Intertemporal pass-through

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Abstract

Forward-looking pricing is at the core of modern macroeconomics, yet a gap remains between its theoretical foundations and their empirical validation. To bridge this gap, we study *intertemporal pass-through* (iPT): the sensitivity of firms' desired prices to changes in their *expected* future marginal costs, a micro building block of foresight in aggregate inflation. On the empirical side, we obtain direct iPT estimates by combining UK firm-level survey data with idiosyncratic news shocks from a natural experiment: the March 2019 announcement of a *future* tariff schedule in the event of a "No-Deal" Brexit. We find iPT to be largest among firms with the lowest frequency of price adjustment and those expecting the cost shock to arrive earlier. In addition, iPT is smaller among firms with state-dependent pricing and for larger shocks. On the theory side, we derive iPT in a model with heterogeneous adjustment frequencies and perceived shock horizons, formally reconciling our empirical findings on the drivers of iPT differences. We also use our setup to assess the general equilibrium consequences of iPT heterogeneity. In particular, we show that the sensitivity of *aggregate* inflation to changes in future costs is *convex* in non-adjustment frequencies and perceived shock horizons. As a result, iPT heterogeneity *amplifies* the degree of forward-lookingness of macroeconomic aggregates. Thus, announcements of future policies have contemporaneous effects, and heterogeneity in pricing decisions increase their magnitude.

Keywords: price-setting, expectations, survey data, tariffs, Brexit

JEL Codes: C83, D25, D84, E31

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1 Introduction

Equilibrium theories of inflation commonly feature forward-looking pricing behavior due to adjustment frictions. Such pricing foresight underpins many fundamental objects in macroeconomics, most notably the forward-looking Phillips Curve, with its distinct implications for optimal policy conduct.[1] Despite the early evidence supporting forward-looking inflation in aggregate data (Galí and Gertler, 1999; Sbordone, 2002), formal empirical validation of the degree of foresight in pricing remains challenging. In fact, the influential review by Mavroeidis et al. (2014) on empirical estimates of inflation expectations in the Phillips Curve concludes that "the literature has reached a limit on how much can be learned from aggregate macroeconomic time series." This paper takes an alternative approach and directly investigates the structural micro building block of forward-looking inflation, given by the sensitivity of a firm's *current* reset price to its expected *future* marginal cost. We label this object *intertemporal pass-through* (iPT):

$$\text{iPT} \quad \equiv \quad \frac{\partial \log P_t^*}{\partial \log \mathbb{E}_t MC_{t+d}} \tag{1}$$

where P_t^* is a firm's current reset price, and $\mathbb{E}_t MC_{t+d}$ is that firm's expectation of its own marginal cost $d > 0$ periods ahead.[2] Our goal is threefold: to obtain direct firm-level iPT estimates, to use the estimates for testing equilibrium pricing theories, and to pin down the implications of the observed firm-level iPTs for the forward-lookingness of aggregate inflation.

Obtaining direct estimates of iPT is challenging even with access to firm-level pricing data, since it additionally requires plausibly exogenous cross-sectional variation in firms' expectations of their own future marginal costs. We overcome this challenge by exploiting a unique natural experiment: the UK government's unexpected announcement in March 2019 of a temporary tariff schedule in the event of a "No-Deal" Brexit. The announcement implied that in the event no free trade deal is reached with the EU, UK firms that import inputs from the EU would unilaterally face tariffs that are substantially lower than the likely alternative of Most Favored

[1] As discussed in Clarida et al. (1999), an environment with a forward-looking Phillips Curve creates possible welfare *gains from commitment* for the policy maker. In particular, promises to be tough on inflation in the future improve the contemporaneous trade-off between inflation and the output gap. Such considerations become even more important whenever there is an effective lower bound on the nominal interest rate, a case very relevant for advanced economies in recent decades (Adam and Billi, 2007).

[2] In the canonical time-dependent pricing model of Calvo (1983), a first-order approximation of the (log) reset price satisfies $p_t^* = (1 - \beta\alpha) \sum_{d \geq 0} (\beta\alpha)^d \mathbb{E}_t mc_{t+d}$, where α is the probability of non-adjustment and β is the discount factor. Therefore, horizon-d iPT admits a closed-form expression of $(1 - \beta\alpha)(\beta\alpha)^d$. Explicit closed-form expressions for iPT are also available for analytically tractable state-dependent pricing models, such as Dotsey et al. (1999) and Gertler and Leahy (2008). It is such explicit link to model primitives that motivates our treatment of iPT as the structural micro building block of forward-looking inflation.

Nation (MFN) rates. Since the proposed reductions differ across products, the announcement generates *sectoral* variation in *future* effective tariffs, depending on the composition of intermediate inputs. We go a step further and create *firm-level* variation by using a confidential and representative survey of UK firms (the Decision Maker Panel). In particular, we combine the sectoral variation with firm-level data on (i) the perceived probability of a "No-Deal" outcome, and (ii) the cost share of intermediate inputs imported from the EU. As a result, we obtain firm-level news shocks, representing exogenous shifters to own expectations of future marginal costs. We then estimate firm-level iPT by regressing survey-based non-zero price changes after the announcement on the constructed news shocks.

Our empirical strategy allows us to estimate the average iPT among firms in the sample, as well as to analyse the cross-sectional heterogeneity in pass-through. We explicitly consider several potential drives of iPT differences. First, we test for the role of the average frequency of price adjustment, as well as for the relevance of the perceived Brexit date. We use measures of adjustment frequency from the official UK PPI/CPI microdata, which we further cross-check with explicit survey-based responses on durations of price spells. Our survey also contains data on firms' perceived Brexit date, which we use to construct firm-level measures of the perceived horizon of the tariff news shock. We find that, *ceteris paribus*, iPT increases monotonically with the average price spell duration. In particular, we find no statistically significant positive pass-through for firms with average price durations of up to 10 months, with statistically significant increases of 0.44 and 0.61 for the durations of 10-20 months and 20+ months, respectively.[3] In other words, firms that, on average, change prices less frequently are more forward-looking in their price setting decisions conditional on adjustment. Second, we find that, all else equal, expecting Brexit to occur later reduces iPT. In quantitative terms, we estimate that holding a belief that Brexit will occur in 2020 (as opposed to 2019) delivers a statistically significant drop in iPT by 0.48.

As a third possible dimension of iPT heterogeneity, we test for the relevance of firms' motivation behind price changes, and for the degree to which iPT depends on the size of the future marginal cost shifter. The survey contains data on whether firms report to be changing prices in response to specific events, or simply at regular intervals. We treat those responses as indicators of whether firms engage in state- or time-dependent pricing, respectively. Our estimation finds that, *ceteris paribus*, state dependence in price setting leads to a statistically significant

[3] Here and for the rest of the paper we express iPT as % price change following a 1% change in expected future marginal cost. For example, iPT=0.5 corresponds to half of the expected future cost change getting passed-through to current prices, whereas iPT=1 corresponds to full pass-through.

drop in iPT. Sectors with above median fraction of firms reporting to be state-dependent have iPT lower by 0.34, on average. In other words, firms that engage in time-dependent pricing are more forward-looking conditional on adjustment. As for any size dependence in iPT, we find that as the tariff news shock becomes bigger, the impact on price adjustment changes *less* than proportionally in magnitude.[4] Quantitatively, for every 1 percentage point increase in the magnitude of the shock, the estimated iPT drops by a statistically significant amount of 0.05.

We assess our empirical findings through the lens of a general equilibrium model with heterogeneity in price setting. Our setup allows for a general pricing hazard, which nests the time-dependent (Calvo, 1983) and the state-dependent menu cost (Golosov and Lucas, 2007) benchmarks as corner cases. For our first set of theoretical results, we analytically derive iPT under small shocks, relying on a time-invariant adjustment probability setup as a first-order approximation for a possibly richer pricing problem. In line with our econometric findings, we show that the theoretical iPT falls in the perceived horizon of the future cost shifter, and that it rises with the average price spell duration for the empirically relevant range of adjustment frequencies.[5] Moreover, we show that iPT is *convex* is the perceived shock horizon and in the relevant set of values for the non-adjustment frequencies. As we then formally deduce, the micro convexity implies that heterogeneity in the perceived shock horizons and adjustment frequencies *amplifies* the impact response of aggregate inflation to anticipated future shocks. This is in sharp contrast to the existing results for *realized* shocks, which suggest that heterogeneity in price rigidity *dampens* aggregate price movements (Carvalho, 2006; Nakamura and Steinsson, 2010). Our findings therefore suggest that micro heterogeneity in pricing matters for the behavior of aggregates in response to shocks or policies which affect future activity, such as forward guidance or anticipated changes in government spending.

In our second set of theoretical results, we study iPT under large shocks. To capture any potential size dependence in the model-based iPT, we consider fully non-linear numerical solutions to a menu cost model with random free adjustment opportunities in the spirit of Nakamura and Steinsson (2010), estimated to match UK pricing moments. For small shocks the computed iPT are very similar to those obtained analytically under time-dependent pricing,

[4]This is in sharp contrast to prior empirical studies of size dependence of aggregate price indices following *realized* shocks, which find that as the shock increases the price response goes up *more* than proportionally (Alvarez et al., 2017; Ascari and Haber, 2022).

[5]As shown by Auclert et al. (2024), for small shocks, the aggregate price behavior in state-dependent models is well approximated by a Calvo (1983) setup calibrated to a higher frequency of price adjustment. To the extent such numerical equivalence extends to the behavior of optimal reset prices, our theoretical finding that iPT under small shocks falls in the frequency of adjustment is consistent with our empirical evidence that iPT is smaller for firms with state-dependent pricing.

with the magnitudes of price changes growing linearly with the size of the shock. In contrast, as the cost shifters get large, the impact on price changes grow *less* than proportionally in magnitude, delivering iPT that is falling with the size of the shock. This model-based finding is consistent with the size dependence of iPT established empirically. The model mechanism behind it is as follows: while for small shocks adjustment decisions are mainly governed by the random free opportunities, for larger cost shifters they are predominantly based on the decision to pay the menu cost to avoid big losses from mispricing. After a very large news shock, firms endogenously increase their expected adjustment probability in the period when the shock actually arrives. As a result, they pass through less of the cost increase on impact, lowering their iPT. In the limit, as the news shock gets infinitely large, firms update their probability of adjustment in the future to one and simply wait for the shock to arrive, delivering an iPT of zero.

Contribution to the literature We contribute to at least four strands of the literature. **First**, our work adds to the literature on forward-looking pricing and the role of inflation expectations. Seminal studies by Galí and Gertler (1999) and Sbordone (2002) econometrically assess structural forward-looking Phillips Curves with aggregate US data, finding a substantial role for inflation expectations. However, subsequent papers point to econometric issues, such as model misspecification and weak identification, which cast doubt on the degree to which inflation expectations are important for explaining the behavior of aggregates (Lindé, 2005; Rudd and Whelan, 2005, 2007; Mavroeidis et al., 2014).[6] Given the difficulty in estimating the role of forward-looking pricing with aggregate time series, more recent work uses micro data in order to pin down model-free elasticities between firms' decisions and expectations. Seminal papers by Coibion et al. (2018) and Coibion et al. (2020), using surveys from New Zealand and Italy respectively, identify firm-level responses to exogenous shifts in their own inflation expectations, finding the response of quantities to be substantial, with a more muted effect on prices.[7] As an important theoretical benchmark, Werning (2022) derives the pass-through of inflation expectations to current inflation for a broad class of pricing models and without imposing rational expectations.

Our novel concept of *intertemporal pass-through* (iPT) and its estimates provides a con-

[6]For a response to the concerns, see Gali et al. (2005). More recent work by Barnichon and Mesters (2020) addresses some of the econometric issues with a novel technique for estimating forward-looking macro equations, and the Phillips Curve in particular, using aggregate time series and identified shocks as instruments.

[7]Also see Enders et al. (2022) and Savignac et al. (2024) for evidence using German and French firm-level data, respectively. Of note is also the paper by Delgado et al. (2024), which uses Colombian data to identify firm-level responses to shifts in their own exchange rate expectations.

nection between the older literature assessing structural forward-looking Phillips Curves with aggregate time series, and the modern approach of pining down model-free elasticities between decisions and expectations using firm-level data. In particular, while relying on firm-level data to strengthen identification, our iPT estimates have a clear structural interpretation in the context of theoretical models, allowing an assessment of whether they produce realistic forward-looking behavior. We find that the canonical time-dependent pricing model of Calvo (1983), calibrated to observed adjustment frequencies, produces iPT that is quantitatively in line with our econometric estimates, averaged across shock sizes. Moreover, the state-dependent CalvoPlus model of Nakamura and Steinsson (2010), when solved fully non-linearly, produces iPT that falls in shock size, consistent with our econometric findings.

Second, we contribute to the literature on micro heterogeneity in price setting and its implications for aggregate dynamics. Pioneering work by Bils and Klenow (2004), Klenow and Kryvtsov (2008) and Nakamura and Steinsson (2008) measures frequency of price adjustment, as well as other pricing moments, for a broad range of product categories in the US, documenting substantial heterogeneity.[8] A separate strand of the literature instead uncovers heterogeneity in the responses of sectoral price indices to identified aggregate and idiosyncratic shocks (Maćkowiak et al., 2009; Boivin et al., 2009). More recent papers focus on estimating differences in the pass-through of costs to firms' prices, by either combining product-/firm-level data with externally identified shocks (Gopinath et al., 2010; Amiti et al., 2019; Dedola et al., 2021; Auer et al., 2021; Alexander et al., 2024) or more directly using surveys (Dogra et al., 2023; Godl-Hanisch and Menkhoff, 2024; Bunn et al., 2024a). On the theoretical front, Carvalho (2006), Nakamura and Steinsson (2010) and Afrouzi and Bhattarai (2023) show that micro heterogeneity in price setting frequency dampens the response of aggregate price indices to monetary shocks, thus amplifying the degree of monetary non-neutrality.

We make both an empirical and a theoretical contribution to this literature. To the best of our knowledge, we are the first to measure the firm-level pass-through of *future* costs to *current* prices *and* to identify several dimensions of heterogeneity of such pricing foresight. Our empirical results on the non-linearity of iPT are also novel and stand in sharp contrast to the prior results on size dependence under *realized* shocks (Alvarez et al., 2017; Ascari and Haber, 2022). We also obtain novel theoretical results, which suggest that heterogeneity in adjustment

[8] For the US, see Gorodnichenko and Weber (2016) and Pasten et al. (2020) for the most recent and granular measures of frequency of price adjustment, as well as Hong et al. (2023) for granular sectoral estimates of higher-order moments. Also see Bunn and Ellis (2012a,b) for estimates of pricing moments in the UK, and Gautier et al. (2024) for the most recent and extensive set of measures for the Euro Area.

frequencies, through the effect on iPT, can *amplify* the response of aggregate price indices to future shocks, very much in contrast to existing results for realized shocks, which instead find dampening. Our fully non-linear solutions for model-based iPT under state-dependent pricing are also both novel and in line with our own empirical evidence on size dependence.

Third, our work relates to the literature on news shocks, reviewed by Beaudry and Portier (2014). Theoretically, Beaudry and Portier (2004) and Jaimovich and Rebelo (2009) show how anticipated changes in future productivity can drive business cycles and, crucially, induce co-movement across the major aggregate and sectoral variables. At the same time, in the context of a fully estimated general equilibrium model, Schmitt-Grohé and Uribe (2012) uncover a small role for news about future productivity, while also allocating a much more important role to news about future preferences and wage markups. On the empirical side, the evidence is also mixed: while Beaudry and Portier (2006) and Kurmann and Otrok (2013) find a large role for productivity news shocks, Barsky and Sims (2011) and Crouzet and Oh (2016) find their role to be much more modest.[9] Although the empirical literature tends to focus on news shocks to productivity, there are also influential papers identifying news shocks to government spending (Ramey, 2011) and oil supply (Känzig, 2021).

Our empirical contribution amounts to constructing a novel set of firm-level news shocks about future costs. We do that by combining a state-contingent policy announcement about future tariffs with survey-based data on firms' perceived probability of the relevant future state, as well as differences in firms' exposure to imported inputs. On the theoretical front, to the best of our knowledge, we are the first to study news shock propagation in an environment with pricing heterogeneity and non-linearities due to state-dependent pricing.

Fourth, our focus on tariff announcements connects the paper to the broader literature on the implications of trade frictions and protectionism. The comprehensive cross-country empirical study by Barattieri et al. (2021) finds that protectionism in the form of temporary trade barriers creates inflation and output reductions in the short run, with only marginally positive effect on the trade balance.[10] In another cross-country study, Furceri et al. (2022) estimate that tariff imposition leads to contractions in output and productivity, with no significant changes to

[9] For more recent alternative approaches to empirical identification of news shocks, see Cascaldi-Garcia and Vukotić (2022) and Cascaldi-Garcia (2025), who exploit patent data and long-run forecast revisions in the Survey of Professional Forecasters, respectively. Also see Görtz et al. (2022) for recent evidence on the effect of productivity news shocks on financial markets.

[10] In the US context, Barattieri and Cacciatore (2023) also study the propagation of temporary trade barriers through production networks. They find no significant positive effects for the protected industries, while the negative employment effects for the downstream industries are substantial and significant.

the trade balance. Most related to our work is the empirical literature on the pass-through to prices of *implemented* tariffs. For the US, Amiti et al. (2019), Flaaen et al. (2020) and Fajgelbaum et al. (2020) provide evidence of very strong pass-through from imposed tariffs to prices. In the UK context, Bakker et al. (2022) find substantial inflationary effects of the non-tariff measures (NTMs) implemented following Brexit. More recently, Di Pace and Masolo (2025) use synthetic control methods and find that Brexit led to a permanent increase in the UK price *level*, driven by both the initial sterling depreciation and the higher non-tariff barriers since 2021.

Our paper complements this literature by focusing on the impacts of *announcements* of future changes in trade restrictions, in contrast to realized changes. We believe this is a valuable exercise as changes in trade restrictions tend to be announced well in advance of the actual implementation.

Roadmap The rest of the paper is organized as follows. Section 2 describes the institutional context of our natural experiment, as well as the construction of our dataset. Section 3 outlines our econometric strategy and highlights the key empirical results, as well as the robustness checks. Section 4 presents the theoretical model, which we use to analyze our empirical results and derive the general equilibrium consequences of iPT heterogeneity. Section 5 concludes and outlines avenues for future work.

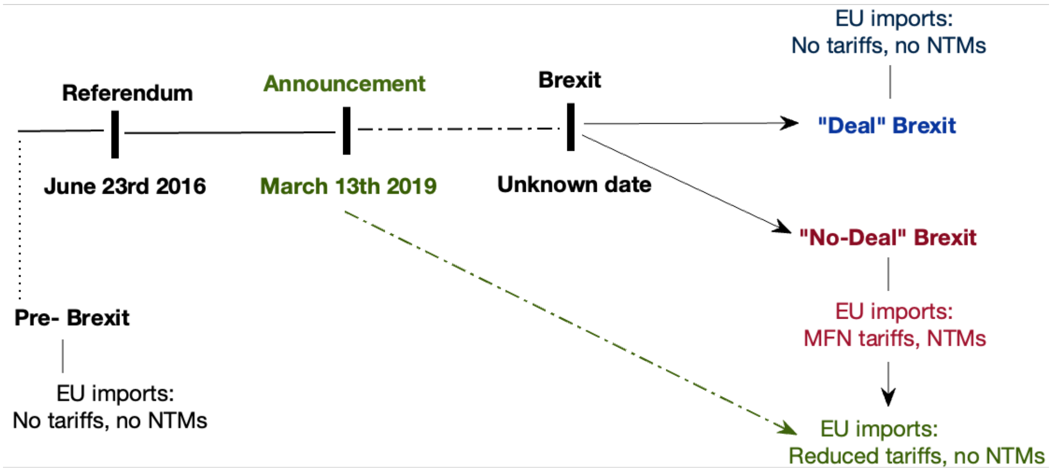
2 Data

We begin by providing a detailed description of our data construction strategy. First, we introduce the institutional context behind the natural experiment that we exploit: the March 2019 announcement of a reduced tariff schedule in the event of a "No-Deal" Brexit. Second, we augment the sectoral variation in expected future tariff generated by the unexpected policy announcement with confidential UK survey data to construct firm-level shifters in expected future marginal costs.

2.1 Institutional context

On the 23rd of June 2016 the United Kingdom (UK) voted to leave the European Union (EU). Among the many uncertainties surrounding the exit process, including its precise timeline and conditions, the issue of post-Brexit trade arrangements with the EU stood out in its economic importance. In the aftermath of the Brexit referendum, the potential future trade arrangements

Figure 1: Timeline of key Brexit tariff scenarios



Notes: This figure shows the timeline of the Brexit process as perceived at the time of the UK Government's March 2019 announcement. See Table A1 for a list of key dates and events associated with Brexit.

could be divided into two categories. The first one corresponded to the various "Deal" Brexit options, implying that the UK would either remain in the European Economic Area (EEA), or in another form of a customs union with the EU. By and large, it would maintain tariff-free trade between the UK and EU, with minimal degrees of frictions such as checks on goods to ensure compliance with regulations.[11] The second set of options implied leaving the EU without a trade agreement, or a "No-Deal" Brexit. Under the latter outcome, the UK would be treated as a third country under the EU's Most Favored Nation (MFN) rules. In the likely retaliation scenario, the UK would apply the MFN external tariff regime to the EU, delivering a substantial increase in the cost of importing from the EU.[12]

Realizing the potentially severe consequences of an immediate transition to MFN tariff rates in the event of a "No-Deal" Brexit, on the 13th of March 2019 the UK Government announced a commitment to a specific temporary tariff regime in case no trade agreement is reached in the future.[13] If introduced, the temporary regime would last for up to 12 months post Brexit completion, and would involve substantial unilateral reductions in tariffs on imports

[11]See Fella (2019). The small exception is the "Canada-style" deal where will be tariffs on a small handful of food products, but in large, broader goods trade would remain tariff-free.

