AN IMPORT(ANT) PRICE OF BREXIT UNCERTAINTY

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ABSTRACT

We estimate the impact of trade policy uncertainty (TPU) on CES import price indices, focusing on the implications of Britain's exit from the European Union (Brexit). Our analysis reveals that an increase in the probability of Brexit increases U.K. import price indices by raising the prices of existing products and by reducing product variety from the E.U. We find evidence that the risk of higher import protection from the 2016 referendum increased current import price indices by more than 10%. This amounted to a 2 log point increase in manufactured goods prices and a 0.6 log point decrease in consumers' real income.

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Introduction

Policymakers have increasingly threatened to increase trade barriers and renegotiate trade agreements, generating uncertainty over future business conditions. There is now a large body of evidence showing that certain agreements reduce trade policy uncertainty (TPU). While most studies focus on the implementation of more predictable and permanent applied trade policies on the value of trade and investments by firms (Handley and Limão, 2022), limited evidence exists on the price effects of TPU. Moreover, no research has examined how uncertainty stemming from threats to leave or renegotiate existing commitments influences trade prices.

We examine the effects of TPU on import prices and welfare in the context of the U.K.'s Brexit referendum. We start by constructing exact constant elasticity of substitution (CES) price indices (year-over-year) for the U.K. by exporter and industry. We then develop a model linking Brexit uncertainty to these price indices through export investments. We estimate the effects on the exporting country's price indices, using model-consistent measures of tariff risk for each industry. Our estimated price effects use time-varying changes in the probability of Brexit in the period before the referendum and the cross-industry variation in the most favored nation (MFN) threat tariffs that would likely apply in a no-deal Brexit.

We estimate that Brexit uncertainty increased the prices of products subject to potentially higher tariffs. A one standard deviation increase in Brexit uncertainty implies a 1.7 log point (lp) increase in annual prices. The price uncertainty elasticities imply that the full referendum shock increased U.K. import prices from E.U. trade partners by 11 lp. When we adjust for general equilibrium effects and the share of imports in firm and household expenditure, this amounts to a 2 lp increase in the aggregate price of manufactured goods and a real income loss of at least 0.6 lp. We provide evidence for the mechanisms, e.g. the results apply only in industries with sunk costs of exports, and test its robustness to both alternative shocks (e.g., exchange rate) and other interpretations.[1]

Trade disagreements are not new and occasionally spill over into hotly contested referendums, unilateral executive action, or legislative votes on decisions to enter, exit, or amend existing international agreements. Beyond Brexit, understanding the policy uncertainty implications of trade disagreements will remain important. For example, the U.S. threatened to exit and later renegotiated major trade agreements with Mexico, Canada, and Korea and countrywide referendums on trade agreements are common.[2] Even the status of the U.K.'s exit under the renegotiated Trade and Cooperation Agreement with the E.U. continued to generate uncertainty for some time due to disagreements over the implementation of customs procedures in Northern Ireland.

Our first contribution is to provide evidence for the price effects of policy uncertainty arising from trade disagreements that increase uncertainty. The Brexit referendum is a recent example where TPU increased because prior trade integration under a previously credible agreement was at greater risk of reversal. Where trade agreements include credible and permanent commitments that reduce TPU, they have also increased welfare through lower prices and gains from new varieties (Handley and Limão, 2022). The research on the price and welfare effects of TPU has focused on reductions in uncertainty from China's WTO accession.

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[1]We find a reduction in both the quantity and value of imports when Brexit uncertainty increases, so the price increases in this period cannot be explained by demand increases such as stockpiling in anticipation of the vote.

[2]These include the Costa Rican referendum to join DR-CAFTA in 2007 (Mendéz and Van Patten, 2022), Dutch and French voters rejecting the E.U. Constitution in 2005, Irish voters initially rejecting the follow-up Treaty of Lisbon in 2008, Dutch voters rejecting a trade deal with Ukraine in 2016, and Swiss voters narrowly approving a trade agreement with Indonesia in 2021.

Handley and Limão (2017) (hereafter HL) find that the reduction in TPU lowered the ideal price index on U.S. imports from China, as do Amiti et al. (2020). Feng et al. (2017) find that Chinese firm export entry to the U.S. is higher in industries where the tariff risk was previously high and these entrants charge lower prices relative to firms that exit.

Second, we provide novel theory and evidence on the mechanisms through which changes in TPU affect prices and welfare. An important potential source of welfare gains from trade agreements are lower prices and increased variety, which may follow from applied tariff reductions or changes in TPU. We show that our findings on prices are consistent with a model of sunk export investments. Higher uncertainty increases prices through the net exit of varieties and foregone investments in upgrading, as in HL. We then extend the literature on TPU to incorporate quality competition. We find evidence consistent with increased TPU leading to selection toward higher productivity firms that compete on quality and charge higher prices. In contrast, most extant research on the welfare effects of trade policy focuses primarily on changes to applied trade barriers; recent examples examining the U.S. trade war include Fajgelbaum et al. (2019) and Amiti et al. (2019).

Third, we expand the set of relevant shocks affecting trade prices to include uncertainty over future trade policy. Realized shocks, such as exchange rate devaluations, often pass through to prices (Goldberg and Campa, 2010; Corsetti, Crowley, and Han, 2022). Research on firm responses to both exchange rate and tariff shocks finds that applied tariff changes have larger trade and extensive margin effects (Fitzgerald and Haller, 2018). We argue that the link between policy uncertainty and prices is also important because firms may partially adjust to changes in the risk of protection.

Our last contribution is to estimate and quantify the import price effects of increased TPU during Brexit. There is evidence for the effects Brexit-related uncertainty on trade flows and firm-level trade participation outcomes. Graziano et al. (2021), hereafter GHL, show that Brexit uncertainty reduces goods trade between the U.K. and the remaining E.U. members and that there are spillovers to trade with the E.U.'s preferential trade agreement (PTA) partners (Graziano et al., 2020). Crowley et al. (2018, 2020) find that reductions in U.K. exports following the referendum are due to reduced firm entry and higher exit on the extensive margin in products where the MFN threat tariffs were higher. Douch et al. (2022) also find that U.K. firms diverted trade to non-E.U. destinations where MFN risk was higher. Ahmad et al. (forthcoming) show that services trade flows are significantly reduced due to Brexit uncertainty.

Our findings also complement the growing evidence that the Brexit referendum increased U.K. import prices, in pounds, due to the exchange rate depreciation. Breinlich et al. (2022) show that this depreciation contributed to inflation and eroded U.K. living standards. Costa et al. (2022) find that industries more exposed to this depreciation experienced a cost shock leading to real wage reductions.[3] Our estimates are robust to controlling for exchange rate depreciation. We also find that exchange rate risk contributed to higher import prices, in euro, similarly to tariff risk.

The rest of the paper proceeds as follows. Section 1 provides background information on Brexit and motivating evidence on the relationship between prices and Brexit uncertainty. Section 3 describes the theoretical model. Section 4 maps the model to an estimating equation and describes the data, results, and robustness checks. We quantify the effects on prices in Section 5 and conclude in Section 6.

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[3]Fernandes and Winters (2021) find that Portuguese exporters to the actually reduce prices relative to non-U.K. destinations after the referendum; they do not disentangle the exchange rate depreciation and TPU channels.

1　The Referendum and Import Prices

There has been support for leaving the E.U. since the U.K. joined in 1973. Just two years after joining, the U.K. renegotiated terms, and the then Labour prime minister argued that "I believe our renegotiation objectives have been substantially though not completely achieved", supporting a vote to remain (H.M. Government, 1975).[4] In the 1975 referendum that followed, "Leave" was supported by one-third of voters. Since then, the support for leaving the E.U. has averaged 40% in public polling (GHL, 2021).

In 2015, Prime Minister Cameron promised a referendum on E.U. membership if the Conservative Party won the general election. Following their victory, the E.U. Referendum Bill was introduced in May 2015. The bill passed the House of Commons in September and the House of Lords in December 2015. In February 2016 the vote was scheduled, and in the June 2016 referendum "Leave" prevailed by 52%.

