0	12.0	.	0.30
		med.	2.8	2.0	0.3	11.1	11.1	35.7	.	0.28

Note: The table presents mean and median estimates of price stickiness for all goods and services as well as at the sectoral level for the full sample (i.e., 2000–2024). Share of sales and the results excluding sales are based on the sales filter proposed by Gautier et al. (2024). Price synchronization is calculated as in Fisher and Konieczny (2000).

Table 2 Mean estimates for price stickiness metrics for selected years

	Frequency of price			Size of price		Average
	change	increases	decreases	increase	decrease	duration
2000–2019	20.3 (16.6)	11.6 (9.7)	8.7 (6.9)	13.0 (12.4)	11.3 (11.0)	4.4 (5.5)
2020	26.7 (21.0)	15.2 (12.2)	11.5 (8.8)	14.6 (13.4)	11.4 (10.7)	3.2 (4.3)
2021	27.4 (21.8)	18.3 (15.4)	9.1 (6.4)	13.9 (12.5)	11.3 (10.5)	3.1 (4.1)
2022	33.3 (28.0)	21.1 (18.6)	12.2 (9.5)	14.5 (13.7)	11.8 (11.5)	2.5 (3.0)
2023	30.2 (24.2)	16.7 (13.6)	13.5 (10.5)	14.3 (13.2)	11.8 (11.3)	2.8 (3.6)
2024	30.1 (22.7)	17.3 (14.0)	12.8 (8.7)	14.9 (13.4)	12.1 (11.3)	2.8 (3.9)

Note: The table presents selected mean price stickiness measures prior to the COVID-19 pandemic (2000–2019) as well as following the outbreak of the COVID-19 pandemic in 2020 and the outburst of the Russian invasion of Ukraine in 2022. In parentheses we report measures calculated on prices following the exclusion of sales with the use of the sales filter by Gautier et al. (2024).

Table 3 Median estimates for price stickiness metrics for selected years

	Frequency of price			Size of price		Average
	change	increases	decreases	increase	decrease	duration
2000–2019	13.0 (10.1)	7.1 (5.6)	4.2 (3.2)	6.6 (5.9)	7.4 (6.9)	7.2 (9.4)
2020	20.7 (14.1)	10.1 (7.6)	6.2 (3.9)	7.2 (6.1)	7.7 (6.7)	4.3 (6.6)
2021	22.0 (14.9)	12.9 (9.0)	5.3 (3.4)	7.5 (6.4)	8.2 (6.9)	4.0 (6.2)
2022	27.9 (20.6)	16.2 (13.6)	5.5 (3.5)	8.7 (8.2)	8.1 (7.6)	3.1 (4.3)
2023	24.5 (17.1)	12.7 (9.9)	7.0 (4.2)	8.2 (7.2)	8.1 (7.6)	3.6 (5.3)
2024	25.8 (14.0)	12.4 (8.3)	7.7 (3.8)	7.8 (6.7)	8.5 (7.3)	3.3 (6.6)

Note: The table presents selected median price stickiness measures prior to the COVID-19 pandemic (2000–2019) as well as following the outbreak of the COVID-19 pandemic in 2020 and the outburst of the Russian invasion of Ukraine in 2022. In parentheses we report measures calculated on prices following the exclusion of sales with the use of the sales filter by Gautier et al. (2024).

Table 4 Distribution of (non-zero) price changes

	Percentiles (in percent)							Percent of
	5th	10th	25th	50th	75th	90th	95th	|dp|≤ 3.5%
Including sales								
CPI	-30.6	-19.9	-6.3	2.0	8.7	20.0	29.1	28.3
Core inflation	-28.8	-18.3	-5.5	1.9	8.6	19.2	28.8	31.0
Food and non-alcoholic beverages	-33.8	-22.4	-8.1	2.2	9.5	21.9	30.9	24.6
Energy	-9.0	-5.8	-2.1	1.0	4.0	9.0	14.7	34.2
Non-energy industrial goods	-29.2	-19.0	-5.9	1.5	7.8	18.2	28.7	31.5
Services	-22.5	-13.4	-1.2	5.7	13.3	24.1	35.5	27.4
Excluding sales								
CPI	-29.0	-18.3	-5.2	2.2	8.3	18.2	28.7	34.1
Core inflation	-27.5	-16.6	-4.6	2.2	8.3	18.2	28.6	36.1
Food and non-alcoholic beverages	-33.7	-21.2	-6.7	2.5	8.7	19.6	29.5	31.5
Energy	-9.1	-5.9	-2.0	1.0	4.1	9.3	15.3	35.0
Non-energy industrial goods	-28.4	-17.2	-5.2	1.7	7.4	16.2	26.2	37.4
Services	-22.4	-12.8	-0.8	5.8	13.5	24.5	35.6	28.2

Note: The table provides selected percentiles of the size of (non-zero) price changes (in percent). Results excluding sales are based on the Gautier et al. (2024) sales filter. The last column denotes the fraction of prices that change by no more than 3.5% in absolute terms.

Table 5 Price stickiness statistics for baskets of flexible and sticky prices

		Frequency of price			Size of price		Implied
		change	increase	decrease	increase	decrease	duration
2000-2024	all	22.1 (11.1)	12.8 (5.9)	9.3 (3.2)	13.3 (6.1)	11.3 (6.9)	4.0 (8.5)
	flexible	33.3 (17.2)	18.6 (9.1)	14.7 (5.4)	10.8 (5.4)	9.8 (6.3)	2.5 (5.3)
	sticky	8.6 (3.8)	5.9 (2.5)	2.8 (0.5)	16.4 (8.9)	13.4 (9.3)	11.1 (25.5)
2000-2019	all	20.3 (10.1)	11.6 (5.3)	8.7 (3.0)	13.0 (6.0)	11.3 (6.9)	4.4 (9.4)
	flexible	31.0 (15,4)	17.2 (8.1)	13.8 (5.3)	10.2 (5.1)	9.4 (6.2)	2.7 (6.0)
	sticky	7.7 (3.3)	5.1 (2.2)	2.6 (0.5)	16.2 (8.4)	13.5 (9.5)	12.5 (29.5)
2020-2024	all	29.6 (15.5)	17.7 (8.7)	11.8 (3.6)	14.5 (7.1)	11.7 (7.3)	2.9 (6.0)
	flexible	42.7 (21.7)	24.5 (12.1)	18.2 (5.8)	13.4 (6.1)	11.3 (6.8)	1.8 (4.1)
	sticky	12.4 (6.1)	8.8 (4.3)	3.6 (0.5)	17.2 (9.5)	13.3 (9.0)	7.6 (16.0)

Note: The table presents selected mean price rigidity measures for baskets of flexible and sticky prices including sales. Mean estimates for given periods are reported. In parentheses we report medians calculated on prices following the exclusion of sales with the use of the sales filter by Gautier et al. (2024).

Table 6 Stylized Phillips curve estimation results

	Headline inflation	Flexible inflation	Sticky inflation
constant	-0.0894	-0.2235	-0.0922
	[-0.4249,0.2461]	[-0.6703,0.2234]	[-0.2954,0.1110]
first lag	0.4429***	0.3447***	0.2779***
	[0.2999,0.5860]	[0.1957,0.4936]	[0.1458,0.4100]
second lag	0.1484*	0.1664**	0.3818***
	[-0.0029,0.2998]	[0.0354,0.2974]	[0.1861,0.5776]
output gap	0.0686***	0.0961***	0.0195*
	[0.0326,0.1046]	[0.0417,0.1505]	[-0.0027,0.0416]
nominal effective exchange rate	-0.1128***	-0.1614***	-0.0207**
	[-0.1481,-0.0775]	[-0.2181,-0.1048]	[-0.0386,-0.0028]
energy prices	0.0323***	0.0470***	0.0053*
	[0.0250,0.0397]	[0.0359,0.0582]	[-0.0005,0.0111]
supply chain disruptions	0.0209*	0.0387**	0.0080
	[-0.0022,0.0439]	[0.0083,0.0691]	[-0.0048,0.0208]
forward looking component	0.1010	-0.0303	0.2017**
	[0.2461,0.3476]	[0.2234,0.2595]	[0.1110,0.3799]
sample size	93	93	93
persistence	0.59	0.51	0.66
overidentifying restrictions	15.9 (0.143)	17.2 (0.103)	15.0 (0.184)

Note: Estimation results for the model specified by equation (11). 95% confidence intervals reported in square brackets. ***, ** and * denote significance at 1%, 5% and 10%, respectively. For the Sargan test of overidentifying restrictions p-values are reported in parentheses.

Table 7 Average frequency of price changes (%)

		Including sales				Excluding sales			
		Frequency of price changes			Share of increases	Frequency of price changes			Share of increases
	Sector	Change	Increase	Decrease		Change	Increase	Decrease	
PL	Overall	17.0	9.2	7.7	61.2	12.3	6.9	5.4	62.5
	Unprocessed food	45.0	24.0	21.1	53.8	31.9	17.2	14.7	54.9
	Processed food	22.6	12.6	10.1	56.6	14.6	8.6	6.1	59.6
	Non-energy industrial goods	19.4	9.7	9.6	51.1	14.3	7.2	7.1	51.6
	Services	6.5	4.2	2.3	73.2	5.7	3.8	1.9	74.2
EA	Overall	13.6	7.6	6.1	63.8	9.8	5.9	3.9	67.3
	Unprocessed food	32.0	17.0	15.0	54.5	24.6	13.3	11.2	57.2
	Processed food	16.1	8.9	7.2	56.6	10.9	6.4	4.5	60.7
	Non-energy industrial goods	15.9	7.7	8.1	51.6	9.4	5.2	4.1	58.4
	Services	6.8	4.8	2.0	79.3	6.6	4.7	1.9	80.0
US	Overall	20.2	10.9	9.3	62.0	11.6	7.6	4.0	71.1
	Unprocessed food	43.0	22.8	20.2	53.1	29.7	17.1	12.6	58.4
	Processed food	26.6	14.7	11.9	55.3	12.7	8.3	4.4	66.3
	Non-energy industrial goods	23.5	10.3	13.2	46.9	8.9	5.7	3.2	66.0
	Services	10.0	7.1	2.8	78.9	9.6	7.0	2.6	80.1

Note: Data for euro area and U.S. are sourced from Gautier et al. (2024). Excluding sales figures estimated using sales filter for: Poland, Greece, Luxembourg, Slovakia and Spain. Frequencies excluding sales are corrected for Slovakia.

Table 8 Size of price changes (%)

		Average				Median			
		Including sales		Excluding sales		Including sales		Excluding sales	
	Sector	Increase	Decrease	Increase	Decrease	Increase	Decrease	Increase	Decrease
PL	Overall	15.0	17.4	14.0	16.9	10.8	13.6	9.8	12.9
	Unprocessed food	15.3	16.5	13.8	15.1	11.7	12.8	10.6	11.7
	Processed food	11.4	13.5	9.3	11.8	7.6	9.7	6.0	7.7
	Non-energy industrial goods	15.6	18.2	14.4	17.7	9.6	13.6	8.2	12.6
	Services	16.5	18.9	16.4	19.2	13.2	15.7	13.1	16.0
EA	Overall	13.9	19.2	10.5	15.8	10.4	13.4	7.5	9.5
	Unprocessed food	17.4	19.2	13.4	14.4	13.0	15.2	10.4	11.3
	Processed food	13.1	15.5	8.7	10.3	9.6	12.5	6.1	7.0
	Non-energy industrial goods	20.1	22.7	13.6	16.2	15.8	19.3	9.8	11.9
	Services	8.5	18.0	8.3	18.3	5.9	8.8	5.8	8.5
US	Overall	17.8	21.6	10.6	13.4	.	.	.	.
	Unprocessed food	27.5	30.0	18.9	20.6	.	.	.	.
	Processed food	24.4	28.1	11.5	15.8	.	.	.	.
	Non-energy industrial goods	21.6	26.4	9.8	12.1	.	.	.	.
	Services	9.5	12.8	9.1	11.7	.	.	.	.

Note: Data for euro area and U.S. are sourced from Gautier et al. (2024). Excluding sales figures estimated using sales filter for: Poland, Greece, Luxembourg, Slovakia and Spain.

Table 9 Implied price duration (in months)

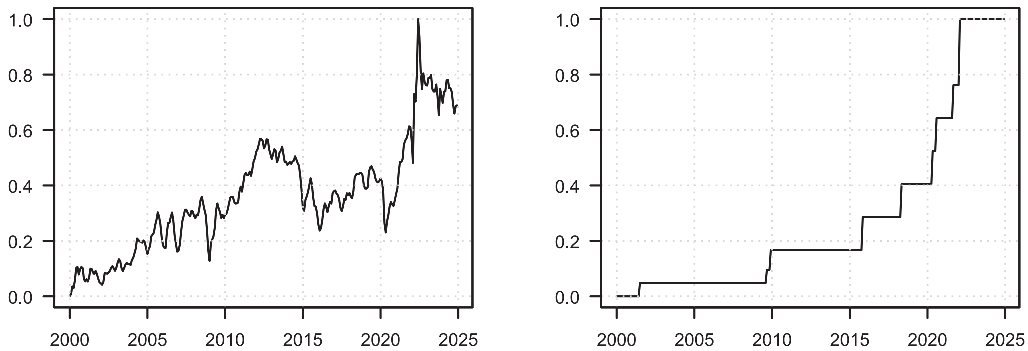
		Average		Median	
	Sector	Including sales	Excluding sales	Including sales	Excluding sales
PL	Overall	5.4	7.6	6.3	8.7
	Unprocessed food	1.7	2.6	1.7	2.8
	Processed food	3.9	6.3	3.9	6.3
	Non-energy industrial goods	4.6	6.5	5.1	7.2
	Services	14.9	17.0	28.3	30.1
EA	Overall	6.8	9.7	8.5	13.4
	Unprocessed food	2.6	3.5	2.8	4.9
	Processed food	5.7	8.6	5.9	10.1
	Non-energy industrial goods	5.8	10.2	6.1	11.6
	Services	14.2	14.7	24.6	25.3
US	Overall	4.4	8.1	4.9	11.2
	Unprocessed food	1.8	2.8	1.6	2.9
	Processed food	3.2	7.4	3.3	8.3
	Non-energy industrial goods	3.7	10.7	3.8	12.7
	Services	9.5	9.9	16.3	16.9
AT	Overall	6.7	9.4	8.9	12.7
BE		6.4	7.0	14.6	14.6
ES		6.2	9.1	6.1	10.9
FR		6.2	7.7	6.1	9.9
DE		6.9	9.3	10.0	17.0
GR		8.3	13.1	10.3	15.2
IT		8.2	15.3	10.5	16.7
LV		4.2	9.0	6.6	16.3
LT		6.1	7.8	7.0	8.9
LU		6.5	10.7	8.4	15.4
SK		6.5	10.3	9.6	12.2

Note: Implied price durations for euro area and U.S. based on the frequencies reported by Gautier et al. (2024). Excluding sales figures estimated using sales filter for: Poland, Greece, Luxembourg, Slovakia and Spain. Median durations are computed based on median frequencies.

