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The Future in Today's Prices: Evidence from a Survey of U.S. Firms

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Abstract:
Do firms adjust prices to realized costs, expected costs, or both? We address this question using a new survey of U.S. businesses that separately measures realized cost changes since the last price adjustment and expected cost changes over the subsequent year, including portions attributable to 2025 trade policies. Using perceived tariff exposure as an instrument, we identify the causal effects of realized and expected costs on prices. Reset prices incorporate almost 70 percent of current costs and nearly 45 percent of expected costs over the next year. The importance of these channels varies significantly across firms. Frequent price adjusters respond mainly to current costs, while sticky-price firms weight expectations more heavily. Goods producers adjust contemporaneously, whereas service firms are more forward looking, as are firms with a high labor share or facing high trade uncertainty. This evidence favors endogenous pricing frameworks in which uncertainty reshapes the reset-price kernel across horizons or imperfect-information models in which uncertainty amplifies the role of expectations over standard time-dependent models.

JEL Classifications: E31, C26, F14
Keywords: Survey, small and medium-sized businesses, price setting, realized cost, pass-through, expected cost

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1 Introduction

A central feature of modern macroeconomic models is that firms set prices in a forward-looking manner. In New Keynesian and related frameworks, inflation dynamics depend on both current marginal costs and firms' expectations about future costs. At the firm level, the optimal reset price reflects a weighted average of expected marginal costs over the price's expected duration. The following reduced-form cross-sectional regression captures this forward-looking logic:

$$\Delta p_i = \alpha + \beta \Delta C_i + \gamma \Delta C_i^e + u_i, \tag{1}$$

where the observed log price change of firm i, Δp_i, responds to the firm's realized log cost change since the last price adjustment, ΔC_i, and to the future log cost changes that are expected during the period when the price may remain unchanged, ΔC_i^e. The residual u_i captures other shocks that are not directly related to costs.

This paper provides new micro-level evidence on how firms adjust prices in response to *both* realized and expected cost changes. The evidence comes from a new quarterly survey of U.S. small and medium-sized businesses (SMBs), with about 500 to 600 firms per quarter, using data from the 2025:Q3 wave. The survey is designed to map directly onto equation (1). It elicits three measures, each captured at the time of price adjustment: (1) the firm's realized price change, (2) the contemporaneous cost change since the previous adjustment, and (3) the expected cost change over the subsequent year. These objects are rarely observed jointly in standard firm-level data sets. Moreover, by focusing on SMBs rather than large multiproduct firms, our empirical setting is closer to the single-product benchmark underlying much of the theoretical pricing literature.[1]

Using these data, we estimate the causal effects of realized and expected cost changes on price adjustments. Our identification strategy exploits unexpected trade-policy changes as a source of cross-sectional variation. We use variation generated by heterogeneity in firms' exposure to tariff changes announced in early 2025 to construct instruments for realized and expected cost changes—heterogeneity that may be related to import intensity, sector, age, and supply-chain structure. Specifically, we directly elicit the portion of realized and expected cost changes attributable to trade policy and thus obtain distinct instruments for contemporaneous and expected cost movements, allowing us to isolate their independent effects on prices.

Our baseline estimates indicate that both channels play an economically meaningful role. We estimate the implied pass-through from realized cost changes to prices to be 0.68. In addition,

[1] While our survey uses a distinct question format and focuses specifically on firm-level outcomes, it is influenced by other recent surveys of U.S. firms such as the Survey of Firms' Inflation Expectations (see Federal Reserve Bank of Cleveland, 2026; Candia et al., 2024) or the Business Inflation Expectations (BIE) survey (see Federal Reserve Bank of Atlanta, 2026). See Section D of the Online Appendix for a detailed comparison of our survey with the BIE.

expected future cost increases are partially incorporated at the time of price adjustment, with an estimated coefficient of 0.43. Firms' reset prices thus reflect both realized cost movements and anticipated future costs.

Yet results are quite nuanced: The importance of current and expected costs varies significantly across firms and economic environments. Frequent adjusters display largely contemporaneous pricing, that is, nearly full pass-through of realized costs, with little sensitivity to expected future costs. By contrast, infrequent adjusters place substantially more weight on expected costs, with a muted response to realized cost changes. We also find systematic sectoral differences: Goods-producing firms exhibit near-complete contemporaneous pass-through and limited anticipatory pricing, whereas service-sector firms are significantly more forward looking, loading on both realized and expected cost changes. This result is not simply an artifact of differential price stickiness between sectors. Even within the set of firms with sticky prices, service firms remain more forward looking than goods producers. Cost structure provides another important source of heterogeneity: Firms with high labor cost shares respond to expected cost changes, whereas firms with low labor cost shares respond primarily to realized costs.

We also document that increased uncertainty shifts pricing toward expectations. Using subjective uncertainty inferred from firms' tariff-related cost expectations, we find that firms facing higher uncertainty respond more strongly to expected future costs. This is consistent with more forward-looking pricing when the cost outlook is uncertain.

Finally, we assess the importance of cost expectations following the sufficient-statistics approach of Werning (2022). This provides independent validation of our results. In time-dependent pricing models, a single statistic summarizes the effect of anticipated shocks on current inflation: the ratio of the average ongoing price-spell duration to the average price duration. We examine the average behavior of reset prices across groups of firms that have higher and lower values of this statistic.[2] Firms with higher values of this statistic exhibit systematically greater pass-through of expected future costs into current prices. This monotonic relationship confirms the theory's predictions and reinforces our interpretation: Cross-firm differences in the loading on expected costs reflect the timing structure of price adjustment, not merely mechanical correlations between realized and expected cost changes.

We interpret the heterogeneity in our results through a generalized-hazard framework of price setting. In this framework, a firm's reset price reflects a weighted average of expected marginal costs during periods when the price is likely to remain unchanged. These weights are summarized by the reset-price kernel, which depends on the price's survival profile—or, equivalently, the full path of adjustment hazards. Changes in pricing frictions shift this survival profile by altering adjustment timing, thus reallocating kernel weight toward nearer or more distant horizons. Differences in adjustment frictions thus translate directly into differences in

[2] To assign firms to these groups, we compute the same ratio at the firm level; in our data, the ratio is close to orthogonal to average price stickiness and sectoral classification.

the horizons over which future costs matter when firms reset prices.

This framework rationalizes the heterogeneity we document. Firms that adjust prices frequently place most weight on current costs, while firms that adjust prices infrequently load more heavily on expected future costs, consistent with longer expected price durations and a more back-loaded reset-price kernel. Uncertainty could further tilt pricing toward expectations through the option value of waiting, which tends to increase the duration of price spells. These findings are thus naturally interpreted through menu-cost frameworks, in which differences in frictions and uncertainty map onto differences in the reset-price kernel. This contrasts with fixed-horizon Taylor pricing or Calvo models with invariant hazard.

A complementary interpretation operates through the information structure. When firms form expectations based on signals combining current and future information, certain information environments shift pricing weight toward more persistent signals. This occurs when future cost pressures become relatively more important or when contemporaneous cost realizations become noisier. Firms can then assign a greater weight to signals that are more informative about persistent costs and their future path.

Our findings have direct monetary policy implications, especially given the current environment of elevated uncertainty and multiple adverse cost shocks. Specifically, cost shocks affect inflation through two channels: realized costs and expected future costs incorporated at the time of price adjustment. Consequently, inflation can respond immediately to changes in expected costs, even before realization.

A simple back-of-the-envelope calculation illustrates the magnitude. Our estimates suggest that a uniform 10 percent tariff raising marginal costs by 1 percent implies an annualized inflation increase of roughly 0.6 to 0.7 percentage points through the expectations channel alone, plus about 1 percentage point through contemporaneous costs. These effects materialize within the first quarter. Moreover, this channel is state-dependent. Under higher uncertainty, the pass-through of expected costs rises sharply, from essentially zero to about 0.8 percentage points at reset. Under higher uncertainty, inflation thus becomes increasingly driven by expectations rather than realized costs. This highlights a role for policy communication as an important stabilization tool. By shaping both the expected path and the uncertainty surrounding future costs, policymakers can influence price setting and inflation dynamics in real time.

Related literature: This paper contributes to three strands of literature. The first studies the role of expectations in inflation and price setting. A large macro tradition estimates forward-looking Phillips curves and the link between inflation and expected marginal costs using aggregate data (see Galí and Gertler, 1999; Sbordone, 2002) and evaluates identification and specification challenges in that approach (see Mavroeidis et al., 2014; Barnichon and Mesters, 2020). More recently, a growing micro literature uses firm surveys to study how inflation expectations map onto various firms' nominal and real decisions (see, e.g. Coibion et al., 2018;

Andrade et al., 2022; Boneva et al., 2019; Coibion et al., 2019; Abberger et al., 2024). Several recent papers use survey data to study how firms form expectations about their own production costs (Meyer and Sheng, 2025; Minton and Monnery, 2026) and how changes in these costs affect their pricing decisions (Gödl-Hanisch and Menkhoff, 2023; Bunn et al., 2024; Meyer and Sheng, 2025; Fiore et al., 2025; Ghassibe et al., 2025; Pfäuti and Wu, 2026). These studies typically relate current prices to current costs, or expected prices to expected costs over a common horizon.[3] Fiori and Scoccianti (2025) also use survey data to show that firms' uncertainty about future sales affects their pricing decisions.

We contribute to this literature by combining direct firm-level measures of *both* current and expected cost changes at the time of price adjustment, together with a trade-policy-based identification. This allows us to separately quantify contemporaneous and anticipatory pass-through of costs to prices. Our estimates therefore isolate adjustments of current prices in response to expected future costs and, in doing so, help discipline the pricing mechanisms embedded in forward-looking models. Moreover, we also show that firms' uncertainty about future trade policy makes them react more preemptively to future costs.

A second strand of the literature studies the pass-through of cost shocks into prices, often exploiting large realized shocks or discrete policy events in micro price data. Prominent examples include exchange rate movements (e.g. Gopinath et al., 2010), tax changes (e.g. Karadi and Reiff, 2019), currency changeover (e.g. Hobijn et al., 2006), and tariffs (e.g. Amiti et al., 2019; Fajgelbaum et al., 2020; Cavallo et al., 2021, 2025; Gopinath and Neiman, 2026). We complement this work by using trade policy to identify not only pass-through from realized cost changes but also anticipatory pass-through from expected future costs measured directly at the firm level, building on recent work that exploits survey-based measures of cost expectations (see Meyer and Sheng, 2025; Ghassibe et al., 2025).

Our analysis also contributes to the ongoing debate on the magnitude of pass-through of tariffs to consumer prices. Available estimates vary widely: Using border prices, Amiti et al. (2019) find near-complete pass-through of tariffs into prices paid by U.S. importers, while studies based on firm-product-level retail prices tend to find lower pass-through (see Cavallo et al., 2021, 2025). Our estimates lie between these benchmarks and are notably higher than those from large-retailer data, despite also relying on firm-level information. First, a key distinction is that we measure pass-through conditional on price adjustment, whereas much of the retail microdata literature estimates unconditional pass-through, averaging over adjusting and non-adjusting prices. This alone mechanically attenuates retail estimates relative to ours. Second, we estimate pass-through of overall cost changes based on direct firm-level measures rather than pass-through of changes in statutory or effective tariffs. Since imports are only a fraction of firms' total costs, pass-through of changes in tariffs is smaller than overall cost pass-through.

[3] Ghassibe et al. (2025) examine how expected cost changes affect current prices.

Third, our estimates based on firm-level variation are less affected by firms spreading cost shocks across prices of many of their products, a practice that can understate pass-through estimated from firm-product-level variation.[4] Finally, distribution margins also matter: As Burstein et al. (2003) emphasizes, large local-input-intensive distribution costs create a wedge between border costs and final consumer prices.

The third strand of literature develops generalized-hazard and endogenous-adjustment models, including menu-cost frameworks, that discipline how price sensitivity is distributed across horizons (see Alvarez et al., 2022; Werning, 2022). In these models, pricing frictions shape the reset-price kernel and thus the relative weight placed on current versus expected future marginal costs. We contribute to this literature by articulating the resulting predictions for cost pass-through and bring them to the data, exploiting heterogeneity in adjustment frequency, sector, and uncertainty to assess which pricing frameworks best organize the empirical patterns we observe.

The remainder of the paper proceeds as follows. Section 2 describes the survey data and measurement of prices, costs, expectations, and trade exposure. Section 3 develops a generalized-hazard pricing framework and presents our empirical strategy, and Section 4 presents empirical results. Section 5 interprets results through specific pricing models and explains heterogeneity by adjustment frequency and uncertainty and discusses aggregate implications. Section 6 concludes.

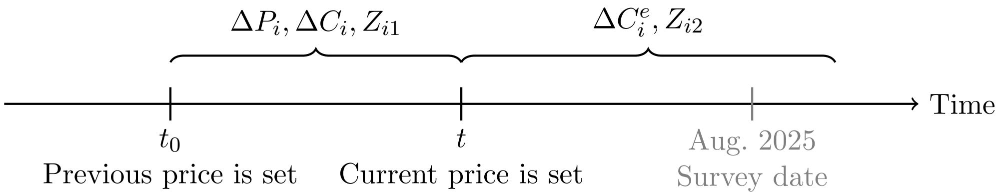
2 Survey Data

We use a new quarterly survey of U.S. small and medium-sized businesses designed to measure firm-level outcomes and expectations among decisionmakers. The survey, administered by Morning Consult, was first fielded in 2022:Q2 and became a permanent quarterly survey in 2024:Q4. Our analysis primarily uses the 2025:Q3 wave, which was fielded between August 20 and August 29, 2025.

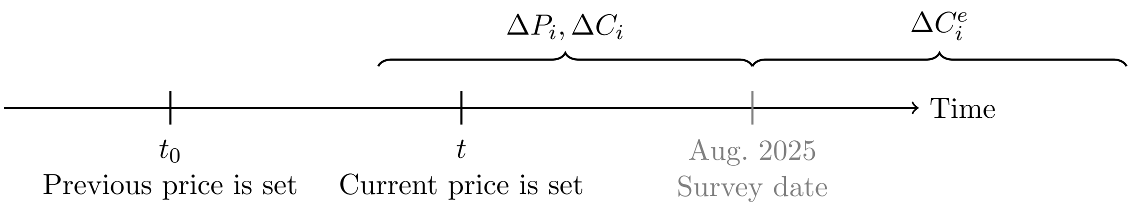
Survey respondents are repeated cross-sections of about 500 to 600 owners and other small and medium-sized business decisionmakers recruited through vetted panel providers, with multiple checks to ensure data quality and to limit inattentive or automated responses.[5] Sampling is stratified by firm size, with equal quotas for sole proprietors, firms with two to 10 employees, 11 to 100 employees, and 101 to 500 employees. We reweight observations to match the national small and medium-sized business population along the region, industry, and firm-size dimensions using iterative proportional fitting.

[4] This practice can be a factor in the spillovers of tariffs to prices of domestically produced goods documented by Cavallo et al. (2025).

[5] For more information on Morning Consult's data collection methodology, see "Morning Consult: Our Methodology," https://pro.morningconsult.com/our-methodology, accessed March 6, 2025, as well as Hajdini et al. (2024).