Before the referendum, U.K. imports from the E.U. were duty-free. However, in the months leading up to the referendum and until late 2020, the prospect of a no-deal Brexit was real.[5] The uncertainty associated with such a regime change would have led E.U. exporters to delay investments in new products and investment to improve or maintain distribution networks in the U.K. The unease over potentially higher tariffs was warranted. Even though a new Trade and Cooperation Agreement was eventually negotiated that maintained duty-free access, changes to origin rules meant that in practice, some U.K. imports from the E.U. did end up facing MFN tariff rates.[6]

Some initial aggregate evidence suggests that this uncertainty affected import prices. To examine this, we compute rolling 12-month changes (year-over-year) in the U.K. exact CES Import Price Index (IPI) by four-digit industry-E.U. origin pairs.[7] We average these indices by month and take the ratio relative to the aggregate IPI from the U.K.'s Office of National Statistics, thus measuring the relative price change of E.U. imports. If Brexit uncertainty increased the price of E.U. imports, then the relative price we construct should co-move with the probability of Brexit. This is what we observe in Figure 1, where we plot this relative price index alongside a measure of Brexit probability—the daily average price of a prediction market contract paying $1 if the "Leave" option won the Brexit referendum in June 2016.[8]

Since the relative import price described above could also reflect changes in other variables, such as the exchange rate, we explore variation in tariff risk across industries. Specifically, we use the threat of the U.K. increasing tariffs to some level that was salient and known: E.U. external MFN tariffs. We separate industries into high- and low-risk groups based on this measure to compute separate averages for the CES IPI as previously described.[9] Figure 2 shows the high- to low-risk price index, which exhibits a co-movement with the probability of Brexit.[10]

Both relative import price indices—the E.U. to ROW and E.U. high-risk versus low-risk industries—track the leave probability. This evidence suggests that imports became costlier due to Brexit uncertainty. In the

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[4]This quote appeared in a 1975 referendum pamphlet mailed to all U.K. households.

[5]In a speech to the Bundestag, Angela Merkel noted "There is still the chance of an agreement... We are continuing to work on it, but we are also prepared for conditions which we cannot accept (Reuters, 2020).

[6]For example, the U.K. Fashion and Textile Association reports that 74% of members faced new tariffs under the agreement (UK Fashion & Textile Assoc., 2021).

[7]We employ monthly trade data from Eurostat that includes bilateral product trade weight and volume as detailed in the data section.

[8]There is a correlation of 0.84 between the 60-day moving average of the price of a Brexit contract and the linearly interpolated E.U. to overall IPI. The prediction market contract price is from PredictIt.org and is described in the data section.

[9]We use the bottom and top quartiles of the distribution of external tariffs to identify these two groups.

[10]The 60-day moving average of the price of a Brexit contract correlates with the linearly interpolated ratio of high- to low-risk import price indices by 0.64.

following sections we outline the theory that formalizes this relationship and provide a regression framework to estimate and quantify its significance.

2   Price Effects of Demand Uncertainty

In this section we provide a model for firm decisions under demand uncertainty. We then discuss its application in the context of Brexit and derive the implications of firm decisions for price indices.

2.1   Firm Decisions

The baseline approach uses the framework of GHL (2021), and thus we provide only its key elements and implications. The key point is that demand uncertainty lowers export entry investments, leading to higher industry ideal price indices, as shown in HL (2017).

Import demand has CES σ over varieties v in each industry V at time t, $q_{vt} = \left[D_t \left(\tau_{Vt} \right)^{-\sigma} \right] p_v^{-\sigma}$. Demand is decreasing in the producer variety price, p_v, and any tax, $\tau_{Vt} \geq 1$. The demand shifter, $D_t = \varepsilon Y_t \left(P_t \right)^{\sigma-1}$, increases with the fraction of importer income spent on differentiated goods, εY_t, and the average price of competing goods reflected in the CES price index, P_t. In the baseline, we assume firms take D_t as given and are independent of uncertainty in any given industry V as long as each industry V is sufficiently small.[11]

The firm observes all relevant information before producing and pricing in a monopolistically competitive market each period. The optimal factory-gate price is the standard mark-up over marginal cost, $p_v/wc_v = \sigma/(\sigma-1)$. The equilibrium operating profit is decreasing in the firm-specific cost parameter and is increasing in the business conditions term, $a_{Vt} = D_t \tau_{Vt}^{-\sigma}$.

$$\pi_{vt} = a_{Vt} c_v^{1-\sigma} \tilde{\sigma}, \tag{1}$$

where $\tilde{\sigma}$ is a constant function of σ. The firm believes that a new a_V' is drawn with probability $\gamma \in [0,1]$ from a distribution $\bar{H}\left(a_V\right)$, independent of the current a_V.

Exporting requires a sunk cost investment, K, if the firm either (i) did not export in the previous period or (ii) exported but its export capital then fully depreciated; depreciation occurs with probability β.[12] The firm faces a dynamic problem, and GHL show that the optimal export entry decision satisfies a cutoff rule. The rule requires the expected value of exporting net of the sunk costs to exceed the value of waiting to enter in a later period. Only firms with costs below the following threshold value enter.

$$c_{Vt}^U = \underbrace{\left[\frac{a_{Vt} \tilde{\sigma}}{(1-\beta)K} \right]^{\frac{1}{\sigma-1}}}_{\equiv c_{Vt}^D} \times \underbrace{\left[1 + \frac{\beta \gamma \left(\bar{\omega}_{Vt} - 1 \right)}{1 - \beta \left(1 - \gamma \right)} \right]^{\frac{1}{\sigma-1}}}_{\equiv U_{Vt}}, \tag{2}$$

where c_{Vt}^D is the deterministic cutoff (reflecting the present discounted value of investment) and $U_{Vt} \in (0,1]$

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[11]Relaxing this assumption does not change the qualitative results or estimation approach, as HL (2017) show, but it can affect the quantification, as we subsequently show in Online Appendix B.

[12]We abstract from fixed costs of exporting, so firms with undepreciated K will export every period.

is an uncertainty factor that depends on a measure of profit tail risk

$$\bar{\omega}_{Vt} - 1 = -\bar{H}(a_{Vt})\frac{a_{Vt} - \mathbb{E}(a_V' \le a_{Vt})}{a_{Vt}} \quad \in (-1, 0]. \tag{3}$$

This measure is the product of the probability of worsening conditions and the expected proportion of profits lost in that event. The uncertainty factor U_{Vt} implies a stricter entry cutoff whenever future conditions are expected to change and there is tail risk ($\bar{\omega}_{Vt} - 1 < 0$)—a sufficient statistic for risk in this setting.

2.2 Policy Risks

GHL model the implementation of Brexit as a change in regime where new business conditions, e.g., from changes in trade policies, are drawn from a distribution H^{BR}. In the period we consider, the change in regime was itself uncertain, which we capture via a probability m_t that implies the following weighted measure of average tail risk:

$$\bar{\omega}_{Vt} = m_t \omega_{Vt}^{BR} + (1 - m_t)\omega_{Vt}^{EU}. \tag{4}$$

Each of the tail risks is defined similarly to (3) but uses their respective distributions. Increases in the probability of Brexit increase tail risk if H^{EU} SSD H^{BR}. The main source of time variation in tail risk before the referendum is unanticipated shocks to the probability that the U.K. leaves the E.U., so our empirical approach exploits this variation and its heterogeneous impact across industries:

$$\Delta\bar{\omega}_{Vt} = \left(\omega_V^{BR} - \omega_V^{EU}\right)\Delta m_t. \tag{5}$$

2.3 Price Indices

We derive the impacts of TPU on ideal price indices. Consumers in the importing country face a price $\tilde{p}_{vt}$ for a variety, which reflects the producer price and any trade costs. The ideal price index aggregates over available varieties, $P_t^{1-\sigma} = \int_{v \in \Omega_t} \tilde{p}_{vt}^{1-\sigma} dG(c)$, where $\Omega_t = \cup \Omega_{jt}$. We model $P_{xV,t}$ as an exporter(x)-industry(V) pair to match the data subsequently used. For now, we omit the exporter subscript for simplicity. In the model's baseline version, TPU increases P_{Vt} by altering decisions to enter the market. We provide the structural relationship between P_{Vt} and TPU that guides the estimation in this setting and discuss extensions in Section 3.4.2 and Appendix A.3.

In a stationary state, only the firms with costs below the current threshold export. The stationary price index is given by the following decreasing function of the cost cutoff:

$$P\left(c_{Vt}^U, \tilde{\tau}_{Vt}\right) = \tilde{\tau}_{Vt} \left[\int_{c^{\min}}^{c_{Vt}^U} c_v^{1-\sigma} dG(c)\right]^{1/(1-\sigma)}, \tag{6}$$

where G is the cost distribution and $\tilde{\tau}_{Vt} \equiv \tau_{Vt}/\left(N_V\tilde{\sigma}\right)^{1/(\sigma-1)}$ is increasing in any marginal costs that are not firm-specific, e.g., tariffs, transport, and wages, and is decreasing in the mass of potential exporters.

We can generalize the relationship in Equation (6) in different ways that are relevant for the estimation. First, after a negative shock, there are transition dynamics due to the death of some firms that do not re-enter. Thus, we must augment the expression in (6) with a legacy term, Λ_t^h, which depends on the history

of recent negative shocks. In Appendix (A.1) we provide the expression for alternative histories and show that the main modification is that the impact of increases in TPU is initially attenuated due to surviving firms. Second, we can capture changes in the prices of continuing firms due to TPU if we incorporate an export marginal cost that requires a sunk cost (see HL, 2017).[13]

3 Estimation

In this section we provide an overview of the baseline econometric approach and review the data used to measure the key covariates. We then estimate the model and investigate the mechanism and robustness of our findings.