Table 10 Welfare cost of business cycles across model specifications

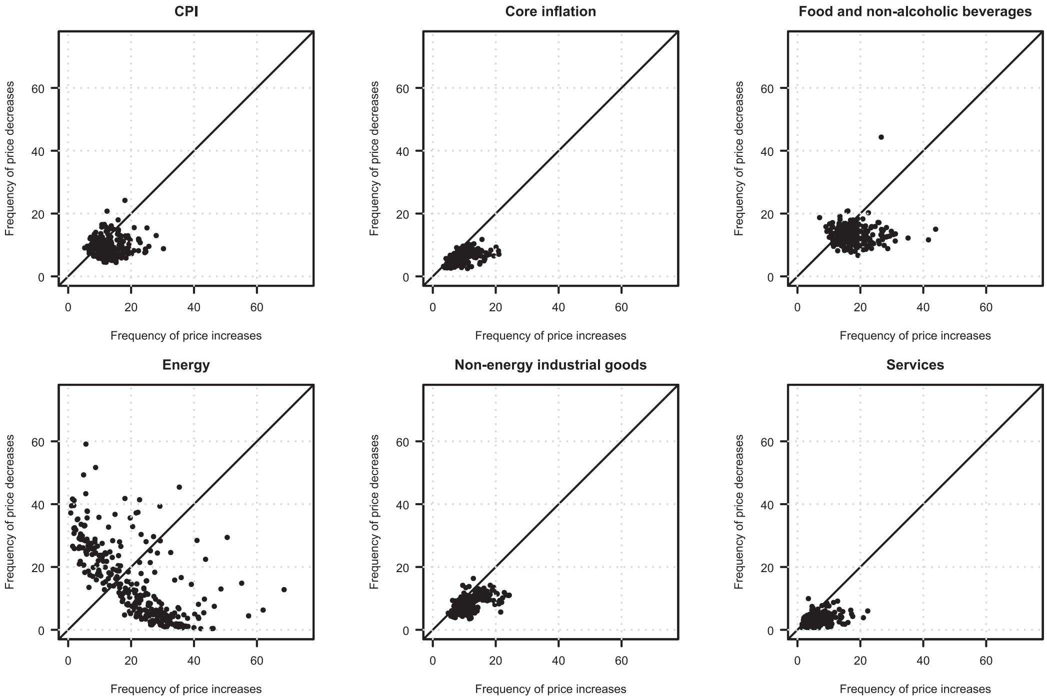
Model Specification	Welfare Cost (in % of Lifetime Consumption)
1-sector model	0.135
2-sector model (perfect labor mobility, $\lambda = \infty$)	0.149
2-sector model (quasi-immobile labor, $\lambda = 0.1$)	0.155

Figure 1 Prices of fuels (left panel) and men's haircut (right panel) at a random point of sale



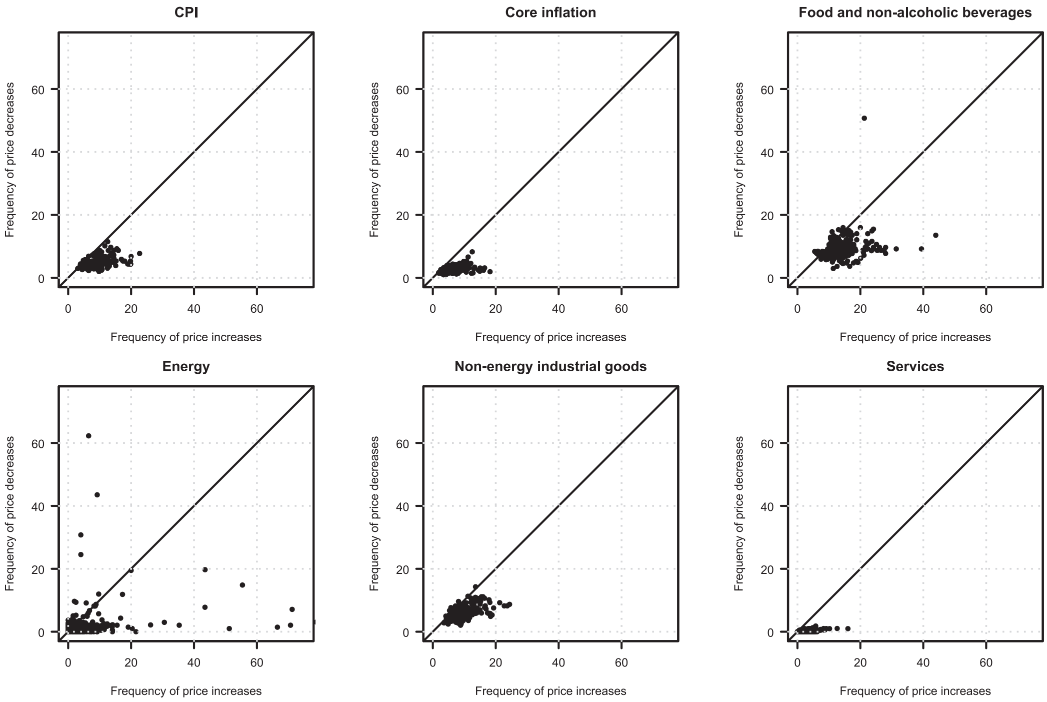
Note: The figure shows prices of two products with a different degree of stickiness at a random point of sale. Fuel price is flexible and changes frequently, while the price for men's haircut is sticky (rigid) and is adjusted only occasionally. Prices plotted on the figure have been normalized.

Figure 2 Price rigidity metrics at the product level (in percent)



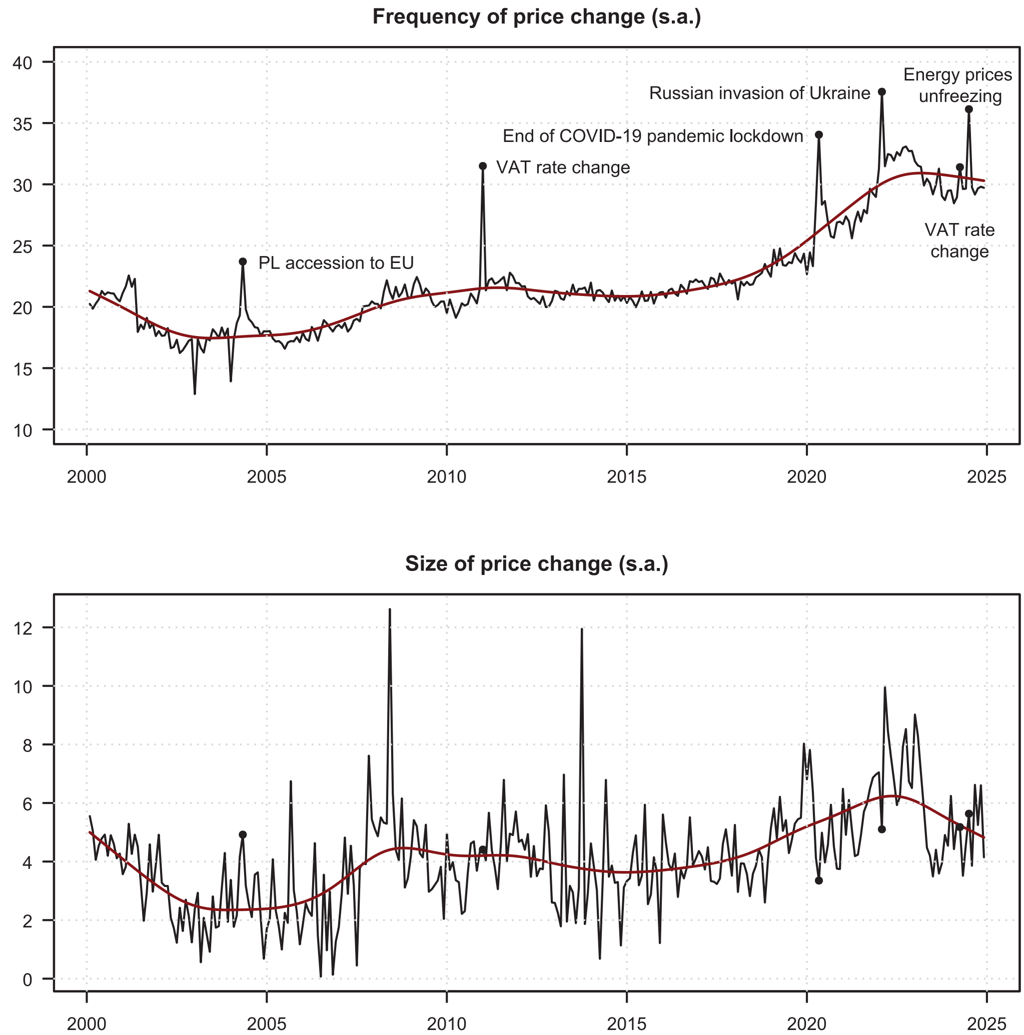
Note: The figure shows the frequency of price increases and decreases at the product level. Results are reported for mean estimates for prices including sales.

Figure 3 Price rigidity metrics at the product level (in percent)



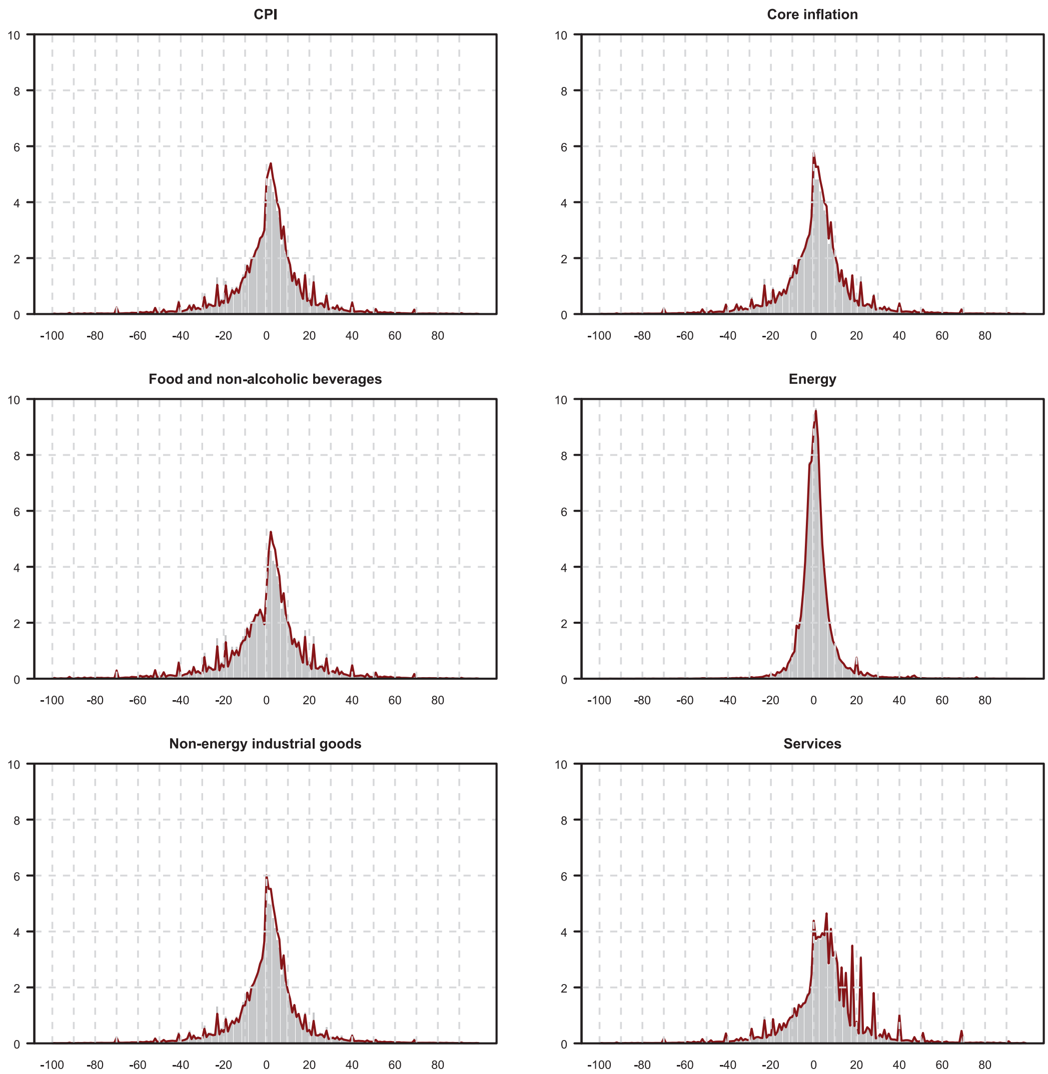
Note: The figure shows the frequency of price increases and decreases at the product level. Results are reported for median estimates for prices including sales.