(a) The Morning Consult SMB Survey, August 2025 Wave



(b) Existing Surveys

Figure 1: Timeline of Price and Cost Changes

After reweighting, the sample aligns well with national benchmarks. In the weighted data, sectoral composition and firm characteristics are broadly consistent with external benchmarks: The most common sectors include retail trade, professional/scientific/technical services, construction, and healthcare and social assistance. The median firm has been operating for more than seven years. The distribution of revenues is comparable to census aggregates, with a modest tilt toward lower revenue ranges, consistent with a small and medium-sized business-focused frame (see Figure A.1 in Section A of the Online Appendix). Demographic characteristics and role characteristics of respondents are reported in Table A.1 of Section A of the Online Appendix; most respondents are C-level decisionmakers, and the survey records race, gender, education, and political affiliation.

A key innovation of our survey is the timing of measurement. The main variables for this paper measure pricing and cost dynamics at the time of firms' price-setting decisions rather than at the survey date. This difference in timing is central to identifying firms' responses to realized and expected cost changes, as we describe in the next section. Figure 1 illustrates this distinction: Panel (a) shows the timeline for our survey, while Panel (b) shows the typical timeline in existing surveys. In our design, when a firm reports its most recent price change, it also reports the cost changes that occurred between the previous price change and this most recent one, as well as the cost expectations it held at the time of the most recent price change.

For the most recent price change, firm i reports the size and direction of the price adjustment ΔP_i. change in costs since the previous price change. Contemporaneous cost changes are reported both in total ΔC_i and separately for the portion attributable to trade policy, which we denote as $Z_{1i} \equiv \Delta(C_i \mid \text{trade policy})$. This decomposition allows us to distinguish trade-related cost pressures from other cost drivers.

The survey also elicits forward-looking beliefs held at the time of price adjustment. Firms report expected total cost changes over the subsequent 12 months, ΔC_i^e, including average, downside, and upside scenarios, and separately report the trade-policy-driven component of those expected cost changes, $Z_{2i} \equiv \Delta(C_i^e \mid \text{trade policy})$. Firms also report expected price changes over the same horizon.

Finally, the survey measures exposure to trade policy directly by asking firms which tariff or trade-policy categories affect their costs. Firms also report the frequency of price changes and the time since the last adjustment. Section C of the Online Appendix contains the main survey questions.

A potential concern arises with our survey design: Because firms report historical values at the time of price changes rather than at the survey date, one might wonder about recall accuracy. Section E of the Online Appendix investigates this possibility in detail and shows that survey responses are quite similar across questions requiring different degrees of recall, suggesting minimal measurement error from this source. That is, we find no evidence that our responses suffer from more recall or approximation error compared with more typical surveys that require no recall.

Another common concern with all surveys is measurement error due to inattentive respondents who do not answer questions to the best of their abilities. To address this concern, we exclude respondents who select "stay the same" across the full block of cost and price expectation questions spanning multiple horizons and categories (for example, unit cost or price developments as well as specific costs across 11 categories like borrowing and credit, labor, utilities, etc.). This filter is designed to remove inattentive response patterns rather than economically meaningful zero changes.

3 Theoretical Framework for Estimation

In this section, we first discuss a general theoretical pricing framework that links price setting to realized and expected cost changes. We then discuss in detail our empirical specification and identification strategy.

3.1 Price Setting and Expected Costs

To derive a theoretical relation between firms' price-setting behavior and both realized and expected cost changes, we build on a general class of linear-quadratic price-setting models such as Alvarez et al. (2022) and Werning (2022). Assume an environment in which firms face idiosyncratic restrictions on their ability to reset prices. Time, denoted by t, is discrete. A firm that resets its price in period t chooses a (log) reset price p_t^*. Between resets, the price remains fixed. We assume that the firm's frictionless (log) target price $p_t^\dagger$, net of time-invariant markups, equals its (log) real marginal cost, mc_t, such that

$$p_t^\dagger = mc_t. \tag{2}$$

Let $\tau > 0$ denote the (random) time until the next reset. The distribution of τ is summarized by its survival function in equation (3):

$$S_j \equiv \Pr(\tau > j \mid t), \qquad j \geq 0. \tag{3}$$

Here, $S_j \in (0,1)$ represents the probability that a price chosen in t remains effective j periods ahead. For now, we assume that the sequence $\{S_j\}$ is given.[6]

A second-order approximation to the firm profit function yields a quadratic objective defined by deviations from the optimal target price $p_t^\dagger$.[7] Future profits are discounted at the firm's discount factor $\beta \in (0,1)$. Conditional on being able to reset its price in period t, the firm thus minimizes the expected discounted squared deviations between its chosen price p_t^* and its desired target prices over the random lifetime of the price set in t:

$$p_t^* = \arg\min_{p_t} \mathbb{E}_t \left[\sum_{k=0}^{\tau-1} \beta^k \left(p_t - p_{t+k}^\dagger \right)^2 \right]. \tag{4}$$

Note that $\mathbb{E}_t(\tau - 1)$ is the expected duration of a price set in t until the next reset. The first-order condition equates the optimal reset price in period t to a discounted average of expected target prices $\mathbb{E}_t p_{t+k}^\dagger$ over the price's random duration:

$$p_t^* = \frac{\mathbb{E}_t \left[\sum_{k=0}^{\tau-1} \beta^k \, p_{t+k}^\dagger \right]}{\mathbb{E}_t \left[\sum_{k=0}^{\tau-1} \beta^k \right]} = \sum_{k \geq 0} \frac{\beta^k S_k}{\sum_{k \geq 0} \beta^k S_k} \mathbb{E}_t p_{t+k}^\dagger. \tag{5}$$

The weight of each future period $t+k$ is a function of the normalized, discounted survival probability of prices into period $t+k$. Since $p_{t+k}^\dagger = mc_{t+k}$, this relationship immediately implies a linear mapping between expected marginal costs at each horizon and the reset price. Proposition 1 summarizes the sensitivity of the optimal reset price in t to expectations over real marginal costs in period $t+k$.

Proposition 1 (General weight formula)**.** *For any exogenous reset process summarized by* $\{S_j\}_{j \geq 0}$, *the sensitivity of the optimal reset price to expected marginal costs at horizon* j *is*

[6] The sequence might itself be implied by a state-dependent or age-dependent hazard rate of adjustment.

[7] See Section F.1 of the Online Appendix for a detailed derivation of the second-order approximation and equation 4.

given by

$$\frac{\partial p_t^*}{\partial\, \mathbb{E}_t mc_{t+j}} \;=\; w_j \;\equiv\; \frac{\beta^j\, S_j}{\sum_{k\geq 0}\beta^k S_k}, \qquad \forall j \geq 0, \tag{6}$$

with $\sum_{j=0}^{\infty} w_j = 1$. *The weight* w_j *assigned to expected marginal costs* j *periods ahead equals their discounted survival probability* $\beta^j\, S_j$, *normalized by the expected discounted lifetime of the price,* $\sum_{k\geq 0}\beta^k S_k$.

See Appendix F.1 for the proof. ■

Equation (6) in Proposition 1 highlights the central intuition: A price set in period t will still be relevant in future periods only if it survives until then, so expected future marginal costs are weighted by the joint effect of discounting and survival. The discount factor β^k captures the standard intertemporal weighting of future periods, while the survival probability S_k adjusts for the probability that the firm's price is still active then. Weights over all periods are normalized such that they sum to unity.

Our survey questions elicit changes in firms' reset prices as well as costs and expected costs. Equation (5) can be used to derive a general mapping between these variables. Using the accounting identity $\mathbb{E}_t mc_{t+j} = mc_t + \mathbb{E}_t \Delta mc_{t,t+j}$, equation (5), and $\sum_{j\geq 0} w_j = 1$, it then follows that

$$p_t^* = \sum_{j=0}^{\infty} w_j \big(mc_t + \mathbb{E}_t \Delta mc_{t,t+j}\big) = mc_t + \sum_{j=1}^{\infty} w_j\, \mathbb{E}_t \Delta mc_{t,t+j}, \tag{7}$$

where $\Delta x_{t,s} = x_s - x_t$. Taking first differences between the current and previous price changes at times t and t_0 (as shown in Figure 1) yields

$$\Delta p^*_{t_0,t} = \Delta mc_{t_0,t} + \mathbb{E}_t\left[\sum_{j\geq 1} w_j\, \Delta mc_{t,t+j}\right] - \mathbb{E}_{t_0}\left[\sum_{j\geq 1} w_j\, \Delta mc_{t_0,t_0+j}\right]. \tag{8}$$

As equations (7) and (8) make clear, today's pricing decisions depend on both realized costs or realized cost changes since the last adjustment and the change in a function of expected future costs. Because we have assumed constant markups, the pass-through of realized cost changes at the point of price reset t should be equal to unity and should not depend on, and hence be uninformative of, any parameters of the model. At the same time, pass-through of future expected cost changes between period t and $t + j$ depends in a more complicated way on our assumptions about the underlying marginal cost processes.

3.2 Empirical Strategy and Identification

We next outline the further assumptions that allow the preceding model-based equation to map exactly onto our survey-based estimating specification and how we causally identify the pass-through parameters.

Mapping between theory and empirical specification: To map expression (8) onto our estimating equation, we make two additional assumptions. First, we define normalized weights over all future horizons by

$$\tilde{w}_j \equiv \frac{w_j}{1-w_0}, \qquad j \geq 1$$

so that $\sum_{j\geq 1} \tilde{w}_j = 1$. We then assume that the survey measure of expected cost changes over the next 12 months, $\Delta C_t^e = \mathbb{E}_t[mc_{t+12}] - mc_t$, provides a measure of the model's normalized weighted expected path of marginal costs relative to the time of the previous price change, that is,

$$\Delta C_t^e \approx \sum_{j\geq 1}^{\infty} \tilde{w}_j \mathbb{E}_t[\Delta mc_{t,t+j}]. \tag{9}$$

It follows that

$$\mathbb{E}_t\left[\sum_{j\geq 1}^{\infty} w_j \Delta mc_{t,t+j}\right] = (1-w_0)\Delta C_t^e.$$

Second, we assume that the expectation term formed at t_0, the time of the last price change,

$$R_{t_0} \equiv \mathbb{E}_{t_0}\left[\sum_{j\geq 1}^{\infty} w_j \Delta mc_{t_0,t_0+j}\right],$$

does not covary systematically with current realized and expected cost changes at time t, that is,

$$\mathrm{Cov}(R_{t_0}, \Delta C_t) = 0, \qquad \mathrm{Cov}(R_{t_0}, \Delta C_t^e) = 0.$$

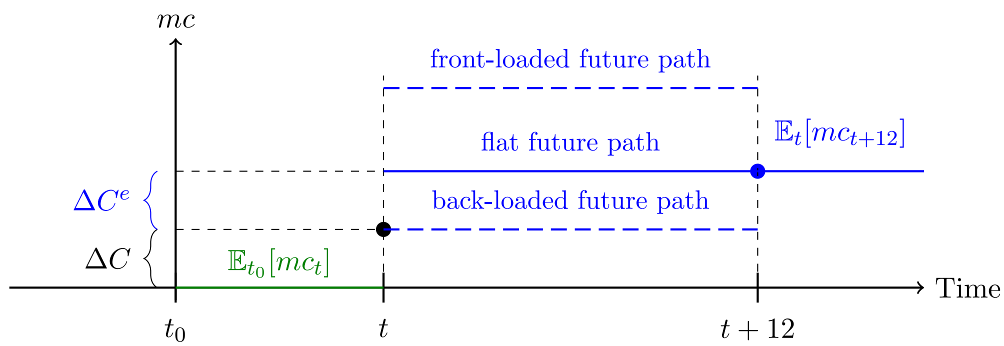
In other words, the expectation term formed at t_0 does not contain information about cost changes realized or expected at time t. When this condition is satisfied, the model implies a reduced-form equation,

$$\Delta p_t = \alpha + \beta \Delta C_t + \gamma \Delta C_t^e + u_t, \tag{10}$$

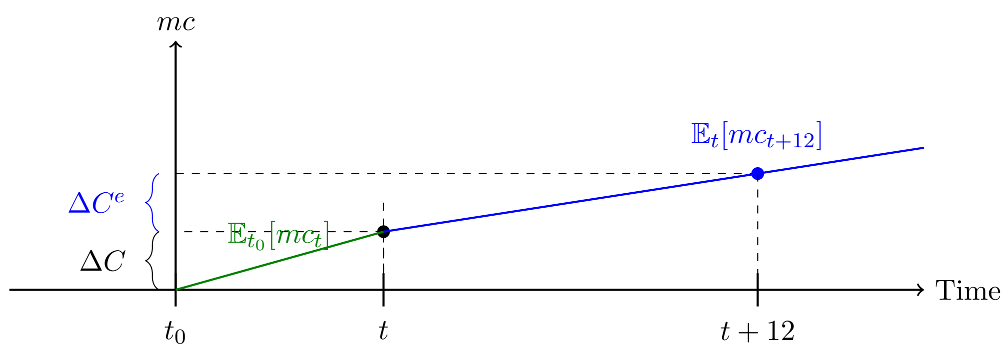
with u_t constituting an error term that satisfies the usual exogeneity condition and benchmark coefficients $\beta = 1$ and $\gamma = 1 - w_0$. This relationship is our main estimating equation, and it maps directly onto the objects the survey measures, as described in the previous survey data section.

Potential identification issues: If the preceding orthogonality condition is violated, our benchmark regression would produce biased coefficient estimates. Figure 2 shows different configurations for the path of expected marginal costs under which this may happen.[8]

[8] We thank Javier Cravino for suggesting this schematic figure.



(a) Future costs: Scenario 1



(b) Future costs: Scenario 2

Figure 2: Future Marginal Cost Paths

Panel (a) illustrates how our measure of expected cost change can reflect distinct expected cost paths. Three prototypical paths are shown for the period between the most recent price change and 12 months into the future: (1) a front-loaded future path; (2) a flat future path (our baseline assumption); and (3) a back-loaded path. From time $t + 12$ on, the expected marginal cost decreases (Case 1), stays the same (Case 2), or increases (Case 3). In Cases 1 and 3, the measure of cost change ΔC_t^e is negatively or positively correlated with the costs expected down the path (Cases 1 and 3). As these expected costs also affect pricing decisions in t, the error term u_t becomes correlated with ΔC_t^e, and estimates of the reduced-form coefficient γ are biased, with $\gamma < 1 - \omega_0$ in Case 1 (and conversely in Case 3).

Panel (b) highlights that the expectations formed in t_0 may already incorporate information about future realized cost changes. In that case, variations in ΔC_t and ΔC_t^e partially reflect changes that were already anticipated in t_0 rather than purely unanticipated cost movements. So, the reduced-form parameters β and γ no longer isolate the effect of surprise changes in current costs and expected future costs, respectively.

Of course, the paths shown are only illustrative, and many such types of paths with nonzero correlations are possible. Both types of cases can arise under policy uncertainty, as was prevalent during our sample time frame. However, our "flat" baseline case is likely a realistic one when we identify the exogenous part of marginal costs and their expectations from trade-policy shocks, as we describe later. Trade-policy changes tend to be persistent, suggesting "flat" expectations like our baseline case. For example, as Figure A.4 in Section A of the Online Appendix illustrates, the 2018 tariff increases on salient imports emphasized in Fajgelbaum et al. (2020)'s analysis remained unchanged for several years and through the next administration.