3.1 Baseline

Our approach is to estimate the impact of TPU on the price index via the cutoff. We provide the derivation in Appendix A.4; here we note two intermediate steps useful in interpreting the key coefficient. We use a first-order log approximation of (6) around a common marginal trade cost and cutoff, $\left(c^U, \tilde{\tau}\right)$. After re-introducing the exporter subscripts relevant for the estimation, this yields

$$\Delta \ln P \left(c_{xVt}^U, \tilde{\tau}_{Vt}\right) = \hat{k} \cdot \Delta \ln c_{xVt}^U + \Delta \ln \tilde{\tau}_{Vt} + u_{Vt}, \tag{7}$$
$$= \hat{k}\tilde{\beta} \left[\left(\omega_{xV}^{BR} - \omega_{xV}^{EU}\right) \Delta m_t\right] + \alpha + e_{Vt},$$

where $\hat{k} < 0$ is the average price index elasticity with respect to the entry cutoff and u is the approximation error.[14] The second line uses the cutoff in (2) and the change in tail risk in (5), and we define $\tilde{\beta} \equiv \frac{1}{\sigma-1} \frac{\beta\gamma}{1-\beta(1-\gamma)}$. The vector α of fixed effects includes exporter-HS4 and exporter-month in the baseline, which captures changes in factors other than TPU (see Appendix A.4).

We next map the policy risk term $\left(\omega_{xV}^{BR} - \omega_{xV}^{EU}\right) \Delta m_t$ in equation (7) into observable measures for our estimating equation. We denote initial business conditions under E.U. terms as a_{Vt}^{EU}, where EU is a state of the policy distribution. Conditional on a policy change arriving, there is a discrete probability that conditions worsen to MFN terms where $a_{xVt}^{MFN} < a_{xVt}^{EU}$.[15] For example, if under Brexit the probability of moving to MFN is η^{BR}, then the tail risk term is simply $\omega_{xV}^{BR} - 1 = \eta^{BR} \left[a_{xVt}^{MFN}/a_{xVt}^{EU} - 1\right]$. Because the probability of moving to MFN is lower under the E.U. distribution, $\eta^{EU} < \eta^{BR}$, we have

$$\left(\omega_{xV}^{BR} - \omega_{xV}^{EU}\right) = \left(\eta^{BR} - \eta^{EU}\right) \left[a_{xVt}^{MFN}/a_{xVt}^{EU} - 1\right] \le 0. \tag{8}$$

To estimate equation (7), we require measures of the tail risk term (8) and the probability of Brexit m_t. We describe the data we use below and then map them into an estimated structural coefficient for our baseline regression.

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[13]Subsequently, we extend the model to allow for a quality decision and show that TPU increases imply higher average measured prices.

[14]Specifically, $\hat{k} \equiv \left(\frac{\partial \ln P\left(c_{Vt}^U, \cdot\right)}{\partial \ln c_{Vt}^U}\right)|_{c^U} = (1-\sigma)^{-1} \frac{d \ln \int_{c^{\min}}^{c_{Vt}^U} c_v^{1-\sigma} dG(c)}{d \ln c_{Vt}^U}|_{c^U}$. If $G(c)$ is exponential, then $\hat{k}$ is independent of c^U and $u_{Vt} = 0$.

[15]It is possible that conditions could worsen to a trade war with much higher tariffs, but we ignore this possibility because it is not significant in GHL. Since only the probability and level of worsening conditions matter to decisions, we do not need to specify the distribution for improving conditions.

3.2 Data and Measurement

In this subsection we describe our method for computing the price index from country-product data and our measures of uncertainty and policy risk.

3.2.1 Price Indices

We construct U.K. import prices using import values and quantities from Eurostat.[16]

A variety is identified by an exporter-product pair, xv. We calculate variety prices as the ratio of (euro) import values to weight at the eight-digit level of the Combined Nomenclature (CN) for each E.U. exporter.[17] We match the monthly time variation in our uncertainty measure by computing rolling 12-month price changes on a year-over-year basis. Therefore, we compare price levels in time t to levels in $t-12$—before the Conservative Party won the 2015 general election that opened up the possibility of a Brexit referendum.".

To construct exact CES price indices, we aggregate varieties to industries—defined at the four-digit level and denoted by V—for each exporter x. These allow for changing varieties as in Feenstra (1994).[18]

$$\Delta \ln P_{xV,t} = \sum_{\Omega_{xV,t}^{cont}} w_{xv,t} \ln \frac{p_{xv,t}}{p_{xv,t-12}} + (\sigma - 1)^{-1} \ln \frac{\psi_{xV,t}}{\psi_{xV,t-12}}. \tag{9}$$

The first term captures the average price index change of continuing varieties—those with positive exports at t and $t-12$—aggregated using the Sato-Vartia weights, $w_{xv,t}$. The lower cost of living due to net variety entry is captured in the second term and is measured by changes in the import shares of continuing varieties, $\psi_{xV,t}$, defined in Appendix A.2.1. We assume a common $\sigma = 4$ as in GHL, and we trim the top and bottom 2.5% of the constructed $\Delta \ln P_{xV,t}$ to reduce the impact of outliers. We test the robustness of both of these choices.

The top panel of Table 1 summarizes the exact CES price index and its two components for high sunk cost industries as defined by GHL.[19] The average price index increased by 1.3 lp, primarily due to an increase in the price of continuing varieties. The variety term decreased on average by 0.2 lp. More importantly, there is substantial variation across industries and time in the price index, which we exploit in the estimation.

3.2.2 Uncertainty

Since we cannot directly measure exporter beliefs on the probability of Brexit m_t, we model them as a function of the observable prediction market contract price. We use the average daily price of a contract traded in PredictIt.org that pays \$1 if a majority voted for Brexit in the referendum of June 2016 and \$0 otherwise. The market opened on May 27, 2015 and closed on June 24, 2016 and is described in detail in GHL.

We refer to the natural log of the monthly average price as a market-based variable, mbv_t. Changes in

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[16]We use the only available consistent quantity measure, weight. Proportional changes in weight are equal to proportional changes in units for any specific firm-level variety.

[17]The CN extends the six-digit product codes of the Harmonized System by adding two extra digits. It identifies approximately 9,500 products.

[18]There are 1,223 four-digit headings, e.g., 0401 is "Non-concentrated milk and cream," and the subcodes 04012011 and 04012091 refer to different fat content.

[19]GHL identify sunk cost industries by estimating those with conditional export persistence.

exporter beliefs about the probability of Brexit are modeled as the first-order log linearization around an initial pre-referendum probability of Brexit m_0:

$$\Delta m_t \approx m_0 \sum_{l=0,\ldots,L} r_l^m \left(mbv_{t-l} - mbv_0 \right). \tag{10}$$

The parameters r_l^m represent the elasticity of exporter beliefs with respect to a change in the market-based variable. We allow for a distributed lag process, so $\sum r_l^m$ is the respective long-run elasticity. In the robustness section, we replace mbv_t with polling averages from Number Cruncher Politics.

The aggregation of the daily information in logs to a monthly level allows us to match it to the monthly price information in the trade data. This aggregation also helps to ameliorate concerns about noise trading. The average Brexit probability was about 0.3, and the average exit share in polls was 0.47 over the sample period, as shown in the second panel of Table 1 (in logs). Figure 1 shows the variation over time.

3.2.3 Policy Risk

Our baseline estimation assumes that the only change in business conditions occurs through tariffs. Thus, the tail risk term $a_{xVt}^{MFN}/a_{xVt}^{EU} - 1$ in (8) varies only by industry and is given by $1 - \left(\tau_V^{MFN}/\tau_V^{EU} \right)^{-\sigma}$. We set the current tariff $\tau_V^{EU} = 1$ and use the E.U. MFN tariffs from the UNCTAD TRAINS database for 2015. We then take simple averages to aggregate the tail risk indicators to the four-digit level and thus match the price index aggregation. Additionally, we assume $\sigma = 4$ consistently with the value used for the variety term in the exact price index.[20]

We summarize the MFN risk statistics in the lower panel of Table 1. The average MFN risk is 0.15, indicating that, all else equal, a tariff increase to MFN rates would reduce the operating profits of E.U. exporters by an average of 15%. The MFN risk measure ranges from 0 (no risk) to 0.72. For a given product v, all E.U. exporters face the same rate in the U.K., resulting in the same (simple) average in V. In the robustness analysis, we consider variation in risk at the exporter-industry level generated by variation in export composition.

3.3 Baseline Estimation Equation and Results

We combine the observable data described above into a baseline estimation equation and interpret the coefficients structurally as the cross-partial elasticity of the probability of Brexit and policy risk. We then estimate this elasticity and show it is robust to controls for unobserved shocks.