[12]The no-retaliation scenario was perceived as very unlikely. This is because, according to the World Trade Organization (WTO) MFN rules, maintaining zero tariffs on EU imports would automatically force the UK to extend the no-tariff regime to the rest of the countries with which it does not have explicit trade agreements.

[13]https://www.gov.uk/guidance/check-temporary-rates-of-customs-duty-on-imports-after-eu-exit

from the EU. The conditions of the proposed regime varied significantly across products: 87% of imports (by value) would be eligible for tariff-free access, while the remaining 13%, primarily in protected industries such as the agriculture and automotive sectors, would remain subject to tariffs, though at levels that are significantly lower than the corresponding MFN rates.

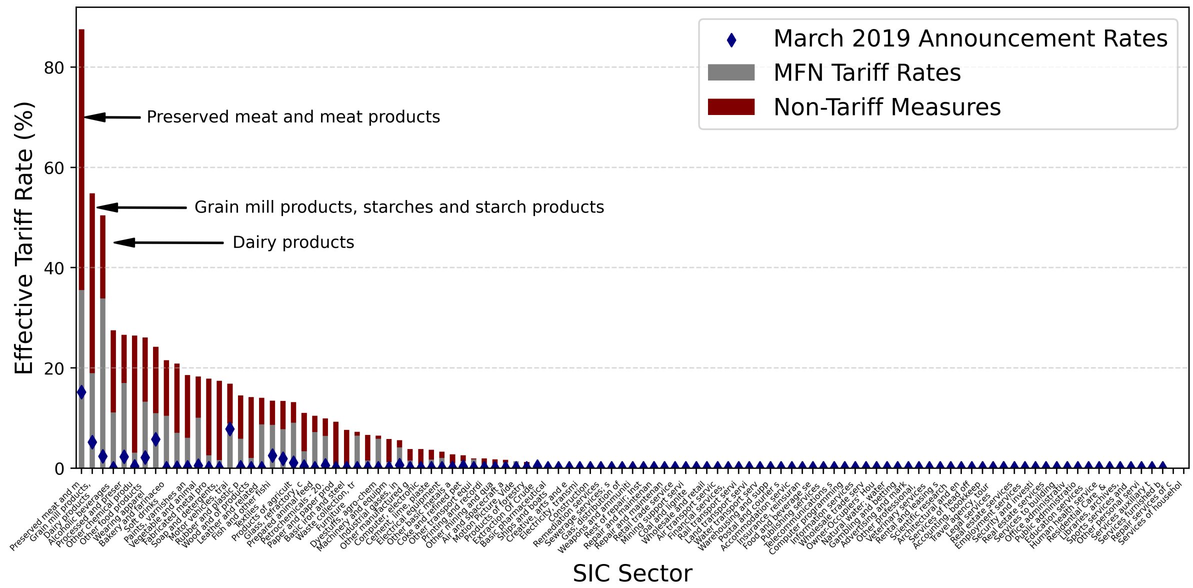
From the perspective of UK firms which, as of March 2019, imported all or part of their production inputs from the EU, such an announcement represented a reduction in their expected future costs. In particular, to the extent an individual firm believed that a "No-Deal" Brexit was possible, the commitment to temporarily lower tariffs on EU imports in such a contingency amounted to a negative-valued news shock to that firm's marginal cost. Naturally, the magnitude of such a news shock depended crucially on a number of firm-specific factors, among them the composition of intermediate inputs, the importance of EU imports in the cost structure, as well as the perceived likelihood of the "No-Deal" outcome. Moreover, the expected date of Brexit, at least as viewed at the time of the announcement in March 2019, pinned down the perceived horizon of the idiosyncratic news shock about future marginal cost. To be clear, we use the tariff announcement in March 2019, rather than the referendum outcome itself, as our main source of variation to study the impact on firm prices. Next, we outline our process of shock construction, paying particular attention to capturing the relevant dimensions of heterogeneity.

2.2 Product- and sector-level tariff variation

The starting point of our shock construction exercise is the trade micro-data on 9,533 products, supplied to us by His Majesty's Revenue & Customs (HMRC). The dataset contains, among other things, the product-specific tariff rates under the temporary scheme that was proposed by the UK Government on the 13th of March 2019. We denote such UK-proposed tariff rate for a product p with τ_p^{UK}.

The trade micro-data also contains the product-specific tariff rates under the MFN external tariff regime, which, by default, would be in place post-Brexit absent a deal or the announcement in March 2019. We denote the MFN tariff rate for a product p with τ_p^{MFN}. Crucially, the MFN rates represent just the direct, pecuniary additional cost associated with absence of a trade deal. As argued by Bakker et al. (2022), equally if not more important costs are brought on by the so-called non-tariff measures (NTMs), both technical and non-technical. Technical NTMs capture sanitary and phytosanitary measures as well as labelling and certification requirements. Non-technical NTMs include contingent trade measures, quantitative restrictions, price controls, and

Figure 2: Effective tariff rates under two main scenarios by sector



Notes: This figure shows MFN tariff rates, non-tariff measures (NTMs, both technical and non-technical), as well as the UK proposed tariff rates that have been aggregated to 2-digit SIC sectoral definitions.

finance measures. We measure the pecuniary costs associated with the NTMs by their product-specific ad valorem tariff equivalents, available from the World Bank.[14] Let τ_p^{NTM} denote such ad valorem equivalent for a product p.

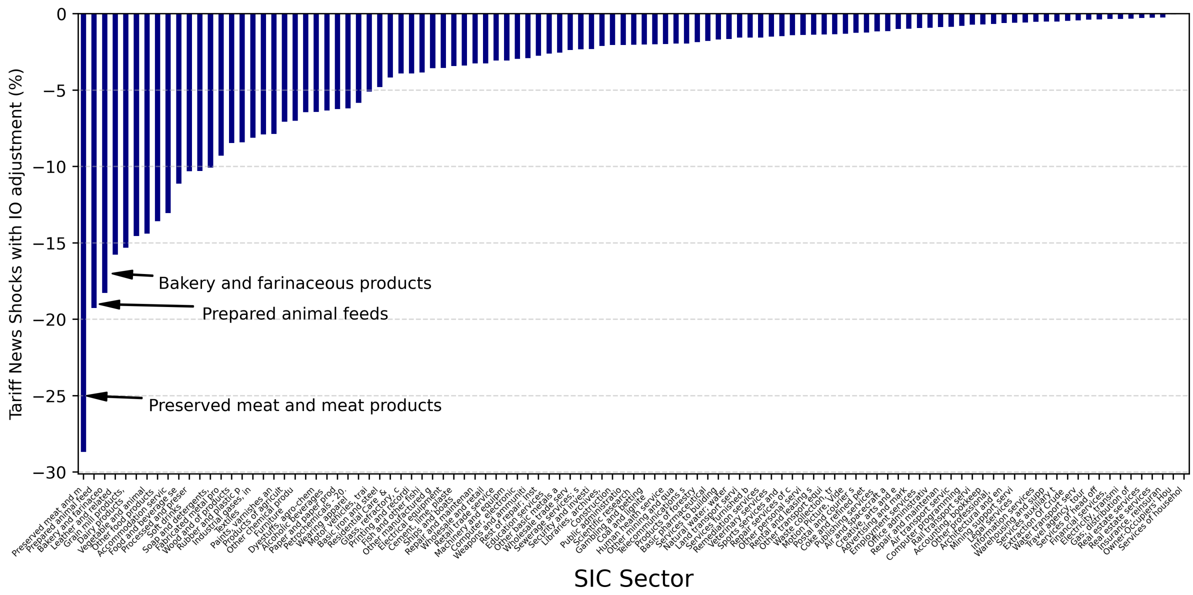
Overall, for imports of product p from the EU, the March 2019 announcement generated the following reduction in the effective tariff rate under the "No-Deal" scenario:

$$\Delta\tau_p \quad \equiv \quad \tau_p^{UK} \quad - \quad \left(\tau_p^{MFN} + \tau_p^{NTM}\right). \tag{2}$$

In order to get a quantitative sense of the magnitudes of the changes, Figure 2 reports the values of τ_p^{UK}, τ_p^{MFN} and τ_p^{NTM} that have been aggregated up to the 2-digit sectoral definitions (SIC 2007/CPA 2008), using relative product import weights. One can see, for example, that for the "Preserved meat and meat products" sector, the combined MFN and NTM effective tariff on EU imports is as large as 85%, which drops considerably down to just 18% under the

[14]For our measure, we include both technical and non-technical NTMs. Those are available at the HS 6-digit product level, and are mapped to the sectoral level following the same approach above. Finally, we note that these NTMs are only available for the EU as an importing country. Hence, we follow Bakker et al. (2022) in assuming that the NTMs are mirrored by the UK after Brexit. The data can be accessed here: https://datacatalog.worldbank.org/search/dataset/0040437/Ad-Valorem-Equivalent-of-Non-Tariff-Measures.

Figure 3: Sectoral tariff news shocks (including input-output adjustment)



Notes: This figure shows the sectoral tariff news cost shocks. These shocks account for technical and non-technical tariff measures and include an input-output table adjustment.

proposal announced in March 2019. For other agriculture-related products, such as "Grain mill products, starches and starch products" and "Dairy products", the drops are from 55% to 7%, and from 50% down to 3%, respectively.

We now account for the fact that firms are differentially exposed to the tariff changes depending on the composition of their intermediate input bundle. In particular, we compute the relevant effective tariff change for a firm in a sector k as:

$$\Delta T_k = \sum_r \overline{\omega}_{kr} \Delta \tau_r, \quad \forall k \tag{3}$$

where $\overline{\omega}_{kr}$ is the cost share of inputs bought in sector r for sector k, which we measure at the 2-digit level using the UK Input-Output accounts published by the Office for National Statistics (ONS)[15], and $\Delta\tau_r$ represents the sector r aggregated tariff change (given by the difference between the diamonds and the bars in Figure 2).

Figure 3 shows the estimated *sectoral* tariff changes $\{\Delta T_k\}_k$ based on (3). Naturally, all the changes are negative, since the March 2019 proposal involved tariff reductions across the board. As one can see, there is also substantial heterogeneity across sectors. The three sectors

[15]The cost shares are expressed as a fraction of total intermediate inputs, so that for any given sector they add up to one by construction: $\sum_r \overline{\omega}_{kr} = 1, \forall k$.

with the largest sectoral changes are "Preserved meat and meat products" (-28%), "Prepared animal feeds" (-19%), and "Bakery and farinaceous products" (-18%).

2.3 Firm-level shifters in expected marginal costs

With the effective sectoral tariff changes at hand, we can now use them to construct firm-level news shocks, representing shifters to their own expected future marginal costs.

Magnitudes of news shocks First, note that the future tariff changes announced in March 2019 are conditional on the "No-Deal" Brexit outcome. As a result, the proposal is only relevant for an individual firm's expected costs to the extent to which it believes the "No-Deal" outcome is probable. Second, the tariff changes matter to the degree to which a given firm imports it intermediate inputs from the EU. All in all, we can write the shifter to the expected future marginal cost of a firm i in sector k as:

$$\varrho_{i,k} \quad = \quad \pi_{i,k}^{\text{NO-DEAL}} \times \delta_{i,k}^{\text{EU}} \times \Delta T_k, \quad \forall i,k \tag{4}$$

where $\pi_{i,k}^{\text{NO-DEAL}}$ is the firm's perceived probability of the "No-Deal" Brexit outcome, and $\delta_{i,k}^{\text{EU}}$ is the share of the firm's total variable costs accounted for by imports of intermediate inputs from the EU. Both variables are at the time of the March 2019 announcement.[16]

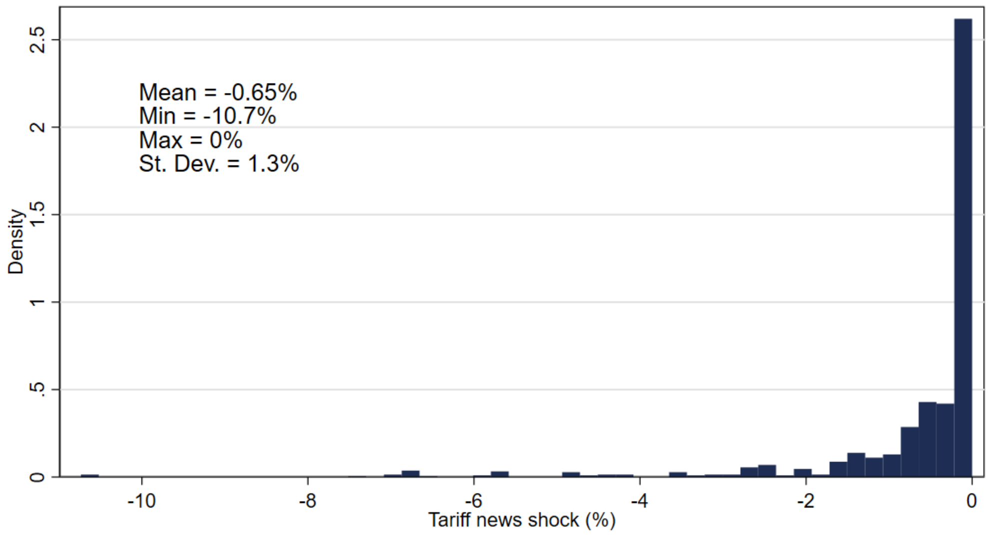
Naturally, it follows that one requires firm-level information in order to construct news shocks according to (4). For that, we use UK firm-level data that comes from the confidential responses underlying the Decision Maker Panel (DMP) survey, jointly constructed and maintained by the Bank of England, University of Nottingham, and King's College London.[17] Launched in 2016, the panel comprises Financial Officers from small, medium and large UK companies operating in a broad range of industries and is designed to be representative of the population of UK businesses (see Bunn et al., 2024b for further details). Approximately 10,000 businesses were part of the panel as of September 2022.

In particular, in order to measure the firm-specific perceived probabilities of the "No-Deal" Brexit outcome, we use direct survey responses to the question of what probability, in percent, do firms attach to a disorderly Brexit, whereby no deal is reached by the end of March 2019. The question was asked to firms between February and April 2018 and again between November 2018 and January 2019, shortly before the March 2019 announcement about future tariffs. We

[16] Naturally, the expression in (4) captures first-order changes in expected future costs, and not higher-order variation, such as changes in expected future cost share of EU imports driven by the announcement.

[17] https://decisionmakerpanel.co.uk/

Figure 4: Distribution of firm-level tariff news shock



Notes: This figure shows the distribution of the firm-level tariff news shock in the main regression sample.

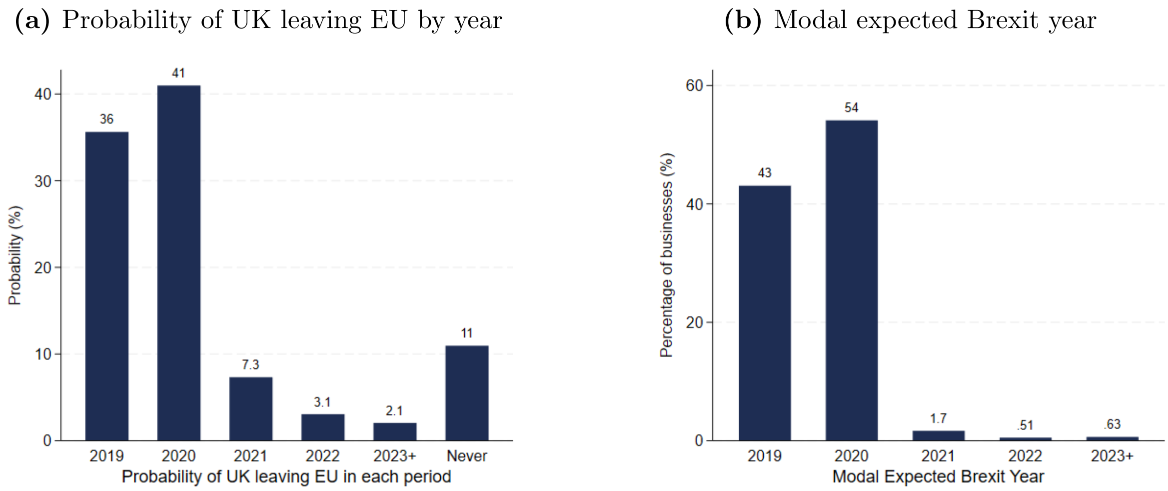
use the latest available response to the question for each firm in our analysis. Figure A2 in the Online Appendix summarizes the firms' responses, exhibiting substantial heterogeneity in the perceived "No-Deal" probabilities. Specifically, the average probability assigned to a "No-Deal" Brexit is 47%, with a standard deviation of 25% across firms.

As for the firm-level cost share of imports from the EU, it is constructed in two steps. First, we use the direct survey responses to the question of what percentage of costs were imports from EU countries pre-Brexit.[18] Those give us EU imports as a fraction of *total costs*, summarized in Figure A3 in the Online Appendix. The mean cost share is 0.17, with a standard deviation of around 0.24, implying a wide heterogeneity across firms in our sample. However, of interest to us are shares of *variable costs*, which are relevant for the marginal cost concept and ultimately for pricing decisions. Since no firm-level data on variable cost shares is available, we approximate it with sectoral measures constructed using Worldscope.[19] Matching firms to 4-digit SIC sectors, we multiply the survey-based cost shares by the sectoral measures of total over variable costs. This gives us firm-level measures of EU imports as a fraction of *variable costs*, which we use in the construction of the shock.

[18]This question has been asked multiple times, including in the introductory survey when a firm joins the sample. For a given firm, the earliest available response is used in the analysis.

[19]Those are constructed as follows. First, for each UK firm available in Worldscope, one constructs the ratio of variable to total costs. Then those variable cost shares are averaged across all firms within a given 4-digit SIC sector. See Figure A5 in the Online Appendix for a histogram across sectors. We thank Maria Ptashkina for sharing the data with us.

Figure 5: Expected Brexit Horizons



Notes: Panel (a) shows the average probability assigned to the UK leaving the EU in each period from 2019 to 2023+, including 'Never'. Panel (b) shows the modal expected Brexit years, meaning the years which firms placed the highest probability on Brexit occurring.

Figure 4 shows the distribution of the firm-level tariff news shock based on equation (4). As previously discussed, those have a clear interpretation as percent changes in firms' expected future marginal costs. Note that all shocks are either zero or negative, since the announcement featured potential tariff reductions in the future. The shock values range from -10.7% to 0, with an average of -0.65% and a standard deviation of 1.3%.[20]

Perceived horizons of news shocks In addition to the magnitudes of the firm-level news shocks, our survey data allows us to measure the perceived *horizon* of each future marginal cost shifter. In particular, one of the survey questions directly asks firms to assign probabilities to potential future dates of Brexit. In our context, the perceived proximity of Brexit completion corresponds to the horizon from which the announced tariff changes could be relevant. Focusing on survey responses closest to the March 2019 announcement, Figure 5 reports the average probability assigned to each potential Brexit year, as well as the distribution of modal Brexit years (year to which the highest probability is assigned by a given firm). We also construct the implied firm-level expected duration between the tariff schedule announcement and Brexit completion. As reported in Figure A7 (Online Appendix), the mean duration is around 7 quarters, with a standard deviation of approximately 2.2 quarters.

[20]Figure A6 shows the equivalent distribution of non-zero firm-level tariff news shocks.

2.4 Firm-level pricing data

Having constructed the idiosyncratic news shocks, we now turn to presenting the data on firms' pricing decisions. Our dataset includes both time series of firm-level price indices, as well as information regarding key attributes of price setting, such as adjustment frequencies and self-reported motivation behind price changes.

Firm-level price indices Our survey contains time series of firms' self-reported growth rates of their own prices over the prior 12 months.[21] The survey design ensures that every firm gets asked about its 12-month price growth rate every quarter, allowing us to create quarterly firm-level price indices. Though the price indices are not comparable across firms in absolute terms, they allow us to track relative price movements at a quarterly frequency.

Frequencies of price adjustment We use two measures of firms' frequency of price adjustment. First, the survey contains a direct questions about how often a given firm changes its price, allowing for seven response options: daily, weekly, monthly, quarterly, half-yearly, annually and never.

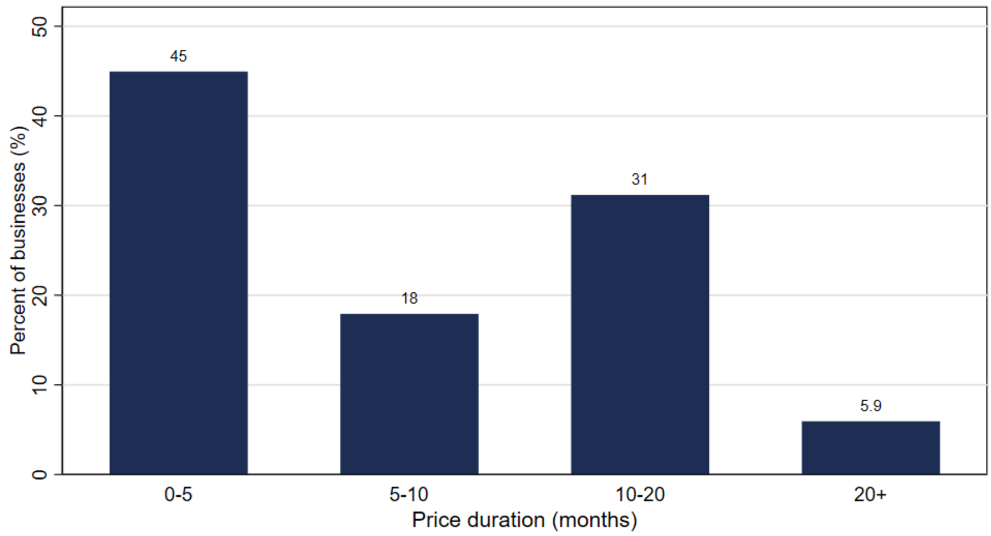
Though useful due to their self-reported nature, the direct survey-based answers only give a qualitative sense of the frequency of price adjustment. To get a more quantitative measure, we rely on the confidential UK microdata, underpinning the official CPI and PPI measures, as collected by the Office for National Statistics (ONS). The dataset contains monthly product-level prices, and we follow Gorodnichenko and Weber (2016) in constructing the frequency of price adjustment for every product by measuring the fraction of months in the sample containing a price change for that product. We then aggregate the product-level frequencies to the 4-digit SIC sector-level. Since the firms in our survey can also be easily mapped to the 4-digit SIC definitions, we use the sectoral measures of frequency of adjustment to get a quantitative approximation of how often a given firm changes its prices.