Identification strategy: We introduce an instrumental variables (IV) strategy that takes advantage of our survey design to isolate exogenous variations in current and expected cost changes. As described in Section 2, we observe firm-level measures of price changes, ΔP_i, realized and expected cost changes, ΔC_i and ΔC_i^e, as well as the components of these changes due to trade policy. This gives us two instruments: Z_{1i}, the trade-policy-induced component of realized cost changes, and Z_{2i}, the trade-policy-induced component of expected cost changes that can be used as instruments for ΔC_i and ΔC_i^e. These instruments satisfy the relevance condition by definition.

As seen on the left side of Figure 3, industries in which the greatest proportion of firms indicate tariff impacts on costs correspond well to those with large import shares. There are also large fractions of firms reporting cost impacts in industries, like healthcare, that have been found in the literature to have large indirect exposures through the production network (see Barbiero and Stein 2025). The right side of the figure shows that industries with more indirect exposures, like many services industries, tend to report larger tariff-related future cost changes than realized cost changes at the time of firms' most recent price changes. As we show in the empirical results section, our instruments correlate strongly with both ΔC_i and ΔC_i^e, likely aided by this cross-sectional variation in actual versus expected future tariff-related costs.

It is also reasonable to assume that firm-specific exposure to trade policy affects price changes only through its impact on current and expected costs, that is, that these instruments satisfy an exclusion restriction: Conditional on controls, they are orthogonal to other firm-specific determinants of price setting, u_i, so that $\mathbb{E}[Z_{1i}u_i] = 0$ and $\mathbb{E}[Z_{2i}u_i] = 0$. Importantly, our IV strategy directly addresses the anticipation problem illustrated in Panel (b) of the preceding Figure 2: To the extent that the 2025 trade-policy changes embodied in the instruments were not forecastable at the firm level at t_0—a reasonable assumption in our setting given uncertainty about the timing, magnitude, and incidence of the 2025 policy changes, as well as heterogeneity in firms' exposure—variation in Z_{1i} captures realized cost changes that are orthogonal to firms' prior information sets.

The first-stage projection of ΔC_i on Z_{1i} and Z_{2i} therefore strips out the anticipatory component embedded in ΔC_i, leaving only exogenous, unanticipated variation in current costs. Analogously, Z_{2i} isolates exogenous variation in expected cost changes that is tied to the policy shock rather than to firms' endogenous beliefs. As a result, the joint IV procedure restores a clean separation between unanticipated current cost movements and policy-driven expected cost

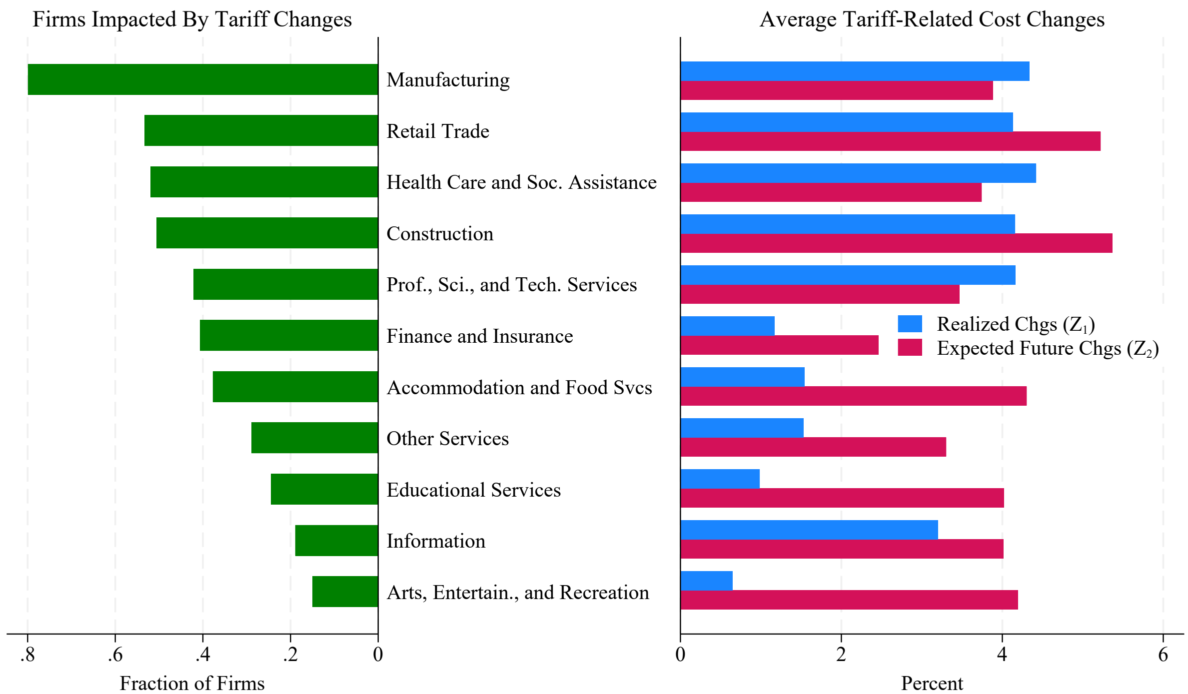


Figure 3: Tariff Exposure and Cost Changes by Industry

Note: The left side of this figure depicts the fraction of firms that indicate having or expecting meaningful tariff-related impacts on costs. The right side shows average realized cost changes and expected future cost changes as of the most recent price change that are directly related to tariffs. The statistics are reported by industry for those with at least 25 respondents.

changes, allowing β and γ to recover their structural interpretations.

Our IV strategy also mitigates other sources of endogeneity. First, salience of trade policy can help reduce classical measurement error in survey-based expectations that otherwise may attenuate coefficient estimates. This likely occurs because price setters pay close attention to the effects of tariffs on costs. Second, estimates of the coefficients on ΔC_t and ΔC_t^e may suffer from omitted variable bias if they are driven by shocks that also directly affect firms' pricing decisions. We assume that, to a first order, trade policy does not systematically shift such omitted variables, such as markups, or that any such effects are limited relative to the direct cost channel captured in the survey. If anything, a plausible concern is that firms partially absorb large tariff-induced cost increases by compressing markups—for example, to preserve market share—implying a negative correlation between cost shocks and markups. To the extent that such a channel is omitted, it would bias our estimates of pass-through downward, suggesting that our estimates should be interpreted as a lower bound on the true pass-through of costs into prices.

Table 1: Price Changes and (Expected) Cost Changes

	(1)	(2)	(3)	(4)
	OLS	IV	IV First Stages	
	ΔP_i	ΔP_i	ΔC_i	ΔC_i^e
ΔC_i	0.598***	0.679***		
	(0.051)	(0.137)		
ΔC_i^e	0.253***	0.427***		
	(0.077)	(0.150)		
Z_{1i}			0.592***	0.024
			(0.088)	(0.067)
Z_{2i}			-0.075	0.588***
			(0.098)	(0.089)
Constant	4.158***	3.063***	2.564***	3.145***
	(0.599)	(1.079)	(0.893)	(0.520)
N	537	539	539	539
R^2	0.37	0.33	0.15	0.27
SW F-stat			55.53	53.46
KP F-stat		25.61		

Note: Columns (1) and (2) contain OLS and IV regressions of price changes on actual contemporaneous cost changes and expected future cost changes, as in equation (10). The instruments used are Z_{1i} and Z_{2i}, described in Section 2. Columns (3) and (4) are the first-stage regressions from the IV. Both OLS and IV estimates are obtained using jackknife procedures that exclude observations that are overly influential to our main coefficients of interest. Robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

4 Results

This section presents our main empirical results. First, using firm-level exposure to the 2025 trade-policy episode as an instrument, we find economically large pass-through from both realized cost changes and expected cost changes: In the full sample, we obtain a coefficient of $\widehat{\beta} \approx 0.68$ for the pass-through of realized cost changes to reset prices and a coefficient of $\widehat{\gamma} \approx 0.43$ for the pass-through of expected cost changes to reset prices. Second, the relative importance of the two channels is highly heterogeneous: Frequent adjusters respond almost entirely to contemporaneous costs, whereas infrequent adjusters load much more on expected costs as do firms with a high labor share. Third, comparing firms with high and low residual trade-related uncertainty reveals an additional reallocation across horizons, with high-uncertainty firms pricing expected cost changes much more strongly than low-uncertainty firms. Finally, consistency with the sufficient-statistics approach of Werning (2022) reinforces our interpretation that the loading on expectations reflects a structural role for anticipated future costs in price setting.

4.1 Estimating Causal Effects

Our benchmark estimation of equation (10), presented in Table 1, reveals that actual price changes have an OLS coefficient on contemporaneous cost changes of about 0.6 and on expected

Table 2: Univariate Regressions of Price Changes on (Expected) Cost Changes

	(1)	(2)	(3)	(4)	(5)	(6)
	OLS		IV		IV First Stages	
	ΔP_i	ΔP_i	ΔP_i	ΔP_i	ΔC_i	ΔC_i^e
ΔC_i	0.679***		0.874***			
	(0.075)		(0.104)			
ΔC_i^e		0.558***		0.764***		
		(0.081)		(0.120)		
Z_{1i}					0.649***	
					(0.090)	
Z_{2i}						0.631***
						(0.087)
Constant	5.432***	5.247***	4.442***	4.038***	2.512***	2.873***
	(0.635)	(0.748)	(0.923)	(0.889)	(0.715)	(0.510)
N	538	539	543	531	543	531
R^2	0.19	0.10	0.22	0.16	0.21	0.27
SW F-stat					51.66	52.50
KP F-stat			51.66	52.50		

Note: Columns (1) and (2) ((3) and (4)) contain OLS (IV) univariate regressions of price changes on either actual contemporaneous cost changes or expected future cost changes. The instruments used are Z_{1i} and Z_{2i}, described in Section 2. Columns (5) and (6) are the first-stage regressions from the IV. Both OLS and IV estimates are obtained using jackknife procedures that exclude observations that are overly influential to our main coefficients of interest. Robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

cost changes over the next year of 0.25. When we instrument with variables constructed from (expected) cost changes due to trade costs, the coefficient on contemporaneous cost changes grows to 0.68 and the coefficient on expected cost changes to 0.43. Both instruments are strong, with high first-stage Sanderson-Windmeijer F-stats and an overall Kleibergen-Paap rk Wald F-stat that well exceeds the rule-of-thumb threshold of 10.

Note that in Table 2, univariate regressions of price changes on either actual or expected cost changes tend to overstate the pass-through of both actual contemporaneous and expected future cost changes. This reflects the identification problem of the positive correlation between cost changes and expectations.

4.2 Heterogeneity of Cost Pass-through

In this subsection, we document substantial cross-firm heterogeneity in the pass-through of realized and expected costs into prices.

We start with some basic facts surrounding sectoral heterogeneity. First, we show that firms with stickier prices set prices with more emphasis on expected future cost movements than contemporaneous ones. Appendix Figure A.3 displays the distribution of average price durations across firms in our sample. We split this distribution at the sample median duration of six months. Column (1) of Table 3 reports the results of our estimation but now includes only SMBs that indicate a price duration of six months or less. We find that only contemporaneous costs matter and, in both cases, have a one-to-one relationship with prices. Column (2) limits the estimation to firms that report typically changing prices every seven or more months. Indeed, the reduced-form relationships indicate that prices for these firms move more with expected future cost movements and less with contemporaneous cost changes. As we show later, this tendency of firms with stickier prices to have a higher loading of price changes on further-ahead expected cost changes is a property shared across several common models of price setting.

Table 3: Heterogeneous Pass-through by Price Duration and Sector

	(1)	(2)	(3)	(4)
	Avg. Price Duration		Sector	
	$\leq$ 6 months	$\geq$ 7 months	Goods	Services
	ΔP_i	ΔP_i	ΔP_i	ΔP_i
ΔC_i	1.017***	0.340*	0.734***	0.438***
	(0.211)	(0.206)	(0.207)	(0.169)
ΔC_i^e	0.254	0.504**	0.122	0.449***
	(0.207)	(0.198)	(0.267)	(0.136)
Constant	2.528*	4.318***	5.119***	3.653***
	(1.386)	(1.460)	(1.523)	(1.046)
N	278	259	110	427
R^2	0.26	0.47	0.31	0.32
SW F-stat ΔC_i	26.12	31.12	16.15	34.62
SW F-stat ΔC_i^e	24.73	34.71	7.60	53.61
KP F-stat	12.46	14.67	4.95	17.17
P-val: Equal γ	0.38		0.27	

Note: This table contains IV regressions (second stage) of price changes on actual contemporaneous cost changes and expected future cost changes, as in equation (10). The instruments used are Z_{1i} and Z_{2i}, described in Section 2. Columns (1) and (2) refer to a sample split of firms with average price durations six months or less versus seven months or more (all results are shown in Appendix Table A.5). Columns (3) and (4) refer to a sample split between goods- and services-producing firms (all results are shown in Appendix Table A.6). IV estimates are obtained using jackknife procedures that exclude observations that are overly influential to our main coefficients of interest. Robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

One may be concerned that these results for firms with different average price durations are particularly sensitive to our assumption that the 12-month cost change expectation is a good proxy of firms' expected path of marginal costs. That is, referring to equation (8), firms with lower price durations likely have $\tilde{w}_j$ closer to zero for $j \geq 12$, so assumption (9) may hold more closely for these firms than for firms with greater price durations. Thus, as a robustness check, we also estimate Column (2) of Table 3, keeping only those firms that also have typical price durations of 12 months or less. Despite the substantially smaller sample of 164 firms compared with 259, the estimated coefficients are little changed and not statistically different: The coefficient on ΔC_i is 0.50 (p-value of 0.045), and the coefficient on ΔC_i^e is 0.43 (p-value of

0.127).

Next, we examine differences between goods- and services-producing firms by estimating our main equation within each of these groups of firms, as shown in Columns (3) and (4) in Table 3, respectively. We document that the pass-through from expected cost changes into prices is larger for services-producing firms. Importantly, we note that in this sector, heterogeneity does not simply reflect a difference due to services (goods) firms typically having higher (lower) price durations. When we focus only on firms with average price durations of seven or more months, services firms within that group show a large and statistically significant coefficient on expected cost changes, while prices of goods firms do not show a statistically significant response to expected cost changes.

To further investigate these sectoral differences, we examine whether services and goods firms differ along a number of dimensions. The greatest difference emerges regarding their cost structures. Specifically, an average of 44 percent of services firms' costs are spent on labor compared with 29 percent for goods firms.[9]

Table 4 estimates our main equation, splitting the sample by firms with lower- or higher-than-median labor cost shares. Those with high labor costs tend to adjust prices in response to changes in cost expectations, while lower labor cost firms tend to adjust only in response to realized cost changes. Notably, this is true even among only services firms, showing that our finding in Table 3 that services firms preemptively adjust prices for expected cost changes is primarily driven by the firms within the services sector that have the highest labor cost shares. One potential explanation for this pattern is that firms less able to adjust their costs when needed, particularly those with large shares of downwardly rigid costs like wages, may face greater losses from being away from their frictionless target prices and may, therefore, be more forward looking in their pricing decisions.