Figure 4 Seasonally adjusted frequency and size of price changes



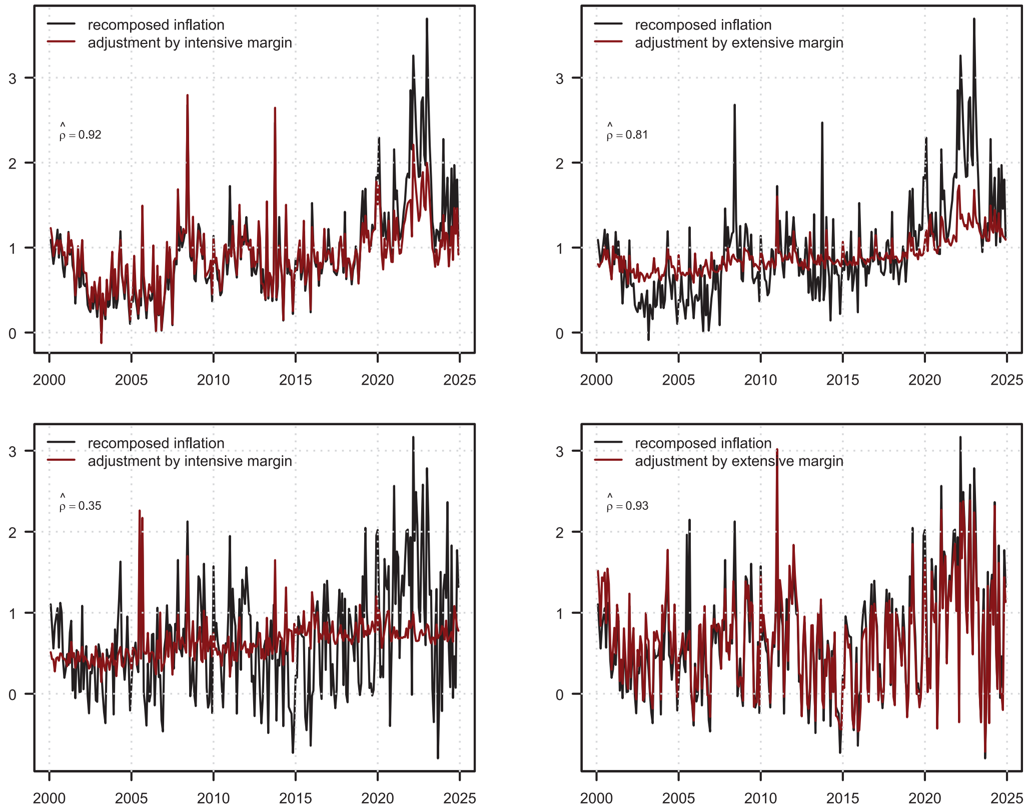
Note: The upper panel of the figure shows the frequency of price changes after adjusting for seasonal patterns along with the trend estimate using the Hodrick-Prescott filter. With dots we denote major events shaping spikes in the frequency of price changes. The lower panel illustrates the seasonally adjusted size of price changes, with dots denoting values corresponding to the events marked in the upper panel.

Figure 5 Distribution of price changes



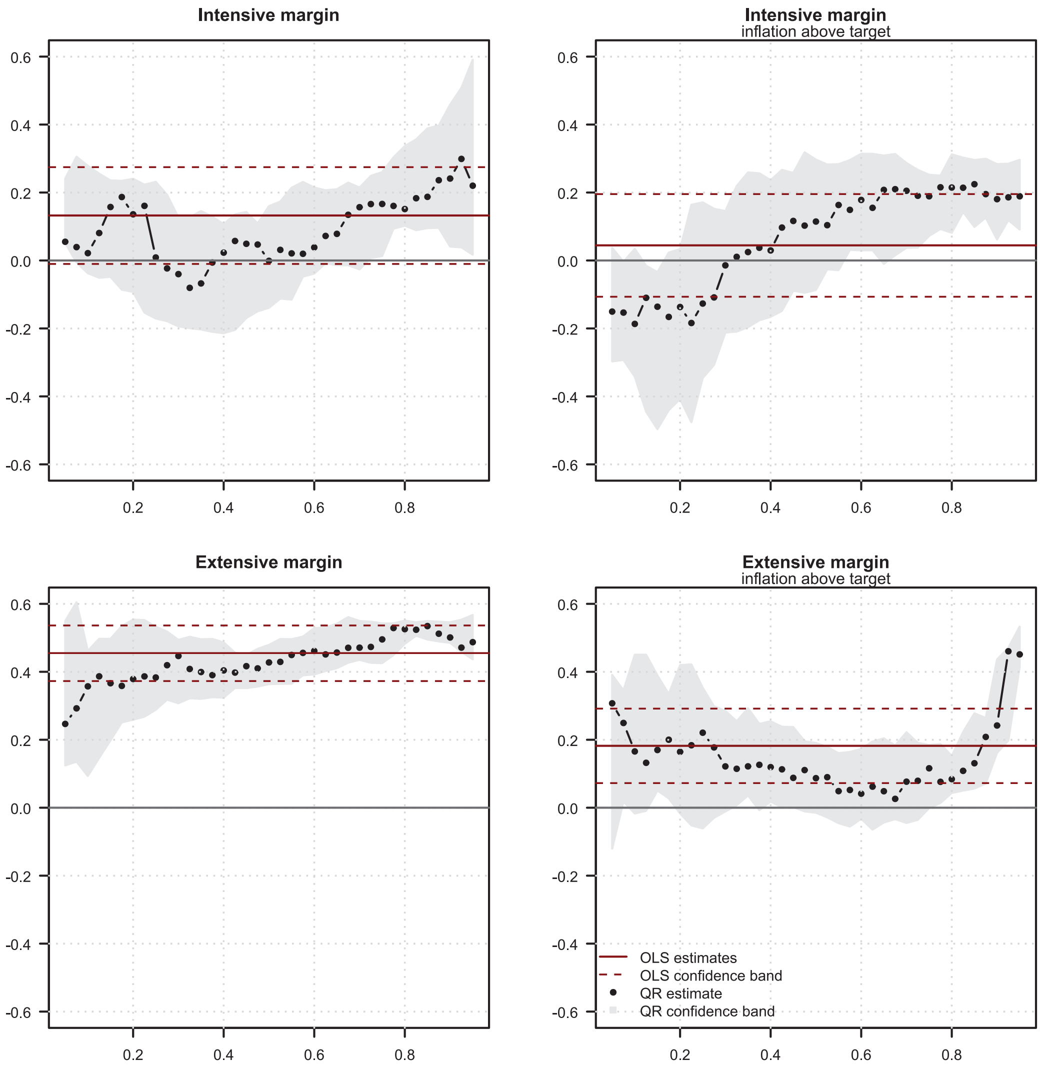
Note: The figure illustrates the distribution of price changes for all goods and services belonging to the CPI basket (upper left panel) as well as at the sectoral level (further panels). Grey histograms plot the distribution for prices including sales, whereas the red line corresponds to the distribution of price changes following the exclusion of sales with the Gautier et al. (2024) sales filter.

Figure 6 Recomposed and counterfactual inflation (in percent)



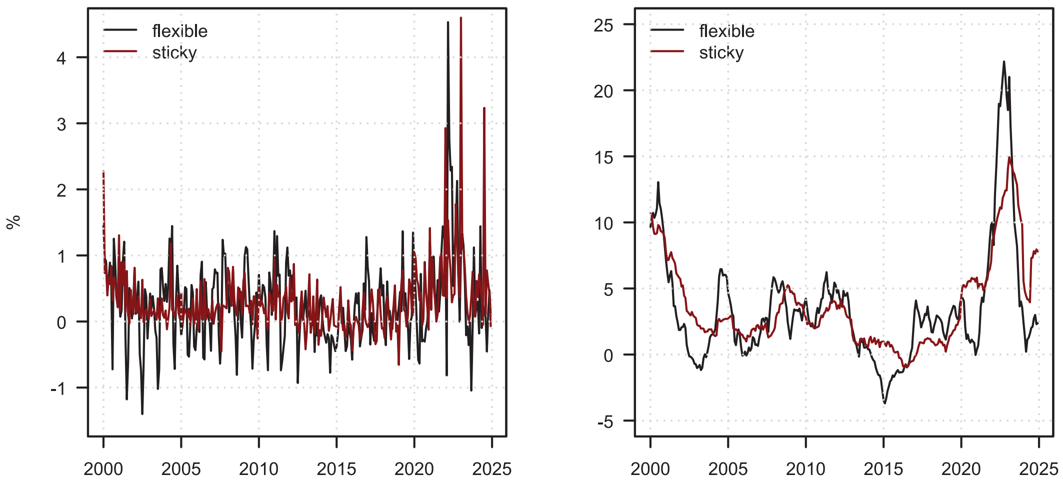
Note: In the upper panel of the figure shows how counterfactual measures of inflation calculated under the assumption of constant frequency of price changes (as in equation (3), red line) and constant size of price changes (as in equation (4), red line) develop against the recomposed monthly inflation calculated as in equation (1). The lower panel illustrates results from a similar exercise with the stipulation that both frequency and size of price increases and decreases are taken separately (as in equations (5) and (6)). Their development is contrasted then to recomposed inflation calculated as in equation (2). Pearson's correlation coefficient between the plotted series is denoted by $\hat{\rho}$.

Figure 7 Quantile regression results



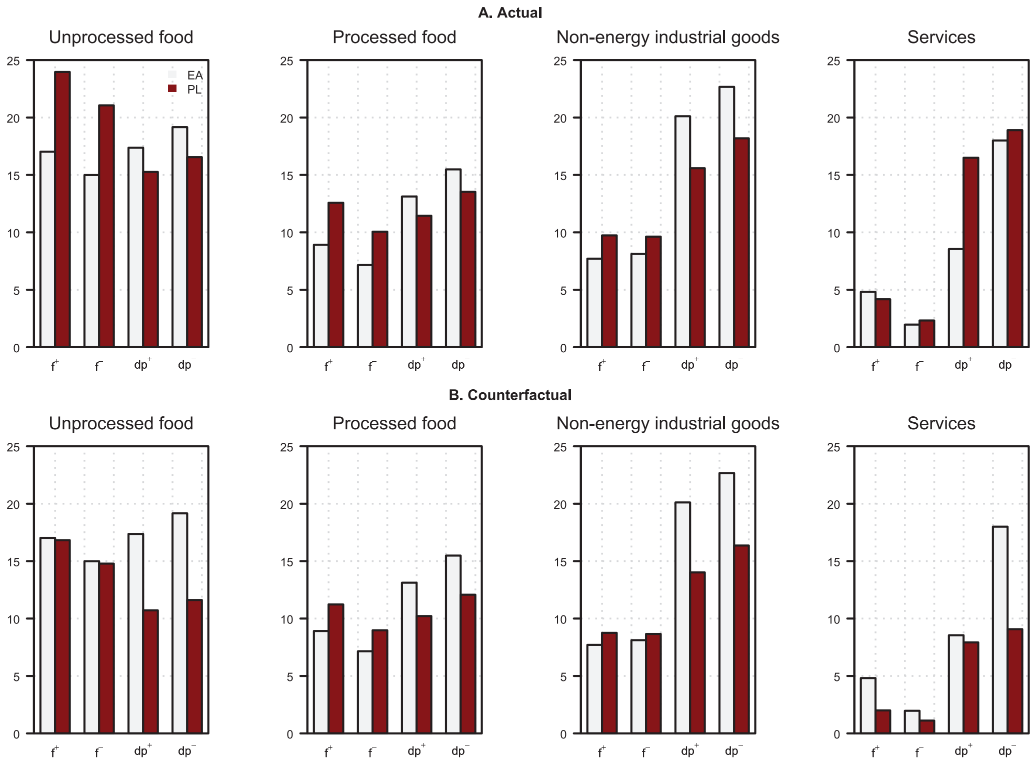
Note: The figure shows the estimation results of the model specified by equation (9). The red solid and dashed lines correspond to the OLS estimates and confidence bands, respectively. The black dots and shaded area denote quantile regression estimates and confidence bands, respectively.

Figure 8 The development in flexible and sticky inflation



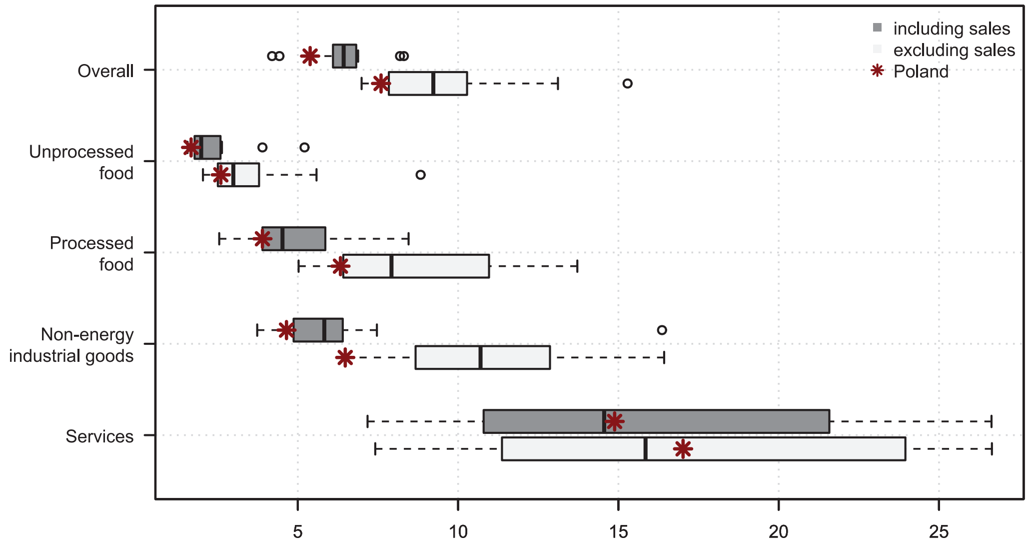
Note: The figure illustrates the development in monthly (left panel) and annual (right panel) rates of change in price indices calculated on baskets of flexible and sticky prices.