For the ease of interpretation, we convert the frequencies of adjustment to average price durations.[22] The median price duration we estimate is 5.8 months. In the estimation, we split the average durations into four bins: 0-5 months, 5-10 months, 10-20 months, and 20+ months. Figure 6 shows the distribution of firms in these four bins. Around 45% are in the first bin with the shortest price durations. But over one-third of firms are in sectors with a price duration

[21] See Figure A11 in the Online Appendix for histograms of the firm-level quarterly year-on-year price growth rates. Figure A12 shows that firms' annual own-price growth is highly correlated with annual CPI inflation in the UK over 2017-2025.

[22] The formula for the conversion is: $\text{Duration} = -1/\log(1 - \text{Frequency})$.

Figure 6: Frequency of price adjustment in main estimation sample



Notes: This figure shows the distribution of estimated durations of product prices in the UK CPI/PPI microdata constructed and maintained by the Office for National Statistics (ONS).

longer than ten months.[23]

Motivation behind price changes A unique feature of our dataset is that firms also explicitly get asked about their motivation behind price changes. More specifically, firms report whether they change their prices in response to specific events, or simply at regular time intervals. Those two options have a natural interpretation as pointing towards state- and time-dependent pricing, respectively.

In order to get a sense of how the motivation behind price adjustment varies across different parts of the economy, Figure A13 reports the share of firms that report to be state-dependent in each 2-digit SIC sector of the UK economy. Notably, four sectors, namely "Programming and broadcasting", "Manufacturing of coke", "Gambling and betting" and "Activities of extraction" have one hundred percent of firms reporting to be state-dependent in their pricing. At the same time, all firms in "Undifferentiated goods", "Remediation activities", "Postal and courier activities", "Manufacturing of leather" and "Activities of households" report to be time-dependent instead. The rest of the sectors lie in-between, with the median fraction of state-dependent firms close to 60 percent (see Bunn et al., 2023 for further details).

[23] In Figure A10 we compare our sectoral average price durations (based on PPI/CPI microdata) to the self-reported survey-based frequencies of price adjustment of firms that belong to those sectors. The two measures show a clear monotonic relationship: higher self-reported frequency is associated with a lower price duration.

2.5 Additional firm-level data

We make use of additional survey-based variables in our empirical analysis, either as control variables, or in order to assess the effect of news shocks on other firm-level outcomes. In particular, we use firms' self reported quarterly employment figures, capital expenditure, sales, as well as time-invariant characteristics, such as exporter and importer status. See Table A2 for summary statistics.

3 Econometric strategy and estimation results

Having constructed the dataset, we now introduce our econometric strategy and describe our key estimation results. Our specification captures possible heterogeneities in the frequency of price adjustment, perceived horizon of the shock, time- or state-dependent pricing, as well as the size of the shock. We find that the iPT is largest for firms with stickiest prices and short perceived shock horizons. Moreover, we find the pass-through to be lower for firms with state-dependent pricing, as well as for large shocks.

3.1 Econometric specification

In order to estimate firm-level iPT, we combine the survey-based quarterly firm-level price indices and the newly constructed news shocks, corresponding to percent changes in expected future marginal costs, in the following linear regression model:

$$p_{i,k,H} \quad = \quad \phi_{i,k,H} \times \varrho_{i,k} \quad + \quad \gamma_H \quad + \quad \boldsymbol{\varphi}'_H \boldsymbol{X}_{i,k} \quad + \quad \varepsilon_{i,k,H} \tag{5}$$

where $p_{i,k,H} \equiv 100 \times [\log P_{i,k,H} - \log P_{i,k,-1}| \neq 0]$, representing non-zero (log) price changes for a firm i in sector k between its most recent price before the March 2019 announcement ($P_{i,k,-1}$) and its price H quarters after the announcement ($P_{i,k,H}$), measured in %. In addition to our constructed firm-level tariff news shock $\varrho_{i,k}$, also measured in %, we allow for quarter-specific fixed effects γ_H, which soak up all general equilibrium effects specific to a particular period after the announcement. The specification also includes a set of firm- and sector-level controls $\boldsymbol{X}_{i,k}$, containing firm-level employment and capital expenditure right before the March 2019 announcement, the firm-level exporter status, the perceived probability of "No-Deal" Brexit, the import cost share from the EU, the perceived horizon of Brexit, as well as the sectoral variable cost share and the sectoral frequency of price adjustment.[24] For inference purposes, the standard errors are clustered at the 2-digit SIC sectoral level.

On the right-hand side of specification equation (5), the main object of interest is $\phi_{i,k,H}$, which directly corresponds to iPT of a firm i in sector k, H periods following the announcement. Indeed, it gives the elasticity of the firm's post-announcement price update to a shifter in that firm's expected future marginal cost, as measured by the idiosyncratic news shock. As the superscript of $\phi_{i,k,H}$ indicates, we would like to capture potential heterogeneity in iPT, as driven by firm- and sector-level characteristics, as well as by the specific period following the announcement.

In addition to the average iPT across all firms ($\phi_{i,k,H} = \phi_H$) and periods after the announcement ($\phi_{i,k,H} = \phi$), we consider several potential dimensions of cross-sectional heterogeneity. First, we analyze how iPT depends on the firm's average frequency of price adjustment. We measure it with the mean price spell duration of products in the granular 4-digit SIC sector to which that firm belongs. The second dimension of heterogeneity we consider is the perceived horizon of the news shock, captured by the expected Brexit date. The perceived horizon is measured directly at the firm level with the survey question which asks respondents to assign probabilities to potential Brexit years. Third, we study how iPT depends on the motivation behind price changes, captured by the fraction of firms in a given sector that report to be state- or time-dependent in their pricing. Finally, given the wide range of magnitudes admitted by our news shock measure, we estimate how iPT varies with the size of the expected marginal cost shifter.

The time dimension of our estimation sample covers three quarters after the 2019 announcement: March-May (2019Q2), June-August (2019Q3) and September-November (2019Q4). We do not study further quarters to avoid overlaps with the Covid crisis which begins in 2020 and complicates inference. All in all, our baseline sample contains 1,010 firm-quarter observations.

3.2 Estimated iPT: Role of adjustment frequency and shock horizon

Our first set of empirical results concerns the relationship between iPT and two characteristics: the average frequency of price adjustment and the perceived horizon of the shifter in the expected future marginal cost. To capture the former, we allocate firms into bins based on the average duration of price spells for products matched to them at the 4-digit level. In particular, we create four bins of average price durations: 0-5 months, 5-10 months, 10-20 months and 20+

[24] Table A2 in the Online Appendix reports summary statistics of these variables in the main estimation sample.

Table 1: Estimated iPT: Interaction with price durations and perceived Brexit horizon

	(1)	(2)	(3)	(4)	(5)
Dependent variable:	100 × Δ log(Price Level)				
Sample:	2019Q2 - 2019Q4				
Tariff news shock$_{i,k}$	-0.107	0.051	0.060	0.054	-0.020
	(0.080)	(0.081)	(0.126)	(0.125)	(0.108)
Tariff news shock$_{i,k}\times$ Price Duration 5-10M$_k$					0.138
					(0.185)
Tariff news shock$_{i,k}\times$ Price Duration 10-20M$_k$					0.435**
					(0.171)
Tariff news shock$_{i,k}\times$ Price Duration 20+M$_k$		0.357***	0.498**	0.481**	0.616***
		(0.108)	(0.218)	(0.221)	(0.224)
Tariff news shock$_{i,k}\times$ Modal Brexit Year=2020$_i$		-0.416***	-0.449***	-0.445***	-0.482***
		(0.124)	(0.160)	(0.162)	(0.139)
Quarter fixed effects	No	No	No	Yes	Yes
Additional firm controls	No	No	Yes	Yes	Yes
Mean of Dependent Variable	1.078	1.078	1.078	1.078	1.078
R^2	0.002	0.012	0.022	0.032	0.036
Observations	1,010	1,010	1,010	1,010	1,010

Notes: The dependent variable is winsorised at the 5th and 95th percentiles. Additional controls include: Perceived probability of No Deal Brexit; Import cost share; Exporter status; Sectoral fixed cost share; Natural logarithms of employment and capital expenditure in March 2019. Standard errors are clustered at the industry (SIC2) level and reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

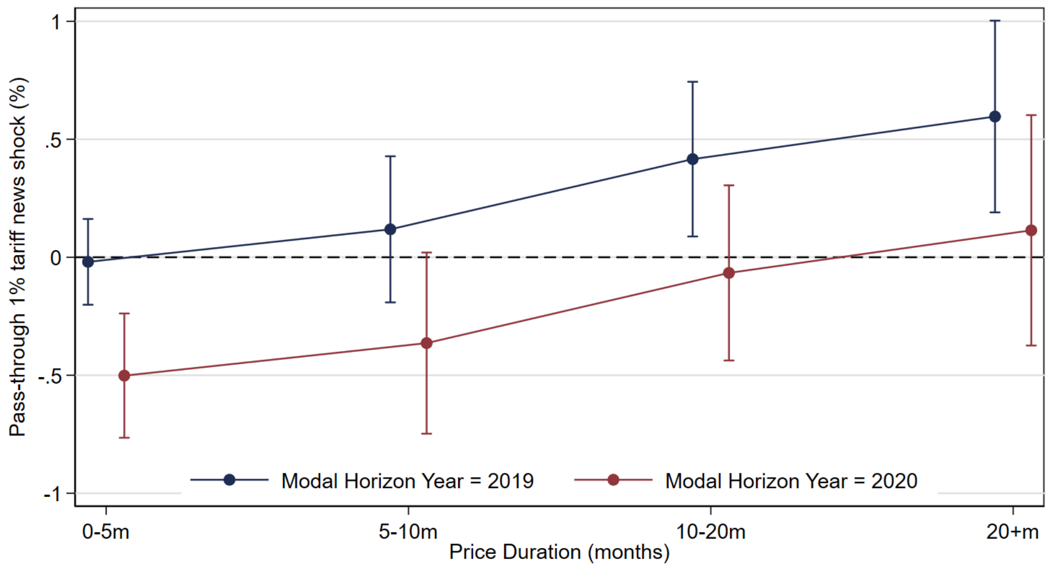
months. As for the perceived horizon of the news shock, we group firms into bins based on their survey-based responses on the probability of Brexit in each year (2019, 2020, 2021, 2022 and 2023+). We use the "modal Brexit year" variable based on the year to which firms assign the highest probability.[25] All in all, we create a dummy variable for each bin of adjustment frequency and perceived horizon, and interact them with our tariff news shock measure. The advantage of this approach based on dummy variables is that it allows to remain agnostic about any non-linearities in the relationship between iPT and the cross-sectional characteristics.[26]

Table 1 presents the results, pooling the estimates across the three quarters after the announcement. Column 1 shows that, on average, when pooled across all firms and quarters, there is no statistically significant intertemporal pass-through. As it turns out, however, this masks substantial heterogeneity across firms. In Column 2, we add interactions with dummy variables for whether a firm belongs to the bin with stickiest prices (20+ months) and whether its modal Brexit year is 2020 (as opposed to 2019). Two key empirical results emerge. First, conditional on price adjustment, belonging to the stickiest price category increases iPT by 0.36,

[25]In our regression sample, we only consider two bins of modal Brexit years: 2019 and 2020. This is because there are very few firms with modal perceived years beyond 2020.

[26]In Table A8 in the Online Appendix we consider an alternative approach and instead of dummy variables use continuous measures of adjustment frequency and perceived probability of Brexit in a given year. Our results are robust to such an alternative specification.

Figure 7: Estimated iPT: Interaction with price durations and perceived Brexit horizon



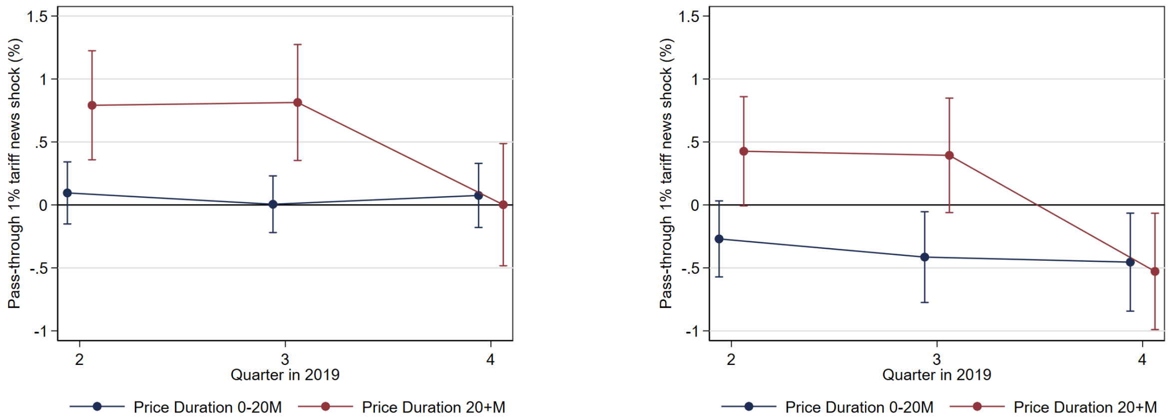
Notes: The figure presents the effect of the tariff import cost shock on firm prices, with interactions for average price duration and expected Brexit date. Standard errors are clustered at the SIC2 level and 90% confidence intervals are reported. The regression results are reported in Column 5 of Table 1.

a highly statistically significant difference. Second, all else equal, believing that Brexit will happen in 2020 as opposed to 2019, so that the shock is perceived to be further away, decreases iPT by a statistically significant amount of 0.42. Columns 3 and 4 further show that the key results above survive the addition of firm-level controls and quarter fixed effects; moreover, the effects become larger in magnitude.

The results in Columns 2-4 of Table 1 single out the firms with the largest average price duration (20+ months). While potentially strengthening the identification, this is a rather extreme distinction, as only around 6% of firms have price duration above 20 months. In Column 5 we allow for a finer breakdown of the average price duration variable, considering four bins: 0-5 (the base category), 5-10, 10-20, and 20+ months. The results suggest that the iPT is increasing monotonically as one moves to bins with larger average price duration, and the interactions are statistically significant for the 10-20 month and 20+ month categories. Crucially, the negative effect on iPT coming from having 2020 as the modal Brexit year remains unaffected by the presence of finer price duration bins. To visualize our key results, Figure 7 plots the estimated iPTs by categories of average price duration and perceived shock horizon, along with 90% confidence intervals. The only statistically significant positive iPT estimates are for firms that perceive Brexit to be most likely in 2019 *and* have average price durations of

Figure 8: Estimated iPT: Effects by quarter, price duration and perceived Brexit horizon

(a) Modal Horizon = 2019 **(b)** Modal Horizon = 2020



Notes: Panel (a) shows the quarter-by-quarter estimated iPT, for firms with different price duration and 2019 modal Brexit. Panel (b) shows the quarter-by-quarter estimated iPT, for firms with different price duration and 2020 modal Brexit The results are based on Column 3 of Table A6.

10-20 months (iPT = 0.42) or 20+ months (iPT = 0.59).

Estimation by sample quarters The specifications above allow for particular dimensions of cross-sectional heterogeneity in iPT, while pooling the estimates across the quarters following the announcement. However, we would also like to check whether our key results on the effect of adjustment frequency and perceived shock horizon hold quarter-by-quarter. In Figure 8 we report quarter-by-quarter iPT estimates, additionally allowing for heterogeneity in adjustment frequency and perceived shock horizon. Panel (a) shows the result for firms with a modal Brexit horizon of 2019; Panel (b) shows the results with a modal Brexit horizon of 2020. Three findings stand out. First, across both panels, we find stronger responses for firms with stickiest prices (red coefficients). Second, the effects are stronger for firms with a 2019 modal Brexit horizon. Third, the positive and statistically significant iPTs are concentrated in the first two quarters following the announcement. By 2019Q4, we no longer find positive statistically significant iPTs. Table A6 in the Online Appendix shows all the estimated coefficients and standard errors used to construct Figure 8.

3.3 Estimated iPT: State- vs. time-dependent pricing

We now investigate the empirical relationship between iPT and firms' motivation behind price changes. Recall that in Section 2.4 we introduced a survey-based measure of state- versus time-dependent pricing, which we summarized at the sectoral level by the fraction of firms of each type in a given 2-digit category. In this subsection we use the sectoral measure of state- vs. time dependence to evaluate whether, all else equal, iPT changes with the type of pricing.[27]

Table 2 presents the estimation results. In Column 1 we add an interaction between our tariff news shock and the sector-specific fraction of state-dependent firms. One can see that having a larger fraction of state-dependent firms leads to a statistically significant reduction in iPT. In particular, for every additional 10 percent of state-dependent firms, the estimated iPT falls by 0.09. The specification in Column 1 imposes a linear relationship between the sectoral measure of state dependence and iPT. We relax this restriction in Columns 2-4 by adding interactions with dummy variables for discrete splits at the 66th, 50th, and 33rd percentiles of the state dependence measure. It follows that even with discrete splits, iPT falls by statistically significant amounts as one increases the degree of state dependence in pricing. Crucially, Column 5 shows that the key result survives adding interactions with the average frequency of price adjustment and the perceived horizon of the shock.

3.4 Estimated iPT: Role of shock size

Recall that in Figure 4 we document a wide heterogeneity in the sizes of our tariff news shocks. While all the shocks are either zero or negative, with the majority being up to -1% in size, some of the firms are experiencing shifters in their future marginal costs that are as large as -7% or -10%. In this subsection we investigate whether there exists a non-linear relationship between iPT and shock size.

We document the estimation results in Table 3. In Column 1, we add an interaction with a dummy variable for whether tariff shocks are below the 25th percentile (i.e., most negative shocks). We find that iPT is significantly lower for these most negative shocks. Column 2 confirms this when isolating just the bottom 15th percentile of shocks, and Column 3 shows that this remains the case even when introducing the additional interactions with the expected shock horizon and the average price duration. Both of those remain statistically significant as in Table 1. In Column 4, we add an interaction with the absolute value of the tariff news shock. This allows us to flexibly test whether the estimated iPT is non-linear. We find a negative statistically significant coefficient on the absolute value of the tariff shock measure. In

[27] The survey question regarding motivation behind price changes was only introduced in 2023, whereas our estimation sample covers the year of 2019. Since the overlap between firms surveyed in 2019 and 2023 is imperfect, we do not use firm-level measures of state- vs time-dependent pricing to avoid losing observations. The sector-level measure allows us to keep the sample unchanged relative to the other empirical exercises, facilitating greater comparability.

Table 2: Estimated iPT: Interactions with state/time dependent pricing

	(1)	(2)	(3)	(4)	(5)
Dependent variable:	100 × Δ log(Price Level)				
Sample:	2019Q2 - 2019Q4				
Tariff news shock$_{i,k}$	0.519*	0.251**	0.185**	0.147	0.253*
	(0.270)	(0.105)	(0.090)	(0.103)	(0.126)
Tariff news shock$_{i,k}$× State-dependent$_k$	-0.860**				
	(0.404)				
Tariff news shock$_{i,k}$× State-dependent$_k^{p66}$		-0.441***			
		(0.096)			
Tariff news shock$_{i,k}$× State-dependent$_k^{p50}$			-0.342***		
			(0.089)		
Tariff news shock$_{i,k}$× State-dependent$_k^{p33}$				-0.259**	-0.228*
				(0.117)	(0.130)
Tariff news shock$_{i,k}$× Price Duration 20+M$_k$					0.443*
					(0.223)
Tariff news shock$_{i,k}$× Modal Brexit Year=2020$_i$					-0.436***
					(0.151)
Quarter fixed effects	Yes	Yes	Yes	Yes	Yes
Additional controls	Yes	Yes	Yes	Yes	Yes
Mean of Dependent Variable	1.087	1.087	1.087	1.087	1.087
R^2	0.025	0.031	0.027	0.024	0.034
Observations	1,000	1,000	1,000	1,000	1,000

Notes: The dependent variable is winsorised at the 5th and 95th percentiles. Additional controls include: Perceived probability of No Deal Brexit; Import cost share; Exporter status; Sectoral fixed cost share; Natural logarithms of employment and capital expenditure in March 2019. Standard errors are clustered at the industry (SIC2) level and reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

particular, for every additional 1% in the magnitude of the shock, the estimated iPT falls by 0.05. In other words, as the news shocks become bigger, the impact price response grows *less* than proportionally in magnitude.

In Figure 9, we provide a quantification exercise based on the coefficient estimates in Column 4 of Table 3. Specifically, we show what the estimated iPT would be, depending on the linear vs. nonlinear specification, for the tariff news shocks in our sample. In this exercise, we focus on firms with a price duration above 20 months and with a 2019 modal Brexit horizon. For smaller shocks, the pass-through grows essentially linearly in the size of the shock, implying that iPT is almost size-independent. However, as the shocks grow in magnitude, the non-linearity becomes meaningful. For the largest shock in our sample (around -10% in size), the linear estimated pass-through to prices is around -9% (iPT = 0.90), whereas the non-linear iPT is around -3.5% (iPT = 0.35).