We next present one of our central findings. Table 5 presents results showing whether pass-through differs in degree according to some additional channels. First, in Columns (1) and (2), we split firms by residual trade-related uncertainty (below versus above the median). This residual uncertainty is the dispersion of trade-attributed cost expectations after controlling for the cross-sectional relationship with realized trade-related costs and the mean expected trade-cost change. The estimates show a sharp reallocation across channels: Low-uncertainty firms (Column 1) exhibit close to one-for-one pass-through of contemporaneous costs and essentially no response to expected costs, whereas high-uncertainty firms (Column 2) load strongly on expected cost changes and much less on contemporaneous costs. The difference in the response to expected cost changes is statistically significant with a p-value of 0.02 across these two groups.

[9] Appendix Table A.3 shows that firms in these two sectors are largely similar in firm size, age, customer relationships, operating geographies, and international trade activity. We also find that services have slightly longer average price durations but do not express more difficulty in changing prices than goods firms (see Appendix Table A.4).

Table 4: Heterogeneous Pass-through by Labor Cost Share

	(1)	(2)	(3)	(4)
	All Firms		Service Firms	
	Low labor costs	High labor costs	Low labor costs	High labor costs
	ΔP_i	ΔP_i	ΔP_i	ΔP_i
ΔC_i	0.922***	0.484**	1.032**	0.365
	(0.258)	(0.225)	(0.417)	(0.376)
ΔC_i^e	0.085	0.435**	0.009	0.636**
	(0.215)	(0.209)	(0.318)	(0.249)
Constant	4.454***	3.347***	4.482**	2.593
	(1.658)	(1.224)	(1.998)	(2.106)
N	247	255	183	214
R^2	0.27	0.47	0.25	0.44
SW F-stat ΔC_i	22.84	26.77	12.37	15.87
SW F-stat ΔC_i^e	28.88	25.66	19.52	18.51
KP F-stat	10.87	13.04	6.46	7.04
P-val: Equal γ	0.24		0.12	

Note: This table contains IV regressions (second stage) of price changes on actual contemporaneous cost changes and expected future cost changes, as in equation (10). The instruments used are Z_{1i} and Z_{2i}, described in Section 2. Columns (1) and (2) refer to a sample split of firms with low (below full-sample median) or high (above full-sample median) labor costs. Columns (3) and (4) refer to the same low versus high labor cost split but only among services firms. The sample in this table only includes firms that report itemized costs over the previous fiscal year. To address concerns about reporting error, we also exclude firms that report costs exceeding two times their total revenue. IV estimates are obtained using jackknife procedures that exclude observations that are overly influential to our main coefficients of interest. Robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

This pattern is consistent with state-dependent pricing models in which uncertainty reshapes the effective reset-price kernel across horizons, and it is harder to reconcile with exogenous-hazard models in which uncertainty does not reallocate the timing weights.

Lastly, we estimate a sample split informed by Werning (2022). This paper derives a sufficient statistic summarizing the pass-through of a shock to expectations into current inflation that is valid for any time-dependent model. Under certain conditions, the statistic is the ratio of the average duration of ongoing price spells to the average price duration. Werning (2022) derives that, among a sample of Calvo-pricing firms, this ratio would be 1.0, while among a sample of Taylor-pricing firms, the ratio would be 0.5. In a menu-cost model (specifically, the Sheshinski-Weiss model) in which frequency of price adjustment is fixed in the very short run, this ratio would be between zero and 0.5. The precise form of pass-through derived in Werning (2022) is that of a shock to overall inflation expectations, holding constant expectations of future real marginal costs while we examine the pass-through of expected nominal marginal costs that contains both types of shocks.

To assess whether we find estimates consistent with this theoretical result, we divide our

Table 5: Heterogeneous Pass-through by Uncertainty and Duration Ratio

	(1)	(2)	(3)	(4)
	Trade Cost Uncertainty		Ongoing-Completed Duration Ratio	
	Low	High	$< .75$	$\geq .75$
	ΔP_i	ΔP_i	ΔP_i	ΔP_i
ΔC_i	0.949***	0.252	0.673**	0.576***
	(0.163)	(0.423)	(0.337)	(0.178)
ΔC_i^e	-0.036	0.799**	-0.093	0.592***
	(0.179)	(0.315)	(0.320)	(0.181)
Constant	4.139***	2.979	5.935**	3.356***
	(1.225)	(1.824)	(2.417)	(1.095)
N	327	205	180	323
R^2	0.29	0.12	0.18	0.36
SW F-stat ΔC_i	34.50	11.49	18.90	25.27
SW F-stat ΔC_i^e	21.43	21.73	13.83	55.75
KP F-stat	11.33	5.65	7.97	12.59
P-val: Equal γ	0.02		0.06	

Note: This table contains IV regressions (second stage) of price changes on actual contemporaneous cost changes and expected future cost changes, as in equation (10). The instruments used are Z_{1i} and Z_{2i}, described in Section 2. Columns (1) and (2) split our sample between firms with below- or above-median residual trade-cost uncertainty (all results appear in Appendix Table A.7). Columns (3) and (4) split our sample between firms whose time since their last price change as a ratio of their average price duration is below or above 0.75 (all results are shown in Appendix Table A.8). IV estimates are obtained using jackknife procedures that exclude observations that are overly influential to our main coefficients of interest. Robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

sample into two subgroups with low versus high ratios of average ongoing price-spell duration to average price duration using proxies for this ratio at the firm level. We compute these ratios using the time since the most recent price change and the elicited average duration of price spells, also excluding firms that report a (denominator) average price-change duration of zero. We split the sample at 0.75, the midpoint of the Calvo and Taylor benchmarks of 1.0 and 0.5, respectively. The ratios of average ongoing to average completed durations are 0.38 and 1.15 within each of these groups. Werning (2022) derives that these ratios should reflect the pass-through of expectations into total inflation, which itself is increasing in the pass-through into reset prices within these models, so we would expect our γ estimate to be higher in the high-ratio group.

Column (3) shows that the low ongoing–completed ratio group reacts only to contemporaneous cost changes, not expected future cost changes, while the high ongoing–completed ratio group reacts to both contemporaneous and expected future cost changes. The difference in the response to expected cost changes is marginally statistically significant with a p-value of 0.06 across these two groups. This result is notably not an artifact of this sufficient statistic reflecting other properties of these firms; within our sample, the firm-level ongoing–completed ratio is unrelated to average price duration or the firm's sector.

5 Pricing Models

To interpret our empirical evidence, we discuss in this section three leading pricing models—Calvo pricing, Taylor pricing, and a menu-cost model—within the generalized linear-quadratic firm problem described in Subsection 3.1. Rather than providing a fully specified quantitative analysis of our empirical results, we focus on qualitative insights derived from these models. We complement this discussion by considering how imperfect information environments can account for the role of uncertainty and the pass-through of expected future costs. Finally, we discuss policy implications of our results.

5.1 Price Rigidity and Specific Hazard Structures

Our main result in Table 1 is a coefficient on cost-change expectations significantly above zero. Here, we show that this result aligns with price-setting models in the presence of pricing rigidities. We begin with benchmark environments in which the timing of price resets is exogenous (that is, Taylor and Calvo pricing) and then consider a generalized-hazard menu-cost model in which the adjustment hazard depends on deviations from desired prices.

First, under Taylor pricing (see Taylor, 1980), a firm can reset its price on a fixed schedule every N periods. There is no uncertainty around the timing of price resets and, thus, the probability of price survival. Proposition 2 solves for the pricing weights on expected future marginal costs, as defined in Proposition 1. It follows from our empirical evidence of $\hat{\gamma} > 0$ that prices are not fully flexible and $N > 1$.

Proposition 2 (Taylor contracts)**.** *In the Taylor pricing model, prices are fixed for N periods. It follows that $S_j = 1$ for $0 \leq j \leq N-1$ and $S_j = 0$ thereafter. Substituting into equation (6) yields*

$$w_j = \begin{cases} \frac{(1-\beta)\,\beta^j}{1-\beta^N} & \text{if } 0 \leq j \leq N-1, \\ 0, & \text{if } j \geq N. \end{cases} \tag{11}$$

Each horizon up to $N-1$ enters with a constant geometric decay due to discounting. The limiting cases $N = 1$ correspond to flexible prices, while $N \to \infty$ approximates full price rigidity.

See Appendix F.1 for the proof. ■

Second, in the Calvo pricing model (see Calvo, 1983), the timing of price resets is random, with the probability of not being able to reset prices in every period equal to θ. Proposition 3 describes the weights defined in Proposition 1. Again, $\hat{\gamma} > 0$ suggests that price rigidities are present, implying that $\theta > 0$.

Proposition 3 (Calvo pricing)**.** *In the Calvo pricing model, the probability of not resetting the price in any given period is constant and equal to $\theta \in (0,1)$. It follows that $S_j = \theta^j$. Substituting into equation (6) yields*

$$w_j = (1-\beta\theta)(\beta\theta)^j \qquad \forall j \geq 0.$$

Weights decline geometrically with horizon j. The degree of price rigidity, governed by θ, determines how strongly the firm internalizes expected future marginal costs. The limiting cases $\theta = 0$ correspond to flexible prices, while $\theta \to 1$ approximates full price rigidity.

See Appendix F.1 for the proof. ■

Third, following the generalized-hazard approach of Alvarez et al. (2022), we consider a menu-cost model represented by a generalized adjustment hazard $\lambda(x_t)$ and defined over $x_t \equiv p_t - p_t^{\dagger}$, the price gap between the posted price and the frictionless benchmark. In this class of models, the probability of price adjustment is a function of the price gap, that is, it is state dependent. We assume that this hazard function is symmetric around zero and increasing in $|x|$; the price-adjustment hazard rises in the size of the price gap. The hazard function is locally linear in $|x|$ for small and medium-sized gaps, that is, $\lambda(x) \approx \lambda_1 |x|$ for $\lambda_1 > 0$. We also assume the price gap x_t evolves as a Gaussian random walk, $x_{k+1} = x_k + \varepsilon_{k+1}$, with normally distributed errors $\varepsilon_k \sim \mathcal{N}(0, \sigma^2)$.

To connect this state-dependent environment to the reset-price kernel w_j, we use a first-order mean-field approximation in which the random, state-dependent adjustment hazard at each horizon is replaced by its expectation. This yields a horizon-dependent adjustment probability governed by the expected size of the price gap at each horizon and, therefore, a non-Calvo reset-price kernel in which the weight placed on future marginal costs reflects the state-dependent nature of adjustment without solving the full path-dependent problem.

Proposition 4 (Menu cost)**.** *Suppose price adjustment is summarized by a generalized hazard $\lambda(x)$ that is symmetric, increasing in $|x|$, and locally linear, $\lambda(x) \approx \lambda_1 |x|$. Let the price gap follow a Gaussian random walk. Then the reset-price kernel implied by the linear–quadratic pricing problem takes a stretched-exponential form,*

$$w_j = \frac{\beta^j e^{-Aj^{3/2}}}{\sum_{k \geq 0} \beta^k e^{-Ak^{3/2}}}, \qquad A = \frac{2}{3} \lambda_1 \sigma \sqrt{\frac{2}{\pi}},$$

where σ is the standard deviation of innovations to the price gap x_t.

See Appendix F.1 for the proof. ■

5.2 The Role of Price Rigidity

We next discuss the role of price rigidity under each of the models described earlier. Table A.5 suggests that the pass-through of expectations into prices is higher for firms with more rigid pricing, $\hat{\gamma}^{rigid} > \hat{\gamma}^{flex}$, a pattern that maps directly onto the model's implication that greater rigidity can shift weight toward longer horizons.

Corollary 1 shows the ambiguous effect of price rigidity on weights in the Taylor and Calvo pricing models.

Corollary 1 (Exogenous pricing models)**.** *Under Taylor and Calvo pricing, the sensitivity weights on expected marginal costs $w_j(N)$ satisfy*

$$\frac{\partial w_j^{Taylor}(N)}{\partial N}\begin{cases}\leq 0 & \forall 0 \leq j \leq N-1\\ \geq 0 & \forall j > N-1\end{cases} \qquad \frac{\partial w_j^{Calvo}(\theta)}{\partial \theta}\begin{cases}\leq 0 & \forall j < \frac{\beta\theta}{(1-\beta\theta)}\\ \geq 0 & \forall j \geq \frac{\beta\theta}{(1-\beta\theta)}\end{cases}. \tag{12}$$

See Appendix F.2 for the proof. ■

The predictions in the corollary align well with the empirical results. In particular, as Table 3 shows, an increase in price rigidity increases the coefficient on expected costs: The coefficient on expected cost changes is 0.25 and not significantly different from zero for firms with relatively flexible prices. Instead, it goes up to 0.50 and is highly statistically significant for firms with relatively sticky prices. These changes align with the predictions of Corollary 1: Coefficients w_j for a period sufficiently in the future are predicted by both models to increase in price stickiness. Increases in pricing frictions shift weight away from today's cost changes toward expected cost changes. Is this likely the case under a reasonable parameterization? To compare with our empirical results where expectations are 12 months ahead, we assume that w_1 in equation (7) corresponds to the estimated coefficient. Given an annual Calvo parameter of about 0.18 (based on a quarterly parameter of 0.65) and a rate of time preference of 0.97, the cutoff for j equals 0.194, comfortably below a one-year horizon.

Corollary 2 discusses the effect of price rigidity on weights for the menu-cost model. Here, we utilize the microfoundation that rationalizes generalized hazards in Alvarez et al. (2022). In this case, random menu costs ψ are distributed according to a cumulative distribution function G, following Alvarez et al. (2022), where $G(\cdot)$ is well defined as the probability that a menu-cost draw falls below a given adjustment gain. Such adjustment occurs when a random cost draw ψ falls below adjustment gain $U(x)$. We index menu costs by a parameter κ such that higher κ corresponds to uniformly higher adjustment costs $\psi_\kappa = \kappa\psi$.

Corollary 2 (Menu-cost model)**.** *For small changes in menu costs and small price gaps for* *each $j \geq 0$,*

$$\frac{\partial w_j}{\partial \kappa} = w_j \left(\frac{\partial \log S_j}{\partial \kappa} - \sum_{m \geq 0} w_m \frac{\partial \log S_m}{\partial \kappa} \right), \tag{P}$$

and there is a cutoff $j^\star(\kappa)$ such that weights w_j decrease at short and increase at long horizons:

$$\frac{\partial w_j}{\partial \kappa} < 0 \quad \text{for } j < j^\star(\kappa), \qquad \frac{\partial w_j}{\partial \kappa} > 0 \quad \text{for } j > j^\star(\kappa).$$

See Appendix F.3 for the proof.

Intuitively, an increase in menu costs raises the threshold for price adjustment at any given price gap, making adjustment less likely and lowering the state-dependent hazard. As a result, prices are expected to remain fixed for longer, increasing the relevance of more distant horizons for optimal price setting. This change does not uniformly dampen responsiveness across horizons. Rather, higher menu costs reallocate sensitivity away from short horizons—where adjustment becomes less likely—and toward longer horizons that become increasingly relevant as expected price durations lengthen. Notably, these predictions of increased pricing frictions align with those of the Calvo and Taylor pricing models and with our empirical findings.