Figure 9 Actual and counterfactual frequency and size of price changes



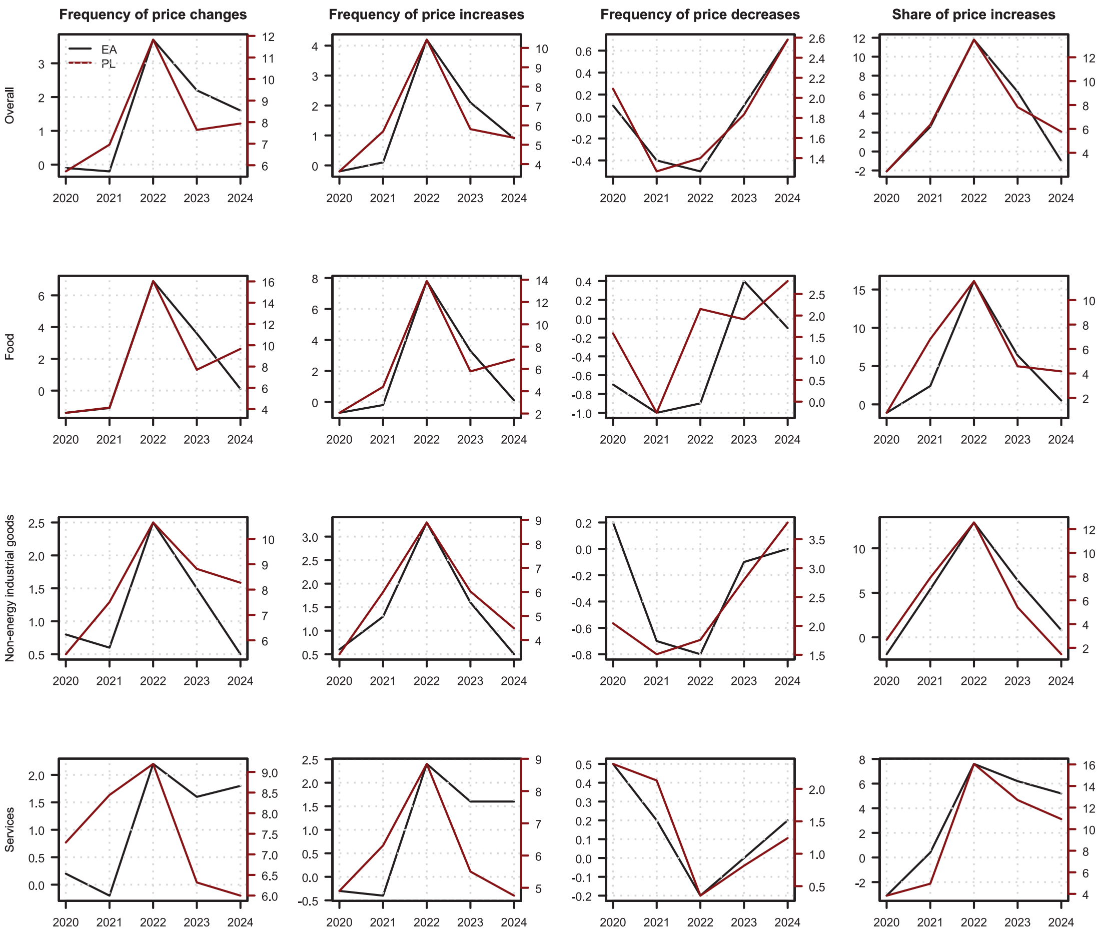
Note: Panel A shows actual rigidities calculated on the available sample, while Panel B shows counterfactual rigidities adjusted for the difference in inflation level. Adjusted frequencies and sizes are equalised to sectoral inflation level in the euro area using scale factors $a_j = \sqrt{\frac{\tilde{\pi}_{j,EA}}{\tilde{\pi}_{j,PL}}}$, where $\tilde{\pi}_j$ is recomposed average inflation for sector j. Normalisation assumes proportionality for increases and decreases.

Figure 10 Implied price duration in Poland and other countries (in months)



Note: The figure shows price duration (in months) for all goods and services as well as at the sectoral level. Dark shaded bars denote the results for prices including sales, whereas light shaded bars denote the range for duration once sales are accounted for by the filter proposed by Gautier et al. (2024). Results for Poland are marked with a red asterisk. Price durations for euro area and the U.S. are based on the frequencies reported by Gautier et al. (2024). Excluding sales figures are estimated using sales filter for: Poland, Greece, Luxembourg, Slovakia and Spain.

Figure 11 Change in frequencies after COVID-19 in Poland and the euro area

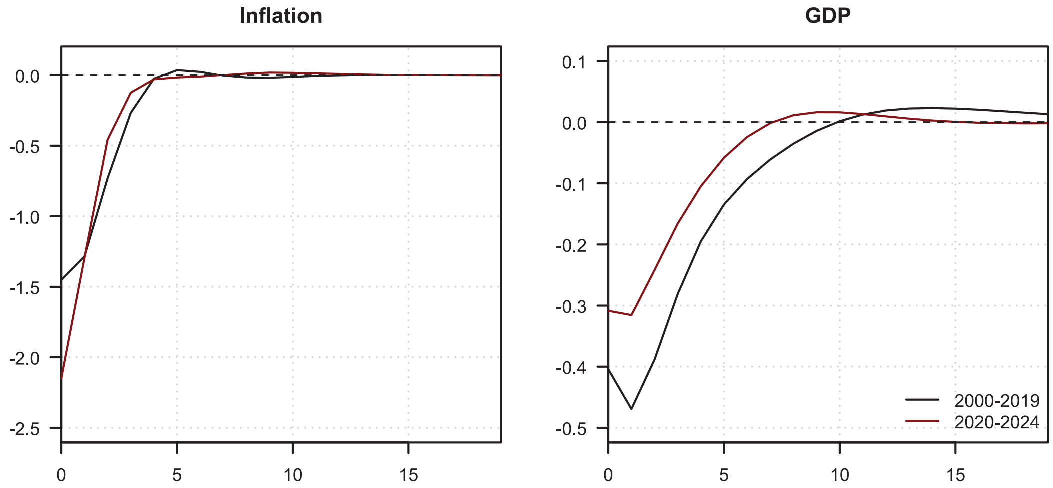


Note: The figure shows how frequency of price changes, increases, decreases and the share of price increases has changed since the outbreak of the COVID-19 pandemic compared to the pre-COVID-19 average. Results for Poland are plotted on secondary axis. Data for euro area are sourced from Gautier et al. (2026).

Figure 12 Monetary policy IRFs in one- and two-sectors frameworks

Note: Impulse responses to a one-percentage-point increase in the policy interest rate from an otherwise standard quarterly two-sector DSGE model with sticky- and flexible-price sectors and CES labour aggregation, allowing for either perfect or quasi-immobile labour. Calvo probabilities are calibrated from the data using median implied price durations excluding sales, computed separately for each sector, with $\theta_f = 0.435$ in the flexible-price sector and $\theta_s = 0.882$ in the sticky-price sector. The implied one-sector Calvo parameter is the weighted average of the two sectors, with the flexible-sector weight set to 0.55, yielding $\theta = 0.636$.

Figure 13 Monetary policy IRFs in high- and low-price stickiness framework



Note: Impulse responses to a one-percentage-point increase in the policy interest rate from an otherwise standard quarterly two-sector DSGE model with flexible- and sticky-price sectors and CES labour aggregation allowing for perfect labour mobility. Calvo probabilities are calibrated from the data using median implied price durations excluding sales, computed separately for each sector. In the baseline simulation (2000–2019), the Calvo parameters equal $\theta_f = 0.498$ in the flexible-price sector and $\theta_s = 0.898$ in the sticky-price sector. In the recent simulation (2020–2024)—corresponding to the post-COVID and full-scale war period—the degree of price rigidity declines, with Calvo parameters equal to $\theta_f = 0.265$ and $\theta_s = 0.813$, respectively.

Appendix A. Calculating measures of price stickiness

This study is based on data from Statistics Poland used for the calculation of the Consumer Price Index, which follows the COICOP classification. Within this framework, information is collected for approximately 1,500–2,000 representative products, with price data typically gathered from several hundred outlets across the country. These representative products are subsequently grouped and aggregated into so-called elementary groups using equal weights. Weights reflecting household expenditure patterns are assigned at the level of elementary groups. By applying these weights, statistics at different levels of the COICOP classifications can be calculated. This data structure has also shaped our procedure for computing measures of price stickiness, which we describe below.

Frequency of price change. Let P_{ijt} denote the price of ith individual product at time t assigned to the representative item j. First, we generate x_{ijt}, an indicator variable that informs whether the P_{ijt} and P_{ijt-1} were observed. Consequently:

$$x_{ijt} = \begin{cases} 1 & \text{if } P_{ijt} \text{ and } P_{ijt-1} \text{ are observed,} \\ 0 & \text{if } P_{ijt} \text{ and } P_{ijt-1} \text{ are not observed.} \end{cases} \qquad (A.1)$$

It should be noted that in practice due to the indicator x_{ijt}, any of the statistics can be computed starting from $t = 2$. Next, we calculate z_{ijt}. Again, it is a binary variable denoting whether the price of the ith product assigned to the representative item j has changed at time t with respect to $t - 1$. Hence,

$$z_{ijt} = \begin{cases} 1 & \text{if } P_{ijt} \neq P_{ijt-1} \wedge \ x_{ijt} = 1, \\ 0 & \text{otherwise.} \end{cases} \qquad (A.2)$$

To account for price increases and decreases we make use of two additional binary variables, w_{ijt} and s_{ijt}. They are generated as follows:

$$w_{ijt} = \begin{cases} 1 & \text{if } P_{ijt} > P_{ijt-1} \wedge x_{ijt} = 1, \\ 0 & \text{otherwise.} \end{cases} \qquad (A.3)$$

$$s_{ijt} = \begin{cases} 1 & \text{if } P_{ijt} < P_{ijt-1} \wedge x_{ijt} = 1, \\ 0 & \text{otherwise.} \end{cases} \qquad (A.4)$$

Using these binary variables from equations (A.1)-(A.4) we can calculate the frequency of price changes, increases and decreases. Let f_{jt} denote the frequency of price changes, f_{jt}^{+} the frequency of price increases and f_{jt}^{-} the frequency of price decreases for the representative item j at time t.

We calculate them as in equations (A.5)-(A.7):

$$f_{jt} = \frac{\sum_{i=1}^{n_j} z_{ijt}}{\sum_{i=1}^{n_j} x_{ijt}}, \qquad (A.5)$$

$$f_{jt}^{+} = \frac{\sum_{i=1}^{n_j} w_{ijt}}{\sum_{i=1}^{n_j} x_{ijt}}, \qquad (A.6)$$

$$f_{jt}^{-} = \frac{\sum_{i=1}^{n_j} s_{ijt}}{\sum_{i=1}^{n_j} x_{ijt}}. \qquad (A.7)$$

In all cases n_j denotes the number of individual products assigned to the representative item j at time t.

The measure of the frequency of price changes, increases and decreases for a given time period, $1, \dots, T^*$, is calculated as a simple average. Hence:

$$\bar{f}_j = \frac{\sum_{t=1}^{T^*} f_{jt}}{T^*}, \qquad (A.8)$$

$$\bar{f}_j^{+} = \frac{\sum_{t=1}^{T^*} f_{jt}^{+}}{T^*}, \qquad (A.9)$$

$$\bar{f}_j^{-} = \frac{\sum_{t=1}^{T^*} f_{jt}^{-}}{T^*}. \qquad (A.10)$$

Finally, it can be noted that measures of the frequency of price changes can be interpreted as the probability of the price change as they show the (average) fraction of prices that changed, increased or decreased over a given time period.

Size of the price change. Having defined measures related to the frequency of price changes, we turn the attention to statistics on the average size of the price changes, increases and decreases. Let dp_{ijt} denote the change in the price of the good or service i from $t-1$ to t, naturally assuming that $x_{ijt} = 1$:

$$\mathrm{dp}_{ijt} = \frac{P_{ijt}}{P_{ijt-1}}. \qquad (A.11)$$

We calculate the size of price changes, increases and decreases at time t for the representative item j as follows:

$$\mathrm{dp}_{jt} = \frac{\sum_{i=1}^{n_j} \mathrm{dp}_{ijt} z_{ijt}}{\sum_{i=1}^{n_j} z_{ijt}}, \qquad (A.12)$$

$$\mathrm{dp}_{jt}^{+} = \frac{\sum_{i=1}^{n_j} \mathrm{dp}_{ijt} w_{ijt}}{\sum_{i=1}^{n_j} w_{ijt}}, \tag{A.13}$$

$$\mathrm{dp}_{jt}^{-} = \frac{\sum_{i=1}^{n_j} \mathrm{dp}_{ijt} s_{ijt}}{\sum_{i=1}^{n_j} s_{ijt}}. \tag{A.14}$$

where n_j again denotes the number of products assigned to the representative item j at time t. Average size of the price changes, increases and decreases are again defined as a simple average over a given period of time:

$$\overline{\mathrm{dp}}_j = \frac{\sum_{t=1}^{T^*} \mathrm{dp}_{jt}}{T^*}, \tag{A.15}$$

$$\overline{\mathrm{dp}}_j^{+} = \frac{\sum_{t=1}^{T^*} \mathrm{dp}_{jt}^{+}}{T^*}, \tag{A.16}$$

$$\overline{\mathrm{dp}}_j^{-} = \frac{\sum_{t=1}^{T^*} \mathrm{dp}_{jt}^{-}}{T^*}. \tag{A.17}$$

Aggregation. As a first step, statistics are computed for elementary groups, which represent the lowest units of classification to which Statistics Poland assigns weights. Given that an elementary group typically comprises several representative products, metrics of frequencies and sizes of prices changes are derived using a simple arithmetic average.

$$f_{kt} = \frac{\sum_{j=1}^{J} f_{jt}}{J}, \tag{A.18}$$

$$\mathrm{dp}_{kt} = \frac{\sum_{j=1}^{J} \mathrm{dp}_{jt}}{J}, \tag{A.19}$$

where J denotes the number of the representative items j belonging to the elementary group k at time t.