Table 3: Estimated iPT: Effect of tariff news shock size

	(1)	(2)	(3)	(4)
Dependent variable:	$100 \times \Delta \log$(Price Level)			
Sample:	2019Q2 - 2019Q4			
Tariff news shock$_{i,k}$	1.089**	0.638	0.859*	0.451
	(0.474)	(0.469)	(0.502)	(0.283)
Tariff news shock$_{i,k}\times$ Tariff news shock $<$ 25 Pctile	-1.185**			
	(0.446)			
Tariff news shock$_{i,k}\times$ Tariff news shock $<$ 15 Pctile		-0.721*	-0.842*	
		(0.408)	(0.433)	
Tariff news shock$_{i,k}$ $\times$ | Tariff news shock$_{i,k}$|				-0.047*
				(0.026)
Tariff news shock$_{i,k}\times$ Modal Brexit Year=2020$_i$			-0.478**	-0.528**
			(0.181)	(0.197)
Tariff news shock$_{i,k}\times$ Price Duration 20+M$_k$			0.402*	0.400*
			(0.236)	(0.226)
Quarter fixed effects	Yes	Yes	Yes	Yes
Additional controls	Yes	Yes	Yes	Yes
Mean of Dependent Variable	1.078	1.078	1.078	1.078
R^2	0.027	0.025	0.036	0.034
Observations	1,010	1,010	1,010	1,010

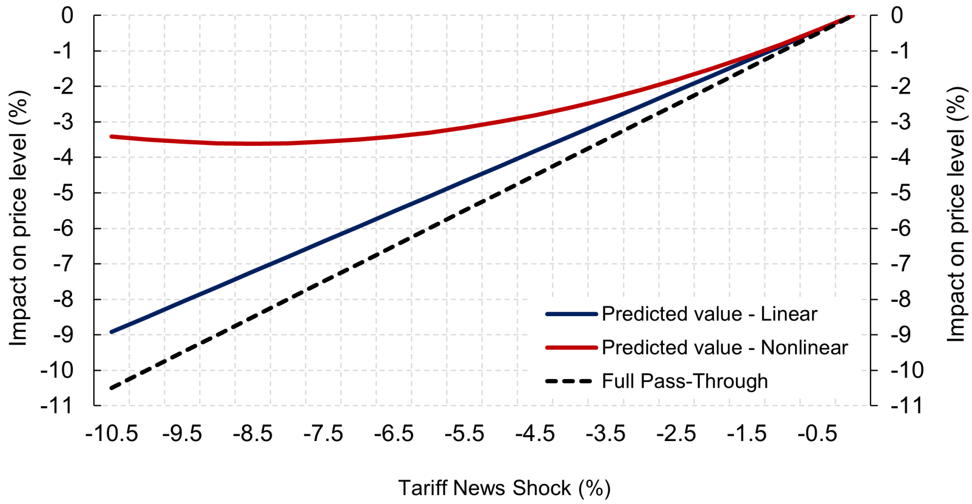
Notes: The dependent variable is winsorised at the 5th and 95th percentiles. Additional controls include: Perceived probability of No Deal Brexit; Import cost share; Exporter status; Sectoral fixed cost share; Natural logarithms of employment and capital expenditure in March 2019. Standard errors are clustered at the industry (SIC2) level and reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

3.5 Additional results and robustness checks

We perform several additional exercises and robustness checks, which we summarize here. The detailed results are reported in the additional figures and tables in the Online Appendix.

Responses of firm-level quantities While our main results concern the pass-through of tariff news shocks to firm-level prices, we also estimate the responses of firm-level quantities. This is useful, since it both provides novel evidence on how quantities respond to news shocks, as well as strengthens our interpretation of tariff news shocks are pure shifters to expected future marginal costs. In Table A4, we analyze the effect of the tariff news shocks on firms' employment, capital expenditure and real sales. The dependent variable is the natural logarithm of these responses, and we control for its level before the March 2019 announcement in each specification. We find that for the firms with stickiest prices, an increase in the expected *future* marginal cost leads to statistically significant decreases in *current* real sales. We also find a negative coefficient on employment, though not statistically significant. Combined with our main results this suggests that, conditional on adjustment, firms with lowest adjustment

Figure 9: Estimated iPT: Linear vs. non-linear effects



Notes: The figure presents the linear and nonlinear predicted effect of the tariff import cost shock on firm prices. The results are based on Column 4 from Table 3. The predicted values are for firms who have a modal expected Brexit year of 2019.

frequency increase current prices following a rise in the expected future costs, which in turn leads to a contemporaneous drop in their real sales and employment. We believe this provides further evidence that our tariff news shock indeed capture idiosyncratic cost shifters.[28]

Heterogeneity by firm size and industry concentration Several papers have emphasized the importance of firm size and competition for the pass-through of cost shocks (Amiti et al., 2019; Godl-Hanisch and Menkhoff, 2024). In Table A11, we test for the heterogeneity in iPT by various measures of firm size, *in addition to* the heterogeneity by price duration and expected Brexit horizons. We find no significant interaction with firm employment (Column 1-4), firm sales (Column 5), or capital expenditure (Column 6), all measured before the tariff announcement. Meanwhile, the effects of Brexit horizons and price duration remain highly significant across all specifications. In Table A12, we test whether there is any additional heterogeneity by sectoral concentration. We use industry sales concentration (HHI) indices from Savagar et al. (2024) for 2018 as an additional interaction. We do not find a significant difference in iPT by sectoral concentration, while the main effects remain robust. Overall, we conclude that there do not appear to be significant differences in iPT by firm size or industry concentration.

Accounting for non-EU imports The construction of our tariff news shock focuses on the share of firm costs imported from the EU. However, strictly speaking, the tariff schedule

[28]Table A5 estimates the results on employment, capital expenditure, and real sales by quarter. Figures A14, A15, and A16 present the corresponding results with coefficient plots.

announcement in March 2019 would have applied to all trading partners (who did not have an existing trade deal with the UK) in the case of a no-deal Brexit. Therefore, even for imports from those countries which were subject to the MFN tariff already, there would have been a disinflationary "news shock" component in the event of a "No-Deal" Brexit. Quantifying this is challenging, because it requires data on firm-level import cost shares from each individual non-EU country which did not have an existing trade deal with the UK. To test how sensitive our results may be to non-EU imports, in Table A13 we exclude firms with non-zero imports from non-EU countries. On average, firms estimate that 11.7% of their costs were imports from non-EU countries in 2016Q1 (versus 16.8% from EU countries). By excluding these, we focus only on the component of the news shock which originated from EU imports. This decreases our sample size, but the main results on price duration and expected Brexit horizon remain statistically significant and with a similar magnitude.

Pre-trends test We conduct additional exercises in order to check whether the shocks that we construct are indeed random and unpredictable at the time of the announcement. As a first exercise, in Table A7 we estimate the effect of our tariff news shocks on prices in the three quarters *before* the March 2019 announcement (2018Q1-2018Q4). Crucially, we find no statistically significant pass-through, regardless of the average frequency of price adjustment or the perceived Brexit horizon. In Tables A9 and A10 we further study whether the key ingredients used in shock construction, namely the perceived probability of "No-Deal" outcome and the perceived horizon of Brexit, are predictable by any of the firm-level characteristics before the announcement.

Placebo tests We also conduct a placebo test, which aims to pick up whether our estimates truly capture the effect of tariff news shocks, as opposed to spurious effects in a finite sample. In particular, we construct a large number of synthetic datasets, where we randomly reassign the tariff news shocks across firms. We then re-estimate our baseline specification for each of the synthetic datasets. Figures A17, A18, A19 and A20 show the distributions of estimated coefficients and t-stats on the price duration indicator (20 months+ or less), the perceived shocks horizon (modal Brexit year of 2020 or not), the measure of state dependence in pricing and the size of the shock across the synthetic datasets. Crucially, all the distributions are centered at zero, suggesting that no effect is found on average when shocks are randomly reassigned. Moreover, our actual estimates under the correctly assigned shocks are in the tail of the distributions, indicating that is is very unlikely that they are obtained spuriously by chance.

4 General equilibrium setup

We now turn to theoretical analysis of iPT in the context of a general equilibrium model that captures the empirically-relevant dimensions of heterogeneity. First, we analytically derive model-based iPT for small shocks, and compare their properties with those of empirical pass-through. Second, we study how micro heterogeneity in iPT affects the forward-lookingness of macroeconomic aggregates, such as CPI inflation. We find that heterogeneity in adjustment frequencies and perceived shock horizons *amplifies* the degree of aggregate foresight. Third, we study iPT under large shocks using a fully non-linear solution to a state-dependent pricing setup. Our results suggest that, consistent with the empirically observed size dependence, iPT falls for larger shocks. All proofs are given in Online Appendix C.

4.1 Model overview

Time is discrete, with outcomes in $t = -1$ exogenously given, and outcomes in $t \geq 0$ determined endogenously. There are three types of agents. First, a continuum of infinitely lived households. Second, a continuum of monopolistically competitive firms, owned by the households, where each firm belongs to one, and only one, of the N sectors; let the set of all firms in sector i be $\Phi_i, \forall i = 1, 2, ...N$. Third, a government, comprising of a central bank setting monetary policy, and a fiscal authority which collects taxes from firms and rebates them to households in lump-sum fashion.

4.2 Agent optimization and market clearing

Let us first introduce the problem solved by each of the agent types.

Households A continuum of infinitely lived households populates the economy and owns all the firms. The representative household makes choices to maximize the expected lifetime utility:

$$\max_{\{C_t, L_t, B_{t+1}\}_{t\geq 0}} \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t u(C_t, L_t) \tag{6}$$

subject to period-by-period budget constraints:

$$P_t^c C_t + \mathbb{E}_t\left[\Xi_{t,t+1} B_{t+1}\right] \quad \leq \quad B_t + W_t L_t + \sum_{i=1}^{N} \int_{j \in \Phi_i} \Pi_{it}(j) dj + T_t, \quad \forall t \tag{7}$$

where $\beta \in (0,1)$ is the discount factor, C_t is aggregate consumption, P_t^c is consumption price index (defined below), L_t is labor supply, B_{t+1} is the payoff of securities purchased at time t, $\Xi_{t,t+1}$ is the stochastic discount factor between periods t and $t+1$, $\Pi_{it}(j)$ denotes nominal profits of firm j in sector i and T_t are lump-sum transfers from the government.

The total consumption C_t is in turn given by an aggregator over sectoral varieties:

$$C_t = \mathcal{C}(C_{1,t}, ..., C_{N,t}) \tag{8}$$

where $\mathcal{C}(.)$ is homogeneous of degree one and increasing in all the inputs. Minimizing the total expenditure on sectoral varieties subject to (8) delivers the ideal price index for households, or the Consumption Price Index (CPI): $P_t^C = \mathcal{P}^{\mathcal{C}}(P_{1,t}, ..., P_{N,t})$, where $P_{i,t}$ is the price index of a sector i (defined below).

As for the final consumption of a sectoral variety, it is given by an aggregator over final demands for the output of each firm in that sector:

$$C_{i,t} = \left\{ \int_0^1 \left[\zeta_{i,t}(j) C_{i,t}(j)\right]^{\frac{\epsilon-1}{\epsilon}} dj \right\}^{\frac{\epsilon}{\epsilon-1}} \tag{9}$$

where $\epsilon > 1$ is the within-sector elasticity of substitution and $\zeta_{i,t}(j)$ is a firm-specific quality process which follows a random walk in logs: $\log \zeta_{i,t}(j) = \log \zeta_{i,t-1}(j) + \sigma_i \varepsilon_{i,t}(j)$. Expenditure minimization on within-sector varieties subject to (9) delivers the firm-specific demand schedule: $\zeta_{i,t}(j) C_{i,t}(j) = \left(\frac{\overline{P}_{i,t}(j)}{\zeta_{i,t}(j) P_{i,t}}\right)^{-\epsilon} C_{i,t}$, as well as the sectoral price index: $P_{i,t} = \left[\int_0^1 \left\{\overline{P}_{i,t}(j)/\zeta_{i,t}(j)\right\}^{1-\epsilon} dj\right]^{\frac{1}{1-\epsilon}}$, where $\overline{P}_{i,t}(j)$ is the nominal price of firm j in sector i at time t.

Firms On the production side, our economy consists of N sectors, indexed by $i \in \{1, 2, ..., N\}$, and a continuum of monopolistically competitive firms, indexed by j, that each belongs to one sector only. The production function of firm j that operates in sector i is given by:

$$\zeta_{i,t}(j) Y_{i,t}(j) = L_{i,t}(j), \tag{10}$$

where $Y_{i,t}(j), L_{i,t}(j)$ are total output and labor input respectively, whereas $\zeta_{i,t}(j)$ is the idiosyncratic quality process introduced before. Standard cost minimization problem subject to (10) delivers the firm-specific marginal cost function: $MC_{i,t}(j) = \zeta_{i,t}(j)W_t$.

The nominal profit of a firm j in sector i at time t can be written as:

$$\Pi_{i,t}(j) = \left[(1-\tau_i)\overline{P}_{i,t}(j) - \{1+\varrho_{i,t}(j)\}MC_{i,t}(j)\right] Y_{i,t}(j) \tag{11}$$

where τ_i is the time-invariant sectoral sales tax and $\varrho_{i,t}(j)$ is an idiosyncratic cost shifter, which can also be interpreted as a firm-specific time-varying payroll tax.[29] Letting $P_{i,t}(j) \equiv \overline{P}_{i,t}(j)/\zeta_{i,t}(j)$ be the firm's *quality-adjusted* price, one can rewrite nominal profits as:

$$\Pi_{i,t}(j) = [(1-\tau_i)P_{i,t}(j) - \{1+\varrho_{i,t}(j)\}W_t] \times (P_{i,t}(j)/P_{i,t})^{-\epsilon} \times Y_{i,t}. \tag{12}$$

Resetting the *nominal* price involves paying a possibly random idiosyncratic menu cost $\kappa_{i,t}(j)$, paid in units of labor. If the nominal price remains unchanged, the quality-adjusted price evolves according to (letting $p_{i,t}(j) \equiv \log P_{i,t}(j)$):

$$p_{i,t}(j) = p_{i,t-1}(j) + \log\left(\frac{\overline{P}_{i,t-1}(i)}{\zeta_{i,t}(j)}\right) - \log\left(\frac{\overline{P}_{i,t-1}(j)}{\zeta_{i,t-1}(j)}\right) = p_{i,t-1}(j) - \sigma_i\varepsilon_{i,t}. \tag{13}$$

Consider a firm with a real quality adjusted price p at the end of period t. Letting $p_+ \equiv p - \sigma_i\varepsilon_{i,t+1}(j)$, one can describe the firm's value with the following Bellman equation:

$$\begin{aligned} V_{i,t}(p) &= \Pi_{i,t}(p) + \\ &+ \mathbb{E}_t\Xi_{t,t+1}\left[\{1-\eta_{i,t+1}(p_+)\}V_{i,t+1}(p_+) + \eta_{i,t+1}(p_+)\left(\max_{p'} V_{i,t+1}(p') - \kappa_{i,t+1}(j)\right)\right], \end{aligned} \tag{14}$$

where $\eta_{i,t}(.)$ is the *pricing hazard* function, which gives the (endogenous) probability of price adjustment as a function of the aggregate/sectoral state and the firm-level quality-adjusted price. The optimal reset price is the one which maximizes the firm's value: $p^* = \arg\max_p V_{i,t}(p)$.

Government policy The central bank conducts monetary policy by setting the path of money supply $\{M_t\}_{t\geq 0}$, which enters the representative household's problem through a cash-in-

[29]We set $\{\tau_i\}_{i=1}^N$ to normalize each sectoral price index to one in steady state.

advance constraint:

$$P_t^C C_t \le M_t. \tag{15}$$

The fiscal authority collects sales taxes from firms and rebates them to households in a lump-sum fashion: $T_t = \sum_{i=1}^N \tau_i \int_0^1 \overline{P}_{i,t}(j) Y_{i,t}(j) dj.$

Market clearing In addition to the optimality conditions above, equilibrium in the economy is characterized by market-clearing conditions in the asset market: $B_t = 0$; the labor market: $L_t = \sum_{i=1}^N \int_{j\in\Phi_i} [L_{i,t}(j) + \kappa_{i,t}(j)\eta_{i,t}\{p_{i,t}(j)\}]dj$; and the goods markets: $Y_{i,t}(j) = C_{i,t}(j), \quad \forall i, \forall j \in \Phi_i$.

4.3 iPT under small shocks

For our first set of results, we analytically derive the model-based iPT for small shocks. To do that, we rely on the purely time-dependent Calvo (1983) setting (for some value of the exogenous time-invariant probability of adjustment) as a first-order approximation for a broader range of pricing models, corresponding to different functional forms of $\eta_{i,t}(p)$ and menu costs $\kappa_{i,t}(j)$. For this subsection, we therefore impose the following assumption on $\eta_{i,t}(p)$ and the menu costs $\kappa_{i,t}(j)$:

Assumption 1 (Calvo, 1983)**.** *The pricing hazard $\eta_{i,t}(p)$ and the menu cost $\kappa_{i,t}(j)$ are:*

$$\eta_{i,t}(p) = 1 - \alpha_i, \quad \forall i, t, p \qquad\qquad \kappa_{i,t}(j) = 0, \quad \forall i, t \tag{16}$$

where $\{\alpha_i\}_{i=1}^N$ are fixed sector-specific parameters, each belonging to $[0, 1)$.

In other words, at the firm-level, all price adjustments are free and occur randomly with a sector-specific and time-invariant probability. In such a setup, it is possible to obtain a closed-form expression for the optimal reset price in every sector. Moreover, after log-linearizing around a deterministic symmetric zero-inflation steady state, the (log) optimal reset price for a firm j in sector i is:

$$p_{i,t}^*(j) \quad = \quad (1 - \beta\alpha_i) \sum_{d\ge 0} (\beta\alpha_i)^d \mathbb{E}_t(\varrho_{i,t+d}(j) + w_{t+d}) \tag{17}$$

where $\varrho_{i,t}(j)$ is the exogenous i.i.d. mean-zero idiosyncratic cost shifter and w_t is the aggregate nominal wage (in log deviations from steady state).

Now suppose that the firm j in sector i receives the following news about its idiosyncratic cost shifter at time t: between t and $t+d_i-1$ it stays at zero, between $t+d_i$ and $t+d_i+D$ it rises to one, after which from $t+d_i+D+1$ onward it stays at zero forever. Noting that the aggregate nominal wage (now or in the future) remains unaffected by the idiosyncratic news shock, we can formally derive the firm-level iPT following such a future cost shifter between periods $t+d_i$ and $t+d_i+D$:

Proposition 1 (iPT for small shocks)**.** *Suppose Assumption 1 holds. Then iPT of a firm in sector i after a unit shifter in its future marginal costs between periods d_i and d_i+D is:*

$$\text{iPT}_{d_i|d_i+D} = (\beta\alpha_i)^{d_i} - (\beta\alpha_i)^{d_i+D+1}. \tag{18}$$

Further, it follows that:

$$\frac{\partial \text{iPT}_{d_i|d_i+D}}{\partial d_i} < 0 \qquad \frac{\partial^2 \text{iPT}_{d_i|d_i+D}}{\partial d_i^2} > 0 \tag{19}$$

$$\frac{\partial \text{iPT}_{d_i|d_i+D}}{\partial \alpha_i} > 0, \quad \alpha_i < \overline{\alpha}_i \qquad \frac{\partial^2 \text{iPT}_{d_i|d_i+D}}{\partial \alpha_i^2} > 0, \quad \alpha_i < \tilde{\alpha}_i \tag{20}$$

where $\overline{\alpha}_i \equiv \frac{1}{\beta}\left[\frac{d_i}{d_i+D+1}\right]^{\frac{1}{D+1}}$ *and* $\tilde{\alpha}_i \equiv \frac{1}{\beta}\left[\frac{d_i(d_i-1)}{(d_i+D+1)(d_i+D)}\right]^{\frac{1}{D+1}}$.

The proposition above establishes two sets of results. First, it formalizes how (small) changes in the perceived shock horizon and frequency of non-adjustment affect iPT. Intuitively, as the shock is perceived to be further out (higher d), iPT monotonically falls. This happens since firms put less weight on cost changes that occur later, since the probability that the current reset price stays fixed all the way until that period is lower. The relationship between iPT and the probability of non-adjustment is more nuanced. As the probability of non-adjustment rises from zero up to a particular threshold value $\overline{\alpha}$, the iPT is rising. At the same time, beyond that threshold value iPT starts falling.[30]

Second, Proposition 1 also derives the *convexity* of iPT in the shock horizon and the probability of non-adjustment. It follows that iPT is always *convex* is the perceived shock horizon. In other words, as the shock horizon increases, the pass-through falls more than

[30]The non-monotonic relationship is driven by the temporary nature of the cost shifter. In particular, as α_i rises, there are two effects: on one hand, firms start putting a higher weight on marginal costs in periods between $t+d_i$ and $t+d_i+D$, which increases iPT; on the other hand, they also start putting a higher weight on periods $t+d_i+D$ and beyond, where the marginal costs are at their steady-state values, which lowers iPT. For $\alpha_i < \overline{\alpha}_i$, the first effect dominates and *vice versa*. Note that as the news shock becomes permanent ($D \to \infty$), iPT unambiguously rises for all $\alpha_i \in (0,1)$, since the second effect disappears.

linearly, suggesting that iPT is disproportionally large for small shock horizons. At the same time, iPT is convex in the non-adjustment probability as long as it is not too large, as pinned down by the threshold $\tilde{\alpha}$.[31] In other words, as the probability of non-adjustment rises from zero up to the threshold value, iPT increases more than linearly.

The formal results for the convexity of iPT in the shock horizon and probability of adjustment are key for pinning down the effect of heterogeneity in pass-through for the behavior of aggregates, such as CPI inflation. We elaborate on that point in the subsection below.