We substitute equations (8) and (10) into the second line of (7). After replacing the tail risk terms with their discretized observable tariff counterparts from above, we obtain

$$\Delta \ln P \left(c_{xVt}^U, \tilde{\tau}_{Vt} \right) = \sum_{l=0,\ldots,L} \mathcal{E}_{t-l} \cdot \left[mbv_{t-l} \left(1 - \left(\tau_V^{MFN} \right)^{-\sigma} \right) \right] + \alpha + \tilde{e}_{xVt}. \tag{11}$$

The main coefficient of interest is $\mathcal{E}_{t-l} \equiv \hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU} \right) \times m_0 r_l^m$, which is the cross-partial of (11) with respect to the Brexit probability and policy risk. The sum of these cross-partials $\mathcal{E} = \sum_l \mathcal{E}_{t-l}$ is the permanent cross-elasticity of uncertainty and risk. According to the model, $\mathcal{E}$ depends on the following

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[20]We also use MFN tariffs applied by other developed countries to construct instruments at the same aggregation level.

parameters: the price elasticity to entry $\hat{k} < 0$, the expected duration of the next policy state $\tilde{\beta} > 0$, the elasticity of substitution, and the baseline Brexit probability in logs $m_0 < 0$. Therefore, the model predicts that the price index is increasing in the TPU cross-elasticity $\mathcal{E} > 0$ unless $\eta^{BR} \leq \eta^{EU}$, i.e., firms believe the probability of MFN barriers is unaffected by Brexit, or $\sum_l r_l^m = 0$, i.e., the probability measure does not permanently affect beliefs about the likelihood of a Brexit state.

Recall that the price index, $\Delta \ln P \left(c_{xVt}^U, \tilde{\tau}_{Vt} \right)$, measures the 12-month, year-over-year change in prices where the base $t - 12$ month is in the period from May 2014 to June 2015 before the Conservative Party won the 2015 General Election. The policy tail risk is multiplied by the time-varying probability of Brexit, mbv_{t-l}. The latter is also relative to a common value mbv_0 from the period before May 2015.[21] We include exporter-industry and exporter-time fixed effects in the baseline $\alpha \equiv \alpha_{xt} + \tilde{\alpha}_{xV}$.[22]

In Table 2, we estimate equation (11) and find evidence that increases in the probability of Brexit raised U.K. import prices from the E.U. in products with MFN risk. In column 1 we employ OLS and control for exporter-HS4 factors (xV) and monthly time effects (t), revealing a positive and significant estimate. However, the MFN risk measure is potentially subject to measurement error due to agents' uncertainty about whether the U.K. would apply the existing E.U. MFN tariff schedule after Brexit or choose new rates. To address this, we instrument MFN risk with the median MFN tariff risk in each product across other developed countries.[23] The interaction of this median risk with the Brexit probability serves as the excluded instrument for our uncertainty measure. In column 2 we find that the estimate using the IV is more than double the OLS estimate, suggesting that the OLS estimates may be attenuated, as GHL (2021) find for exports.

The above specifications do not control for unobserved heterogeneity due to exporter shocks, such as exchange rate shocks that could affect price indices and correlate with our uncertainty measure. Moreover, individual exporter shocks may introduce bias to our estimates, depending on the risk distribution in their export product baskets to the U.K. To address these concerns, we include exporter-time effects in columns 3 and 4, resulting in positive and significant estimates for both OLS and IV estimates. In columns 5 and 6, we also address industry shocks by controlling for section-time fixed effects and find similar estimates. We employ the specification in column 4 as the baseline.

3.4 Mechanisms

We now test alternative mechanisms through which TPU increases the price index via export investments. Specifically, we show that the TPU effect is only significant in industries with export sunk costs. We then extend the model to incorporate quality competition.

3.4.1 Sunk Cost Industries and Price Index Decomposition

The model predicts that TPU only affects industries with sunk costs of exporting to the U.K. Therefore, we use these industries in the baseline results presented in Table 2. Table 3 demonstrates the absence of TPU

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[21] We generalize this to accommodate other shocks through exchange rate dynamics in the robustness checks.

[22] These industry effect $\tilde{\alpha}_V$ absorbs the terms $\alpha_V - \hat{k}\tilde{\beta}m_0 \times \left(\eta^{BR} - \eta^{EU} \right) \left(\left(\tau_V^{MFN} \right)^{-\sigma} - 1 \right) \sum_l r_l^m mbv_0$ in addition to any other unobservable industry heterogeneity. The composite, idiosyncratic error term, $\widetilde{e}_{xVt}$, incorporates the approximation error in the price index and beliefs about Brexit probability.

[23] We calculate the median MFN risk factor across Australia, Canada, Japan, and the E.U. and aggregate to the HS4 level.

effects in industries without these sunk costs. In column 1 we pool all industries and find a positive and significant effect of Brexit uncertainty. This effect is smaller than for the baseline high sunk cost subsample, as replicated in column 2. In column 3 we find no effect for the low sunk cost sample; the estimate is 60% smaller and statistically insignificant.

Table 3 also presents a complete decomposition of the total effect into the continuing and new variety channels. TPU does not affect either channel in industries where sunk costs are low (third column for each group), but it does so for the industries where they are high. In the latter case the continuing varieties channel accounts for over 80% of the full price index elasticity.

One explanation for the smaller share of the variety effect is the product-level aggregation. We identify a reduction in variety only when all firms in a E.U. country stop exporting an eight-digit product to the U.K. Nonetheless, this variety measure contains other relevant information for the model's mechanism. Specifically, GHL find that Brexit uncertainty reduced entry and increased exit for exporter eight-digit varieties. Moreover, that effect is smaller for exit than entry, which is consistent with sunk costs causing hysteresis.

3.4.2 Extension to Quality Competition

How can the model account for higher prices in continuing varieties due to TPU? The baseline version, where uncertainty affects only firm export entry, cannot do so because it predicts constant prices for incumbents. Therefore, we extend the baseline model to allow for quality competition, which is novel in the context of TPU models and, importantly, generates testable predictions.[24]

The key point is that TPU increases the entry productivity cutoff, and the resulting selection changes the average observed prices that we measure. This selection from TPU implies that the average productivity of the continuers is higher, but the implications for their average price depend on whether there is quality competition. In the baseline model with price competition, the continuers are those firms with lower average prices, so the prediction is that TPU lowers the average price over continuing firms. However, when firms can choose quality, this selection effect can lead to the opposite prediction for average observed prices.

To model this quality effect, we build on the insight from static models where firms compete in terms of quality-adjusted prices (cf. Johnson, 2012; Baldwin and Harrigan, 2011). In the static models, a deterioration in current export conditions leads the firms with the highest quality-adjusted prices to exit and implies a lower average price over the continuing firms if they are adjusted for quality but not necessarily if they are not. In the appendix, we incorporate this quality mechanism into our setting with TPU and show that all our results can be re-derived in terms of quality-adjusted prices. Moreover, we derive the condition for TPU to increase the average observed prices, which is that the observed firm price is negatively correlated with its unobserved quality-adjusted price. In industries where this occurs, the continuing firms after a TPU increase are those with higher observed prices (due to higher quality).[25]

To test if TPU increases the average price of continuing varieties in industries where quality competition is more important, we first identify those industries and then re-estimate by quality subsample. We follow Baldwin and Ito (2011) to classify the importance of quality competition. Specifically, we use panel data

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[24]An alternative extension of the model can also generate this price prediction. Suppose that incumbents can invest in a technology with a lower marginal cost. If that investment is sunk, then TPU will reduce it, leading to higher marginal costs and prices of incumbents (see Handley and Limão, 2017). However, these data do not allow us to test this directly.

[25]As noted above, GHL find evidence for exporter eight-digit exit when uncertainty increases.

for each exporter eight-digit product (e.g., German widget exports) to different destinations. We then run separate log-linear regressions of free-on-board unit values on distance and other controls. More distant markets are costlier to reach, so if the free-on-board price distance elasticity for German widgets is positive, then it indicates quality competition, and if it is negative, it indicates price competition.

We employ export data for the top three E.U. exporters—Germany, France, and the Netherlands—annually from 2012 to 2016. The destination markets are the rest of the E.U. countries (not the U.K.) plus the remaining OECD countries as well as Brazil, Russia, India, and China. We estimate an average of about 8,000 coefficients for each of the three exporters. Out of the 7,200 overlapping eight-digit estimates, over 2,000 have significant distance elasticities, and almost 1,900 are also positive, indicating quality competition.

We take these 1,900 industry indicators, 1_v^q, and aggregate them to the four-digit heading level to generate an index of the prevalence of quality competition. Specifically, for each four-digit heading, we compute the share of U.K. imports from the E.U. over all the eight-digit products with $1_v^q = 1$. Finally, we split the baseline sample's four-digit industries based on values of this index.

The results in Table 4 show the role of quality in explaining the TPU impact on average continuing variety prices. Column 1 replicates the coefficient for the high sunk cost sample from Table 3. We find no TPU effect on average prices for the subsample of industries with a quality competition index below the median. In contrast, the effect is positive and significant for industries above the median. Moreover, industries in the top quartile of quality competition show an even larger effect, as seen in column 4.[26]

The quality model does not have a differential prediction for the variety term. This is consistent with the estimates in columns 7–9 that show no significant difference across subsamples.