5.3 Uncertainty and Price Setting

A key finding from our empirical analysis is that uncertainty increases the importance of expected cost changes for today's pricing decisions. Exogenous pricing models like Calvo and Taylor do not conventionally entail such a prediction, but endogenous pricing models can, as can models that feature information frictions.[10]

As is well known, uncertainty about future marginal costs in the Calvo- and Taylor-pricing frameworks does not affect the pass-through of expected costs into reset prices, as adjustment probabilities are exogenously fixed and independent of the volatility of shocks. As a result, higher uncertainty does not affect firms' pricing decisions through the duration structure itself. The survival profile and thus the relative weights assigned to different horizons remain unchanged. This finding usually emerges from the certainty equivalence of the linear-quadratic approximation of the pricing problem underlying the weights in Proposition 1.

By contrast, our empirical finding that uncertainty matters for pricing points to state-dependent pricing models if one wants to account for the effect of future uncertainty on today's prices. In such state-dependent settings, two well-known, potentially opposing effects are generally at work. First, a volatility effect raises the hazard of adjustment and shortens the expected

[10] A separate though related mechanism is the precautionary pricing channel documented by Oh (2020); Cho et al. (2021) in higher-order solutions of the nonlinear pricing problem. Here, while firms raise prices in a precautionary manner in the face of higher uncertainty, the pass-through of expected marginal cost changes into prices is unchanged to the first order.

lifetime of the current price. Consequently, more weight is placed on costs at the immediate horizon, and the kernel is front loaded away from future horizons. Second, higher uncertainty also increases the option value of waiting and can widen the inaction bands, which tends to increase the expected duration of a price spell and to place relatively more weight on costs at future horizons.

This latter channel is emphasized in recent menu-cost models with news about future volatility, in which firms anticipate greater future dispersion before the dispersion is realized and therefore delay adjustment today (see Aruoba et al., 2024). Given these two forces, pass-through rises at a given horizon when the strength of the real option effect exceeds the rate at which the kernel shifts mass away from that horizon. For short horizons, the volatility effect may dominate, reducing weights, while for sufficiently distant horizons, the effect may reverse, increasing pass-through. Our empirical findings suggest that, in our setting, the option-value channel is strong enough to reallocate the reset-price kernel toward expected future costs.

A complementary mechanism that can rationalize this pattern operates through the information structure rather than adjustment frictions. Suppose that firms do not observe the current realization of marginal costs perfectly but instead receive signals that jointly contain information about current and future fundamentals. When these signals are relatively more informative about persistent or policy-driven future cost components than about the contemporaneous realization, firms optimally place greater weight on expected future costs in their pricing decisions, effectively reweighting the reset-price kernel toward longer horizons. Higher uncertainty about contemporaneous cost realizations becoming relatively noisier or persistent future cost components becoming more important can amplify this effect, as either force increases firms' reliance on signals that better isolate persistent future costs.

5.4 Policy Implications

A natural question concerns the policy implications of our findings for the cost-price pass-through. Our results imply a substantial inflationary impact of trade-policy shocks arising not only from contemporaneous cost increases that are passed through to consumer prices. Expected future cost changes also matter, implying that even if current costs do not adjust immediately, shifts in beliefs about future costs can generate inflationary pressures without much delay, limiting potential inflation inertia due to sluggish current cost reactions to trade-policy shocks. In addition, our findings point to an important role for policy communication as a potential policy tool: Because expected costs affect price setting and because the uncertainty surrounding those expectations shapes how strongly they are weighted, communication that influences both the expected path and the uncertainty of future costs could meaningfully affect inflation dynamics.

The impact of tariffs on prices: A back-of-the-envelope calculation suggests a sizable effect on inflation arising from expected future cost changes induced by trade-policy shocks, in addition to contemporaneous cost movements. To perform this calculation, we abstract from substitution effects, general-equilibrium responses, and potential changes in trade patterns. We consider a uniform 10 percent tariff on all imports that is fully passed through into (expected) input prices. Using a foreign value-added share of 10 percent, as in Barbiero and Stein (2025), this implies a 1 percent increase in firms' (expected) marginal costs. To map this cost shock onto inflation, we assume that consumer prices adjust gradually over time, as captured by a Calvo price-adjustment process with a quarterly parameter of 0.65.

Then, the following inflationary impact arises: Based on Table 1, if the tariff operates through changes in expected future costs, our estimates imply an annualized inflation increase of approximately 0.69 percentage points in the first month and 0.60 percentage points over the first quarter. If instead the tariff operates through current costs, price adjustment is more front loaded: The annualized inflation impact rises to about 1.09 percentage points in the first month and 0.95 percentage points over the first quarter. In a more realistic scenario of a tariff implementation, both the expected and realized costs might be affected. In that case, the total effect might be a combination of both channels discussed earlier.

Notably, the estimated inflationary effect is heterogeneous along several policy-relevant dimensions. Across sectors, goods-producing firms exhibit largely contemporaneous pass-through, with inflation effects of about 1.18 percentage points in the first month (1.03 over the first quarter) from current costs but only modest contributions from expected costs (0.20 and 0.17, respectively). By contrast, service-sector firms display a much stronger forward-looking component, with expected costs contributing about 0.72 percentage points in the first month (0.63 over the first quarter), exceeding the effect of contemporaneous costs (0.70 and 0.61).

Heterogeneity is even more pronounced across uncertainty regimes: In low-uncertainty environments, inflation is driven almost entirely by current costs (1.52 percentage points in the first month and 1.33 percentage points over the first quarter), with a relatively limited role for expectations. By contrast, in high-uncertainty environments, contemporaneous pass-through is negligible, while expected costs dominate, generating inflation increases of about 1.28 percentage points in the first month and 1.12 percentage points over the first quarter. These patterns highlight that—beyond our baseline case—both the timing and the magnitude of inflationary responses to policy shocks depend critically on sectoral composition and the degree of uncertainty surrounding future costs.

The role of communication: Our findings point directly to another important policy implication: Communication about the future has the potential to serve as a policy tool to affect inflation. This potential arises because expected costs affect price setting and because the uncertainty regime surrounding those expectations shapes how heavily expected future costs are weighted. In practice, policymakers could act on both margins in several ways—by providing clearer and more detailed forward guidance on the expected path and persistence of cost-relevant policies, by reducing ambiguity around the policy outlook, and by coordinating communication across policy institutions to limit dispersion in firms' beliefs.

This interpretation is consistent with evidence that improved information can both shift expectations and compress their dispersion (see Coibion et al., 2022) and, in particular, with our finding that the pass-through of expected costs depends on the uncertainty regime. Importantly, while the literature has often focused on how communication can affect individual beliefs about aggregate variables, particularly inflation (see Andrade et al., 2019; Coibion et al., 2022, 2023), our results suggest that what matters more is how communication affects firms' expectations about their own (nominal) costs. Taken together, these results suggest that policies that not only anchor the level of expected future costs but also reduce the uncertainty surrounding them might help attenuate the short-term inflationary impact of trade-policy shocks.

6 Conclusion

This paper provides new evidence on how firms incorporate current and expected costs into price-setting decisions and how pricing frictions shape the relative importance of these channels. Using firm-level data that directly measure realized and expected cost changes—and their trade-policy components—we document a novel empirical pattern: Prices respond both to contemporaneous costs and to anticipated future costs but in systematically different ways across firms. Frequent adjusters (and goods producers) exhibit largely contemporaneous pass-through, while sticky-price firms (and service producers, independently of pricing frictions) load more heavily on expectations, with uncertainty tilting pricing further toward forward-looking considerations.

These patterns are difficult to reconcile with price-setting models in which the adjustment hazard is exogenous and independent of firms' economic states. By contrast, they arise naturally in frameworks such as menu-cost models, in which both the timing of adjustment and the reset-price kernel are endogenous to the price gap and firms' expectations. More generally, our findings provide direct microeconomic evidence that pricing frictions affect how they set prices conditional on price adjustment. More generally, our findings provide direct microeconomic evidence that pricing frictions not only affect *when* firms adjust prices but also *how* they set prices conditional on adjustment.

One of our empirical findings in particular highlights an important puzzle and points to a promising avenue for future research: The estimated pass-through of realized marginal cost changes into prices is significantly below unity for firms with long average price durations. This finding appears at odds with the predictions of the linear-quadratic pricing model in Subsection 3.1, which implies full pass-through from realized marginal cost shocks to optimal reset prices, regardless of adjustment frequency.

Two complementary explanations could explain this incomplete pass-through. First, strategic complementarities in price setting—whereby firms' pricing decisions respond not only to their own marginal costs but also to competitors' prices—naturally attenuate adjustments to cost shocks (see Albagli et al., 2025; Alvarez et al., 2023). Under strategic complementarities, even flexible-price firms would exhibit incomplete pass-through, but the effect should be amplified among sticky-price firms that anticipate muted responses by their competitors. Second, imperfect observation of marginal costs implies that firms optimally adjust prices only partially to noisy cost signals, filtering out transitory fluctuations. Recent work by Afrouzi et al. (2024) and Gagliardone and Tielens (2026) uses survey data to disentangle the effects of nominal rigidities from those of imperfect information on pricing decisions.

Extending this approach to our survey setting—which uniquely combines measures of realized costs, expected costs, and price-setting behavior—would allow us to quantify the relative importance of strategic complementarities versus information frictions in explaining incomplete pass-through. This decomposition is crucial for understanding the aggregate implications of microeconomic pricing behavior and represents a natural direction for future research.

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A Tables and Figures

Table A.1: Owner Characteristics

	%		%
Race		**Urban vs. Rural**	
White	79.0	Urban	30.1
Black	15.0	Suburban	50.6
Other	6.1	Rural	19.3
Ethnicity		**Census Region**	
Non-Hispanic	90.8	Northeast	19.4
Hispanic	9.2	Midwest	18.5
		South	35.8
Gender		West	26.2
Male	58.1		
Female	41.9	**Role**	
		C-Level	72.7
Education		Vice President	1.5
less than college	47.2	Director	8.8
bachelor's degree	33.6	Manager	17.0
postgraduate degree	19.2		
Political Ideology			
Liberal	25.6		
Moderate	39.1		
Conservative	35.3	N=599	

Note: The table shows weighted percentages of respondents in different demographic groups and categories.

Table A.2: Industry Distribution

	Survey	Census 2023 ABS
Group 1	**32.0**	**32.2**
Professional, Scientific, and Technical Services	16.2	14.4
Finance and Insurance	5.7	4.0
Real Estate and Rental and Leasing	5.1	5.9
Management of Companies and Enterprises	2.8	0.2
Information	2.2	1.5
Administrative and Support and Waste Management and Remediation Services	0.0	6.2
Group 2	**26.0**	**24.6**
Health Care and Social Assistance	8.3	11.2
Education Services	6.6	1.7
Arts, Entertainment, and Recreation	5.6	2.3
Accommodation and Food Services	5.5	9.3
Group 3	**20.3**	**18.8**
Retail Trade	16.3	10.5
Wholesale Trade	3.9	4.4
Transportation and Warehousing	0.1	3.8
Group 4	**17.7**	**17.4**
Construction	10.8	12.8
Manufacturing	4.0	4.0
Agriculture, Forestry, Fishing and Hunting	2.2	0.3
Utilities	0.7	0.1
Mining, Quarrying, and Oil and Gas Extraction	0.0	0.2
Group 5	**4.1**	**7.0**
Other Services (except Public Administration)	4.1	7.0

Note: The table shows the distribution of firms in our sample across sectors. The survey is weighted to be representative at the industry group level.

Table A.3: Firm Characteristics, by Sector

	Services	Goods
Full-Time Employees		
1	9.4	7.8
2 to 5	49.6	51.4
6 to 10	16.9	16.5
11 to 20	11.6	12.4
21 to 50	7.5	7.9
51 to 100	2.6	1.8
101 to 250	1.6	1.5
251 to 500	0.7	0.6
Firm Age		
Less than 2 years	7.9	11.6
2-9 years	49.5	54.3
10 or more years	42.6	34.1
Main Customers		
Consumers	49.5	49.8
Other businesses	15.6	4.9
Both businesses and consumers	34.8	45.3
Operating Geographies		
Within local community	91.8	96.7
Within state	95.3	93.1
Within multiple states	48.3	54.3
Within all US states	36.8	21.7
Internationally	22.3	11.5
Intl Trade Activity		
Only exports	6.2	7.1
Only imports	8.4	7.0
Both exports and imports	16.7	14.2
Neither exports nor imports	68.6	71.6

Note: The table shows weighted percentages of respondents within services and goods firms selecting different responses (not mutually exclusive for "Operating Geographies").

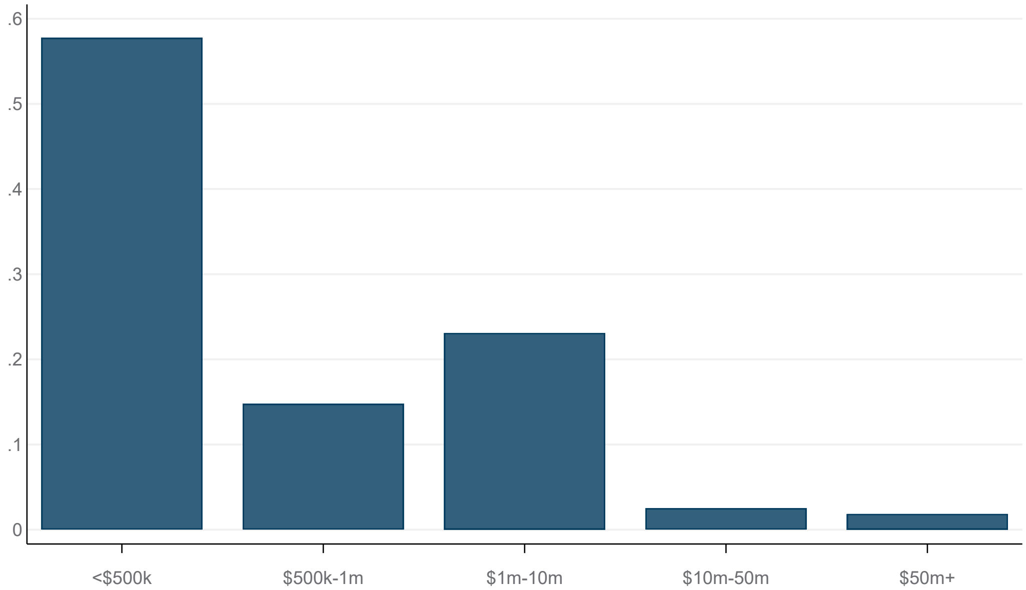
Table A.4: Pricing and Cost Characteristics, by Sector

	Services	Goods
Ease of Changing Prices		
Very easy	15.1	22.6
Somewhat easy	47.4	36.8
Somewhat difficult	25.5	24.3
Very difficult	6.2	6.7
Don't know / No opinion	5.9	9.6
Average Price Duration (Months)		
25th Percentile	4	4
Median	10	7
75th Percentile	13	13
Mean	10.9	9.8
Mean Labor Share of Costs	43.8	29.3
	(2.5)	(3.6)

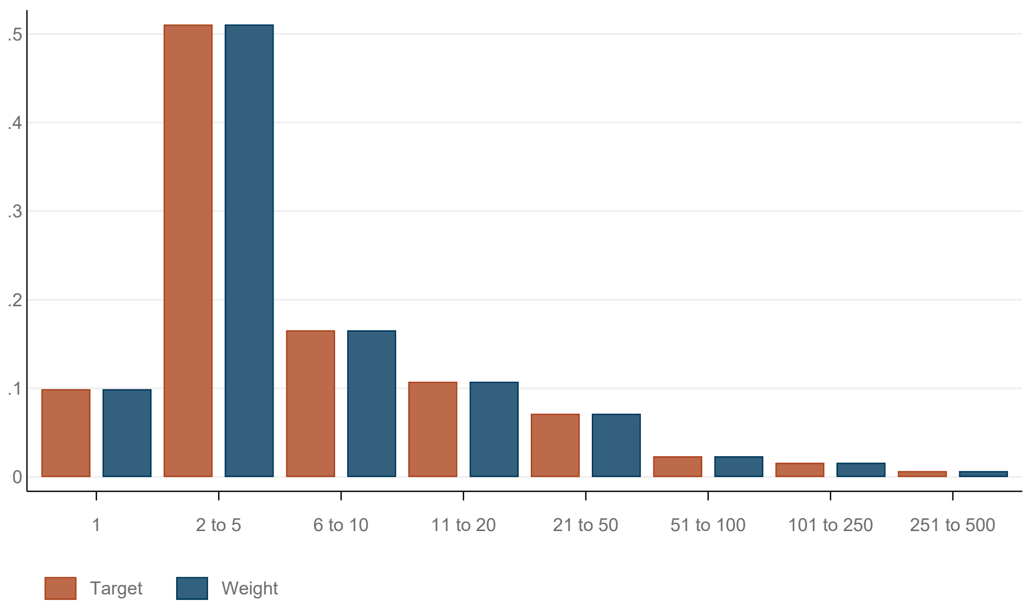
Note: The first block of this table shows the weighted percentages of respondents within services and goods firms selecting different responses regarding whether it was easy for the firm to change prices. The next block shows summary statistics for the average price duration (in months). The last block shows the mean labor share of costs along with robust standard errors in parentheses.