Micro to macro We now study the behavior of *aggregate* price indices following a common shifter to the expected future marginal costs. Suppose that *all* firms in sector i experience a news shock to their cost shifter at the same time. In this case, we can no longer abstract from the behavior of aggregates, such as the nominal wage, and need to take a stance on the form of household preferences and the response of monetary policy. For the former, we follow the menu cost literature and consider log-linear preferences over consumption and labor as in Golosov and Lucas (2007):

Assumption 2 (Golosov-Lucas preferences)**.** *The preferences over consumption and labor are given by:* $u(C_t, L_t) = \log C_t - L_t$.

As for monetary policy, we assume that the central bank stabilizes money supply in every period, $M_t = 1, \forall t$, or $m_t = 0$ in log-linear terms.

Assuming the economy is in steady state at $t-1$, one can write the first-order response of aggregate CPI inflation $\pi_t^C \equiv \Delta p_t^C$ to a common unit cost shifter with sector-specific perceived horizons as:

$$\pi_t^C \quad = \quad \sum_{i=1}^{N} \overline{\omega}_i^C p_{i,t} \quad = \quad \sum_{i=1}^{N} \overline{\omega}_i^C (1-\alpha_i) \left[(\beta\alpha_i)^{d_i} - (\beta\alpha_i)^{d_i+D+1} \right], \tag{21}$$

where $\overline{\omega}_i^C$ is the steady-state final consumption share for sector i.

We can now use the formal results on the *convexity* of iPT in the perceived shock horizons and adjustment frequency to pin down the effect of iPT heterogeneity on the degree of forward-lookingness of aggregate CPI inflation. First, note that since iPT is always convex in perceived shock horizons, a simple application of Jensen's inequality implies that heterogeneity in perceived horizons *amplifies* the impact response of aggregate CPI inflation:

[31] As before, the move from convexity to concavity is driven by the temporary nature of the cost shifter, delivering two opposing effects. As the cost shifter becomes permanent ($D \to \infty$), iPT is convex in all $\alpha \in (0, 1)$.

Table 4: Calibration of the model (UK, monthly data)

Parameter	Description	Value	Source/Target
N	Number of sectors	63	Data availability
β	Discount factor	0.997	Real rate of 3.6%
D	Duration of news shock	12	Temporary cost shifter
$\{d_i\}_{i=1}^N$	Horizon of news shock		DMP Brexit horizon
$\{\alpha_i\}_{i=1}^N$	Sectoral Calvo parameters		UK PPI microdata
$\{\overline{\omega}_i^C\}_{i=1}^N$	Steady-state final consumption shares		ONS IO Tables

Proposition 2 (Heterogeneity in news horizons)**.** *Suppose Assumption 1 and 2 hold, and further assume that $\alpha_i = \alpha, \forall i$. Then heterogeneity in perceived horizons* ***amplifies*** *the contemporaneous response of aggregate CPI to a common shifter in expected future marginal costs.*

Intuitively, the convexity of iPT in the horizon implies that the pass-through is disproportionally larger for shocks that are expected to arrive very soon. A means-preserving spread of perceived horizons exactly delivers those sectors with shocks that are expected to arrive much sooner than average, which then disproportionally drive the aggregate response.

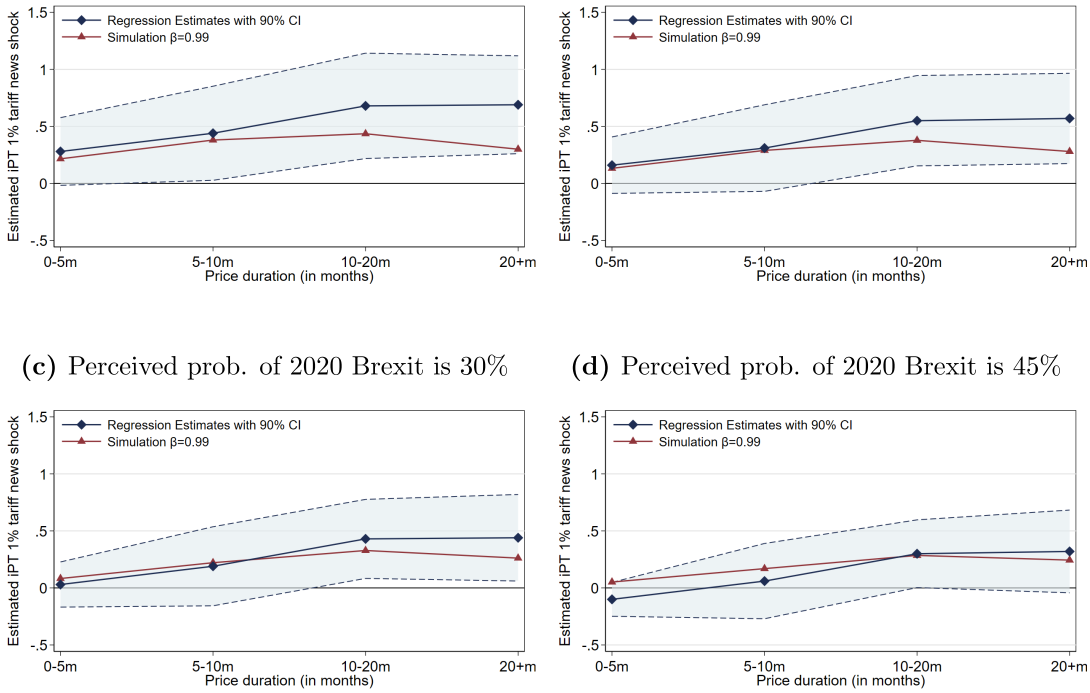
As for the effect of heterogeneity in adjustment frequencies, the result in more nuanced, since iPT is convex in resetting probabilities as long as they are not too low. Luckily, the ambiguity vanishes in the limit of permanent shifters ($D \to \infty$):

Proposition 3 (Heterogeneity in adjustment frequencies)**.** *Suppose Assumption 1 and 2 hold, and further assume that $d_i = d, \forall i$. Moreover, suppose that the future cost shifters are perceived to be permanent ($D \to \infty$) and $\alpha_i \in \left(0, \frac{d-1}{d+1}\right), \forall i$. Then heterogeneity in adjustment frequencies* ***amplifies*** *the contemporaneous response of aggregate CPI to a common shifter in expected future marginal costs.*

The theoretical results above apply to the broad class of models that can, up to first order, be approximated with the purely time-dependent pricing setup for some values of adjustment probabilities. As a next step, we would like to ascertain whether the canonical Calvo (1983) model, calibrated to the *actual* sector-specific frequencies of adjustment in the UK, is able to match the estimated iPT. Moreover, we would like to quantify the extent to which observed heterogeneity in adjustment frequencies and perceived horizons amplifies the forward-lookingness of aggregate inflation.

Quantitative exercises In order to quantify our key results, we calibrate our model to monthly data for $N = 63$ sectors of the UK economy. We calibrate the horizons $\{d_i\}_{i=1}^N$ of the

Figure 10: Estimated vs theoretical iPT



(a) Perceived prob. of 2020 Brexit is 0%

(b) Perceived prob. of 2020 Brexit is 15%

(c) Perceived prob. of 2020 Brexit is 30%

(d) Perceived prob. of 2020 Brexit is 45%

Notes: This figure reports estimated iPT versus model-based iPT under UK calibration, conditional on average price duration and the perceived probability of Brexit happening in 2020. The estimated iPT is based on Column 1 of Table A8. The theoretical iPT values apply the calibrated parameters from Table 4 (transformed to quarterly equivalents) to the iPT formula in Proposition 1.

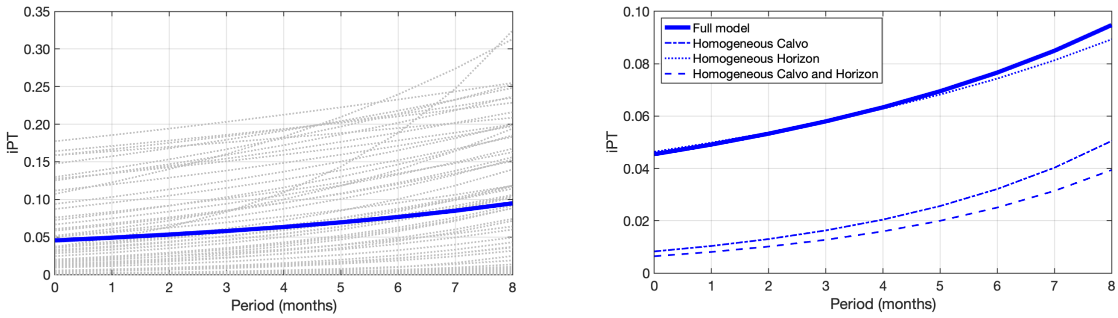
common unitary marginal cost shifter to match the sectoral averages of perceived number of months between March 2019 and Brexit in the DMP survey; the duration D of the news shock is set to 12 months, since the UK-proposed temporary tariff schedule was announced to be in effect for up to a year. As for the sector-specific frequencies of adjustment, we use the UK CPI/PPI product-level pricing data, just as in the econometric exercises. Table 4 summarizes our calibration strategy.

For our first quantitative exercise, we compare the theoretical iPT in the calibrated model with our econometric estimates. In Figure 10, we report our econometrically estimated iPT for different perceived probabilities of Brexit in 2020 and average price durations, as well as their theoretical equivalents in the calibrated model.[32] Notably, all the theoretical iPT values lie within the 90% confidence bands of their estimated equivalents. Such results suggest that the canonical time-dependent pricing model of Calvo (1983), when calibrated to observed frequencies of adjustment, generates iPT that cannot be rejected based on the econometric estimates.

[32]To construct econometrically estimated iPT, we use the specification in Column 1 of Table A8, which allows for a continuous measure of the perceived probability of 2020 Brexit. The theoretical iPT apply the calibrated parameters to the iPT formula in Proposition 1.

Figure 11: Heterogeneity in micro iPT

(a) Sector-specific iPTs and their average **(b)** Role of micro heterogeneity for average iPT



Notes: Panel (a) shows the responses of sector-specific optimal reset prices to the common cost shifter under the calibration in Table 4. Panel (b) reports the average optimal reset price under different dimensions of heterogeneity.

Crucially, the iPT estimates in Figure 10 condition on cross-sectional characteristics, while pooling across shock sizes. We defer theoretically matching the size dependence properties of the estimated iPT to Section 4.4.

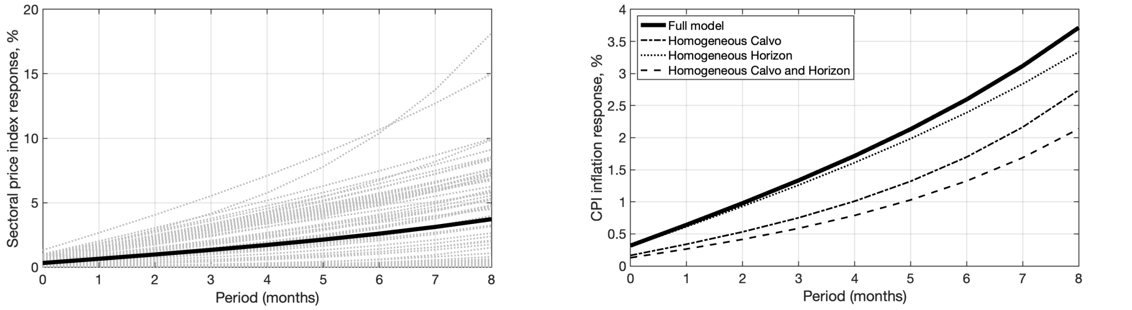
Our second exercise quantifies the implications of cross-sectional hetorogeneities for firm-level iPT. In Panel (a) of Figure 11 we show the firm-level iPT for the 63 sectors of our economy.[33] As one can see, there is a wide heterogeneity in iPT, varying between close to zero and almost 0.20 on impact, with an average of around 0.05. In Panel (b) of Figure 11 we quantify the role of heterogeneity in frequencies of price adjustment the perceived shock horizons. It follows that an otherwise identical economy with homogeneous frequencies of adjustment has iPT that is lower by a factor of five. At the same time, heterogeneity in the perceived horizons has a minor quantitative role.

For our third quantitative exercise, we look at the implications of micro iPT heterogeneity for the forward-lookingness of aggregate CPI inflation. In Panel (a) of Figure 12 we show the responses of sectoral price indices, as well as the aggregate CPI, of our economy to the expected future cost shifters. As one can see, there is wide heterogeneity in the responses, with the average sectoral price index response given by 0.5% on impact. In Panel (b) of Figure 12 we quantify the role of heterogeneity in frequencies of price adjustment the perceived shock horizons. It follows that an otherwise identical economy with homogeneous frequencies of adjustment has iPT that is lower by a factor of two. At the same time, heterogeneity in the perceived horizons has a minor quantitative role.

[33]Crucially, iPT is common for all firms within a given sector, due to the common sector-specific optimal reset price. Therefore, any differences in iPT across sectors are attributed to sector-specific dimensions of heterogeneity, such as the frequency of price adjustment and the perceived horizon of the news shock.

Figure 12: Aggregate pass-through

(a) Sectoral price indices and aggregate CPI **(b)** Role of micro heterogeneity for aggregate CPI



Notes: Panel (a) shows the responses of sectoral price indices to the common cost shifter under the calibration in Table 4. Panel (b) reports the response of aggregate CPI inflation under different dimensions of heterogeneity.

4.4 iPT under large shocks

We now turn to the analysis of iPT under large shocks. Instead of relying on a time-dependent model as a first-order approximation, we now obtain fully non-linear solutions to a random menu cost model. In particular, we abstract from heterogeneity across sectors and consider a one-sector ($N = 1$) version of our model with CalvoPlus pricing (Nakamura and Steinsson, 2010):

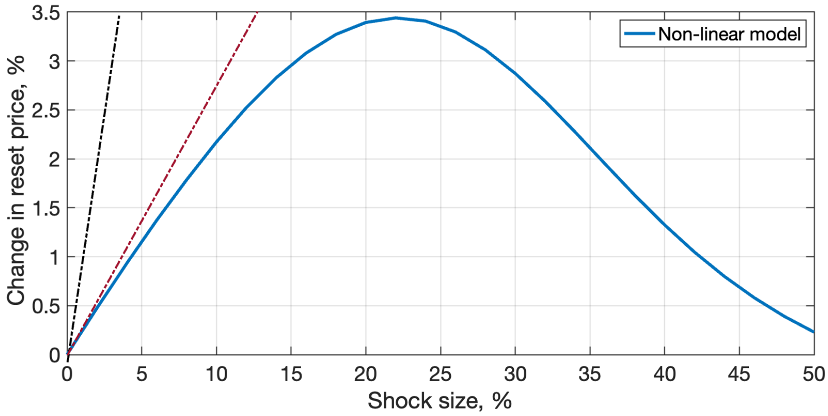
Assumption 3 (Nakamura and Steinsson, 2010)**.** *The pricing hazard $\eta_t(p)$ and the menu cost $\kappa_t(j)$ are:*

$$\eta_t(p) = \mathbf{1}\left\{\mathcal{L}_t(p) > 0\right\}, \quad \forall t, p \qquad\qquad \kappa_t(j) = \begin{cases} 0 & \text{with prob.} \quad \ell \\ \overline{\kappa} & \text{with prob.} \quad 1-\ell \end{cases} \qquad \forall t \tag{22}$$

where $\mathcal{L}_t(p) \equiv \max_p V_t(p) - V_t(p) - \kappa_t(j)$ is the (net) value gain from choosing to adjust the price and ℓ is the probability of drawing a zero menu cost.

To calibrate the model for one sector, we need to pin down three parameters of the pricing problem: the size of the non-zero menu cost κ, the probability of drawing a zero menu cost ℓ and the standard deviation of the quality innovation σ. We estimate those parameters by jointly matching three aggregate pricing moments in steady-state: frequency of adjustment, standard deviation of non-zero price changes and kurtosis of non-zero price changes. We consider the values of those pricing moments for the UK as reported in Blanco et al. (2024). In this setup, we also need to take a stance on the within-sector elasticity of substitution, which we set to $\epsilon = 3$. In order to obtain fully non-linear solutions, we construct perfect foresight transitions in the space of sequences, as in Karadi and Reiff (2019) and Ghassibe and Nakov (2024).

Figure 13: iPT under large shocks



Notes: This figure shows the response of the optimal reset price to 6-month ahead news shocks about future costs of different sizes. The responses are generated under a fully non-linear perfect foresight solution to the one-sector model with CalvoPlus pricing. The blue line is the fully non-linear optimal reset price, the red line is optimal reset price under linearization, the black line corresponds to full pass-through.

To assess the degree of size dependence in model-based iPT, we compute contemporaneous responses of the optimal reset price to 6-months ahead news about future permanent cost shifters of different sizes. The results are summarized in Figure 13. One can see that for shock sizes of up to 3% or so, the optimal reset price change increases linearly with the shock size, implying a constant iPT of around 0.20. For large shocks, the response of the optimal price starts to flatten out very sharply and in fact reaches a plateau for a shock of around 23%. Beyond that point, large shocks lead to *decreases* in the contemporaneous pass-through, reaching almost zero response for a 50% shock. This is because for large shocks firms endogenously revise their perceived probability of adjustment upwards in the period when the shock actually arrives. In the limit, as the shock becomes extremely large, they revise the probability upwards to one, delivering iPT of zero.

Overall, our results suggest that the decline of iPT with shock sizes that we document empirically can be rationalized in the context of a standard random menu cost model that is solved fully non-linearly.

5 Conclusion

Forward-looking pricing and inflation expectations have long been central to macroeconomic theory and policy. Some of the key prescriptions for central banking, such as benefits of commitment to an inflation target, rest firmly on foresight in price setting. Our work investigates the structural foundations of forward-looking pricing by studying *intertemporal pass-through* (iPT): the sensitivity of firms' desired price adjustments to their expectations of own marginal costs. We view our approach as complementary to existing methodologies for assessing the role of inflation expectations, such as the estimation of structural Phillips Curves with aggregate time series (Galí and Gertler, 1999), or the construction of model-free elasticities between firms' decisions and inflation expectations using micro data (Coibion et al., 2018, 2020). The principal advantage of our approach is the structural mapping between iPT estimates and their theoretical equivalents. This allows a direct assessment of whether conventional pricing models produce realistic forward-looking behavior. All in all, we find that they largely do.

We make three key contributions on our way to reaching the principal conclusion. First, a crucial ingredient for estimating firm-level iPT is a set of exogenous shifters to their expectations of own *future* marginal costs. We construct those using a natural experiment: the announcement in March 2019 of a future tariff schedule in the event of a "No-Deal" Brexit by the UK government. Combined with confidential UK survey data on managers' perceived probabilities of leaving without a deal and exposure to inputs imported from the EU, we construct a set of firm-level news shocks, generating the required exogenous variation in cost expectations. Second, we obtain firm-level iPTs, allowing for cross-sectional heterogeneity and non-linearity in our estimation. Crucially, we find that, all else equal, iPT is higher among firms with lower average frequency of price adjustment and for those that perceive the cost shock to be arriving earlier. In addition, iPT is lower among firms with state-dependent pricing and for larger shocks. Third, we show that our empirical iPT estimates are consistent with those generated by a conventional pricing model, once the latter is disciplined to match the observed firm characteristics and solved non-linearly for large shocks.

Our results create implications both for future research and for policy making. First, while we have been fortunate to rely on a natural experiment to estimate iPT, such exogenous variation is generally hard to come by. We believe a fruitful avenue for future research is to estimate iPT using data from surveys and experiments, featuring direct measurement and exogenous variation in firms' expected marginal costs. Second, our theoretical results imply that microeconomic heterogeneity in price setting matters for the contemporaneous aggregate effects of future policy announcements, such as those under forward guidance or fiscal plans. We believe that further investigation of such micro-to-macro connection can be a promising avenue both for researchers and for policy makers. Third, the non-linearity that we find in the pass-through of news shocks can have potentially far-reaching implications for the optimal *size* of promises about future policy interventions.