3.5 Robustness

In this subsection, we examine if the results are robust to alternatives measures of risk, outliers, aggregation, and alternative demand shocks.

3.5.1 Trade Policy Risk Measurement

In Table 5, we test our MFN risk measure. In column 1 we replicate the baseline, which uses the average MFN risk across six-digit products—the level at which we have the MFN tariff information. In column 2, we instead weight the MFN risk by using U.K. import shares within each HS4 over the 2012–2014 period. The coefficient estimate is very close to the baseline.

A related source of measurement error from averaging could occur if countries export different varieties within industries. To address this, we calculate exporter-specific MFN risk averages. The results are similar, as shown in column 3.

A similar measurement issue could arise if the elasticity of substitution is very heterogeneous across varieties within an HS4 relative to our choice of $\sigma = 4$. Thus, we restrict the sample to HS4 industries with a median σ between 2 and 6, as estimated by Broda and Weinstein (2006). The point estimate in column 4 is higher in this case but not significantly different from the baseline.

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[26]We find similar results with an alternative classification that aggregates the eight-digit index using exporter-specific shares (available upon request).

3.5.2 Uncertainty Measurement

We employ the simple average of the daily log contract price to construct the baseline uncertainty measure. In Table 6, we provide two alternative measures. First, we calculate the weighted average by using the daily number of transactions; this may be useful if days with high trading volume reflect more precise information about referendum beliefs. Column 2 shows that the estimates are the same when normalized to one standard deviation unit of the baseline measure. Second, we employ monthly polling averages about the referendum outcome; the standardized TPU elasticity is the same as in the baseline (column 3).

3.5.3 Price Index Outliers

In the baseline sample we excluded 5% of outlier observations—the top and bottom 2.5% of the observations of the dependent variable, $\Delta \ln P_{xV,t}$. We provide robustness to this decision in Table A1. In column 1, we remove any restriction on outliers and find a positive and significant estimate, which is slightly higher than the baseline. In columns 2–5, we use alternative samples varying the outlier restriction.[27] All four estimates are positive and significant and are of similar magnitude.

3.5.4 Aggregation Level

The baseline price indices are constructed at the four-digit heading level, and we now consider alternative levels of aggregation while using the same variety definition. We first consider indices at the two-digit level (by exporter-month). The MFN risk is now constructed by taking the simple average of the rates in each two-digit industry. In Table A2, columns 1–3, we include the coefficient estimates and find they are all positive. The impact on the variety term is also positive but noisier, as in the baseline. The magnitude of the coefficients is also higher.

Next, we disaggregate to a lower level, six-digit sub-headings, and use the MFN risk variables reported at that level. Columns 7–9 show that all three coefficients are positive and significant but are lower in magnitude.

3.5.5 Other Import Demand Shocks

We examine if the positive price elasticity is due to alternative U.K. demand shocks, which would threaten the identification if they were correlated with the MFN risk.

Sector level. If the demand shocks were common to different groups of four-digit industries, then we may be able to control for them. The last two columns of Table 2 control for this possibility by including separate time fixed effects for each of the 20 sections that group similar four-digit industries. The estimates are similar to the baseline.

Stockpiling. Suppose that firms in the U.K. responded to Brexit uncertainty by stockpiling goods in anticipation of higher protection. If so, then they would do so more in industries with a higher MFN risk. If

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[27]Specifically, we exclude the top and bottom 1%, 2.5% (the baseline), and 5% as well as the extreme outliers, which are defined as observations that are above (below) the third (first) quartile by more than three times the interquartile range.

this increase in demand was present in 2016, then it could explain the higher prices, but it would also imply an increase in quantities. We test this in Table A3 by separately regressing the log change in both import values and a quantity index on the regressors used in the baseline. Column 1 shows a negative elasticity of import values to Brexit uncertainty (similar to GHL, 2021 but at a different level of aggregation). Column 3 shows the results for a quantity index computed by subtracting the log change in the price index from the log change in import value. The quantity elasticity to Brexit uncertainty is negative and significant, ruling out stockpiling import demand shocks in this period, even if they remain plausible later in the process and in other settings (see, e.g. Alessandria et al., 2019).

3.6 Exchange Rate Shocks and Uncertainty

We examine how pound exchange rate shocks may influence our baseline estimates, which addresses identification issues and provides insights into the interpretation of the magnitudes for our quantification exercises.

3.6.1 Robustness to Exchange Rate Shocks

The pound/euro rate depreciated about 10% over the sample period, i.e., before the referendum vote. In Figure A1 we see this exchange rate co-moved with the Brexit probability contract, since the beginning of 2016. We examine if this correlation generates an omitted variable bias.

The baseline estimation controls for exchange rate changes, at least partially, through the exporter-time effects, but we perform two robustness checks. First, we allow the exchange rate to have differential effects across MFN risk by including an additional interaction of this variable with the time-varying bilateral exchange rate. In Table 7, column 2 we use this interaction instead of the contract price and find that it is statistically insignificant. In column 3 we include both MFN risk interactions and find that the one using the exchange rate is insignificant, whereas our baseline measure coefficient is unchanged. Second, we allow the exchange rate to have heterogeneous pass-through rates by interacting it with section dummies, and in column 4 we find it does not affect the baseline coefficient.

3.6.2 Robustness to Exchange Rate Uncertainty

The pound depreciated by almost 7% immediately after the referendum result was known, followed by an additional depreciation in 2017 when Brexit became more likely. In the context of our model, the exchange rate enters demand and thus profits exactly in the same way as tariffs. A depreciation increases the cost for U.K. consumers and thus lowers their demand.

To incorporate exchange rates ξ_{xt}, we must replace τ_{Vt} with $\tau_{Vt}\xi_{xt}$ everywhere in the model. Deriving the full tail risk with respect to both shocks requires using the approach of Carballo et al. (2022). We simplify their model and assume a discrete distribution where there is a single state with higher probability of both higher barriers and a depreciation. In that case, the proportional reduction in profits is $a_{Vxt}^{MFN}/a_{Vxt}^{EU} = \left(\frac{\xi_x^{MFN}}{\xi_x^{EU}}\frac{\tau_V^{MFN}}{\tau_V^{EU}}\right)^{-\sigma}$. We can then use the same estimation method as in our baseline but with a new risk measure that also reflects the potential depreciation.

How do our results change when we allow for this type of exchange rate risk? If all E.U. exporters to the U.K. set prices in a common currency, euros, then the risk increases proportionally for all xV, so it is

already reflected in the baseline estimation coefficient. If exporters price in their own currency, then the risk will vary by xV, and we capture this by constructing a bilateral exchange rate augmented risk, which we refer to as "MFN-XR" risk. This requires taking a stand on the long-term rate, ξ_x^{MFN}, that firms in the pre-referendum period expected if the MFN state materialized. However, this measure is not readily available. Our approach is to take the average realized depreciation in the period after the referendum (in 2017) relative to the period before for each country to proxy for ξ_x^{MFN}/ξ_x^{EU}.[28] The average "MFN-XR" risk is about 3.5 times higher than the MFN risk alone, as seen in Table 1. Moreover, the baseline MFN varies only over industries, but the MFN-XR also varies across exporters with different currencies. Figure A2 illustrates the distribution of the combined MFN-XR measure along with the baseline MFN risk measure without depreciation risk.

In Table 8, we use the MFN-XR risk to replace MFN risk in the baseline. We obtain the same sign and significance as in the baseline. The magnitude of the coefficients is higher, but in the next section we show that it yields the same quantitative impacts for MFN risk alone as the baseline.

4   Quantification

In this section we use the baseline estimates to quantify the different effects of Brexit uncertainty on U.K. import prices from the E.U.

4.1   Average Effects on Import Prices

We calculate the average effect of Brexit uncertainty as follows. Using the structural approximation in (7), our risk measure, and the definition of the estimation coefficients, we obtain the impact of a permanent change in probability, Δmbv, on the stationary price change as

$$\mathbb{E}_{xV} \ln \frac{P_{xV}\left(m',.\right)}{P_{xV}\left(m,.\right)} = \sum_l \mathcal{E}_{t-l} \cdot \Delta mbv \cdot \mathbb{E}_{xV}\left(1 - \frac{a_{xV}^{MFN}}{a_{xV}}\right), \tag{12}$$

where the last term evaluates the impact at the average risk.

Permanent uncertainty impacts.   Figure 3(a) plots the expression in equation (12) for alternative probability shocks at the average MFN tariff risk in the sample. In this case, the average price increase is 1.64 lp after a permanent one standard deviation increase in the probability measure, indicated by the first vertical line at 12.1 lp. The maximum shock in panel (a) is 123 lp and corresponds to the probability of going from the sample mean to unity. This is an order of magnitude larger than the standard deviation shock.