Figure A.1: Firm Characteristics I

(a) Revenues



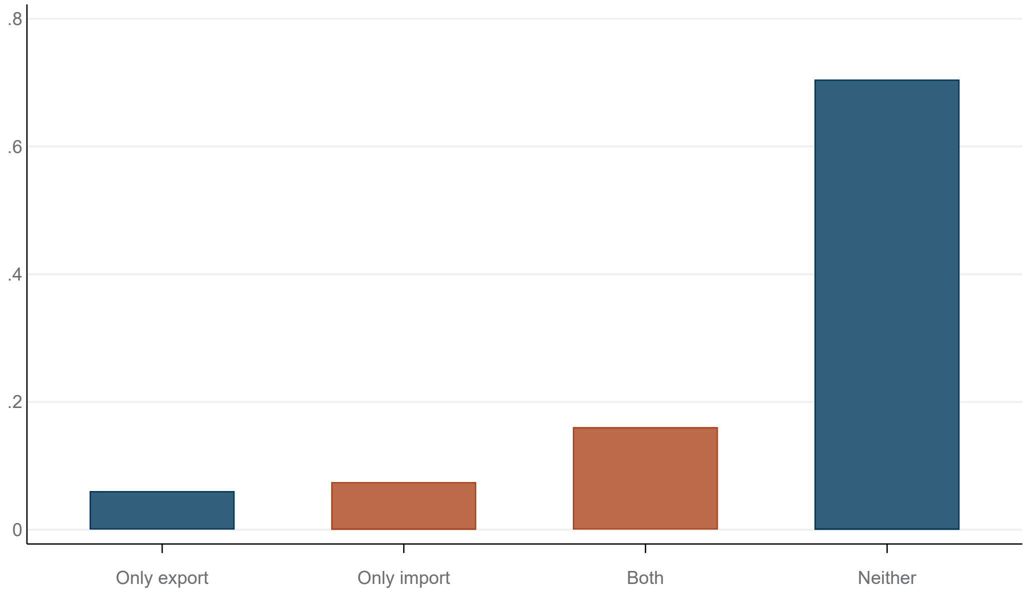
(b) Employees



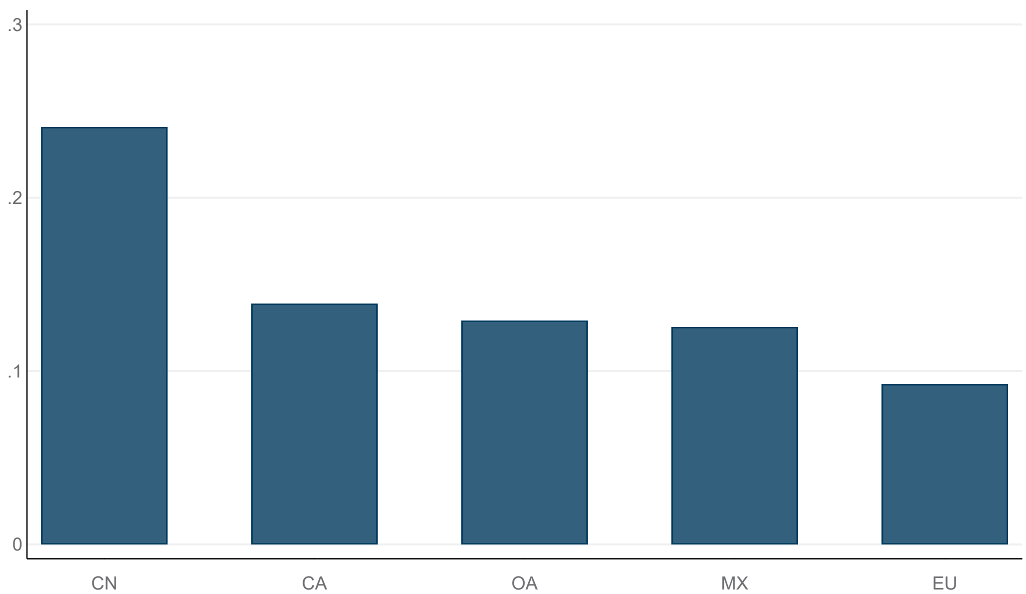
Note: The figure displays the distribution of annual revenue (Panel a) and the number of employees (Panel b) across firms in the sample.

Figure A.2: Firm Characteristics II

(a) Trade status

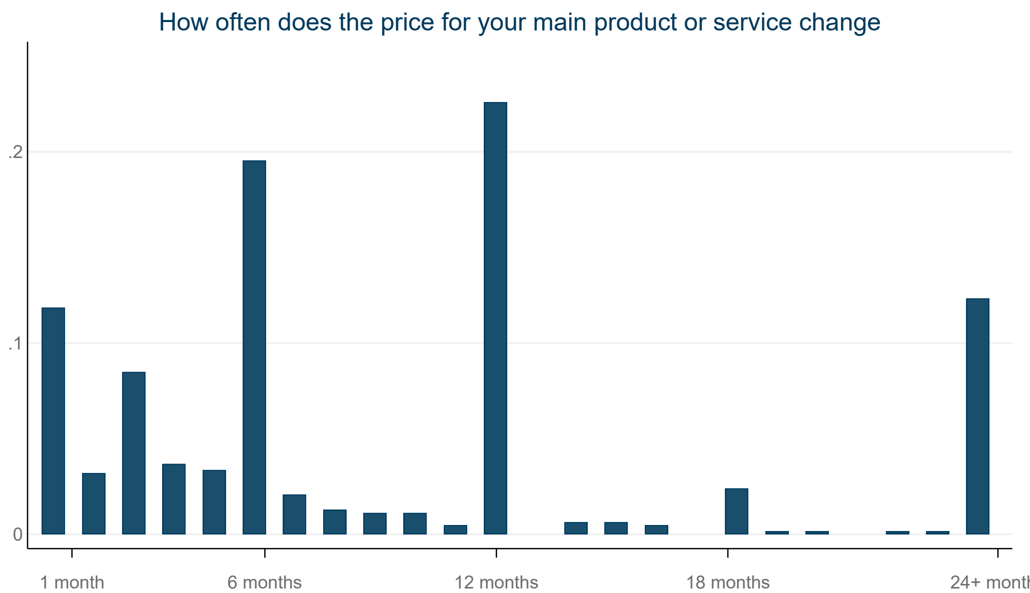


(b) Tariff impact



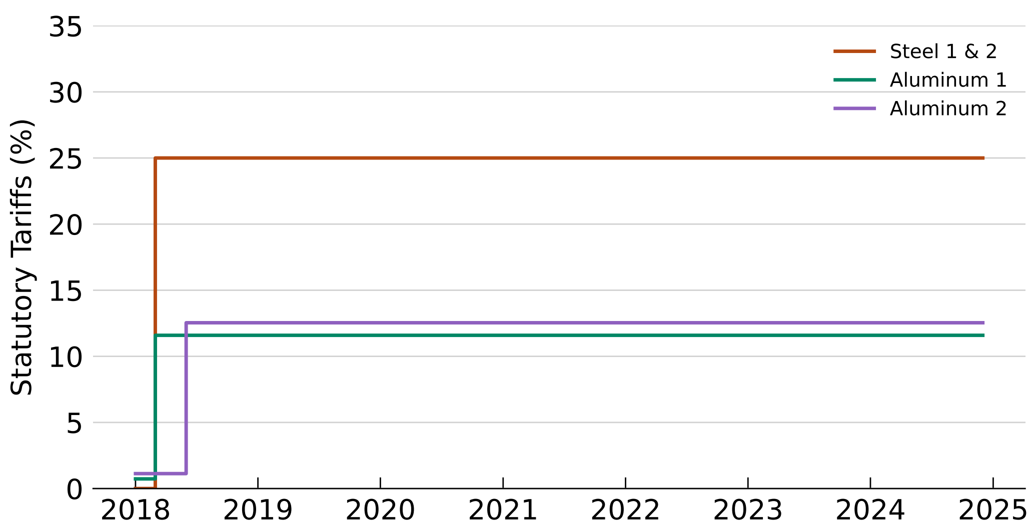
Note: The figure displays the distribution of trade status (Panel a) and impact of tariffs (Panel b) across firms in the sample.

Figure A.3: Price Duration



Note: The figure displays the distribution of average price durations across firms in the sample.

Figure A.4: U.S. Tariffs on Selected Imports (Average of Statutory Rates)



Source: Fajgelbaum et al. (2020) and authors' calculations using orders from Federal Registers.

Table A.5: Price Changes and (Expected) Cost Changes, by Average Price Duration

Panel A) Price duration of six months or less

	(1)	(2)	(3)	(4)
	OLS	IV	IV First Stages	
	ΔP_i	ΔP_i	ΔC_i	ΔC_i^e
ΔC_i	0.886***	1.017***		
	(0.096)	(0.211)		
ΔC_i^e	-0.045	0.254		
	(0.133)	(0.207)		
Z_{1i}			0.570***	0.130
			(0.097)	(0.091)
Z_{2i}			-0.001	0.602***
			(0.075)	(0.131)
Constant	4.526***	2.528*	1.586***	2.115***
	(1.042)	(1.386)	(0.505)	(0.668)
N	279	278	278	278
R^2	0.31	0.26	0.33	0.41
SW F-stat			26.12	24.73
KP F-stat		12.46		

Panel B) Price duration of seven months or more

	(1)	(2)	(3)	(4)
	OLS	IV	IV First Stages	
	ΔP_i	ΔP_i	ΔC_i	ΔC_i^e
ΔC_i	0.389***	0.340*		
	(0.129)	(0.206)		
ΔC_i^e	0.649***	0.504**		
	(0.129)	(0.198)		
Z_{1i}			0.658***	0.022
			(0.137)	(0.078)
Z_{2i}			-0.081	0.575***
			(0.199)	(0.107)
Constant	3.327***	4.318***	3.236**	3.929***
	(0.689)	(1.460)	(1.629)	(0.770)
N	257	259	259	259
R^2	0.37	0.47	0.12	0.19
SW F-stat			31.12	34.71
KP F-stat		14.67		

Note: Columns (1) and (2) contain OLS and IV regressions of price changes on actual contemporaneous cost changes and expected future cost changes, as in equation (10). The instruments used are Z_{1i} and Z_{2i}, as described in Section 2. Columns (3) and (4) are the first-stage regressions from the IV. The two panels are estimated on subsamples defined by firms' reported frequency of price changes. Both OLS and IV estimates are obtained using jackknife procedures that exclude observations that are overly influential to our main coefficients of interest. Robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A.6: Price Changes and (Expected) Cost Changes, by Sector

Panel A) Goods-producing firms	(1)	(2)	(3)	(4)
	OLS	IV	IV First Stages	
	ΔP_i	ΔP_i	ΔC_i	ΔC_i^e
ΔC_i	0.962***	0.734***		
	(0.125)	(0.207)		
ΔC_i^e	-0.214	0.122		
	(0.181)	(0.267)		
Z_{1i}			0.662***	0.071
			(0.151)	(0.089)
Z_{2i}			0.003	0.496**
			(0.120)	(0.193)
Constant	6.217***	5.119***	0.823	1.884***
	(0.934)	(1.523)	(0.857)	(0.664)
N	113	110	110	110
R^2	0.46	0.31	0.47	0.41
SW F-stat			16.15	7.60
KP F-stat		4.95		

Panel B) Services-producing firms	(1)	(2)	(3)	(4)
	OLS	IV	IV First Stages	
	ΔP_i	ΔP_i	ΔC_i	ΔC_i^e
ΔC_i	0.503***	0.438***		
	(0.105)	(0.169)		
ΔC_i^e	0.350***	0.449***		
	(0.098)	(0.136)		
Z_{1i}			0.545***	0.029
			(0.093)	(0.060)
Z_{2i}			0.004	0.698***
			(0.110)	(0.073)
Constant	3.756***	3.653***	2.696***	3.149***
	(0.676)	(1.046)	(0.999)	(0.564)
N	423	427	427	427
R^2	0.29	0.32	0.13	0.35
SW F-stat			34.62	53.61
KP F-stat		17.17		

Note: Columns (1) and (2) contain OLS and IV regressions of price changes on actual contemporaneous cost changes and expected future cost changes, as in equation (10). The instruments used are Z_{1i} and Z_{2i}, as described in Section 2. Columns (3) and (4) are the first-stage regressions from the IV. The two panels are estimated on subsamples of only goods- or services-producing firms. Both OLS and IV estimates are obtained using jackknife procedures that exclude observations that are overly influential to our main coefficients of interest. Robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table A.7: Price Changes and (Expected) Cost Changes, by Residual Trade-cost Uncertainty

Panel A) Low Residual Trade-cost Uncertainty

	(1)	(2)	(3)	(4)
	OLS	IV	IV First Stages	
	ΔP_i	ΔP_i	ΔC_i	ΔC_i^e
ΔC_i	0.657***	0.949***		
	(0.110)	(0.163)		
ΔC_i^e	0.110	-0.036		
	(0.107)	(0.179)		
Z_{1i}			0.585***	0.036
			(0.101)	(0.093)
Z_{2i}			0.033	0.526***
			(0.088)	(0.128)
Constant	4.768***	4.139***	1.659***	3.455***
	(0.800)	(1.225)	(0.475)	(0.598)
N	330	327	327	327
R^2	0.29	0.29	0.24	0.20
SW F-stat			34.50	21.43
KP F-stat		11.33		