To arrive at price stickiness measures at the aggregated level, we employ the weighing system that is also used in the compilation of the CPI. This type of aggregation is common to studies on price rigidity (e.g., Dhyne et al., 2006; Gautier et al., 2024, Gautier et al., 2026). This aggregation can be described as follows:

$$\bar{f}_t = \sum_{k=1}^{K} \omega_{kt} f_{kt}\,, \tag{A.20}$$

$$\overline{\mathrm{dp}}_t = \sum_{k=1}^{K} \omega_{kt} \mathrm{dp}_{kt} \,. \qquad (A.21)$$

where ω_{kt} denotes the weight of elementary group k at time t. Conceptually, $\bar{f}_t$ and $\overline{\mathrm{dp}}_t$ denote here measures of the frequency and size of price changes for any group at a higher level of aggregation than the elementary group level at time t. As already noted, in generating the results solely for Poland, we make use of CPI weights that are in line with the consumption patterns (as reported by Statistics Poland). In turn, for the purpose of international comparisons, we use HICP weights, ω_k, averaged over the same period as in the literature (Gautier et al., 2024, Gautier et al., 2026) with the aim of obtaining comparable results.

Again, the measure of the frequency and sizes of prices changes for a given time period, $1, \dots, T^*$, is calculated as a simple average. Hence:

$$\bar{f} = \frac{\sum_{t=1}^{T^*} \bar{f}_t}{T^*}, \qquad (A.22)$$

$$\overline{\mathrm{dp}} = \frac{\sum_{t=1}^{T^*} \overline{\mathrm{dp}}_t}{T^*}. \qquad (A.23)$$

Duration. A simple rationale would be to calculate the implied duration for any representative item j as the inverse of the average price frequency: $\bar{T}_j = \bar{f}_j^{-1}$. However, this calculation method implicitly assumes that retailers change their prices at most once per month. We follow other contributions (e.g., Bils and Klenow, 2004; Baudry et al., 2007; Nakamura and Steinsson, 2008) and estimate the implied price duration under the implicit assumption of equal probability of price change over the course of the month (so that prices are determined in continuous time).[32] Therefore, the average duration of a price is given by:

$$\bar{T}_j = -\frac{1}{\ln(1 - \bar{f}_j)}. \qquad (A.24)$$

Since for some representative items we can expect that the average frequency of price change may be zero or close to zero, $\bar{f}_j \approx 0$, thus leading to extremely large values in implied duration

[32] Specifically, we also do not calculate direct measures of duration, based on price spells and trajectories. Baudry et al. (2007) discusses why using indirect (frequency) approach offers several advantages over inferring duration directly. Further discussion related to assumptions behind calculating various duration measures can be found there.

on the aggregate level, we prefer to calculate duration from aggregated statistics, such as the average frequency for the overall CPI:

$$\bar{T} = -\frac{1}{\ln\left(1-\bar{f}\right)}. \qquad (A.25)$$

Price synchronization. We assess the degree of price synchronization at any group k using the ratio proposed by Fisher and Konieczny (2000):

$$S_k = \left(\frac{1}{T}\frac{\sum_{t=1}^{T}\left(f_{kt}-\bar{f}_k\right)^2}{\bar{f}_k \times \left(1-\bar{f}_k\right)}\right). \qquad (A.26)$$

This measure is bounded between 0 (perfect staggering) and 1 (perfect synchronization).

Appendix B. New Keynesian Model

B1. The model

Households. Households are forward-looking. In each period t they choose consumption c_t and labor supply n_t to maximize expected lifetime utility:

$$\max_{\{c_t(i),n_t(i)\}} \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left\{ \frac{(c_t(i) - \xi c_{t-1})^{1-\sigma_c}}{1-\sigma_c} - \psi \frac{n_t^{1+\sigma_n}}{1+\sigma_n} \right\}, \qquad (B.1)$$

where $\beta \in (0,1)$ is the discount factor, σ_c is the inverse of the intertemporal elasticity of substitution, σ_n is the inverse of the Frisch elasticity of labour supply, $\psi > 0$ scales the disutility from labor, and $\xi \in [0,1)$ captures external habit formation in consumption.

Households supply labour to firms in both the flexible and sticky sectors, $j \in \{f, s\}$, with aggregate labour determining total disutility given by:

$$n_t = \left(\omega_f^{-\frac{1}{\lambda}} n_{f,t}^{\frac{1+\lambda}{\lambda}} + \omega_s^{-\frac{1}{\lambda}} n_{s,t}^{\frac{1+\lambda}{\lambda}} \right)^{\frac{\lambda}{1+\lambda}}, \qquad (B.2)$$

where $\lambda > 0$ determines the degree of labor mobility between sectors. Perfect labor mobility corresponds to $\lambda \to \infty$, while $\lambda \to 0$ implies that labor is effectively non-substitutable across sectors.

Households earn income from supplying labour at the nominal wage $W_{j,t}$ in each sector j from renting their fixed capital stock k at the nominal rental rate $R_{k,t}$, and from firm dividends Div_t. They can also save by purchasing nominal bonds B_t that pay the gross nominal interest rate R_t. The nominal budget constraint is given by:

$$P_t c_t + B_t = \sum_{j\in\{f,s\}} W_{j,t}\, n_{j,t} + R_{t-1}B_{t-1} + R_{k,t}k + Div_t, \qquad (B.3)$$

where P_t denotes the aggregate price level.

Firms. There are two sectors in the economy: one with relatively flexible prices and the other with sticky prices, $j \in \{f, s\}$. A final good producer combines sectoral goods into a homogeneous final consumption good using a CES technology:

$$y_t = \left(\omega_f^{\frac{1}{\eta}} y_{f,t}^{\frac{\eta-1}{\eta}} + \omega_s^{\frac{1}{\eta}} y_{s,t}^{\frac{\eta-1}{\eta}} \right)^{\frac{\eta}{\eta-1}}, \qquad (B.4)$$

with weights satisfying:

$$\omega_f + \omega_s = 1, \qquad (B.5)$$

and η denoting the elasticity of substitution between between goods produced in the flexible-price sector Y_t^f and sticky-price sector Y_t^s.

Each sector j is itself a Dixit-Stiglitz aggregator of a unit measure of differentiated varieties $i \in [0,1]$:

$$y_{j,t} = \left[\int_0^1 y_{j,t}\,(i)^{\varepsilon_j - 1/\varepsilon_j} di \right]^{\frac{\varepsilon_j}{\varepsilon_j - 1}}, \qquad (B.6)$$

where ε_j denotes elasticity of substitution between individual varieties within sector j.

Solving the final good producer problem we get the aggregate price index:

$$P_t = \left(\omega_f P_{f,t}^{1-\eta} + \omega_s P_{s,t}^{1-\eta} \right)^{\frac{1}{1-\eta}}, \qquad (B.7)$$

where the sectoral price index is given by:

$$P_{j,t} = \left[\int_0^1 P_{j,t}\,(i)^{1-\varepsilon_j} di \right]^{\frac{1}{1-\varepsilon_j}}. \qquad (B.8)$$

Demand for each variety is:

$$y_{j,t}(i) = \left(\frac{P_{j,t}(i)}{P_{j,t}} \right)^{-\varepsilon_j} Y_{j,t}. \qquad (B.9)$$

Each sector consists of monopolistically competitive firms producing differentiated goods. Firm-level production follows a Cobb–Douglas technology:

$$y_{j,t}(i) = z_t k_{j,t}(i)^{\alpha} n_{j,t}(i)^{1-\alpha}. \qquad (B.10)$$

Prices are set under Calvo stickiness. In each sector, a sector-specific fraction θ_j of firms cannot reoptimize each period. Reoptimizing firms choose prices to maximize expected discounted profits, leading to a New Keynesian Phillips Curve with partial indexation.

B2. Closing the model.

Monetary Authority. The central bank follows a Taylor-type rule with interest rate smoothing:

$$\frac{R_t}{R} = \left(\frac{R_{t-1}}{R}\right)^{\phi_R} \left[\left(\frac{\pi_t}{\pi^{ss}}\right)^{\phi_\pi} \left(\frac{y_t}{y}\right)^{\phi_y}\right]^{1-\phi_R} e^{v_t^R}. \qquad (B.11)$$

Market Clearing. Markets clear each period.

Factor markets:

$$k_{s,t} + k_{f,t} = k_t, \qquad (B.12)$$

$$n_{s,t} + n_{f,t} = n_t, \qquad (B.13)$$

with price dispersion defined as:

$$\Delta_{j,t} \equiv \int_0^1 \left(\frac{P_{j,t}(i)}{P_t^j}\right)^{-\varepsilon_j} di, \qquad (B.14)$$

which can be expressed recursively:

$$\Delta_t \;=\; (1-\theta)\left(\frac{P_t^*}{P_t}\right)^{-\varepsilon} \;+\; \theta\, \pi_t^{\varepsilon}\, \Delta_{t-1}. \qquad (B.15)$$

Goods market.

$$c_t = y_t. \qquad (B.16)$$

Definition of GDP.

$$p_{1,t}c_{1,t} + p_{2,t}c_{2,t} = \tilde{y}_t. \qquad (B.17)$$

B3. Social Welfare

Household welfare is defined as the expected discounted sum of per-period utilities:

$$W_0 = \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t u(c_t - \xi c_{t-1}, n_t), \qquad (B.18)$$

where $\beta \in (0,1)$ is the subjective discount factor, and $u(\cdot)$ denotes the instantaneous utility function.

To compare the welfare in the business-cycle equilibrium with welfare in the steady state, we compute the consumption-equivalent welfare compensation. Specifically, we define a constant proportional change m (expressed in percentage terms) in consumption that would make households in the business-cycle equilibrium as well off as in the steady state. Formally, m satisfies:

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t u\left((1+m)(c_t^{\mathrm{A}} - \xi c_{t-1}^{\mathrm{A}}), n_t^{\mathrm{A}}\right) = \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t u((1-\xi)\bar{c}, \bar{n}), \qquad (B.19)$$

where $\bar{c}$ and $\bar{n}$ denote steady-state consumption and labour, respectively, and superscript A refers to variables in the business-cycle equilibrium.

The solution m represents the consumption-equivalent compensation—that is, the permanent proportional increase in consumption required in the business-cycle equilibrium for households to attain the same level of welfare as in the steady state.

B4. Parametrization

The model is parametrized at a quarterly frequency. The values of structural parameters are chosen to be consistent with standard targets in the New Keynesian and DSGE literature and to ensure that the steady state of the model reflects key empirical regularities of the Polish economy.

Table B1 summarizes the parameters and their calibration targets.

Specifically, the discount factor $\beta = 0.99$ implies a steady-state annual real interest rate of approximately 4%. The capital share in production is set to $\alpha = 0.3$, consistent with standard estimates for advanced and emerging European economies[33]. The inverse of the intertemporal elasticity of substitution is $\sigma_c = 1.5$, while the inverse of the Frisch elasticity of labor supply is $\sigma_n = 3$, both in line with typical values used in medium-scale DSGE models.

The parameter $\psi = 44$ governs the disutility of labor and is calibrated so that steady-state hours worked correspond to one third of the available time endowment. The elasticity of substitution between the two consumption goods is set to $\phi = 0.6$, capturing moderate substitutability between flexible- and sticky-price sectors. Habit persistence is parameterized by $\xi = 0.8$, implying a high degree of consumption smoothing, while $\zeta = 0.8$ governs the degree of price indexation to past inflation, consistent with empirical estimates from inflation dynamics.

The parameter $\eta_f = 0.55$ determines the steady-state share of flexible-price goods in aggregate consumption. This value implies that roughly half of household expenditure falls on goods with relatively flexible prices, consistent with micro-level evidence for the Polish economy and similar European markets.

The gross steady-state markup is set to $\mu = 1.1$, implying a 10% price markup over marginal cost. To neutralize the resulting distortion in the steady state, a tax rate τ equal to the markup ($\tau = \mu$) is introduced. The steady-state quarterly inflation rate is set to $\pi^{ss} = 1.0062$, corresponding to an annual inflation rate of approximately 2.5%.

Finally, the Taylor rule parameters are chosen to reflect a standard interest rate reaction function with policy inertia. The interest rate smoothing coefficient is $\phi_R = 0.85$, the response to inflation deviations is $\phi_\pi = 1.5$, and the response to output deviations is $\phi_y = 0.15$. These values are in line with conventional calibrations in the monetary policy literature.

[33] Gradzewicz et al. (2024) estimate the labour share in Poland to lie in the range of approximately 50–60%, depending on the methodology and the treatment of self-employed income. However, the empirical measure of the labour share is not directly equivalent to the capital share parameter in the DSGE production function. In our model, the parameter α is calibrated in line with standard values used in the New Keynesian and RBC literature (typically 0.3–0.35), which ensures comparability with existing studies and preserves standard quantitative properties of the model. We do not interpret this parameter as a direct empirical estimate of Poland's current capital share. Moreover, we verified that increasing the capital share to $\alpha = 0.4$ does not materially affect the impulse responses or the main quantitative conclusions of the model.