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Online Appendix

Intertemporal Pass-Through

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Boromeus Wanengkirtyo, *Bank of England*
Ivan Yotzov, *Bank of England*

A Figures

Figure A1: Sectoral tariff news shocks before input-output adjustment

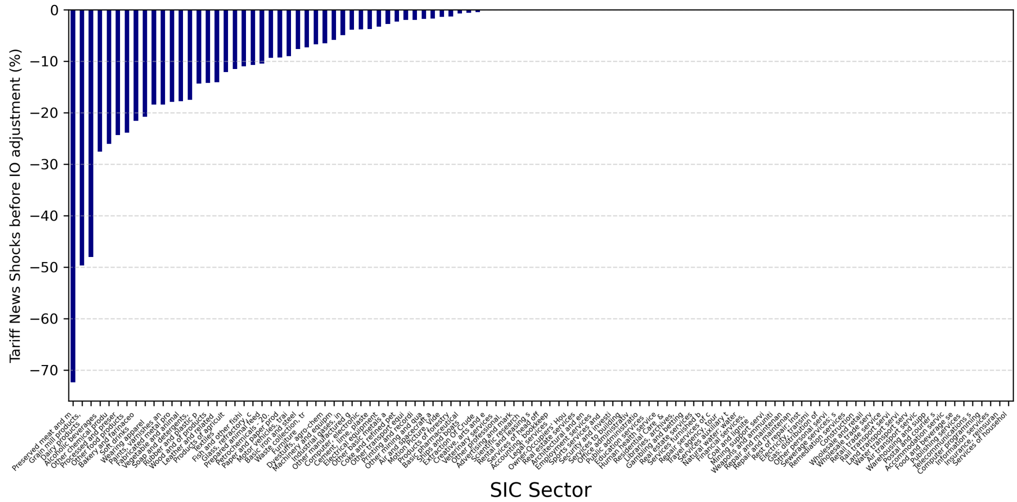
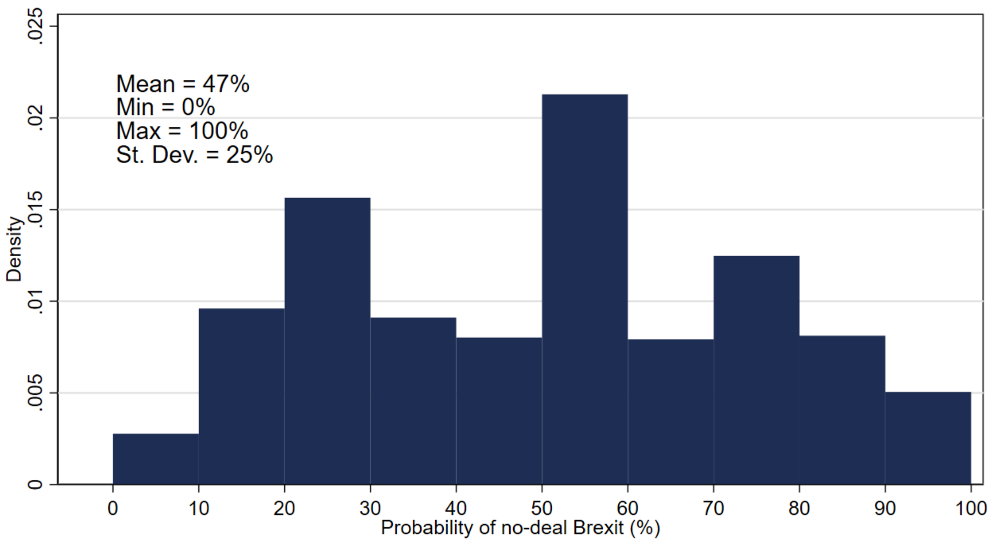
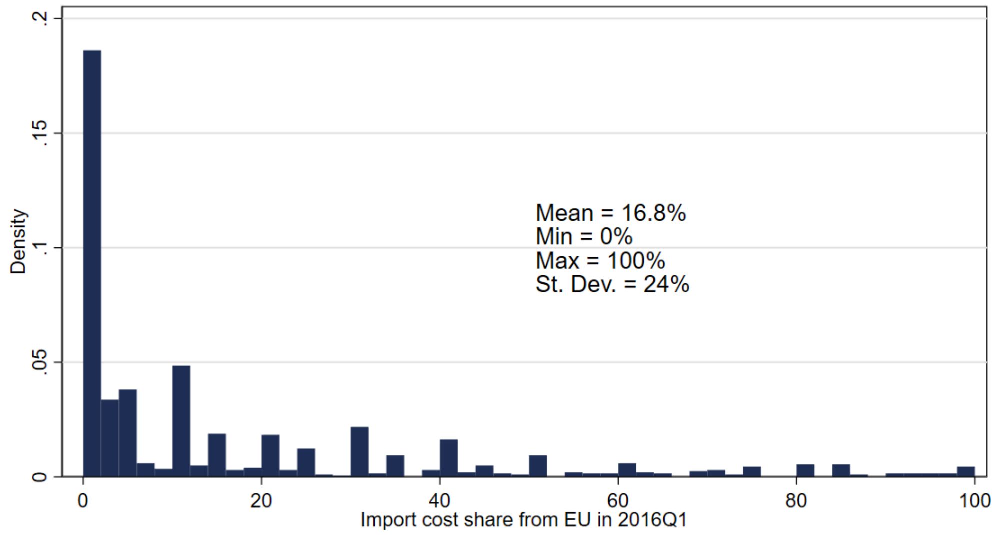


Figure A2: Probability of "No-Deal" Brexit in main estimation sample



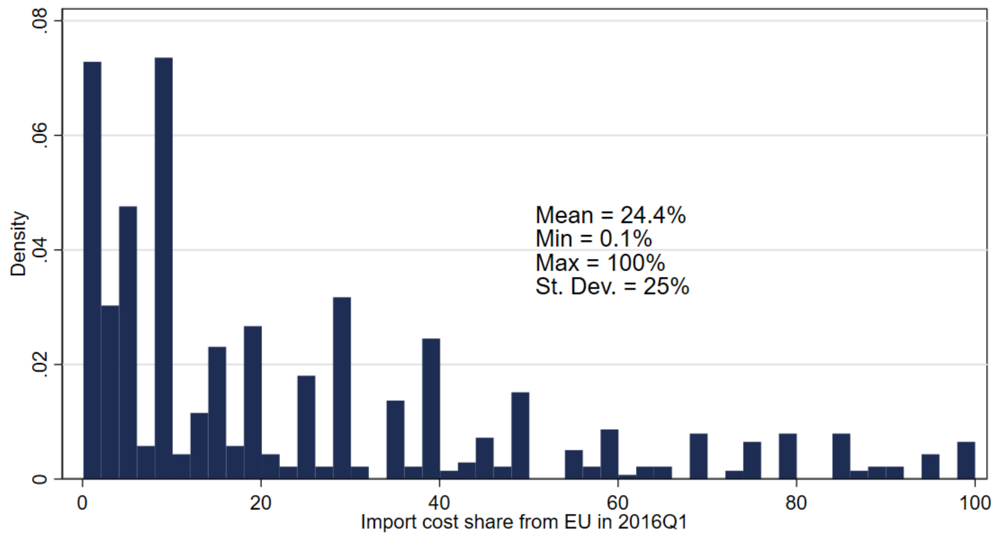
Notes: This figure shows the distribution of perceived probabilities of a "No-Deal" Brexit across firms in the Decision Maker Panel (DMP). The question was asked between February-April 2018 and November 2018-January 2019.

Figure A3: Histogram of import cost shares from EU in main regression sample



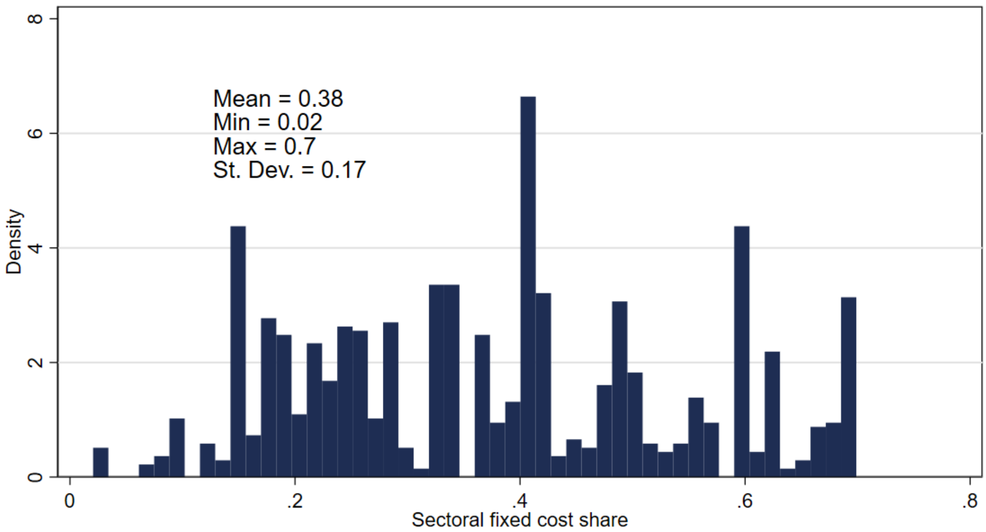
Notes: This figure presents the distribution of firm-level cost shares of intermediate inputs from the EU, as reported in the Decision Maker Panel (DMP). This question has been asked multiple times over the survey sample. For a given firm, the earliest available response is used.

Figure A4: Histogram of import cost shares from EU in main regression sample (excluding zeros)



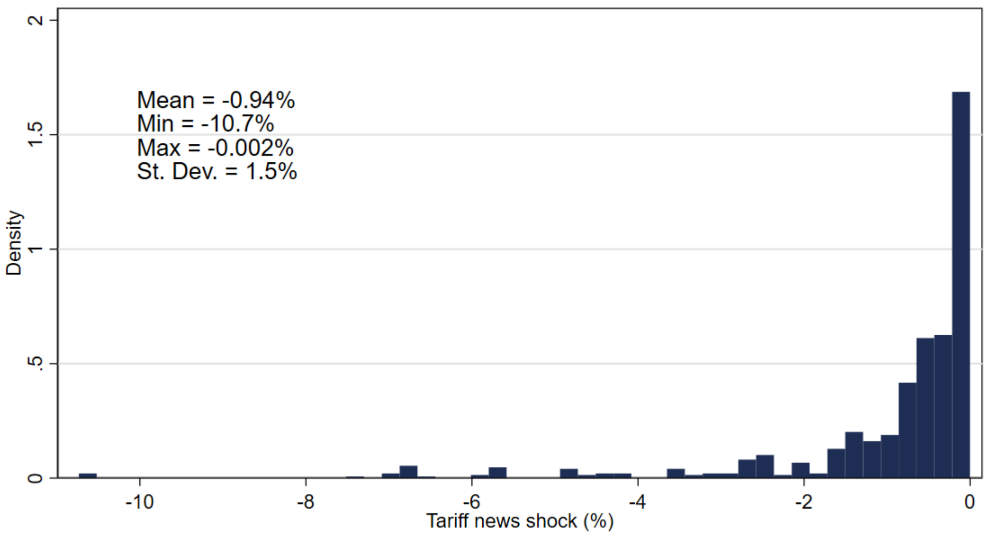
Notes: This figure presents the distribution of non-zero firm-level cost shares of intermediate inputs from the EU, as reported in the Decision Maker Panel (DMP). This question has been asked multiple times over the survey sample. For a given firm, the earliest available response is used.

Figure A5: Histogram of sectoral fixed cost shares in main regression sample



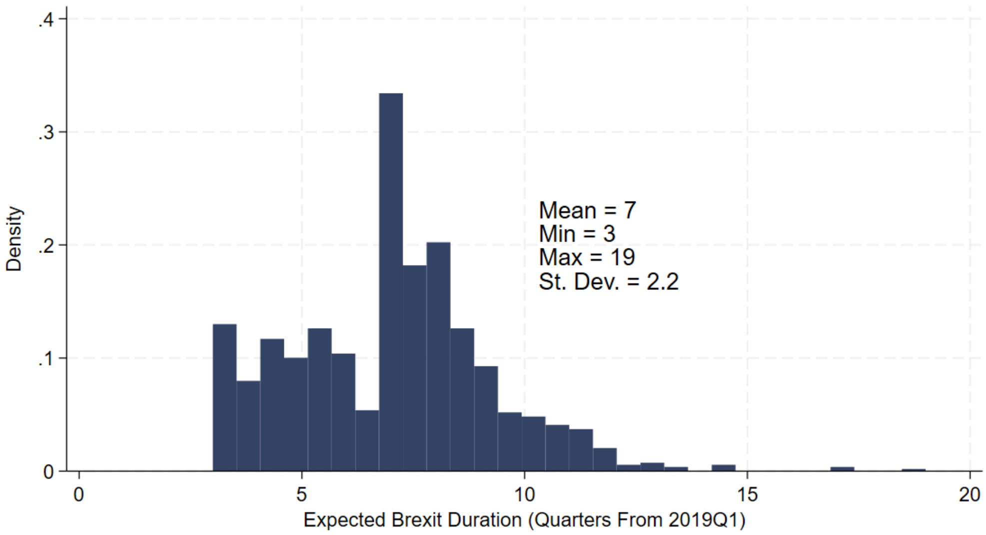
Notes: This figure presents the distribution of sectoral fixed cost shares in the main estimation sample.

Figure A6: Distribution of (non-zero) firm-level tariff news shock



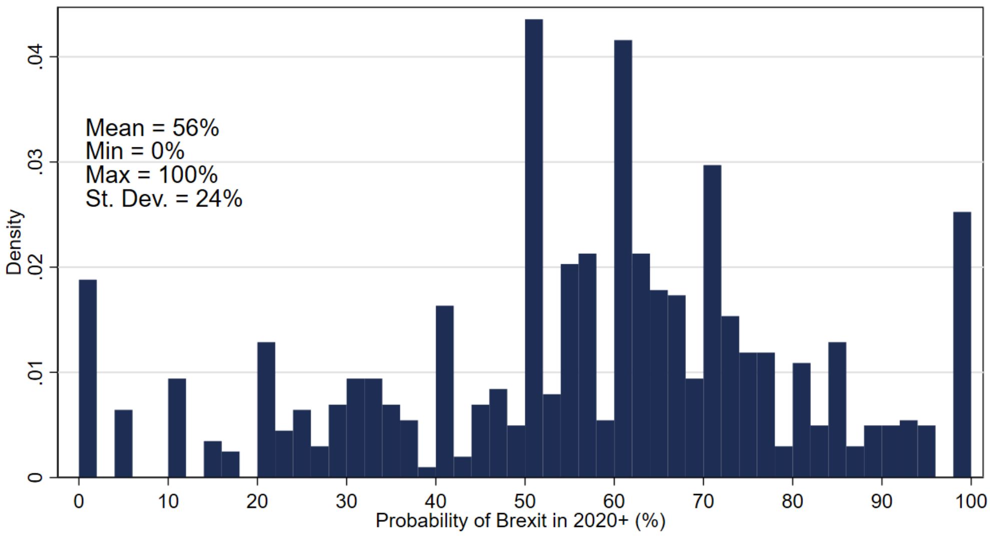
Notes: This figure presents the distribution of (non-zero) firm-level tariff news shocks in the main estimation sample.

Figure A7: Expected Brexit duration (quarters)



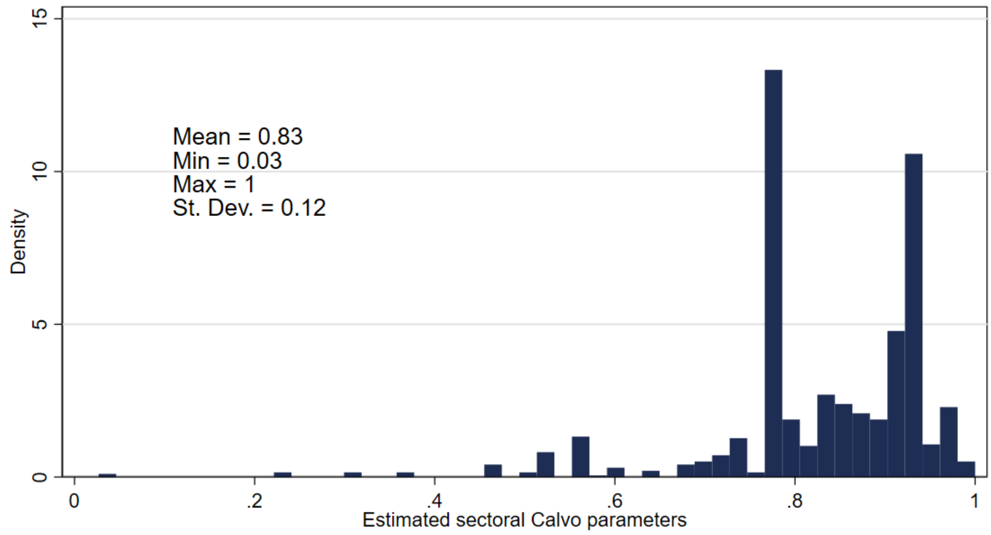
Notes: The figure shows the distribution of the expected duration, presented as quarters from 2019Q1, constructed using survey responses in the Decision Maker Panel (DMP).

Figure A8: Histogram of perceived probability of Brexit occurring in 2020 or later



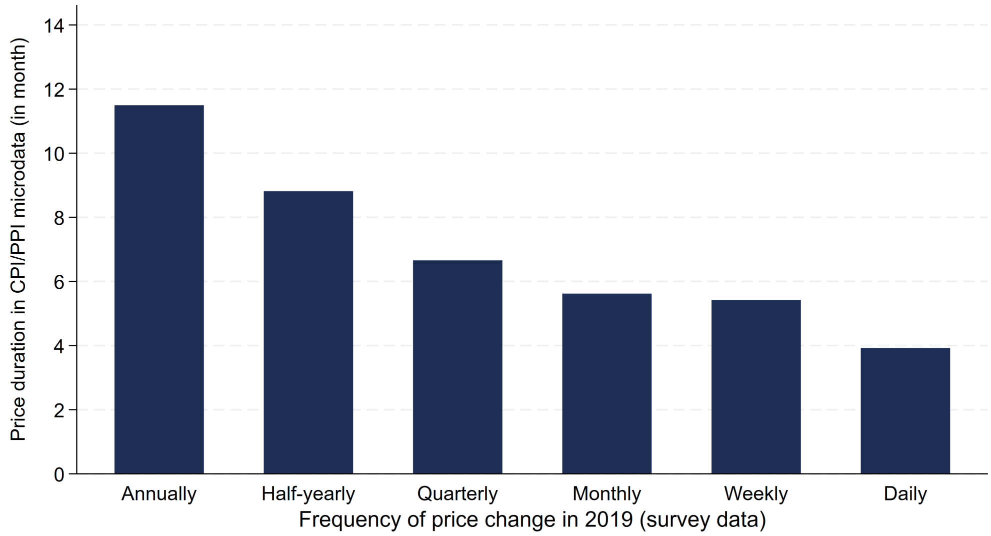
Notes: This figure presents the distribution of perceived probabilities for Brexit occurring in 2020 or later, as opposed to Brexit occurring in 2019.

Figure A9: Histogram of estimated Calvo parameters in main estimation sample



Notes: This figure presents the distribution of estimated frequencies of non-adjustment.

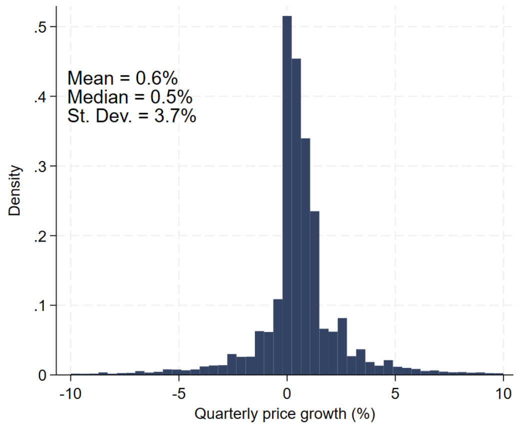
Figure A10: Comparison of CPI/PPI micro data price duration estimates with survey responses on frequency of price change in 2019



Notes: This figure compares DMP survey responses on the frequency of price change in 2019 (horizontal axis) with sectoral price duration estimates from PPI/CPI micro data (vertical axis). The question on frequency of price change in 2019 was asked retrospectively in multiple waves over 2023-2025, and responses are available from 3,098 unique firms.

Figure A11: Distribution of quarterly firm price growth

(a) Full Sample



(b) Full Sample excl. zeros

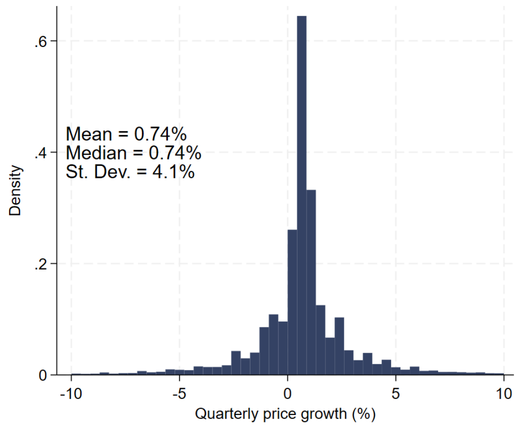
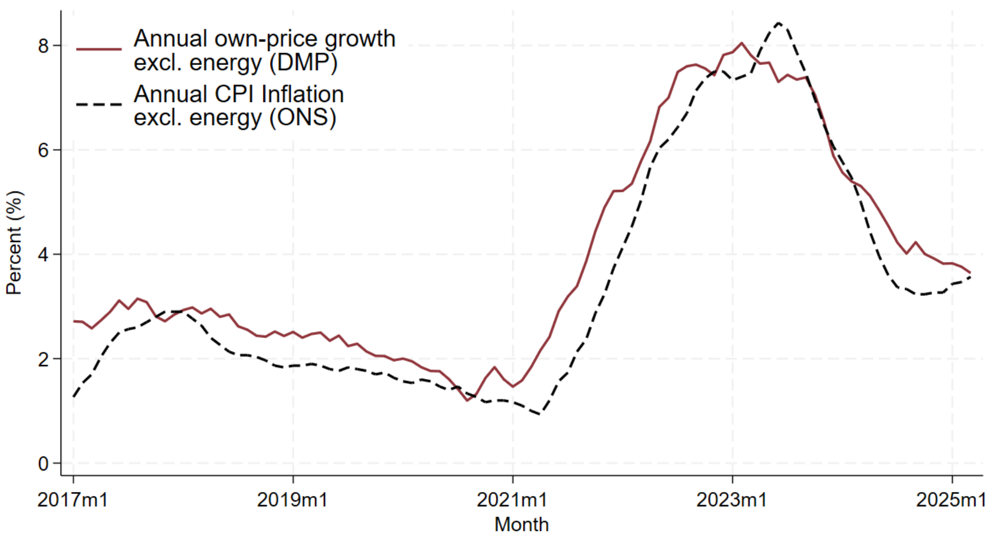


Figure A12: Annual UK CPI inflation and DMP annual own-price growth (excluding energy)



Notes: The data on annual own-price growth are based on data from the Decision Maker Panel, excluding energy firms. The data on annual CPI inflation excluding energy are taken from the Office for National Statistics. The series are three-month moving averages.