Panel B) High Residual Trade-cost Uncertainty

	(1)	(2)	(3)	(4)
	OLS	IV	IV First Stages	
	ΔP_i	ΔP_i	ΔC_i	ΔC_i^e
ΔC_i	0.449***	0.252		
	(0.153)	(0.423)		
ΔC_i^e	0.532***	0.799**		
	(0.144)	(0.315)		
Z_{1i}			0.454***	0.072
			(0.109)	(0.093)
Z_{2i}			0.158	0.736***
			(0.134)	(0.179)
Constant	2.868***	2.979	1.463*	2.078*
	(0.694)	(1.824)	(0.795)	(1.256)
N	204	205	205	205
R^2	0.29	0.12	0.25	0.35
SW F-stat			11.49	21.73
KP F-stat		5.65		

Note: Low-uncertainty (high-uncertainty) firms in Panel A (Panel B) are below (above) the median of residual trade-related uncertainty (dispersion of trade-attributed cost expectations, residualized with respect to realized trade-related costs and the mean expected trade-cost change). Columns (1)–(2) report OLS and 2SLS estimates of (10). Columns (3)–(4) report first stages for ΔC_i and ΔC_i^e. Instruments are Z_{1i} (trade-attributed realized cost change) and Z_{2i} (trade-attributed expected cost change orthogonalized to Z_{1i}). A jackknife procedure is used as in the baseline; robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.8: Price Changes and (Expected) Cost Changes, by the Time since Last Price Change Relative to the Average Price Duration

Panel A) Firms whose ongoing-completed duration ratio of time since last price change to average price duration is below 0.75

	(1)	(2)	(3)	(4)
	OLS	IV	IV First Stages	
	ΔP_i	ΔP_i	ΔC_i	ΔC_i^e
ΔC_i	0.043	0.673**		
	(0.123)	(0.337)		
ΔC_i^e	0.516***	-0.093		
	(0.162)	(0.320)		
Z_{1i}			0.407***	-0.006
			(0.132)	(0.162)
Z_{2i}			0.109	0.560**
			(0.166)	(0.224)
Constant	5.218***	5.935**	2.770***	4.312***
	(1.485)	(2.417)	(1.020)	(1.324)
N	180	180	180	180
R^2	0.17	0.18	0.14	0.15
SW F-stat			18.90	13.83
KP F-stat		7.97		

Panel B) Firms whose ongoing-completed duration ratio of time since last price change to average price duration is at or above 0.75

	(1)	(2)	(3)	(4)
	OLS	IV	IV First Stages	
	ΔP_i	ΔP_i	ΔC_i	ΔC_i^e
ΔC_i	0.727***	0.576***		
	(0.119)	(0.178)		
ΔC_i^e	0.212*	0.592***		
	(0.125)	(0.181)		
Z_{1i}			0.648***	0.052
			(0.139)	(0.063)
Z_{2i}			-0.092	0.660***
			(0.172)	(0.077)
Constant	3.923***	3.356***	2.509*	2.550***
	(0.650)	(1.095)	(1.434)	(0.490)
N	324	323	323	323
R^2	0.35	0.36	0.13	0.33
SW F-stat			25.27	55.75
KP F-stat		12.59		

Note: Panel A (B) estimates our main regressions using the subsample of firms whose ongoing-completed duration ratio of ongoing price spells to average price duration is below (at or above) 0.75. Columns (1)–(2) report OLS and 2SLS estimates of (10). Columns (3)–(4) report first stages for ΔC_i and ΔC_i^e. Instruments are Z_{1i} (trade-attributed realized cost change) and Z_{2i} (trade-attributed expected cost change orthogonalized to Z_{1i}). A jackknife procedure is used as in the baseline; robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

B Additional estimates

B.1 Baseline estimates without using the jackknife procedure

Table B.1: Price Changes and (Expected) Cost Changes (without Jackknife)

	(1)	(2)	(3)	(4)
	OLS	IV	IV First Stages	
	ΔP_i	ΔP_i	ΔC_i	ΔC_i^e
ΔC_i	0.527***	0.795***		
	(0.137)	(0.304)		
ΔC_i^e	0.160	0.440		
	(0.236)	(0.381)		
Z_{1i}			0.654***	-0.074
			(0.119)	(0.058)
Z_{2i}			-0.074	0.644***
			(0.106)	(0.087)
Constant	5.541***	2.625	2.562***	3.575***
	(1.158)	(2.379)	(0.873)	(0.578)
N	549	549	549	549
R^2	0.20	0.11	0.27	0.28
SW F-stat			37.62	74.22
KP F-stat		28.83		

Note: Columns (1) and (2) contain OLS and IV regressions of price changes on actual contemporaneous cost changes and expected future cost changes, as in equation (10). The instruments used are Z_{1i} and Z_{2i}, as described in Section 2. Columns (3) and (4) are the first-stage regressions from the IV. Robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table B.2: Price Changes and (Expected) Cost Changes (with Additional Controls)

	(1) Avg Price Duration	(2) NAICS2 Industry	(3) Resid. Trade Cost Unc.	(4) Duration Ratio
	ΔP_i	ΔP_i	ΔP_i	ΔP_i
ΔC_i	0.650***	0.757***	0.696***	0.645***
	(0.151)	(0.121)	(0.116)	(0.149)
ΔC_i^e	0.631***	0.329**	0.300**	0.503***
	(0.233)	(0.149)	(0.125)	(0.182)
Constant	0.954	0.081	3.661***	2.465
	(1.014)	(5.586)	(0.910)	(1.776)
N	540	539	540	508
R^2	0.12	0.41	0.38	0.18
SW F-stat ΔC_i	56.59	53.78	58.39	51.02
SW F-stat ΔC_i^e	44.30	106.61	71.78	42.49
KP F-stat	18.76	26.08	29.36	22.50

Note: Each column contains IV regressions of price changes on actual contemporaneous cost changes and expected future cost changes, as in equation (10) but with additional control variables. The instruments used are Z_{1i} and Z_{2i}, as described in Section 2. In order from left to right, the additional controls are fixed effects for firm-level average price durations, NAICS2 industry fixed effects, the level of residual trade-cost uncertainty, and the ratio between the duration of ongoing price spells (that is, the time since the last price change) and completed price spells (that is, the average price duration). Estimates are obtained using a jackknife procedure that excludes observations that are overly influential to our main coefficients of interest. Robust standard errors are in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

C Main Survey Questions

This section lists the main survey questions used in the analysis.

C.1 Price Adjustment Frequency

Q1: Please consider your most important product or service, for a typical customer. On average, how often does the price for this product or service change, and when did it last change?

On average, the price changes every ...	▽ 0 months (1) ... more than 24 months (26)
The last time the price changed was ...	▽ 0 months (1) ... more than 24 months (26)

C.2 Trade Policies

Q2: Which of the following trade policies have already meaningfully increased or are expected to meaningfully increase your firm's costs? Please select all that apply.

- ◯ Tariffs on imports from China
- ◯ Tariffs on imports from Canada
- ◯ Tariffs on imports from Mexico
- ◯ Tariffs on imports from the EU
- ◯ Tariffs on imports from other Asian countries
- ◯ Tariffs on steel and aluminum
- ◯ Tariffs on autos and auto parts
- ◯ Ending of the de minimis provision (tariff exemption for low-value packages of $800 or less)
- ◯ Other tariffs or trade policies
- ◯ None of the above. My costs are not affected by trade policies.

C.3 Pricing Decisions

Q3: Please think of your main product or service and the most recent price change for it. By how much did the price change?

- ◯ Price increased by _____ %
- ◯ Price decreased by _____ %

Q4: Now, please think back to the most recent time you adjusted your price. We are interested in how your firm's costs have changed between the most recent and previous price adjustments, and how recent trade policy may have contributed to those cost changes. Think about any changes—upward, downward or none—in your firm's costs over this period, and how much of those changes were related to trade policy. Please enter your responses as percentages. Use 0 if there was no change.

	Costs increased by %	No cost change (please write 0)	Costs decreased by %
Change in costs between the most recent *price* adjustment and the previous one	○ ____	○ ____	○ ____
Change in costs between the most recent *price* adjustment and the previous one, **only due to trade policy**	○ ____	○ ____	○ ____

Q5: Now, please continue thinking back to the most recent time you adjusted your price. What were your expectations back then of how costs and prices would change in the future? Please enter your responses as percentages, using negative values for decreases.

	Average change in %	Lowest possible change in %	Highest possible change in %
As of the time you most recently adjusted your price, expected change in *costs* over the following 12 months	○ ____	○ ____	○ ____
As of the time you most recently adjusted your price, expected change in *costs* over the following 12 months, **only due to trade policy**	○ ____	○ ____	○ ____
As of the time you most recently adjusted your price, expected change in *prices* over the following 12 months	○ ____	○ ____	○ ____

D Comparison with BIE

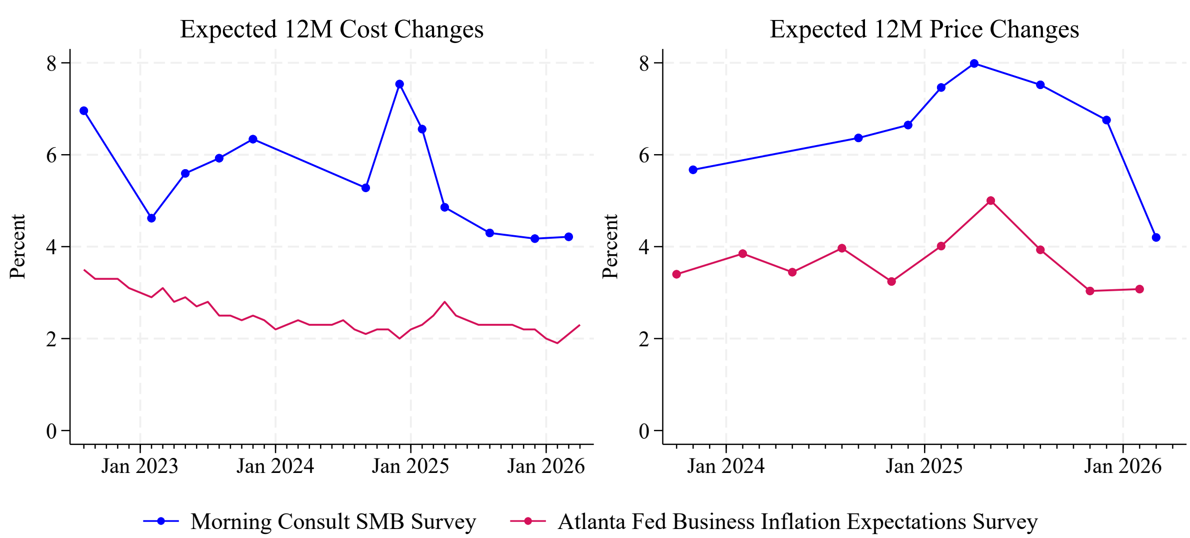
This section compares cost- and price-change expectations elicited in the Morning Consult survey of small and medium-sized businesses (MCSMB) with similar questions from the Atlanta Fed Business Inflation Expectations (BIE) survey.

Figure D.1 shows the mean expected cost and price changes over the next 12 months from both surveys. The overall movement of both variables over this sample corresponds well in the two surveys.

The levels of the variables in the BIE tend to be lower, which may be attributable to two differences in the surveys. The first is that the BIE utilizes a panel of firms that tend to stay in the survey for multiple years. By contrast, the MCSMB has a repeated cross-section design, thus containing respondents who tend not to be as experienced with surveys, the idea being that such inexperienced respondents more accurately reflect the population of firm decision-makers. Kim and Binder (2023) find that consumer inflation expectations converge to being, on average, 2.6 percentage points lower after remaining in a survey for 12 months, a quantity that is very close to the level differences we see in Figure D.1.

The second difference is that the questions eliciting cost-change expectations in the two surveys are different. While the MCSMB directly asks respondents for a percent change, the BIE asks respondents for probabilities that future cost changes will fall into a set of fixed bins. To obtain the value for the mean expectation, the distribution is truncated, with the highest (lowest) bins being coded as 6% (−2%), thus limiting the size of the changes.

Figure D.1: Comparison of Expected Cost and Price Changes with the BIE



Note: This figure compares the mean expected 12-month cost and price changes from the Morning Consult survey of small and medium-sized businesses and the Atlanta Fed Business Inflation Expectations survey.

E Recall Bias

To determine the extent of recall errors in our survey, we compare the distributions and some statistical properties of reported variables of three types: (i) those of distant adjusters, (ii) those of recent adjusters, and (iii) observations made as of the survey date. Distant adjusters are those whose most recent price adjustment was more than six months ago and hence those who had to recall the relatively more distant past. Recent adjusters are those who last adjusted within the past six months.[11] The last set of responses are ones not used in our main pass-through estimates and ask firms to report cost changes since their most recent price change and *current* expectations of cost and price changes over the next 12 months. These are typical questions asked in existing surveys and, as they require no recall, serve as a comparison data set for assessing recall errors.

Figure E.1 shows the distributions of variables recalled from the more distant versus more recent past or not recalled at all. Note that as differing degrees of recall also correspond to different points in time, we should not be surprised to see some differences in the distributions, particularly in the overall level of cost or price changes, but their overall shapes are remarkably similar.[12]

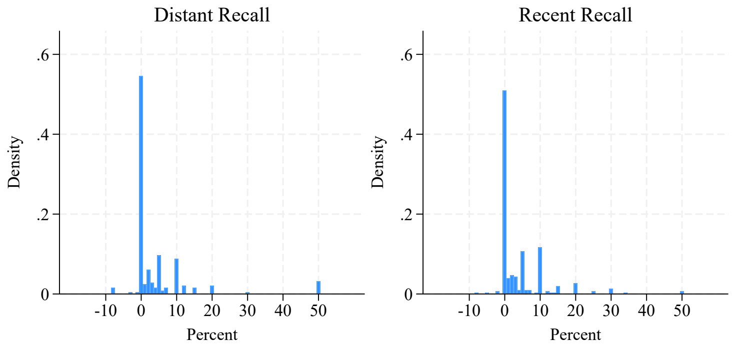
Next, we examine proxies for the degree of approximation in responses that may indicate poor recall. Table E.1 presents two different measures. The first is based on a visible feature of the distributions in Figure E.1, which is the prevalence of round-number responses.[13] Respondents who have trouble recalling past realizations or expectations may give more round-number responses. The table shows that the fraction of round-number responses is, for all cost and price variables, very similar across responses requiring distant, recent, or no recall. So, by this metric, there is no evidence of relatively more approximation in responses that required recall (whether more distant or more recent).

Another measure of approximation in the survey is the extent to which responses tend to be more similar across different variables due to poor recall. This type of recall error is particularly problematic, as it could lead to upward biases in our estimates. To assess this error, we examine the within-respondent standard deviation across all the five variables we use in our estimation $(\Delta P_i, \Delta C_i, \Delta C_i^e, Z_{1i}, Z_{2i})$. The last column of Table E.1 shows that responses are actually more disparate when more recall is required, suggesting no upward bias in our pass-through estimates due to recall issues.