Table B1 Structural parameters

Parameter	Value	Description / Calibration Target
ψ	44	Labor disutility parameter (implies $n = 1/3$ in steady state)
ϕ	0.6	Elasticity of substitution between two consumption goods
ξ	0.8	Habit persistence in consumption
ζ	0.8	Degree of price indexation
σ_n	3	Inverse of Frisch elasticity of labor supply
σ_c	1.5	Inverse of intertemporal elasticity of substitution
α	0.3	Capital share in production
β	0.99	Discount factor (implies 4% annual steady-state rate)
η_f	0.55	Share of flexible-price goods in steady-state consumption
μ	1.1	Steady-state price markup (10% above marginal cost)
τ	1.1	Tax rate offsetting the markup distortion ($\tau = \mu$)
π^{ss}	1.0062	Steady-state quarterly inflation (≈2.5% annual)
ϕ_R	0.85	Taylor rule interest rate smoothing parameter
ϕ_π	1.5	Taylor rule coefficient on inflation
ϕ_y	0.15	Taylor rule coefficient on output gap

B5. Impulse response function to a monetary policy shock

In this subsection we explain the sources of the differences in aggregate responses by examining the sectoral dynamics that underlie them. Figure B1 provides a more detailed view of the two-sector economy. In addition to the responses of aggregate GDP and inflation, it displays sectoral outputs, employment, sectoral inflations, and relative prices.

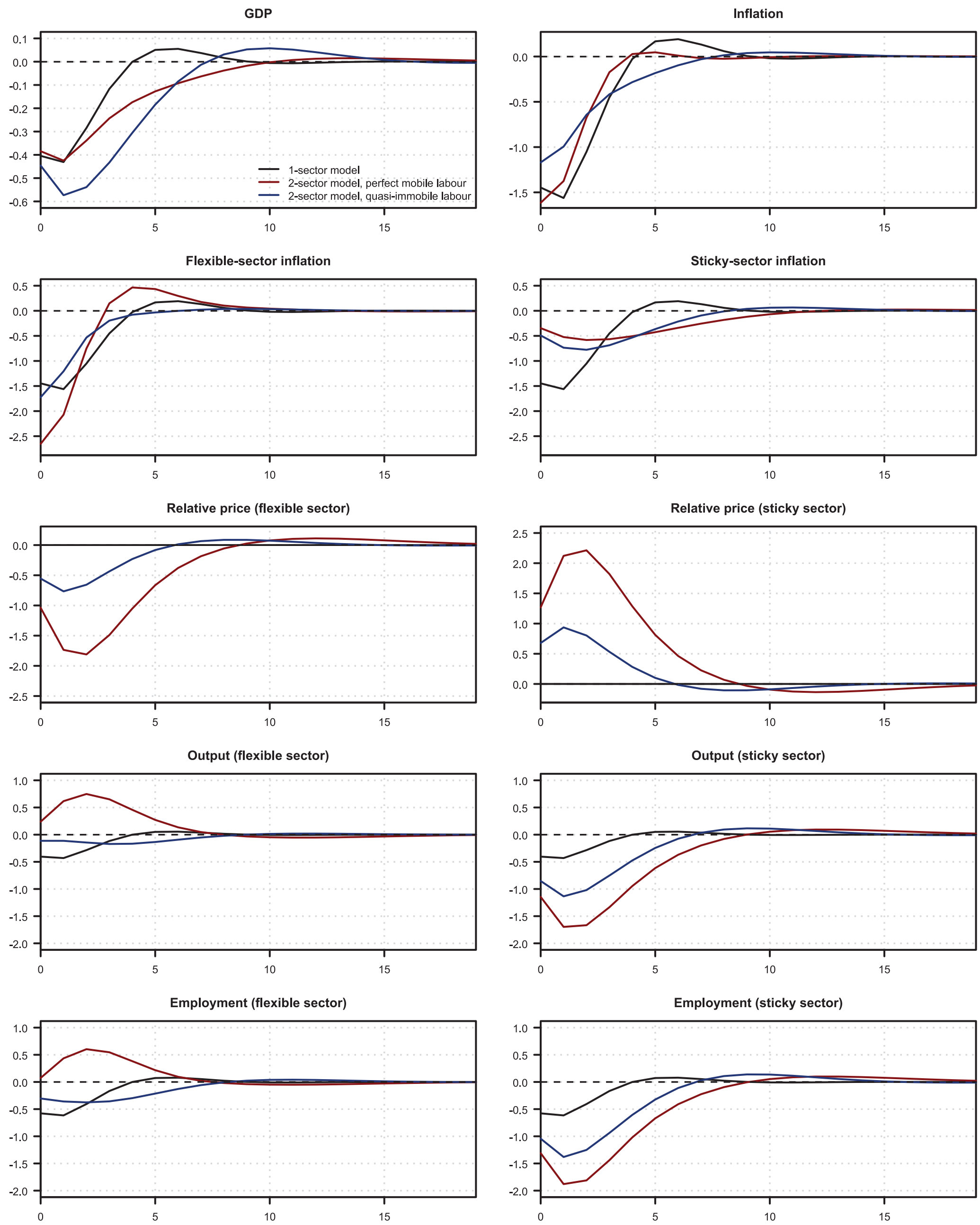
Prices adjust more quickly in the flexible sector, leading to a decline in the relative price of flexible goods, while the relative price of sticky goods rises. This movement in relative prices induces a reallocation of resources toward the flexible sector. When labour is mobile, the fall in the relative price of flexible goods outweighs the contractionary effect of lower aggregate demand, resulting in an increase in output in the flexible sector. However, when labour is immobile, this mechanism only mitigates the overall effect of the decline in aggregate demand. Consequently, output in the flexible sector still falls, though by less, and recovers more rapidly than in the one-sector model, whereas the sticky sector experiences a sharper and more prolonged contraction in both output and employment.

With perfect labour mobility, the adjustment of relative prices is gradual and lasts longer than the convergence of aggregate inflation. This sluggish reversion reflects the persistent dispersion in sectoral price dynamics and slows the normalization of production and employment in the sticky sector. As a result, the aggregate economy remains below its steady state for an extended period even after inflation begins to stabilize.

When labour mobility is restricted, these reallocation mechanisms weaken. Labor cannot easily move from the contracting sticky sector to the expanding flexible sector, amplifying the output loss and making the recovery more gradual. The persistence of relative price misalignments becomes more pronounced, reinforcing the differences in inflation dynamics observed across models. Hence, imperfect labour mobility magnifies the real effects of monetary policy shocks and reduces the speed of aggregate adjustment.

Overall, the two-sector impulse responses highlight that heterogeneity in price stickiness and labour mobility introduces additional propagation channels absent in the one-sector framework. Relative price movements, sectoral reallocation, and labour frictions jointly determine both the strength and persistence of the transmission of monetary policy shocks to the real economy.

Figure B1 Monetary policy IRFs in one- and two-sectors frameworks



Note: Impulse responses to a one-percentage-point increase in the policy interest rate from an otherwise standard quarterly two-sector DSGE model with sticky- and flexible-price sectors and CES labour aggregation, allowing for either perfect or quasi-immobile labour. Calvo probabilities are calibrated from the data using median implied price durations excluding sales, computed separately for each sector, with $\theta_f = 0.435$ in the flexible-price sector and $\theta_s = 0.882$ in the sticky-price sector. The implied one-sector Calvo parameter is the weighted average of the two sectors, with the flexible-sector weight set to 0.55, yielding $\theta = 0.636$.

Appendix C. Further results

Table C1 Price stickiness metrics

			Frequency of price			Size of price		Average	Share of	Price
			change	increase	decrease	increase	decrease	duration	sales	synch.
CPI	Incl. sales	mean	22.1	12.8	9.3	13.3	11.3	4.0	3.4	0.27
		med.	14.5	7.7	4.3	6.7	7.5	6.4	1.7	0.23
	Excl. sales	mean	18.0	10.7	7.3	12.5	11.0	5.0	.	0.26
		med.	11.1	5.9	3.3	6.1	6.9	8.5	.	0.21
By sector										
Core inflation	Incl. sales	mean	14.3	8.6	5.6	15.2	12.6	6.5	2.5	0.26
		med.	8.9	5.3	2.4	7.7	8.3	10.7	1.0	0.23
	Excl. sales	mean	11.6	7.2	4.4	14.7	12.5	8.1	.	0.24
		med.	7.5	4.4	1.9	7.1	8.0	12.8	.	0.20
Food and non-alcoholic beverages	Incl. sales	mean	30.5	17.3	13.2	12.0	11.1	2.7	6.5	0.18
		med.	25.1	13.0	9.4	6.7	7.6	3.5	5.7	0.20
	Excl. sales	mean	22.5	13.2	9.3	10.6	10.4	3.9	.	0.19
		med.	17.2	9.1	5.8	5.6	6.7	5.3	.	0.21
Energy	Incl. sales	mean	36.5	20.7	15.9	7.2	6.2	2.2	1.7	0.46
		med.	12.6	3.7	1.2	3.4	3.5	7.5	0.0	0.40
	Excl. sales	mean	33.6	19.1	14.5	7.2	6.2	2.4	.	0.47
		med.	10.0	3.1	1.1	3.4	3.5	9.5	.	0.35
Non-energy industrial goods	Incl. sales	mean	19.3	11.1	8.2	12.7	10.9	4.7	4.0	0.22
		med.	15.3	8.5	5.3	6.2	7.1	6.0	2.4	0.18
	Excl. sales	mean	14.7	8.7	6.0	11.7	10.5	6.3	.	0.19
		med.	11.5	6.4	3.9	5.7	6.8	8.2	.	0.14
Services	Incl. sales	mean	8.8	6.0	2.8	18.4	14.9	10.9	0.8	0.31
		med.	3.0	2.2	0.5	11.1	10.7	32.8	0.0	0.33
	Excl. sales	mean	8.1	5.5	2.5	18.2	15.0	11.9	.	0.30
		med.	2.8	2.0	0.4	11.1	11.1	35.7	.	0.28

Note: The table presents mean and median estimates of price stickiness for all goods and services as well as at the sectoral level for the full sample (i.e., 2000–2024). Share of sales and the results excluding sales are based on the filter proposed by Nakamura and Steinsson (2008). Price synchronization is calculated as in Fisher and Konieczny (2000).

Table C2 Mean estimates for price stickiness metrics for selected years

	Frequency of price			Size of price		Average
	change	increases	decreases	increase	decrease	duration
2000–2019	20.3 (16.6)	11.6 (9.7)	8.7 (6.9)	13.0 (12.4)	11.3 (11.0)	4.4 (5.5)
2020	26.7 (21.0)	15.2 (12.2)	11.5 (8.8)	14.6 (13.4)	11.4 (10.7)	3.2 (4.3)
2021	27.4 (21.8)	18.3 (15.4)	9.1 (6.4)	13.9 (12.5)	11.3 (10.5)	3.1 (4.1)
2022	33.3 (28.0)	21.1 (18.6)	12.2 (9.5)	14.5 (13.7)	11.8 (11.5)	2.5 (3.0)
2023	30.2 (24.2)	16.7 (13.6)	13.5 (10.5)	14.3 (13.2)	11.8 (11.3)	2.8 (3.6)
2024	30.1 (24.0)	17.3 (14.0)	12.8 (10.0)	14.9 (13.4)	12.1 (11.6)	2.8 (3.6)

Note: The table presents selected mean price stickiness measures prior to the COVID-19 pandemic (2000–2019) as well as following the outbreak of the COVID-19 pandemic in 2020 and the outburst of the Russian invasion of Ukraine in 2022. In parentheses we report measures calculated on prices following the exclusion of sales with the use of the sales filter by Nakamura and Steinsson (2008).

Table C3 Median estimates for price stickiness metrics for selected years

	Frequency of price			Size of price		Average
	change	increases	decreases	increase	decrease	duration
2000–2019	13.0 (10.1)	7.1 (5.6)	4.2 (3.2)	6.6 (5.9)	7.4 (6.9)	7.2 (9.4)
2020	20.7 (14.1)	10.1 (7.6)	6.2 (3.9)	7.2 (6.1)	7.7 (6.7)	4.3 (6.6)
2021	22.0 (14.9)	12.9 (9.0)	5.3 (3.4)	7.5 (6.4)	8.2 (6.9)	4.0 (6.2)
2022	27.9 (20.6)	16.2 (13.6)	5.5 (3.5)	8.7 (8.2)	8.1 (7.6)	3.1 (4.3)
2023	24.5 (17.1)	12.7 (9.9)	7.0 (4.2)	8.2 (7.2)	8.1 (7.6)	3.6 (5.3)
2024	25.8 (15.6)	12.4 (8.3)	7.7 (4.9)	7.8 (6.7)	8.5 (7.6)	3.3 (5.9)

Note: The table presents selected median price stickiness measures prior to the COVID-19 pandemic (2000–2019) as well as following the outbreak of the COVID-19 pandemic in 2020 and the outburst of the Russian invasion of Ukraine in 2022. In parentheses we report measures calculated on prices following the exclusion of sales with the use of the sales filter by Nakamura and Steinsson (2008).