Figure A13: Share of state-dependent firms by SIC2 sector

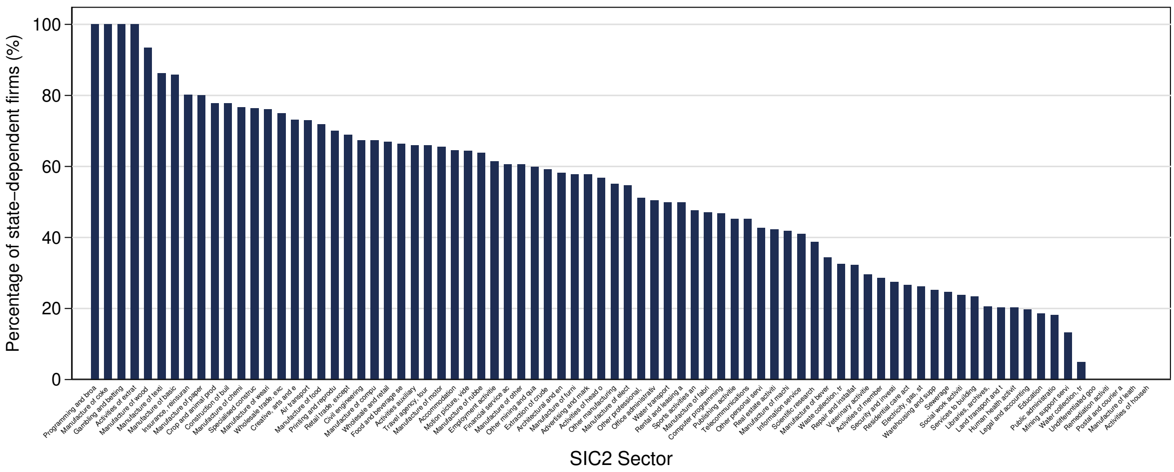
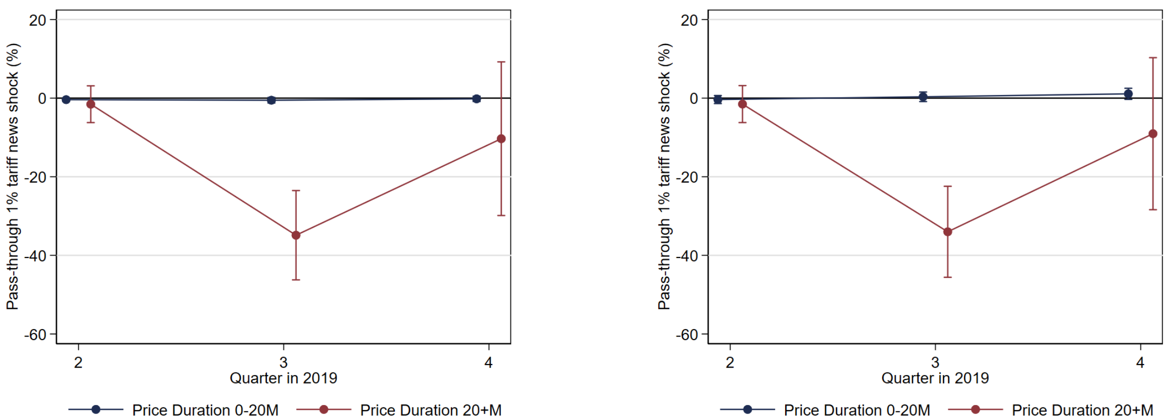


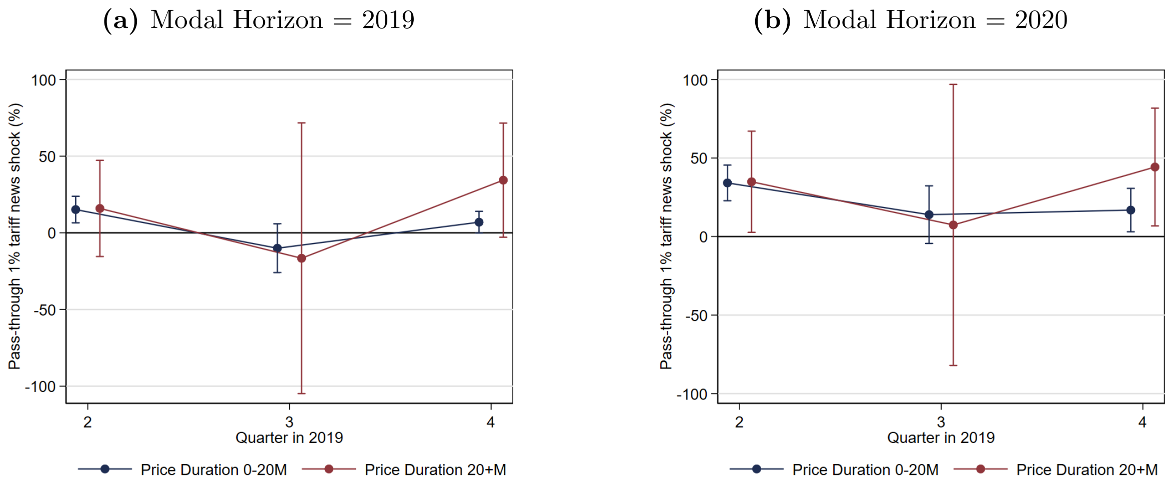
Figure A14: Impact of tariff news shock on employment: Effects by estimation quarter

(a) Modal Horizon = 2019 **(b)** Modal Horizon = 2020



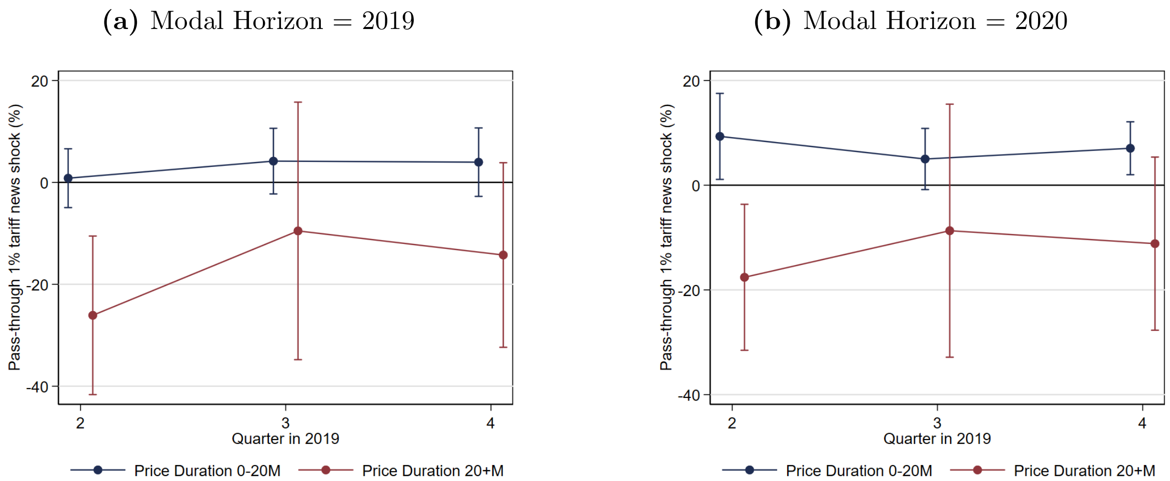
Notes: The results are based on Column 1 of Table A5. Standard errors are clustered at the industry (SIC2) level.

Figure A15: Impact of tariff news shock on capital expenditure: Effects by estimation quarter



Notes: The results are based on Column 2 of Table A5. Standard errors are clustered at the industry (SIC2) level.

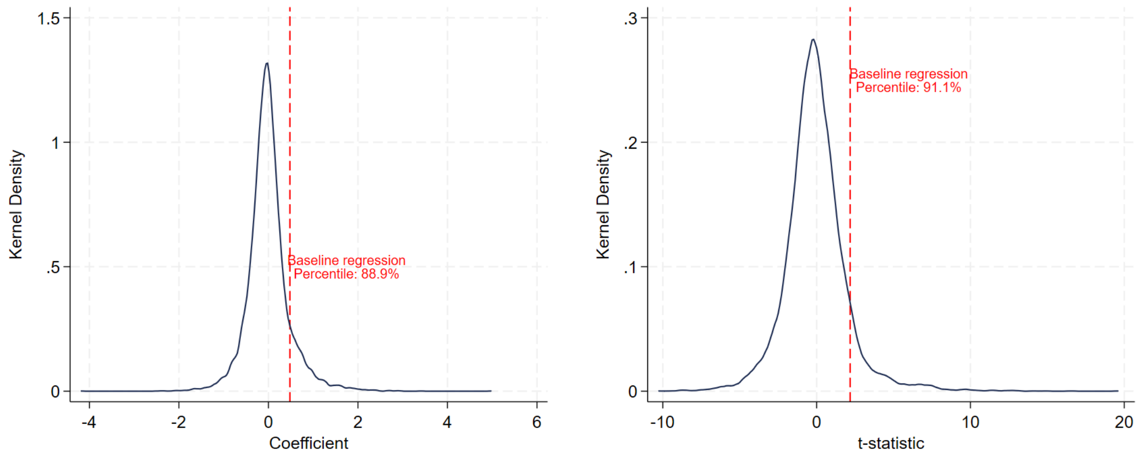
Figure A16: Impact of tariff news shock on real sales: Effects by estimation quarter



Notes: The results are based on Column 3 of Table A5. Standard errors are clustered at the industry (SIC2) level.

Figure A17: Placebo test: iPT and adjustment frequency under randomized news shocks

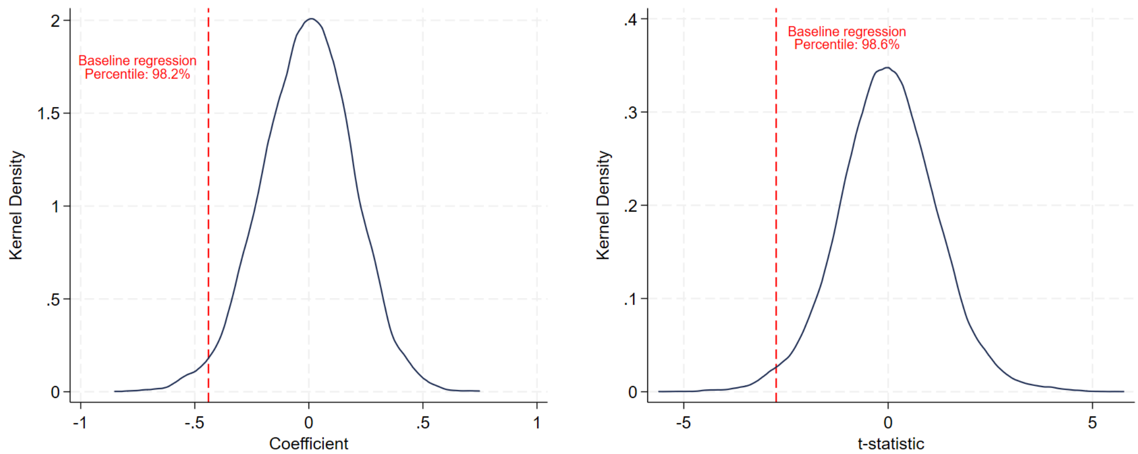
(a) Coefficient **(b)** t-statistic



Notes: For the placebo tests, we use the regression specification in Column 4 of Table 1 (sticky prices defined as longer than 20 months price duration). Standard errors are clustered at the industry (SIC2) level. We take 10,000 replications and plot the kernel density of the estimated coefficients across replications.

Figure A18: Placebo test: iPT and perceived horizon under randomized news shocks

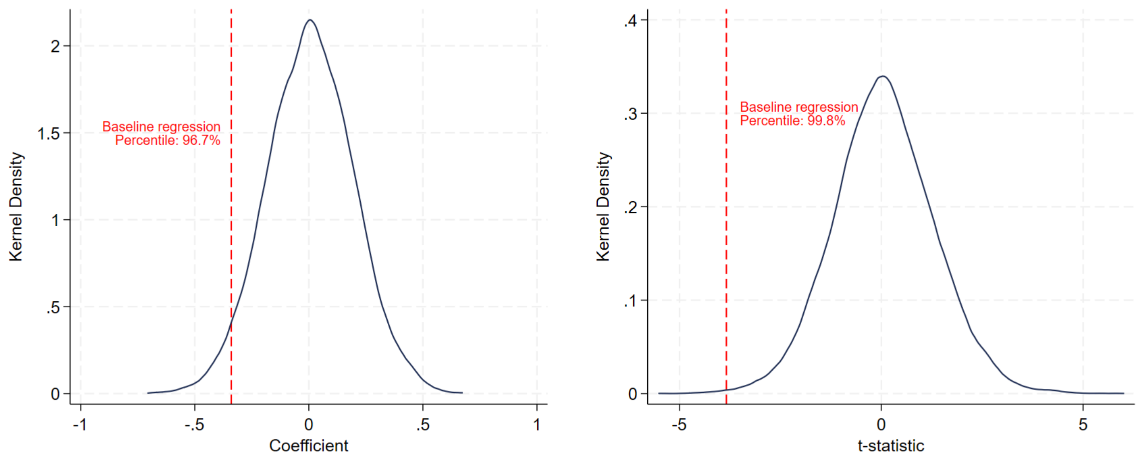
(a) Coefficient **(b)** t-statistic



Notes: For the placebo tests, we use the regression specification in Column 4 of Table 1 (Modal Brexit year = 2020). Standard errors are clustered at the industry (SIC2) level. We take 10,000 replications and plot the kernel density of the estimated coefficients across replications.

Figure A19: Placebo test: iPT and state/time-dependent pricing under randomized news shocks

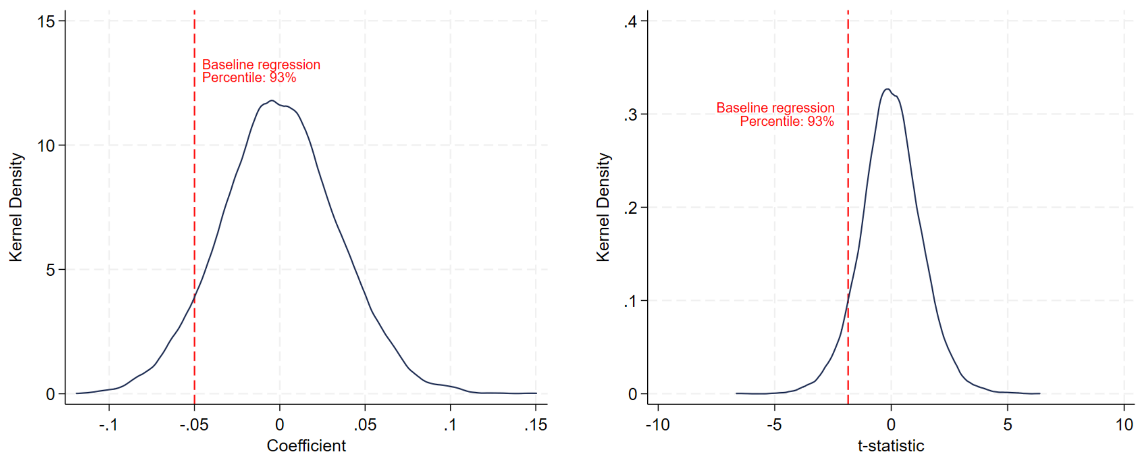
(a) Coefficient **(b)** t-statistic



Notes: For the placebo tests, we use the regression specification in Column 3 of Table 2. Standard errors are clustered at the industry (SIC2) level. We take 10,000 replications and plot the kernel density of the estimated coefficients across replications.

Figure A20: Placebo test: iPT non-linearity under randomized news shocks

(a) Coefficient **(b)** t-statistic



Notes: For the placebo tests, we use the regression specification in Column 4 of Table 3. Standard errors are clustered at the industry (SIC2) level. We take 10,000 replications and plot the kernel density of the estimated coefficients across replications.

B Tables

Table A1: Key (non-exhaustive) relevant dates and official date of UK's departure from the EU

Date	Event	Official Departure Date
23/06/2016	UK referendum on EU Membership	-
29/03/2017	UK PM formally triggers Article 50	29/03/2019
28/02/2018	Draft Withdrawal Agreement published (transition period ends 31/12/2020)	29/03/2019
25/11/2018	EU27 leaders re-confirm endorsement of Withdrawal Agreement	29/03/2019
13/03/2019	*Announcement of temporary tariff regime*	29/03/2019
14/03/2019	House of Commons vote to extend Article 50, granted by EU27 the next day	30/06/2019
05/04/2019	UK PM requests European Council to extend Article 50	30/06/2019
10/04/2019	EU27 agree to extend Article 50	31/10/2019
19/10/2019	Proposed Brexit deal lost on amendment in the House of Commons	31/10/2019
29/10/2019	European Council agree to UK PM request to extend	31/01/2020
31/01/2020	UK formally leaves the EU, enters a transition period	31/01/2020
31/12/2020	UK leaves EU Single Market and customs union	-

Notes: House of Commons Research Briefing and European Council

Table A2: Descriptive statistics for key variables in main estimation sample

	Observations	Mean	St. Dev.	Min.	Max.	25th Pctile	50th Pctile	75th Pctile
Tariff news cost shock	1010	-0.647	1.313	-10.735	0.000	-0.670	-0.126	0.000
=1 Price duration 20+ months	1010	0.059	0.237	0.000	1.000	0.000	0.000	0.000
=1 Modal Brexit Year = 2020	1010	0.568	0.496	0.000	1.000	0.000	1.000	1.000
Sectoral share of state-dependent firms	1010	0.595	0.195	0.186	1.000	0.470	0.669	0.749
Dummy variable =1 for exporter	1010	0.570	0.495	0.000	1.000	0.000	1.000	1.000
Probability of no-deal Brexit	1010	46.811	25.036	0.000	100.000	25.000	50.000	70.000
Percent of costs imported from EU in 2016Q1	1010	16.824	23.747	0.000	100.000	0.000	5.000	25.000
Sectoral fixed cost share	1010	0.376	0.168	0.021	0.699	0.237	0.369	0.490
Capital Expenditure DHS Growth Rate	700	-8.917	122.146	-200.000	200.000	-111.956	0.000	90.098
Real Sales Growth Rate	995	2.842	13.653	-36.000	40.000	-4.300	1.300	9.000
Employment DHS Growth Rate	744	0.280	10.957	-32.500	24.047	-3.774	0.000	5.173

Table A3: Impact of tariff news shock on prices: Interaction with price durations and expected Brexit duration

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	$100 \times \Delta \log$(Price Level)					
Sample:	2019Q2 - 2019Q4					
Tariff news shock$_{i,k}$	-0.107	0.051	0.060	0.054	-0.135	-0.020
	(0.080)	(0.081)	(0.126)	(0.125)	(0.294)	(0.108)
Tariff news shock$_{i,k}\times$ Modal Brexit Year=2020$_i$		-0.416***	-0.449***	-0.445***	-0.684**	-0.482***
		(0.124)	(0.160)	(0.162)	(0.259)	(0.139)
Tariff news shock$_{i,k}\times$ Price Duration 20+M$_k$		0.357***	0.498**	0.481**	1.221**	
		(0.108)	(0.218)	(0.221)	(0.585)	
Tariff news shock$_{i,k}\times$ Price Duration 5-10M$_k$						0.138
						(0.185)
Tariff news shock$_{i,k}\times$ Price Duration 10-20M$_k$						0.435**
						(0.171)
Tariff news shock$_{i,k}\times$ Price Duration 20+M$_k$						0.616***
						(0.224)
Probability No-Deal Brexit$_i$			-0.009*	-0.009*	-0.014*	-0.008*
			(0.005)	(0.005)	(0.007)	(0.005)
Import Cost Share EU$_i$			0.004	0.004	-0.010	0.005
			(0.006)	(0.006)	(0.010)	(0.005)
Industry Fixed Cost Share$_k$			0.478	0.472		0.384
			(0.518)	(0.514)		(0.520)
=1 Exporter$_i$			-0.164	-0.160	0.467	-0.129
			(0.262)	(0.264)	(0.292)	(0.331)
ln(Capex)$_{i,M19}$			0.053	0.055	0.012	0.060
			(0.045)	(0.044)	(0.055)	(0.045)
ln(Employment)$_{i,M19}$			0.023	0.020	0.158	0.002
			(0.112)	(0.110)	(0.140)	(0.112)
Industry (SIC4) fixed effects	No	No	No	No	Yes	No
Quarter fixed effects	No	No	No	Yes	Yes	Yes
Mean of Dependent Variable	1.078	1.078	1.078	1.078	1.081	1.078
R^2	0.002	0.012	0.022	0.032	0.325	0.036
Observations	1,010	1,010	1,010	1,010	1,003	1,010

Notes: The dependent variable is winsorised at the 5th and 95th percentiles. Standard errors are clustered at the industry (SIC2) level and reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A4: Impact of tariff news shock on employment, capital expenditure, and real sales: Interaction with price durations and expected shock horizon

	(1)	(2)	(3)
Dependent variable:	log(Employ)	log(CapEx)	log(Real Sales)
Sample:	2019Q2 - 2019Q4		
Tariff news shock$_{i,k}$	-0.004	0.033	0.031
	(0.003)	(0.048)	(0.036)
Tariff news shock$_{i,k}\times$ Modal Horizon 2020$_i$	0.009	0.182**	0.040
	(0.007)	(0.070)	(0.027)
Tariff news shock$_{i,k}\times$ Price Duration 20+M$_k$	-0.069	0.154	-0.195*
	(0.081)	(0.180)	(0.101)
ln(CapEx)$_{i,M19}$	0.007***	0.443***	0.060***
	(0.002)	(0.057)	(0.019)
ln(Employment)$_{i,M19}$	0.973***	0.630***	0.437***
	(0.010)	(0.105)	(0.064)
ln(Real Sales)$_{i,M19}$			0.398***
			(0.070)
Quarter fixed effects	Yes	Yes	Yes
Additional controls	Yes	Yes	Yes
Mean of dependent variable	4.810	4.681	3.823
R^2	0.978	0.503	0.692
Observations	946	865	938

Notes: The dependent variable is winsorised at the 5th and 95th percentiles. Additional controls include: Perceived probability of No Deal Brexit; Import cost share; Exporter status; Sectoral fixed cost share; Natural logarithms of employment, capital expenditure, and real sales in March 2019. Standard errors are clustered at the industry (SIC2) level and reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A5: Impact of tariff news shock on employment, capital expenditure, and real sales: Interaction with price durations, expected shock horizon, and quarter dummies