Post-referendum, it was not clear if and when the U.K. would leave the E.U. We use the approach of GHL and calculate a scenario where there is Brexit regret after the referendum, i.e., Bregret.[29] We find that Bregret attenuates the maximum probability increase after the referendum, increasing it by only 82 lp. The average price increases by 11.2 lp under this Bregret scenario.

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[28] The average pound/euro exchange rate was 0.88 in 2017, reflecting a pound depreciation of about 16% relative to the sample period.

[29] GHL argue there was a 33% Brexit regret after the referendum: support fell by about 33% after the referendum relative to the relationship between the referendum result and previous polls, and the conditional probability of Brexit after the referendum as measured by the probability of triggering the necessary Article 50 was about two-thirds.

Impact over levels of MFN risk. In Figure 3(b) we plot the expression in equation (12) at a given increase in Brexit probability (82 lp) and varying export risk on the x-axis. We also plot the contribution from continuing varieties, so the difference between the lines represent the variety effect. Several relevant scenarios emerge. First, if the only source of risk was an increase in tariffs from duty-free to MFN levels, then the x-axis can be directly interpreted as $100 \times \ln \tau^{MFN}$. The first vertical line shows the effect at the sample mean tariff risk, equivalent to a threat tariff of 4 lp. It implies an average 11.2 lp price increase, as previously described, with almost 9 lp of this increase attributed to continuing varieties. The second vertical line adds one standard deviation to the mean, implying a price index increase of 20.6 lp, with 16 lp due to the continuing term.

Second, as noted in Section 3.6, the exchange rate depreciation enters the risk term symmetrically, allowing us to evaluate alternative MFN scenarios accompanied by depreciation risk using Figure 3(b). In the hours following the referendum result, the pound depreciated 6.5 lp against the euro, reflecting the news shock. If exporters expected a combined 4 lp tariff and a 6.5 lp depreciation under an MFN scenario, a total shock of 10.5 lp, then the price effect from a higher Brexit probability is about 26 lp.

Third, the same approach can be applied if E.U. exporters believed they would face additional non-tariff barriers under Brexit. The average ad valorem equivalent of E.U. non-tariff barriers plus MFN tariffs on non-members is around 9 lp (Kee et al., 2009), located around the mean plus one standard deviation of MFN tariffs in the sample described in the first scenario.

Summary and robustness. Table 9 summarizes the results above and verifies their robustness. Column 2 computes the average effect for industries with $\sigma \in (2, 6)$, using the estimates in Table 5. The uncertainty elasticity and the average risk in these industries is higher, and therefore the average price effect under Bregret is now higher than the baseline, at 16.6 lp.

Additionally, the risks are heterogeneous across xV, and if they are correlated with imports, then the results in this table would differ from weighted price changes. To examine this, we aggregate the xV price changes using the theory-consistent Sato-Vartia weights.[30] Table 9 shows that this weighted average, in column 3, is nearly identical to the simple average in column 1.

4.2 Aggregate Price Effects

How important are the import price increases we estimate relative to the aggregate data? A full answer requires a specific general equilibrium model. Here we provide bounds for the first-order impacts. First, we compute the direct effect on final consumers, assuming all else is constant, i.e., ignoring any general equilibrium effects. Second, we account for one equilibrium effect of individual prices on the aggregate and show how it attenuates the direct effect.

We start with a simple partial effect calculation using the import share of our sample in total consumption. We assume that expenditure on E.U. imports is Cobb-Douglas with an expenditure share μ. Therefore, all else equal, the change in the aggregate price index is proportional to the expected change in exporter-industry prices: $[\Delta \ln P] = \mu \mathbb{E}_{xV} \left[\ln \frac{P_{xV}(m,.)}{P_{xV}(m',.)} \right]$. Our baseline sample is 94% of total U.K. imports from the E.U. in 2015 (Eurostat), which includes firm and household consumption. In OECD national account data, this

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[30]These weights are $w_{xV} \equiv \frac{(s_{xV,T} - s_{xV,T-12})/(\ln s_{xV,T} - \ln s_{xV,T-12})}{\sum_{x,V}(s_{xV,T} - s_{xV,T-12})/(\ln s_{xV,T} - \ln s_{xV,T-12})}$, where T is August 2015-June 2016.

amounts to 7% of total U.K. consumption in 2015 on a euro basis. Multiplying this share with the import price change under Bregret in Table 9, we obtain an increase of 0.77 lp.

The partial increase described above can be considered an upper bound since it would be partially offset if we consider the general equilibrium feedback into aggregate prices. In Online Appendix B we assume a Pareto distribution of firm productivities and use it to derive a lower bound of 9.1 lp for the aggregate import price effect in Table 9. Using the E.U. import share of 7%, the latter implies a total effect on prices of 0.64 lp ($= 9.1 \times 0.07$).[31] Alternatively, we can use input-output tables to break out expenditure on manufactured goods by firms and households. Imports account for 46% of expenditure of manufactured goods in the U.K., and E.U. imports represent 52% of total U.K. imports in 2015. This implies an increase in the manufacturing price index of at least 2.2 lp ($= 9.1 \times 0.46 \times 0.52$) using our lower bound.

5   Conclusion

Several decades of trade liberalization reduced TPU through credible commitments via the GATT/WTO and agreements such as the E.U. and NAFTA. However, several of these commitments have now been reversed, as evident from the ongoing U.S.-China tariff war and the U.K.'s departure from the E.U. in 2021. TPU emerges as another important feature of deglobalization, incurring significant costs even without actual policy change.

Our paper focuses on the price and welfare effects of policy uncertainty of the U.K.'s Brexit referendum in 2016. We find evidence that import prices rise and welfare falls in response to this Brexit uncertainty, with more pronounced effects observed in industries more exposed to the risk of higher MFN tariffs. Importantly, the effects are only present in industries with higher sunk investment costs of market entry. Our estimates predict that an increase in Brexit uncertainty with a magnitude similar to the referendum increased the price of imports from the E.U. by 9 to 11 lp. After adjusting for the share of E.U. imports in total expenditure on manufactured imports, these import price effects reduced real incomes by more than 0.6 lp.

These findings have broader implications for the renegotiation of trade agreements and rules. Disputes over trade can have trade and welfare effects through TPU even before any applied policies change, warranting careful consideration in future research comparing pre- versus post-negotiation outcomes. The U.K., for example, continues to renegotiate bilateral agreements previously made through E.U. membership, while the U.S. has renegotiated agreements with Canada, Mexico, and Korea and continues to challenge the spirit of WTO rules in its ongoing trade war with China. Such events may erode the credibility of existing agreements, leading to partial disintegration where exporters and importers perceive risks of renegotiation or exit. Future research should also consider the effects on importers' input sourcing strategies and explore the mechanism through which TPU affects consumer prices.

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[31]If we also incorporate data on import shares of household and firm expenditure on manufactured goods using input-output tables, then we obtain a slightly smaller effect of 0.53 lp. See Online Appendix B.

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Figure 1: Brexit Uncertainty and Relative EU Import Price Index
(year-over-year change by month)
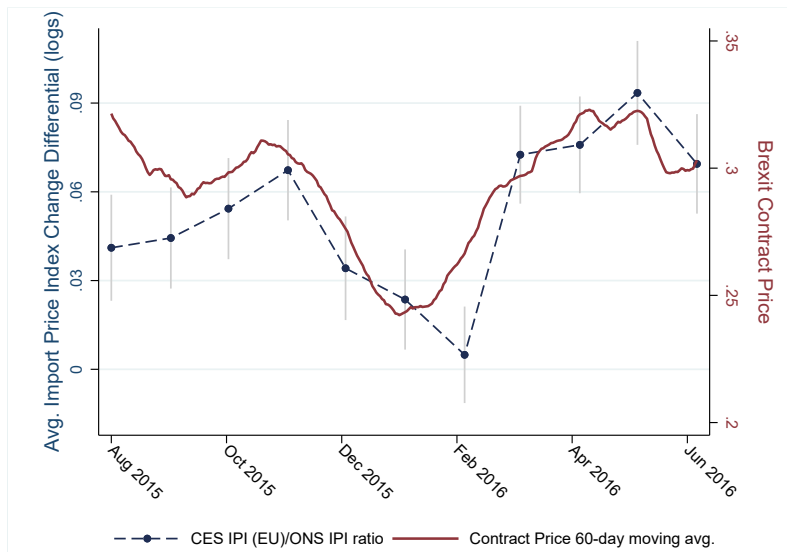
Notes: The blue dashed line is the month-specific difference between the average yearly
ln change of exporter-HS4 specific CES EU Import Price Indices (IPI) and the yearly ln
change of the ONS IPI. The solid line is the 60-day moving average of the price of a
contract (left axis) on PredictIt.org that pays $1 if Britain votes to leave the EU and
zero otherwise.