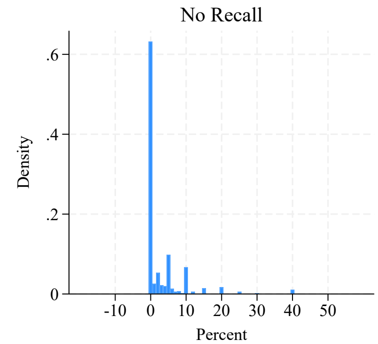
[11] This six-month cutoff splits the sample into two roughly equal groups.

[12] When comparing distributions of price changes, note that Panel (e) shows recalled past *actual* price changes while Panel (f) depicts *expectations* of future price changes, so more differences naturally appear between these two distributions.

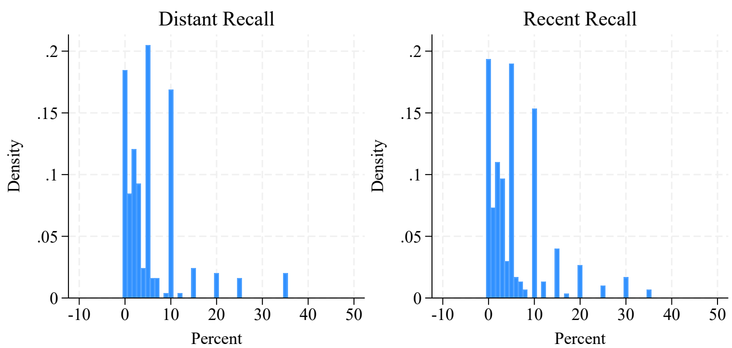
[13] Binder (2017) proposes using round-number responses in inflation expectations to proxy for uncertainty.



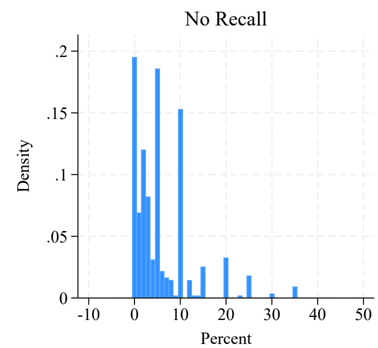
(a) Actual Cost Change between Last Two Price Changes $(C_{i,t_i} - C_{i,t_{o,i}})$



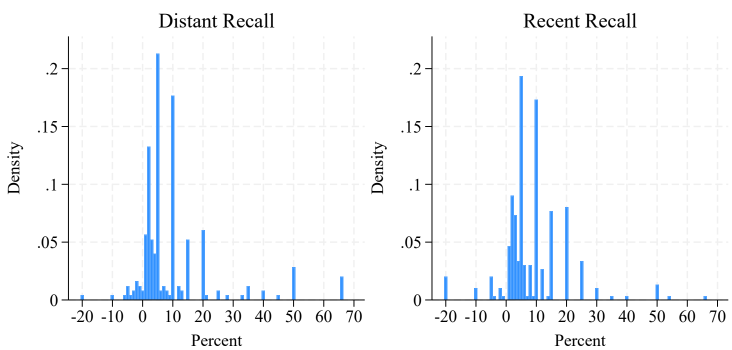
(b) Actual Cost Change since Last Price Change $(C_{i,\tau} - C_{i,t_i})$



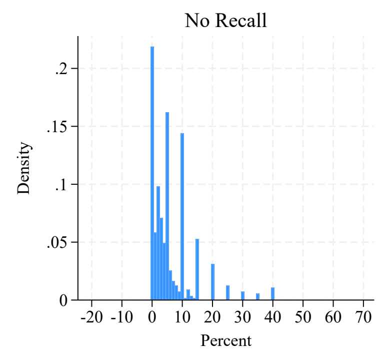
(c) Expected Cost Change as of Last Price Change $(E_{i,t_i}[C_{i,t_i+12} - C_{i,t_i}])$



(d) Expected Cost Change as of Survey $(E_{i,\tau}[C_{i,\tau+12} - C_{i,\tau}])$



(e) Most Recent Price Change $(P_{i,t_i} - P_{i,t_{o,i}})$



(f) Expected Price Change as of Survey $(E_{i,\tau}[P_{i,\tau+12} - P_{i,\tau}])$

Figure E.1: Distributions of Responses.

Table E.1: Measures of the Degree of Approximation in Responses

	(1)	(2)	(3)	(4)
	Fraction of Round-Number Responses			Cross-Variable
	(1) ΔC	ΔC^e	ΔP	Std Dev
Distant Recall	0.80	0.64	0.58	4.98
Recent Recall	0.81	0.64	0.64	4.52
No Recall	0.85	0.62	0.64	3.06

Note: Columns (1)–(3) of this table compare the fraction of round-number responses (defined as multiples of five) for cost changes, cost-change expectations, and price changes that required firms to recall the more distant past or the more recent past or that required no recall at all. For the no-recall case, price changes are expected price changes over the next 12 months. Column (4) compares—across the same distant, recent, and no-recall cases—the averages of the respondent-level standard deviation across five variables: cost changes, expected cost changes, cost changes due to tariffs, expected cost changes due to tariffs, and price changes. Here again, for the distant and recent-recall cases, the price changes are actual ones, while they are expected future changes in the no-recall case.

F Model Appendix

F.1 General LQ problem (FOC and survival representation)

Consider the problem of choosing the reset price in period t:

$$p_t^* \;=\; \arg\min_p \; \mathbb{E}_t\left[\sum_{k=0}^{\tau-1} \beta^k \left(p - p_{t+k}^{\dagger}\right)^2\right],$$

where τ is the random time until the next reset and $\beta \in (0,1)$. Further define the firm's objective $L(p) \equiv \mathbb{E}_t\big[\sum_{k=0}^{\tau-1} \beta^k (p - p_{t+k}^{\dagger})^2\big]$. Differentiating and setting the derivative to zero:

$$\frac{\partial L(p)}{\partial p} = \mathbb{E}_t\left[\sum_{k=0}^{\tau-1} 2\,\beta^k \left(p - p_{t+k}^{\dagger}\right)\right] = 0$$

Hence,

$$\begin{aligned} \mathbb{E}_t\left[\sum_{k=0}^{\tau-1} \beta^k\, p\right] \;&=\; \mathbb{E}_t\left[\sum_{k=0}^{\tau-1} \beta^k\, p_{t+k}^{\dagger}\right]. \\ p_t^* \;&=\; \frac{\mathbb{E}_t\left[\sum_{k=0}^{\tau-1} \beta^k\, p_{t+k}^{\dagger}\right]}{\mathbb{E}_t\left[\sum_{k=0}^{\tau-1} \beta^k\right]}. \end{aligned}$$

Now let $S_k \equiv \Pr_t(\tau > k)$ denote the survival probability. Using

$$\mathbb{E}_t\left[\sum_{k=0}^{\tau-1} \beta^k\right] = \sum_{k\geq 0} \beta^k\, S_k, \qquad \mathbb{E}_t\left[\sum_{k=0}^{\tau-1} \beta^k\, p_{t+k}^{\dagger}\right] = \sum_{k\geq 0} \beta^k\, S_k\, \mathbb{E}_t p_{t+k}^{\dagger},$$

equation (13) becomes

$$p_t^* = \frac{\sum_{k\geq 0} \beta^k\, S_k\, \mathbb{E}_t p_{t+k}^{\dagger}}{\sum_{k\geq 0} \beta^k\, S_k}. \tag{13}$$

With $p_{t+k}^{\dagger} = mc_{t+k}$, differentiating (13) with respect to $\mathbb{E}_t mc_{t+j}$ yields the general weight formula

$$\frac{\partial p_t^*}{\partial\, \mathbb{E}_t mc_{t+j}} \;=\; \frac{\beta^j\, S_j}{\sum_{k\geq 0} \beta^k\, S_k}, \qquad j = 0, 1, 2, \ldots$$

which is used throughout the main text.

Calvo and Taylor. For Calvo pricing, $S_j = \theta^j$, so $\sum_k \beta^k S_k = (1-\beta\theta)^{-1}$ and $w_j = (1-\beta\theta)(\beta\theta)^j$. For Taylor contracts, $S_j = 1$ for $j \leq N-1$ and zero thereafter, giving $\sum_k \beta^k S_k = (1-\beta^N)/(1-\beta)$ and $w_j = (1-\beta)\beta^j/(1-\beta^N)$.

Menu-cost environments. Following Alvarez et al. (2022), adjustment behavior can be summarized by a generalized hazard $\lambda(x)$ giving the instantaneous probability of price adjustment as a function of the price gap x. This representation can be microfounded by a broad class of sticky-price environments, including random menu-cost models and observation-friction models. We impose symmetry and monotonicity in $|x|$, and assume local linearity around the origin or equivalently,

$$\lambda(x) = \lambda_1|x| + o(|x|).$$

We also assume that the price gap evolves as a Gaussian random walk,

$$x_{k+1} = x_k + \varepsilon_{k+1}, \qquad \varepsilon_k \sim \mathcal{N}(0, \sigma^2), \qquad x_0 = 0.$$

Rather than tracking realized path-dependent adjustment risk, we use a first-order mean-field approximation and replace the hazard along a path $\{x_k\}$ with its horizon-by-horizon expectation,

$$\bar{\lambda}_k \equiv \mathbb{E}[\lambda(x_k)].$$

Under the random walk and local linearity,

$$\bar{\lambda}_k \approx \lambda_1 \mathbb{E}|x_k| = \lambda_1 \sqrt{\frac{2}{\pi}}\sigma\sqrt{k}.$$

Then, the implied survival probability up to horizon j is

$$S_j \approx \exp\left(-\sum_{\ell=0}^{j-1} \bar{\lambda}_\ell\right).$$

Since $\sum_{\ell=1}^{j} \sqrt{\ell} = \frac{2}{3} j^{3/2} + O\big(j^{1/2}\big)$, survival decays at a stretched-exponential rate,

$$S_j \simeq e^{-Aj^{3/2}}, \qquad A = \frac{2}{3}\lambda_1\sigma\sqrt{\frac{2}{\pi}}.$$

Using the general linear–quadratic pricing result, reset prices place weights

$$w_j = \frac{\beta^j S_j}{\sum_{k\geq 0} \beta^k S_k}$$

on expected future marginal costs, yielding the expression in Proposition 4.

F.2 Proof for Corollary 1

Taylor From Proposition 2 weights $w_j = \frac{(1-\beta)\,\beta^j}{1-\beta^N} \forall j \leq N-1$ and $w_j = 0 \forall j > N-1$ It follows that

$$\frac{\partial w_j}{\partial N} = (1-\beta)\beta^j \frac{\beta^N \ln \beta}{(1-\beta^N)^2}, \qquad j \leq N-1,$$

Due to $\beta < 1$ it follows that $ln(\beta) < 0$ and thus $\frac{\partial w_j}{\partial N} < 0 \forall j \leq N-1$.

Calvo From Proposition 3 weights $w_j = (1-\beta\theta)(\beta\theta)^j \forall j \geq 0$. It follows that

$$\begin{aligned}\frac{\partial w_j}{\partial \theta} &= -\beta(\beta\theta)^j + j(\beta\theta)^{j-1}\beta(1-\beta\theta) \\ &= \beta(\beta\theta)^{j-1}\left[j - (1+j)(\theta\beta)\right].\end{aligned}$$

Thus, $\frac{\partial w_j}{\partial \theta} > 0$ if $\theta < j/[(1+j)\beta]$.

F.3 Proof for Corollary 2

Given the random menu-cost microfoundation that rationalizes generalized hazards in Alvarez, Lippi, and Oskolkov (2022), adjustment occurs when a random cost draw ψ falls below the adjustment gain $U(x)$, implying a generalized hazard of the form

$$\lambda(x;\kappa) = G\left(\frac{U(x)}{\kappa}\right),$$

where $\psi_\kappa = \kappa\psi$ with $\psi \sim G$.

For a locally linear adjustment gain $U(x) = u_1|x| + o(|x|)$ implied by small price gaps and with $u_1 > 0$, a first-order expansion around the origin yields

$$\lambda(x;\kappa) = G\left(\frac{u_1(\kappa)}{\kappa}|x| + o(|x|)\right) = g(0)\,\frac{u_1(\kappa)}{\kappa}|x| + o(|x|),$$

so the local slope satisfies

$$\lambda_1(\kappa) = g(0)\,\frac{u_1(\kappa)}{\kappa}.$$

Differentiating,

$$\lambda_1'(\kappa) = g(0)\left(\frac{u_1'(\kappa)}{\kappa} - \frac{u_1(\kappa)}{\kappa^2}\right) = \frac{g(0)}{\kappa^2}\big(\kappa u_1'(\kappa) - u_1(\kappa)\big).$$

In an environment with small and medium price gaps, for small changes in κ, the gain of adjusting does not change much, that is $\frac{dU}{d\kappa} = \frac{dv(x)}{d\kappa} - \frac{dv(0)}{d\kappa} \approx 0$, that is, $u_1'(k) \approx 0$.
Therefore,

$$\lambda_1'(\kappa) < 0.$$

Thus higher menu costs reduce adjustment intensity at small and medium price gaps through a lower λ_1.

Under the mean-field and random walk approximation, the horizon-k expected hazard is

$$\bar{\lambda}_k(\kappa) \equiv \mathbb{E}[\lambda(x_k;\kappa)] \approx \lambda_1(\kappa)\mathbb{E}|x_k|.$$

Since $\lambda_1'(\kappa) < 0$ and $\mathbb{E}|x_k| > 0$ for all $k \geq 1$, it then follows that

$$\frac{\partial \bar{\lambda}_k(\kappa)}{\partial \kappa} < 0, \quad k \geq 1.$$

The implied survival profile

$$S_j(\kappa) \approx \exp\left(-\sum_{\ell=0}^{j-1} \bar{\lambda}_\ell(\kappa) \right)$$

therefore satisfies $\partial S_j/\partial\kappa > 0$ for all $j \geq 1$.

How do weights w_j change? First note that

$$w_0(\kappa) = \Big(\sum_{k\geq 0} \beta^k S_k(\kappa) \Big)^{-1},$$

so it follows that $\partial w_0/\partial\kappa < 0$.

Then

$$\frac{\partial}{\partial\kappa} \log S_{j+1}(\kappa) - \frac{\partial}{\partial\kappa} \log S_j(\kappa) = -\frac{\partial\bar{\lambda}_j(\kappa)}{\partial\kappa} \approx -\lambda_1'(\kappa)\,\mathbb{E}|x_j| > 0,$$

so $j \mapsto \partial_\kappa \log S_j(\kappa)$ is strictly increasing. Consequently, since $\sum_{j\geq 0} w_j(\kappa) = 1$, there exists a cutoff $j^\star(\kappa)$ such that

$$\frac{\partial w_j(\kappa)}{\partial\kappa} < 0 \quad \text{for } j < j^\star(\kappa), \qquad \frac{\partial w_j(\kappa)}{\partial\kappa} > 0 \quad \text{for } j > j^\star(\kappa),$$

(with equality at $j = j^\star(\kappa)$ if the cutoff is an integer). In particular, higher menu costs shift mass in $\{w_j\}$ from short horizons to long horizons and lower the contemporaneous weight w_0.