Table C4 Data coverage by country

	Available country data		Common subsample		
	% of EA basket	Number of COICOP5 groups	% of EA basket	Number of COICOP5 groups	Percent of sales
Poland	98.0	276	58.2	165	3.6
Euro area	.	.	58.9	166	3.8
United States	88.3	254	58.5	163	7.4*
Austria	89.2	230	57.5	156	5.1
Belgium	42.5	95	36.5	81	1.1
France	83.2	231	56.9	162	3.2
Germany	86.9	237	58.1	163	4.1
Greece	64.0	200	51.4	150	3.8
Italy	61.1	174	56.2	159	4.3
Latvia	92.5	206	56.8	148	10.7
Lithuania	82.3	228	57.8	161	2.3
Luxembourg	97.0	242	58.9	164	4.6
Slovakia	94.1	220	56.7	153	4.9
Spain	52.4	129	50.4	124	5.1

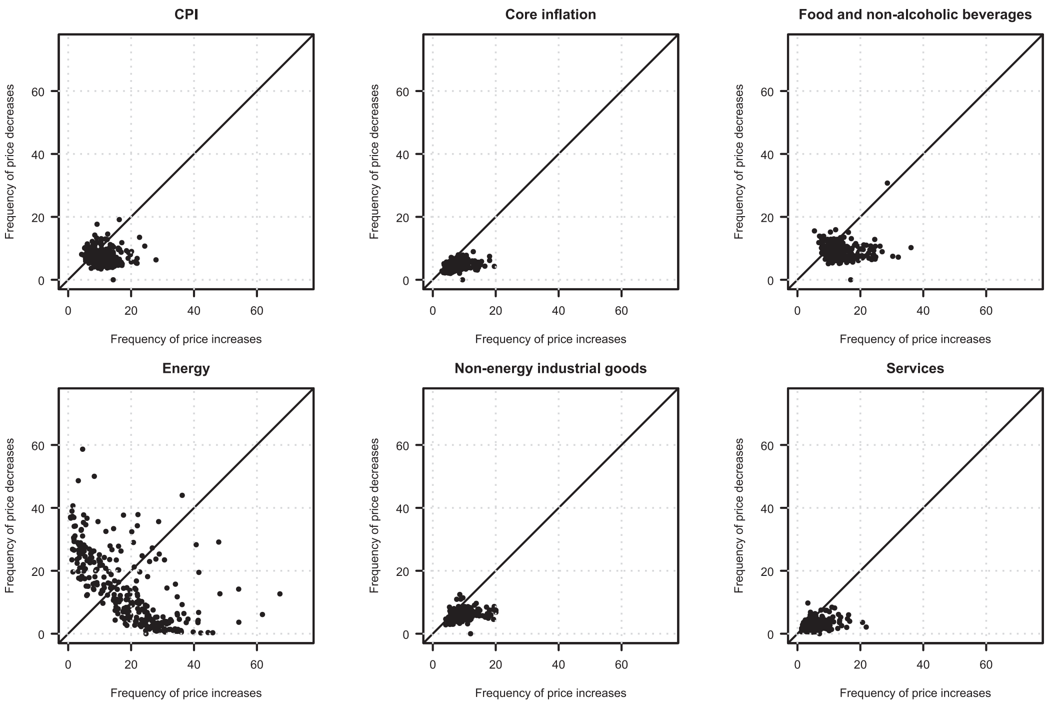
Note: *Figure reported in Nakamura and Steinsson (2008). Data for euro area countries taken from Gautier et al. (2024). Percent of sales is based on a sales filter for Poland, Greece, Luxembourg, Slovakia and Spain.

Table C5 Median frequency of price changes (%)

	Including sales				Excluding sales			
	Frequency of price changes			Share of increases	Frequency of price changes			Share of increases
	Change	Increase	Decrease		Change	Increase	Decrease	
PL	14.6	7.6	7.9	56.9	10.9	5.4	4.7	60.3
EA	11.1	6.3	4.4	58.4	7.2	4.8	2.3	63.5
US	18.3	10.3	8.4	56.6	8.5	5.7	2.7	70.7
AT	10.6	7.3	4.2	58.4	7.6	4.8	2.1	62.8
BE	6.6	4.5	2.7	61.2	6.6	4.4	2.7	64.2
ES	15.1	7.7	6.5	56.1	8.8	5.7	3.7	59.5
FR	15.1	7.7	6.6	57.5	9.6	6.5	3.9	59.7
DE	9.5	5.2	3.9	57.4	5.7	3.5	2.0	63.5
GR	9.2	5.7	4.0	60.3	6.4	4.1	2.4	63.8
IT	9.1	5.6	3.6	61.1	5.8	4.4	1.8	67.4
LV	14.0	7.4	6.9	55.2	6.0	4.2	1.7	71.2
LT	13.2	7.2	6.3	55.6	10.6	6.5	4.1	64.3
LU	11.2	7.4	3.6	67.0	6.3	4.7	1.1	79.5
SK	9.9	5.6	4.4	57.8	7.9	4.5	3.2	61.3

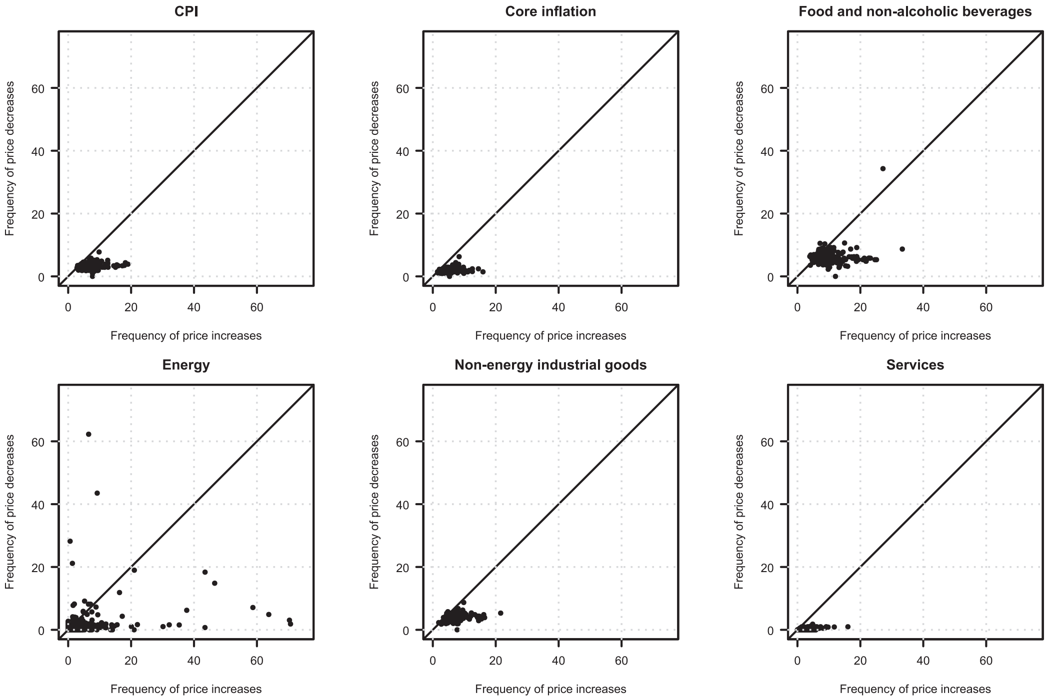
Note: Data for euro area and U.S. are sourced from Gautier et al. (2024). Excluding sales figures estimated using sales filter for: Poland, Greece, Luxembourg, Slovakia and Spain. Frequencies excluding sales are corrected for Slovakia.

Figure C1 Price rigidity metrics at the product level (in percent)



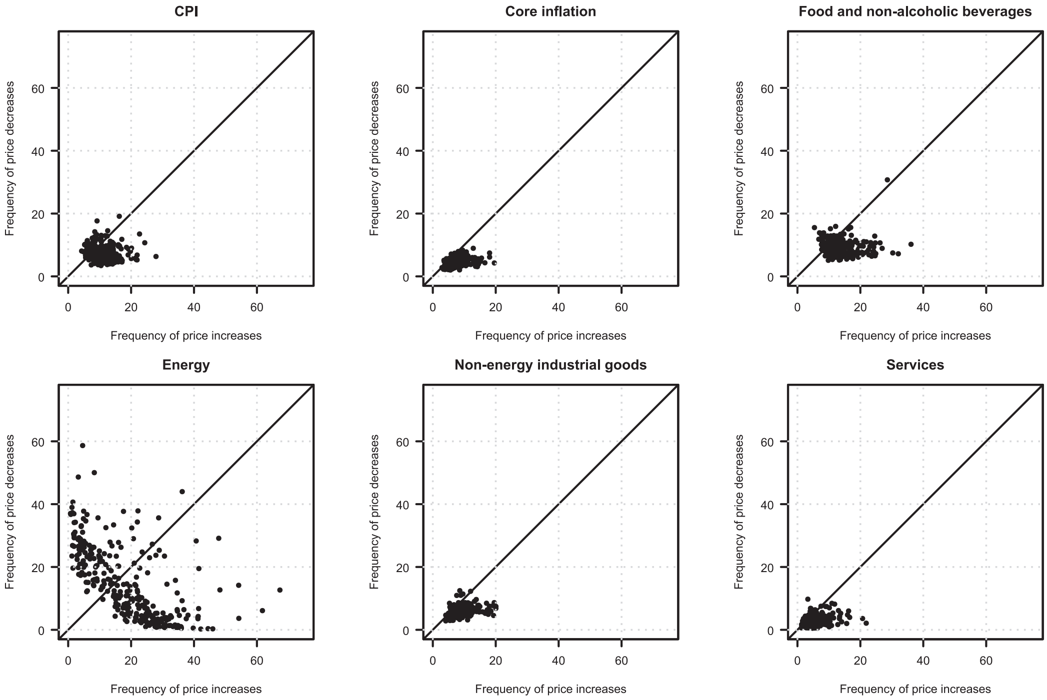
Note: The figure shows the frequency of price increases and decreases at the product level. Results are reported for mean estimates for prices excluding sales in line with the Gautier et al. (2024) approach.

Figure C2 Price rigidity metrics at the product level (in percent)



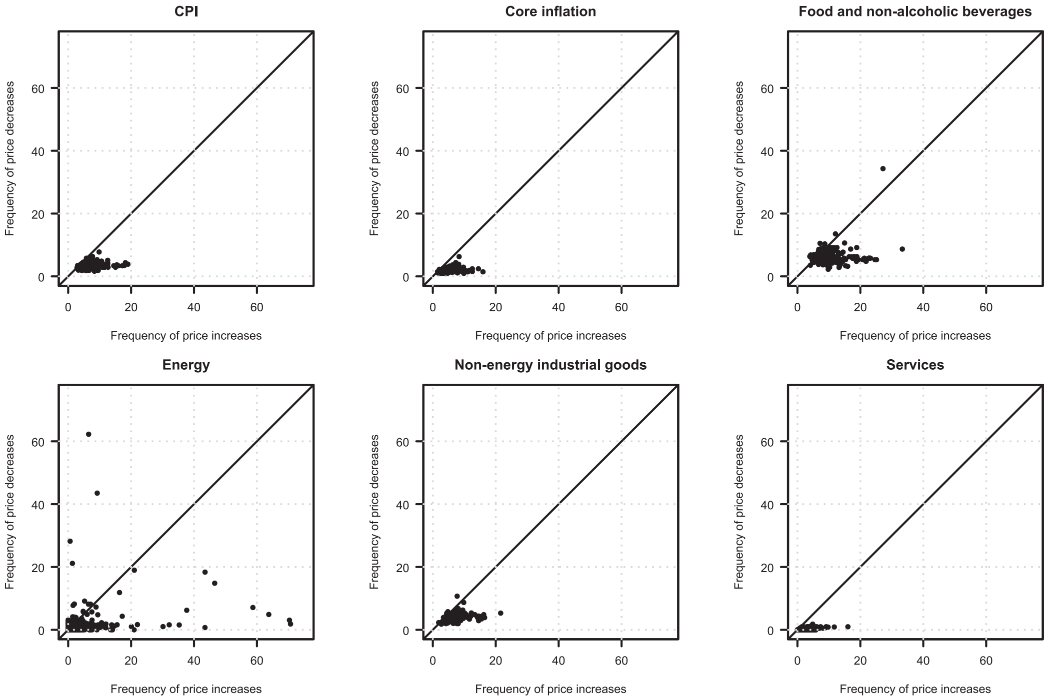
Note: The figure shows the frequency of price increases and decreases at the product level. Results are reported for median estimates for prices excluding sales in line with the Gautier et al. (2024) approach.

Figure C3 Price rigidity metrics at the product level (in percent)



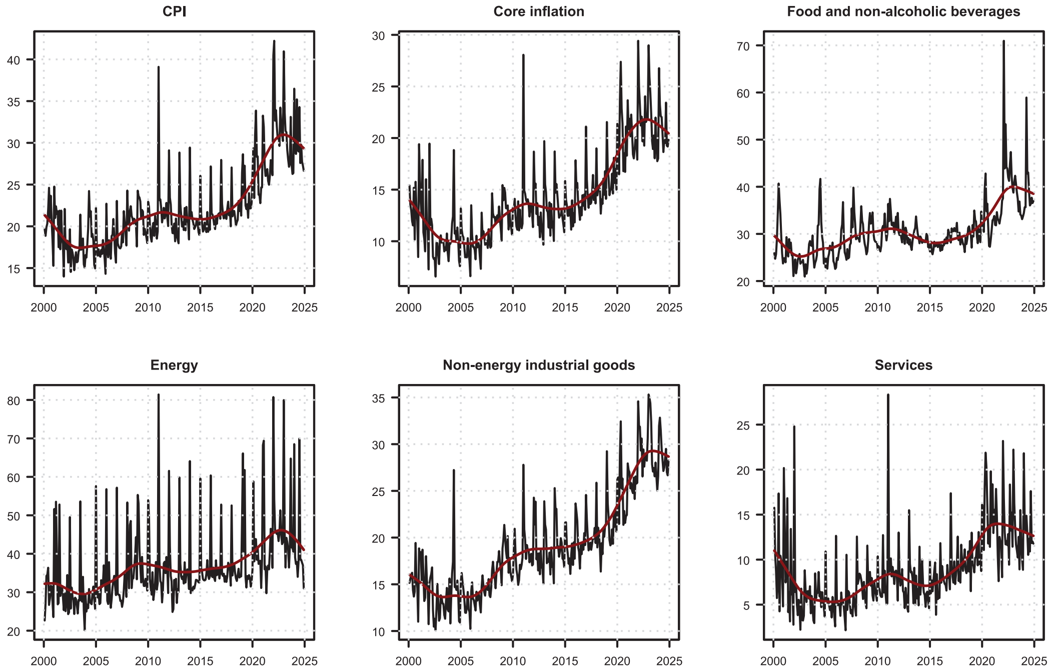
Note: The figure shows the frequency of price increases and decreases at the product level. Results are reported for mean estimates for prices excluding sales in line with the Nakamura and Steinsson (2008) approach.