Figure 2: CES Price Index Change (year-over-year by month)—High vs.
Low MFN Risk Industries
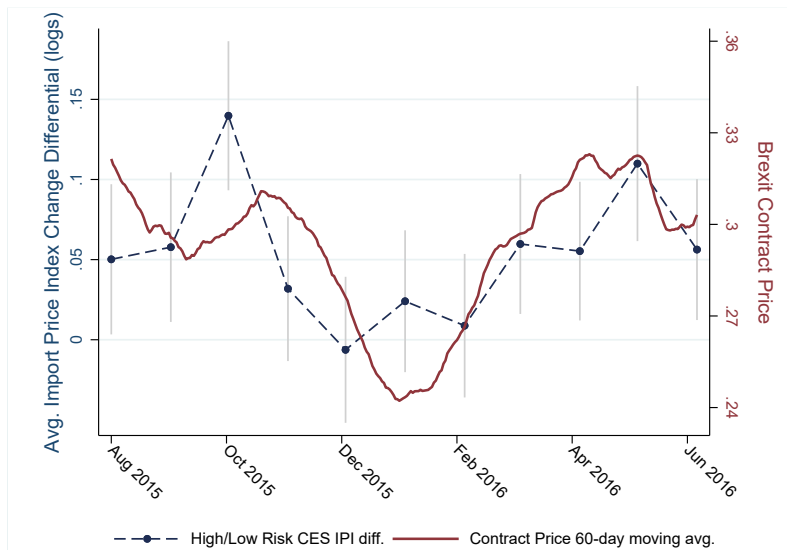
Notes: We define high MFN risk exporter by HS4 industries using the top quartile of
MFN risk distribution and low MFN risk using the bottom quartile. The blue dashed
line is the year-over year, month-specific average difference between the ln CES price
indices of high vs low MFN risk by exporter-HS4. The solid line is the 60-day moving
average of the price of a contract (left axis) on PredictIt.org that pays $1 if Britain
votes to leave the EU and zero otherwise.

Figure 3: Quantification of Price Index Response to Brexit Uncertainty

(a) Response of Price Index to Probability of Brexit at Mean MFN Risk
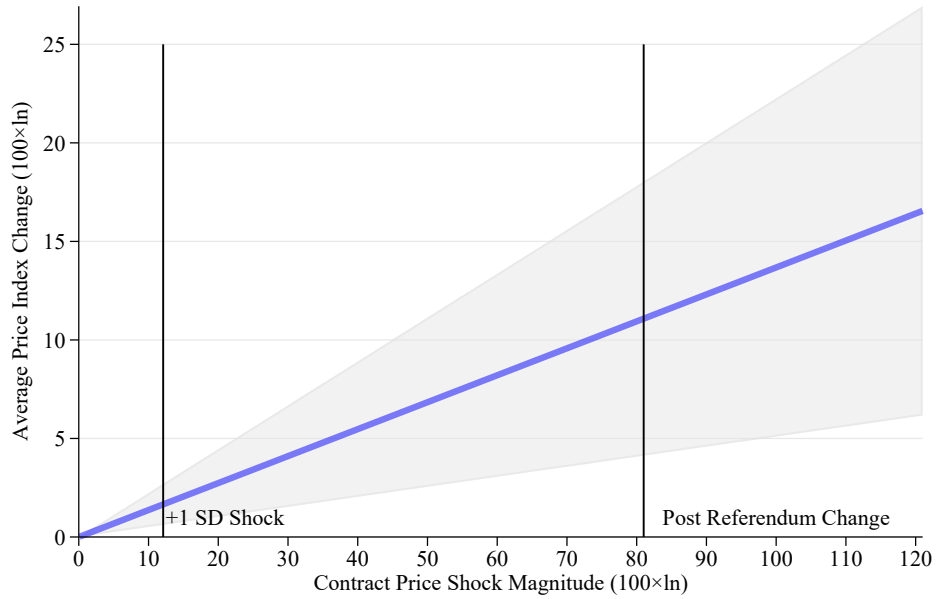
(b) Response of Price Index over MFN Risk Level for large Brexit Uncertainty Shock
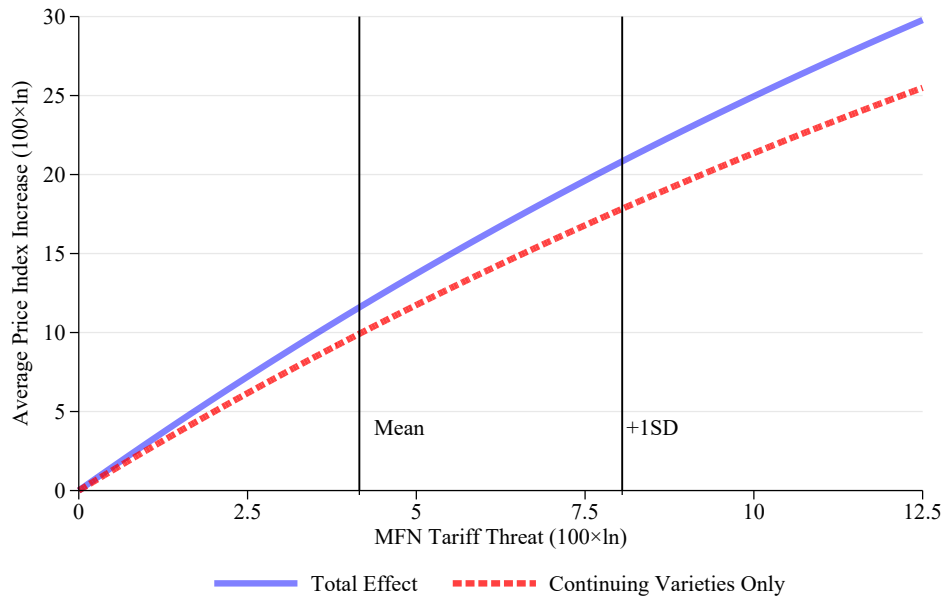
Notes: Panel (a) uses the IV estimate of the cross-elasticity from Table 3 to compute the average change in the price index at the mean MFN risk factor over the range of a log change in the contract price ($100 \times \Delta mbv$) from 0 to 120. Panel (b) holds the change in the log contract price fixed at 82 log points, i.e. the implied shock assuming 1/3 Bregret as shown in Table 9, and then increases the potential MFN tariff from 0 to 12.5 log points. See section 5 for additional details on the computation.

Table 1: Summary Statistics - Regression Sample

	Mean	SD	Median	Min	Max	N
		Price Variables				
Ideal Price Index	0.013	0.540	-0.006	-1.849	1.981	72,448
Continuing varieties	0.016	0.525	-0.005	-3.315	3.643	72,448
Variety term	-0.002	0.177	0.000	-2.535	2.977	72,448
		Brexit Uncertainty Measures				
ln Pr(Brexit)	-1.227	0.121	-1.191	-1.501	-0.985	72,448
Transation Vol. Weighted	-1.214	0.146	-1.192	-1.480	-0.854	72,448
ln Poll Share (Exit)	-0.760	0.029	-0.769	-0.801	-0.691	72,448
		Risk Measures				
MFN Risk	0.148	0.125	0.115	0.000	0.721	72,448
MFN Risk IV	0.094	0.097	0.064	0.000	0.409	72,448
MFN-XR Risk	0.534	0.071	0.520	0.377	0.848	72,448
MFN-XR Risk IV	0.504	0.057	0.490	0.377	0.703	72,448

Notes: Ideal price index reflects yearly changes in exporter-HS4-month level for EU exporters to the UK (2015 membership). Pr(Brexit) defined as the monthly average (ln) MBV, the leave referendum prediction market contract price. MFN risk defined as $1-(\tau^{MFN})^{-\sigma}$, where $\sigma=4$ and $\tau^{MFN}=1+$MFN advalorem/100. MFN-XR risk defined as $1-(\tau^{MFN}\hat{e})^{-\sigma}$, where $\hat{e}$ is the expected pound depreciation as measured by e^{2017}/e with e^{2017} the exporter-specific average exchange rate in 2017 and e the exporter-specific average exchange rate over the sample period. Poll Share is the ln of monthly average across polls of the fraction of voters supporting exit.

Table 2: MFN Risk Price Effects. High Sunk Cost Industries.

Yearly Changes in Ideal Price Index by month (ln) 8/15-6/16

	1	2	3	4	4	5
	OLS	IV	OLS	IV	OLS	IV
Pr(Brexit) × MFN Risk	0.461	1.074	0.405	0.923	0.646	1.730
	(0.212)	(0.300)	(0.209)	(0.296)	(0.283)	(0.495)
Exporter-HS4 FE	Yes	Yes	Yes	Yes	Yes	Yes
Month FE	Yes	Yes	No	No	No	No
Exporter-Month FE	No	No	Yes	Yes	Yes	Yes
Section-Month FE	No	No	No	No	Yes	Yes
N	72,448	72,448	72,448	72,448	72,448	72,448
R2	0.232	n/a	0.238	n/a	0.241	n/a
IV weak ID KP_F-stat	n/a	272.2	n/a	296.1	n/a	93.22

Notes: Includes all industries with export sunk costs as defined in the text. Dependent variable yearly changes in ideal price index defined at the exporter-HS4-month level for EU exporters to the UK (2015 membership). All estimations include exporter-HS4 effects. Pr(Brexit) defined as the monthly average of the (ln) Brexit contract price from PredictIt.org. MFN risk defined as $1-(\tau^{MFN})-\sigma$, where $\sigma=4$ and $\tau^{MFN}=1+$MFN ad valorem/100. In columns 2, 4,6 we instrument the risk variables by their respective median HS4-specific risk across four large countries (Australia, Canada, Japan and US). Coefficients report the sum of current and two monthly lags. Robust standard errors clustered at the exporter-HS4 level below estimates.