	(1)	(2)	(3)
Dependent variable:	log(Employ)	log(CapEx)	log(Real Sales)
Sample:	2019Q2 - 2019Q4		
Tariff news shock$_{i,k}$	-0.004	0.152***	0.008
	(0.002)	(0.051)	(0.034)
Tariff news shock$_{i,k}\times$ 2019Q3	-0.001	-0.252**	0.033
	(0.003)	(0.113)	(0.022)
Tariff news shock$_{i,k}\times$ 2019Q4	0.002	-0.082	0.031**
	(0.003)	(0.059)	(0.015)
Tariff news shock$_{i,k}\times$ Price Duration 20+M$_k\times$ 2019Q2	-0.012	0.007	-0.269***
	(0.027)	(0.190)	(0.094)
Tariff news shock$_{i,k}\times$ Price Duration 20+M$_k\times$ 2019Q3	-0.343***	-0.065	-0.137
	(0.069)	(0.511)	(0.140)
Tariff news shock$_{i,k}\times$ Price Duration 20+M$_k\times$ 2019Q4	-0.101	0.274	-0.182*
	(0.115)	(0.220)	(0.100)
Tariff news shock$_{i,k}\times$ Modal Horizon 2020$_i\times$ 2019Q2	0.000	0.190***	0.085**
	(0.006)	(0.055)	(0.034)
Tariff news shock$_{i,k}\times$ Modal Horizon 2020$_i\times$ 2019Q3	0.009	0.240*	0.008
	(0.007)	(0.133)	(0.039)
Tariff news shock$_{i,k}\times$ Modal Horizon 2020$_i\times$ 2019Q4	0.013	0.099	0.031
	(0.009)	(0.078)	(0.039)
ln(CapEx)$_{i,M19}$	0.007***	0.444***	0.060***
	(0.002)	(0.057)	(0.019)
ln(Employment)$_{i,M19}$	0.973***	0.625***	0.438***
	(0.010)	(0.105)	(0.064)
ln(Real Sales)$_{i,M19}$			0.398***
			(0.070)
Quarter fixed effects	Yes	Yes	Yes
Additional controls	Yes	Yes	Yes
Mean of dependent variable	4.810	4.681	3.823
R^2	0.978	0.506	0.692
Observations	946	865	938

Notes: The dependent variable is winsorised at the 5th and 95th percentiles. Additional controls include: Perceived probability of No Deal Brexit; Import cost share; Exporter status; Sectoral fixed cost share; Natural logarithms of employment, capital expenditure, and real sales in March 2019. Standard errors are clustered at the industry (SIC2) level and reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A6: Impact of tariff news shock on prices: Interaction with price durations, expected shock horizon, and quarter dummies

	(1)	(2)	(3)
Dependent variable:	$100 \times \Delta \log$(Price Level)		
Sample:	2019Q2 - 2019Q4		
Tariff news shock$_{i,k}$	0.123	0.095	0.095
	(0.146)	(0.146)	(0.146)
Tariff news shock$_{i,k} \times$ 2019Q3	-0.110	-0.090	-0.090
	(0.088)	(0.093)	(0.093)
Tariff news shock$_{i,k} \times$ 2019Q4	-0.085	-0.019	-0.020
	(0.110)	(0.131)	(0.132)
Tariff news shock$_{i,k} \times$ Price Duration 20+M$_k \times$ 2019Q2	0.741***		0.696***
	(0.239)		(0.244)
Tariff news shock$_{i,k} \times$ Price Duration 20+M$_k \times$ 2019Q3	0.823***		0.808***
	(0.253)		(0.264)
Tariff news shock$_{i,k} \times$ Price Duration 20+M$_k \times$ 2019Q4	-0.118		-0.074
	(0.219)		(0.234)
Tariff news shock$_{i,k} \times$ Modal Horizon 2020$_i \times$ 2019Q2		-0.357*	-0.365*
		(0.185)	(0.183)
Tariff news shock$_{i,k} \times$ Modal Horizon 2020$_i \times$ 2019Q3		-0.409**	-0.420**
		(0.177)	(0.179)
Tariff news shock$_{i,k} \times$ Modal Horizon 2020$_i \times$ 2019Q4		-0.545**	-0.530**
		(0.254)	(0.254)
Tariff news shock$_{i,k} \times$ Price Duration 20+M$_k$		0.473**	
		(0.221)	
Tariff news shock$_{i,k} \times$ Modal Horizon 2020$_i$	-0.444***		
	(0.163)		
Quarter fixed effects	Yes	Yes	Yes
Additional controls	Yes	Yes	Yes
Mean of dependent variable	1.078	1.078	1.078
R^2	0.033	0.033	0.033
Observations	1,010	1,010	1,010

Notes: The dependent variable is winsorised at the 5th and 95th percentiles. Additional controls include: Perceived probability of No Deal Brexit; Import cost share; Exporter status; Sectoral fixed cost share; Natural logarithms of employment and capital expenditure in March 2019. Standard errors are clustered at the industry (SIC2) level and reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A7: Impact of tariff news shock on prices: Interaction with price durations and expected shock horizon (placebo test)

	(1)	(2)
Dependent variable:	$100 \times \Delta \log$(Price Level)	
Sample:	2019Q2 - 2019Q4	2018Q1 - 2018Q4
Tariff news shock$_{i,k}$	0.054	-0.080
	(0.125)	(0.089)
Tariff news shock$_{i,k} \times$ Modal Brexit Year=2020$_i$	-0.445***	-0.342
	(0.162)	(0.212)
Tariff news shock$_{i,k} \times$ Price Duration 20+M$_k$	0.481**	0.403
	(0.221)	(0.342)
Quarter fixed effects	Yes	Yes
Additional controls	Yes	Yes
Mean of Dependent Variable	1.078	-1.198
R^2	0.032	0.076
Observations	1,010	1,212

Notes: The dependent variable is winsorised at the 5th and 95th percentiles. Additional controls include: Perceived probability of No Deal Brexit; Import cost share; Exporter status; Sectoral fixed cost share; Natural logarithms of employment and capital expenditure in March 2019. Standard errors are clustered at the industry (SIC2) level and reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A8: Impact of tariff news shock on prices: Interaction with price durations and expected shock horizon (continuous vs. categorical measures)

	(1)	(2)	(3)	(4)
Dependent variable:	$100 \times \Delta \log$(Price Level)			
Sample:	2019Q2 - 2019Q4			
Tariff news shock$_{i,k}$	-0.020	0.282	-0.069	0.244
	(0.108)	(0.182)	(0.108)	(0.192)
Tariff news shock$_{i,k} \times$ Modal Brexit Year=2020$_i$	-0.482***		-0.473***	
	(0.139)		(0.158)	
Tariff news shock$_{i,k} \times$ Prob(Brexit in 2020+)$_i$		-0.008***		-0.008**
		(0.003)		(0.003)
Tariff news shock$_{i,k} \times$ Price Duration 5-10M$_k$	0.138	0.155		
	(0.185)	(0.203)		
Tariff news shock$_{i,k} \times$ Price Duration 10-20M$_k$	0.435**	0.394**		
	(0.171)	(0.166)		
Tariff news shock$_{i,k} \times$ Price Duration 20+M$_k$	0.616***	0.411**		
	(0.224)	(0.201)		
Tariff news shock$_{i,k} \times$ Price Duration$_k$			0.029**	0.019*
			(0.012)	(0.011)
Quarter fixed effects	Yes	Yes	Yes	Yes
Additional controls	Yes	Yes	Yes	Yes
Mean of Dependent Variable	1.078	1.078	1.078	1.078
R^2	0.036	0.031	0.034	0.028
Observations	1,010	1,010	1,010	1,010

Notes: The dependent variable is winsorised at the 5th and 95th percentiles. Additional controls include: Perceived probability of No Deal Brexit; Import cost share; Exporter status; Sectoral fixed cost share; Natural logarithms of employment and capital expenditure in March 2019. Standard errors are clustered at the industry (SIC2) level and reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A9: Determinants of perceived no Deal Brexit probability

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Dependent variable:	Perceived No Deal Brexit Probability									
Sample:	2019Q2 - 2019Q4									
Import Cost Share EU_i	0.046									0.055
	(0.037)									(0.038)
=1 $Exporter_i$		0.109								-1.146
		(1.581)								(1.723)
=1 Price Duration 5-10M_k			1.897							1.682
			(2.214)							(2.307)
=1 Price Duration 10-20M_k			1.388							1.706
			(1.828)							(2.107)
=1 Price Duration 20+M_k			-3.141							-4.995
			(3.640)							(3.596)
=1 State-dependent$_k^{p33}$				-2.761						-4.201**
				(1.688)						(1.775)
=1 Expected Brexit Date 2020_i					-2.298					-2.397
					(1.596)					(1.592)
Industry Fixed Cost $Share_k$						-1.791				-5.697
						(4.815)				(5.206)
Price $Growth_{i,M19}$							0.655***			0.657***
							(0.200)			(0.209)
$\ln(\text{Employment})_{i,M19}$								-3.053***		-2.392**
								(0.875)		(1.052)
$\ln(\text{Capex})_{i,M19}$									-1.089***	-0.654
									(0.326)	(0.401)
Mean of Dependent Variable	46.811	46.811	46.811	46.811	46.811	46.811	46.811	46.811	46.811	46.811
R^2	0.002	0.000	0.002	0.003	0.002	0.000	0.012	0.013	0.011	0.041
Observations	1,010	1,010	1,010	1,010	1,010	1,010	1,010	1,010	1,010	1,010

Notes: The dependent variable is the perceived "No-Deal" Brexit probability in the main estimation sample. Robust standard errors are reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A10: Determinants of expected Brexit horizon

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Dependent variable:	=1 Expected Brexit Date 2020									
Sample:	2019Q2 - 2019Q4									
Import Cost Share EU_i	0.000									0.001
	(0.001)									(0.001)
=1 Exporter_i		-0.098***								-0.088**
		(0.031)								(0.034)
=1 Price Duration 5-10M_k			-0.062							-0.035
			(0.044)							(0.046)
=1 Price Duration 10-20M_k			0.075**							0.077*
			(0.036)							(0.041)
=1 Price Duration 20+M_k			-0.059							-0.069
			(0.069)							(0.073)
=1 State-dependent$_k^{p33}$				0.006						0.012
				(0.033)						(0.037)
No Deal Brexit Probability					-0.001					-0.001
					(0.001)					(0.001)
Industry Fixed Cost Share_k						0.021				-0.018
						(0.092)				(0.099)
Price $\text{Growth}_{i,M19}$							-0.000			-0.001
							(0.004)			(0.004)
$\ln(\text{Employment})_{i,M19}$								0.021		0.024
								(0.016)		(0.019)
$\ln(\text{Capex})_{i,M19}$									-0.004	-0.012
									(0.007)	(0.008)
Mean of Dependent Variable	0.568	0.568	0.568	0.568	0.568	0.568	0.568	0.568	0.568	0.568
R^2	0.000	0.010	0.011	0.000	0.002	0.000	0.000	0.002	0.000	0.024
Observations	1,010	1,010	1,010	1,010	1,010	1,010	1,010	1,010	1,010	1,010

Notes: The dependent variable is a dummy variable for the expected modal Brexit date being 2020. in the main estimation sample. Robust standard errors are reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A11: Impact of tariff news shock on prices: Interaction with price durations and expected Brexit duration and firm size

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	$100 \times \Delta \log$(Price Level)					
Sample:	2019Q2 - 2019Q4					
Tariff news shock$_{i,k}$	0.071	0.081	0.159	0.359	-0.209	0.176
	(0.122)	(0.127)	(0.237)	(0.513)	(0.488)	(0.314)
Tariff news shock$_{i,k}\times$ Modal Brexit Year=2020$_i$	-0.421**	-0.404**	-0.407**	-0.422**	-0.389**	-0.417**
	(0.164)	(0.161)	(0.150)	(0.172)	(0.147)	(0.156)
Tariff news shock$_{i,k}\times$ Price Duration 20+M$_k$	0.444***	0.415***	0.413**	0.429***	0.308**	0.407**
	(0.119)	(0.119)	(0.172)	(0.121)	(0.122)	(0.156)
Tariff news shock$_{i,k}\times$ Employment$_{i,M19} \geq 250$	-0.084					
	(0.161)					
Tariff news shock$_{i,k}\times$ Employment$_{i,M19} \geq 100$		-0.080				
		(0.176)				
Tariff news shock$_{i,k}\times$ Employment$_{i,M19} \geq 50$			-0.162			
			(0.262)			
Tariff news shock$_{i,k}\times$ log(Employment$_{i,M19}$)				-0.064		
				(0.098)		
Tariff news shock$_{i,k}\times$ log(Sales$_{i,M19}$)					0.026	
					(0.055)	
Tariff news shock$_{i,k}\times$ log(Capital Expenditure$_{i,M19}$)						-0.030
						(0.078)
Quarter fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Additional firm controls	Yes	Yes	Yes	Yes	Yes	Yes
Mean of Dependent Variable	1.078	1.078	1.078	1.078	1.078	1.078
R^2	0.032	0.032	0.032	0.032	0.040	0.032
Observations	1,010	1,010	1,010	1,010	1,010	1,010

Notes: The dependent variable is winsorised at the 5th and 95th percentiles. Measures of employment, sales, and capital expenditure are based on the most recent observations at the firm level prior to the March 2019 announcement. Standard errors are clustered at the industry (SIC2) level and reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A12: Impact of tariff news shock on prices: Interaction with price durations and expected Brexit duration and HHI

	(1)	(2)	(3)
Dependent variable:	$100 \times \Delta \log$(Price Level)		
Sample:	2019Q2 - 2019Q4		
Tariff news shock$_{i,k}$	0.021	0.040	0.178
	(0.106)	(0.108)	(0.336)
Tariff news shock$_{i,k}\times$ Modal Brexit Year=2020$_i$	-0.417***	-0.411***	-0.398***
	(0.141)	(0.144)	(0.143)
Tariff news shock$_{i,k}\times$ Price Duration 20+M$_k$	0.440***	0.439***	0.424***
	(0.142)	(0.136)	(0.134)
Tariff news shock$_{i,k}\times$ HHI$_k$ > 50pctile	0.110		
	(0.120)		
Tariff news shock$_{i,k}\times$ HHI$_k$ > 75pctile		-0.709	
		(0.610)	
Tariff news shock$_{i,k}\times$ log(HHI$_k$)			-0.032
			(0.075)
Quarter fixed effects	Yes	Yes	Yes
Additional firm controls	Yes	Yes	Yes
Mean of Dependent Variable	1.078	1.078	1.078
R^2	0.032	0.033	0.032
Observations	1,010	1,010	1,010

Notes: The dependent variable is winsorised at the 5th and 95th percentiles. HHI is a measure of industry sales concentration (at the SIC2 level) for 2018 taken from Savagar et al. (2024). Standard errors are clustered at the industry (SIC2) level and reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table A13: Impact of tariff news shock on prices: Interaction with price durations and expected Brexit duration (0% import cost share from non-EU countries)

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	$100 \times \Delta \log$(Price Level)					
Sample:	2019Q2 - 2019Q4					
Tariff news shock$_{i,k}$	-0.137*	0.019	0.086	0.081	0.098	
	(0.077)	(0.089)	(0.194)	(0.189)	(0.174)	
Tariff news shock$_{i,k} \times$ Price Duration 5-10M$_k$					-0.080	
					(0.394)	
Tariff news shock$_{i,k} \times$ Price Duration 10-20M$_k$					0.734***	
					(0.218)	
Tariff news shock$_{i,k} \times$ Price Duration 20+M$_k$		0.512***	0.483***	0.491***	0.771***	
		(0.132)	(0.171)	(0.167)	(0.235)	
Tariff news shock$_{i,k} \times$ Modal Brexit Year=2020$_i$		-0.391***	-0.470*	-0.467*	-0.750***	
		(0.124)	(0.241)	(0.238)	(0.172)	
Quarter fixed effects	No	No	No	Yes	Yes	
Additional firm controls	No	No	Yes	Yes	Yes	
Mean of Dependent Variable	1.274	1.274	1.274	1.274	1.274	
R^2	0.004	0.014	0.025	0.037	0.051	
Observations	447	447	447	447	447	

Notes: The table presents the results from Table 1 for the sub-sample of firms which report 0% import cost share from non-EU countries in 2016Q1. The dependent variable is winsorised at the 5th and 95th percentiles. Standard errors are clustered at the industry (SIC2) level and reported in parentheses, stars indicate *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

C Proofs

Proof of Proposition 1. Given the timing of the unit shifter, the response of the optimal reset price (in log deviations from steady state), and hence the iPT, is given by:

$$\begin{aligned}\mathrm{iPT}_{d_i|d_i+D} = p^*_{i,t}(j) &= (1-\beta\alpha_i)\left[(\beta\alpha_i)^{d_i} + ... + (\beta\alpha_i)^{d_i+D}\right] \\ &= (1-\beta\alpha_i)(\beta\alpha_i)^{d_i}\left[\frac{1-(\beta\alpha_i)^{D+1}}{1-\beta\alpha_i}\right] = (\beta\alpha_i)^{d_i} - (\beta\alpha_i)^{d_i+D+1}. \end{aligned} \tag{23}$$

Further, it follows that for changes in frequency:

$$\frac{\partial \mathrm{iPT}_{d_i|d_i+D}}{\partial \alpha_i} = d_i\beta(\beta\alpha_i)^{d_i-1} - (d_i+D+1)\beta(\beta\alpha_i)^{d_i+D}, \tag{24}$$

so that $\frac{\partial \mathrm{iPT}_{d_i|d_i+D}}{\partial \alpha_i} > 0$ iff $\alpha_i < \frac{1}{\beta}\left[\frac{d_i}{(d_i+D+1)}\right]^{\frac{1}{D+1}}$. Differentiating once again delivers:

$$\frac{\partial^2 \mathrm{iPT}_{d_i|d_i+D}}{\partial \alpha_i^2} = d_i(d_i-1)\beta^2(\beta\alpha_i)^{d_i-2} - (d_i+D+1)(d_i+D)\beta^2(\beta\alpha_i)^{d_i+D-1}, \tag{25}$$

so that $\frac{\partial^2 \mathrm{iPT}_{d_i|d_i+D}}{\partial \alpha_i^2} > 0$ iff $\alpha_i < \frac{1}{\beta}\left[\frac{d_i(d_i+1)}{(d_i+D+1)(d_i+D)}\right]^{\frac{1}{D+1}}$.

As for changes in the horizon of the shifter, one similarly obtains:

$$\frac{\partial \mathrm{iPT}_{d_i|d_i+D}}{\partial d_i} = \underbrace{\log(\beta\alpha_i)}_{<0} \times \underbrace{\left[(\beta\alpha_i)^{d_i} - (\beta\alpha_i)^{d_i+D+1}\right]}_{>0} < 0, \tag{26}$$

and differentiating once again:

$$\frac{\partial^2 \mathrm{iPT}_{d_i|d_i+D}}{\partial d_i^2} = \underbrace{[\log(\beta\alpha_i)]^2}_{>0} \times \underbrace{\left[(\beta\alpha_i)^{d_i} - (\beta\alpha_i)^{d_i+D+1}\right]}_{>0} > 0. \tag{27}$$

□

Proof of Proposition 2. We can express CPI inflation from (21) as $\pi_t^C = \sum_{i=1}^N \varpi_i^C f(d_i)$, where $f(d_i) \equiv (1-\alpha)\left[(\beta\alpha)^{d_i} - (\beta\alpha)^{d_i+D+1}\right]$. From Proposition 1 it follows that $f(d_i)$ is convex in d_i, hence by Jensen's inequality:

$$\pi_t^C = \sum_{i=1}^N \varpi_i^C f(d_i) > f\left(\sum_{i=1}^N \varpi_i^C d_i\right) = \pi_t^C(\bar{d}) \tag{28}$$

where $\pi_t^C(\bar{d})$ is the response of CPI inflation in a counterfactual economy where all firms have a homogeneous perceived horizon $\bar{d} \equiv \sum_{i=1}^N \varpi_i^C d_i$. Therefore, it follows that $\pi_t^C > \pi_t^C(\bar{d})$ and heterogeneity in perceived horizons amplifies the response of aggregate inflation to a common shifter in expected future marginal costs . □

Proof of Proposition 3. When the common marginal cost shifter is permanent, one can write the aggregate CPI inflation as $\pi_t^C = \sum_{i=1}^N \overline{\omega}_i^C g(\alpha_i)$, where $g(\alpha_i) \equiv (1-\alpha_i)(\beta\alpha_i)^d$. Note that

$$g'(\alpha_i) = -(\beta\alpha_i)^d + (1-\alpha_i)d(\beta\alpha_i)^{d-1}\beta, \tag{29}$$

and differentiating once again

$$g''(\alpha_i) = -d(\beta\alpha_i)^{d-1}\beta - d\beta(\beta\alpha_i)^{d-1} + (1-\alpha_i)d(d-1)(\beta\alpha_i)^{d-2}\beta^2. \tag{30}$$

From the latter expression for the second derivative it follows that $g(\alpha_i)$ is convex iff:

$$\alpha_i < \frac{d-1}{d+1}. \tag{31}$$

Therefore, as long as $\alpha_i \in \left(0, \frac{d-1}{d+1}\right)$, $\forall i$, Jensen's inequality implies that:

$$\pi_t^C = \sum_{i=1}^N \overline{\omega}_i^C g(\alpha_i) > g\left(\sum_{i=1}^N \overline{\omega}_i^C \alpha_i\right) = \pi_t^C(\overline{\alpha}), \tag{32}$$

where $\pi_t^C(\overline{\alpha})$ is the response of CPI inflation in a counterfactual economy where all firms have a homogeneous pobability of non-adjustment $\overline{\alpha} \equiv \sum_{i=1}^N \overline{\omega}_i^C \alpha_i$. Therefore, it follows that $\pi_t^C > \pi_t^C(\overline{\alpha})$ and heterogeneity in adjustment frequencies amplifies the response of aggregate inflation to a common shifter in expected future marginal costs . □