Figure C4 Price rigidity metrics at the product level (in percent)



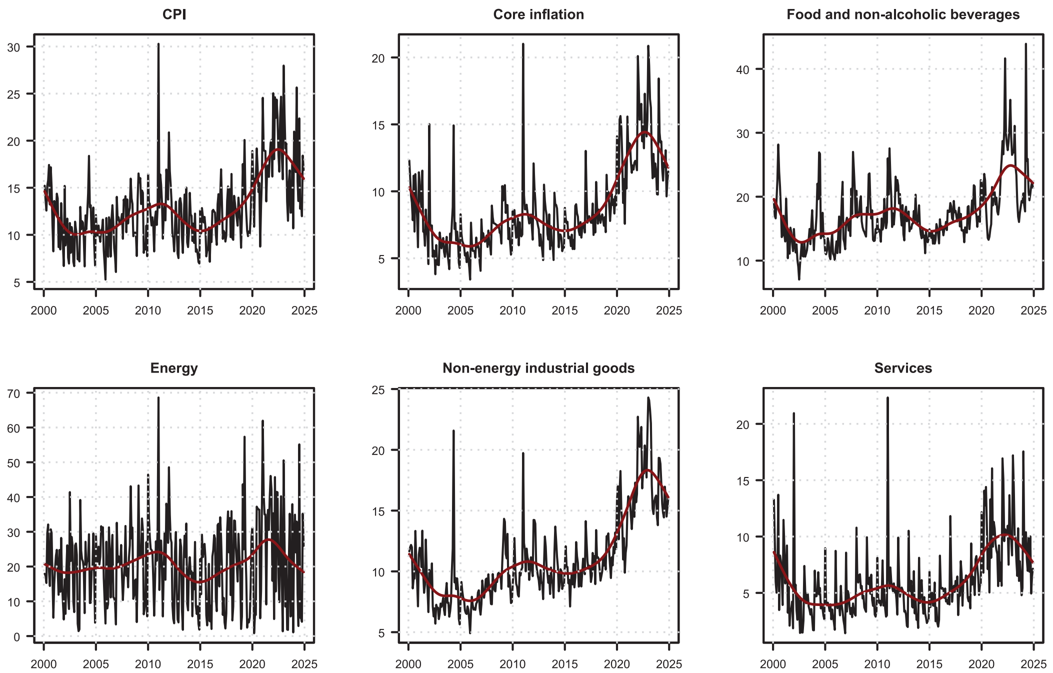
Note: The figure shows the frequency of price increases and decreases at the product level. Results are reported for median estimates for prices excluding sales in line with the Nakamura and Steinsson (2008) approach.

Figure C5 Frequency of price change (in percent)



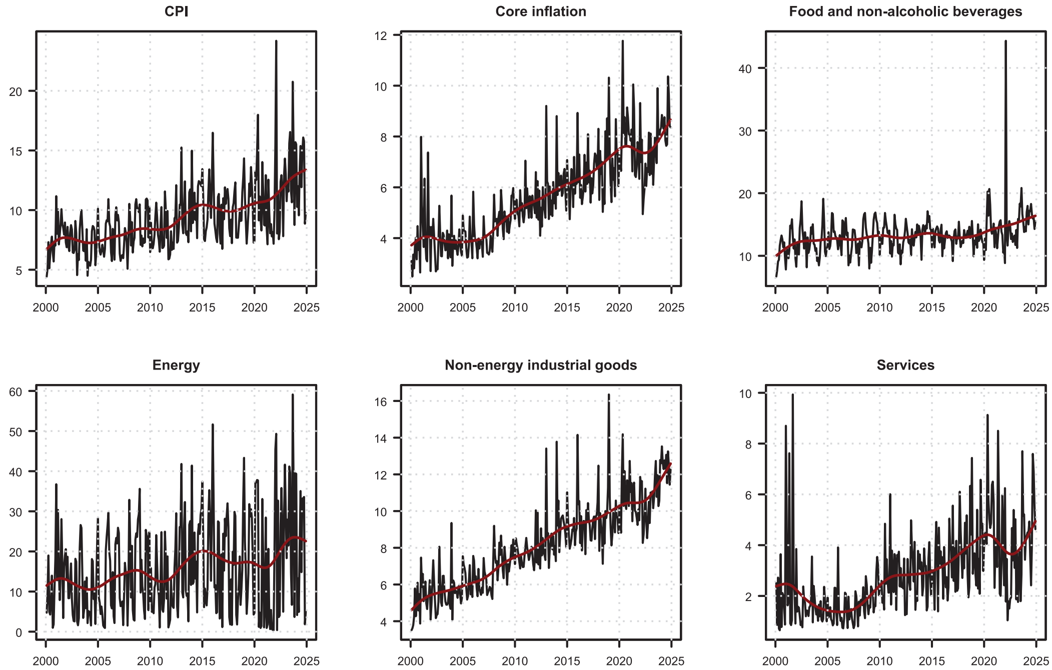
Note: The figure shows the mean frequency of price change for all goods and services as well as at the sectoral level. Red line denotes trend extracted using the Hodrick-Prescott filter. Median estimates or estimates for prices excluding sales are qualitatively and quantitatively similar. They are available upon request.

Figure C6 Frequency of price increase (in percent)



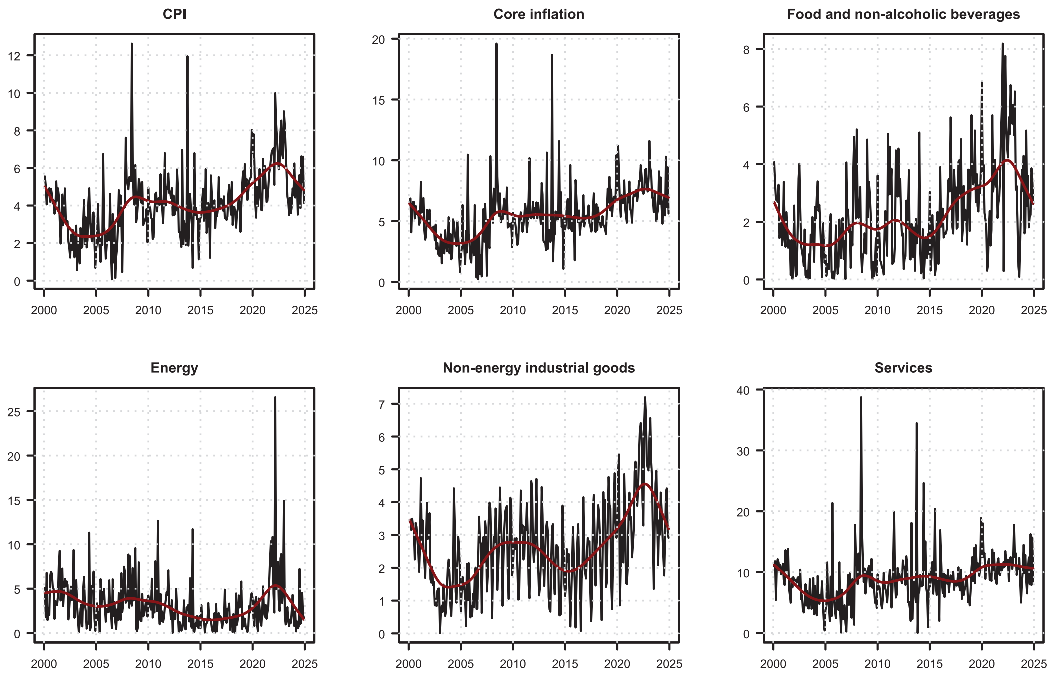
Note: The figure shows the frequency of price increase for all goods and services as well as at the sectoral level. Red line denotes trend extracted using the Hodrick-Prescott filter. Median estimates or estimates for prices excluding sales are qualitatively and quantitatively similar. They are available upon request.

Figure C7 Frequency of price decrease (in percent)



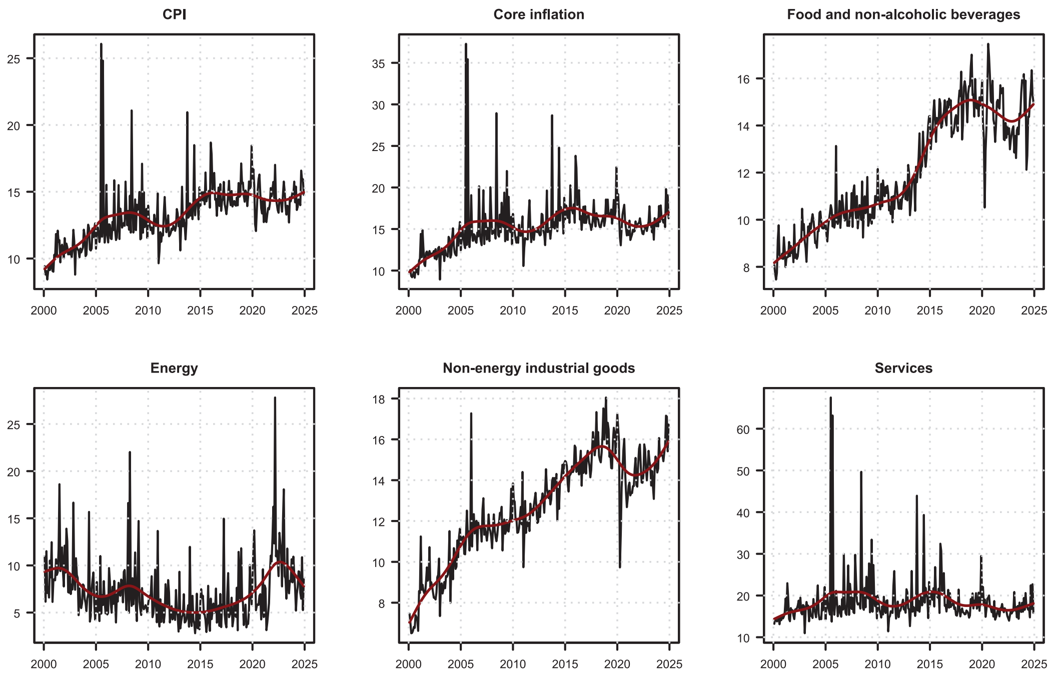
Note: The figure shows the frequency of price decrease for all goods and services as well as at the sectoral level. Red line denotes trend extracted using the Hodrick-Prescott filter. Median estimates or estimates for prices excluding sales are qualitatively and quantitatively similar. They are available upon request.

Figure C8 Size of price change (in percent)



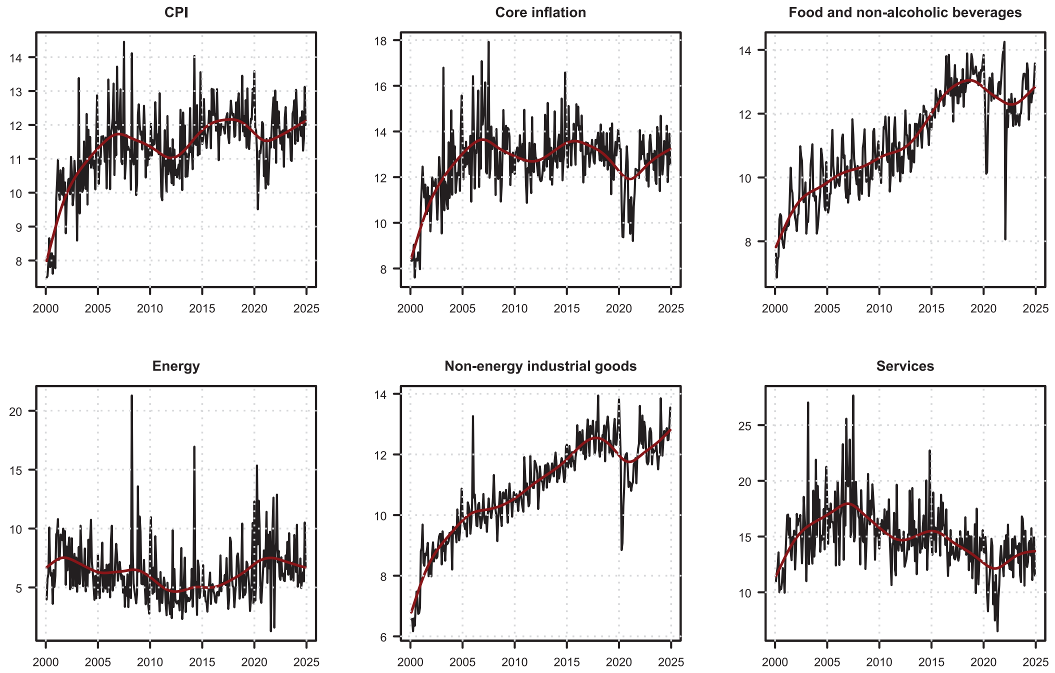
Note: The figure shows the size of price change for all goods and services as well as at the sectoral level. Red line denotes trend extracted using the Hodrick-Prescott filter. Median estimates or estimates for prices excluding sales are qualitatively and quantitatively similar. They are available upon request.

Figure C9 Size of price increase (in percent)



Note: The figure shows the size of price increase for all goods and services as well as at the sectoral level. Red line denotes trend extracted using the Hodrick-Prescott filter. Median estimates or estimates for prices excluding sales are qualitatively and quantitatively similar. They are available upon request.

Figure C10 Size of price decrease (in percent)



Note: The figure shows the size of price decrease for all goods and services as well as at the sectoral level. Red line denotes trend extracted using the Hodrick-Prescott filter. Median estimates or estimates for prices excluding sales are qualitatively and quantitatively similar. They are available upon request.

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