Table 3: MFN Risk Price Effect Decomposition and High vs. Low Sunk Cost Industries
Yearly Changes in Ideal Price Index by month (ln) 8/15-6/16

	Ideal Price Index			Continuing Varieties			Variety term		
	1	2	3	1	2	3	1	2	3
Sunk Cost Sample:	All	High	Low	All	High	Low	All	High	Low
Pr(Brexit) × MFN Risk	0.720	0.923	0.279	0.638	0.790	0.555	0.0815	0.134	-0.276
	(0.287)	(0.296)	(0.997)	(0.275)	(0.281)	(0.975)	(0.079)	(0.086)	(0.229)
N	90,605	72,448	11,895	90,605	72,448	11,895	90,605	72,448	11,895

Notes: Dependent variable yearly changes in ideal price index defined at the exporter-HS4-month level for EU exporters to the UK (2015 membership). Rows labelled "High" correspond to the baseline sample including only industries with high sunk costs of exporting and column 2 replicates the baseline in column 4 of Table 2 so all corresponding notes apply here. "Low" includes the remaining industries and "All" is their union. We instrument the risk variables by their respective median HS4-specific risk across four large countries (Australia, Canada, Japan and US). Coefficients report the sum of current and two monthly lags. Robust standard errors clustered at the exporter-HS4 level below estimates. All estimations include exporter-HS4 and exporter-month fixed effects.

Table 4: MFN Risk Price Effect Decomposition for Quality vs Price Competition
Yearly Changes in Ideal Price Index components by month (ln) 8/15-6/16

	Continuing Varieties				Variety term			
	1	2	3	4	6	7	8	9
Quality Quantile Sample:	All	<50	>50	>75	All	<50	>50	>75
Pr(Brexit) × MFN Risk	0.790	-0.0526	1.066	1.728	0.134	0.132	0.134	0.176
	(0.281)	(0.624)	(0.319)	(0.516)	(0.086)	(0.164)	(0.101)	(0.168)
N	72,448	18,902	53,545	23,752	72,448	18,902	53,545	23,752

Notes: Dependent variable is the two components of the yearly change in ideal price index defined at the exporter-HS4-month level for EU exporters to the UK. To partition the sample by the degree of quality competition, we first define an indicator by HS8 digit products using top three exporters to the UK (Germany, France, Netherlands) using a regression of price on distance with other controls described in the main text. When the coefficient on distance is positive and significant at 5% level for all three top exporters we code the HS8 variety as quality competition. We then aggregate these indicators to the HS4 level using their 8 digit export shares to define the quantile samples. Columns 1 and 6 replicate the pooled high sunk cost sample using the IV approach described in Table 3. We run a similar regression for each of the subsamples in columns 2-4 and 7-9, which include exporter-HS4 and exporter-month fixed effects. We report the sum of current and two monthly lags and the robust standard errors clustered at the exporter-HS4 level.

Table 5: MFN Risk Price Effects--Alternative MFN Risk Measures

Yearly Changes in Ideal Price Index by month (ln) 8/15-6/16

	1	2	3	4
Risk:	Simple Average	Weighted Average	Exporter-specific	Industries with $\sigma \epsilon$ [2, 6]
Pr(Brexit)×MFN Risk	0.923 (0.296)	0.962 (0.301)	0.957 (0.296)	1.253 (0.373)
N	72,448	72,448	72,108	45,747

Notes: Column 1 replicates Table 2 column 4 baseline and all notes in that table apply here as well. Column 1 MFN risk is the simple mean over all possible HS6 within an HS-4. In column 2 we use a weighted average and in column 3 an exporter specific measure (mean only over products it exports). Column 4 excludes industries with sigmas higher than 6 and lower than 2 based on estimations using US import data in Broda and Weinstein (2006). We instrument the risk variables by their respective median HS4-specific risk across four large countries (Australia, Canada, Japan and US). Coefficients report the sum of current and two monthly lags. Robust standard errors clustered at the exporter-HS4 level below estimates. All estimations include exporter-HS4 and exporter-month fixed effects.

Table 6: MFN Risk Price Effects--Alternative Brexit Measures

Yearly Changes in Ideal Price Index by month (ln) 8/15-6/16

	1	2	3
	Average Contract Price		Opinion Polling Averages
	Unweighted	Weighted	
Pr(Brexit) × MFN Risk *Standardized Coef.*	0.118 (0.0377)	0.120 (0.0421)	0.129 (0.0438)
N	72,448	72,448	72,448

Notes: Column 1 replicates Table 2 column 4 baseline and all notes in that table apply here as well. Column 2 weights the daily probabilities with the square root of the volume of daily transactions within month. Column 3 replaces Pr(Brexit) with the (ln) average share of exit voters adjusted by (1 - undecided share of voters). The poll data for July and August 2015 is imputed to be the same as in September when polls begin. We instrument the risk variables by their respective median HS4-specific risk across four large countries (Australia, Canada, Japan and US). Coefficients report the sum of current and two monthly lags. Robust standard errors clustered at the exporter-HS4 level below estimates. All estimations include exporter-HS4 and exporter-month fixed effects.

Table 7: MFN Risk Price Effects--Exchange Rate Robustness
Monthly Export Value (ln) 8/15-6/16

	1	2	3	4
		Exchange Rate as Brexit Measure		Differential XR Pass-Throughs
Pr(Brexit) × MFN Risk	0.923		0.973	0.950
	(0.296)		(0.345)	(0.292)
XR × MFN Risk		-0.230	-0.151	
		(0.688)	(0.807)	
N	72,448	72,448	72,448	72,448
IV weak ID KP_F-stat	296.1	297.4	148.4	299.7

Notes: Column 1 replicates Table 2 column 4 baseline and all notes in that table apply here as well. Column 2 replaces Pr(Brexit) by the (ln) Pound/exporter exchange rate. Column 3 includes both the monthly average of the (ln) Brexit contract price from PredictIt.org and the (ln) Pound/exporter exchange rate. Column 4 includes XR-section fixed effects interactions (not shown). We instrument the risk variables by their respective median HS4-specific risk across four large countries (Australia, Canada, Japan and US). Coefficients report the sum of current and two monthly lags. Robust standard errors clustered at the exporter-HS4 level below estimates. All estimations include exporter-HS4 and exporter-month fixed effects.

Table 8: MFN and Pound Depreciation Risk Price Effects
Monthly Export Value (ln) 8/15-6/16

	1	2	3
	Ideal Price Index	Continuing Varieties	Variety term
Pr(Brexit) × MFN-XR Risk	1.644	1.406	0.237
	(0.538)	(0.513)	(0.155)
N	72,448	72,448	72,448
IV weak ID KP_F-stat	72.08	72.08	72.08

Notes: Dependent variable yearly changes in ideal price index defined at the exporter-HS4-month level for EU exporters to the UK (2015 membership). Pr(Brexit) defined as the monthly average of the (ln) Brexit contract price from predictit.org. MFN-XR risk defined as $1-(\tau^{MFN} \hat{e})^{\wedge}(-\sigma)$, where $\hat{e}$ is the expected pound depreciation as measured by e^{2017}/e with e^{2017} being the exporter-specific average exchange rate in 2017 and e the exporter-specific average exchange rate over the sample period. We instrument the risk variables by their respective median HS4-specific risk across four large countries (Australia, Canada, Japan and US). Coefficients report the sum of current and two monthly lags. Robust standard errors clustered at the exporter-HS4 level below estimates. All estimations include exporter-HS4 and exporter-month fixed effects.

Table 9: Brexit Uncertainty Price Effects Impacts at Average MFN Risk

	1	2	3
	Simple Average Effect		**Weighted Average Effect**
	All	Industries with σ ϵ [2, 6]	
Uncertainty Elasticity	0.137	0.203	0.134
A. One SD shock	0.017	0.025	
B. Referendum shock (1/3 Bregret)	0.112	0.166	0.110

Notes: Calculations in column "All" employ IV coefficients using the baseline in Table 2, column 4 and the MBV measure and summary statistics from Table 1. Row A uses the standard deviation of the MBV measure. Row B uses a shock of (2/3) the change in average ln(MBV) in the pre-referendum period to its post referendum value.

A Appendix

A.1 Theory: Negative Shocks

To allow for non-stationary periods we rewrite the general price expression as $P_{Vt}^{1-\sigma} = P\left(c_{Vt}^U, .\right)^{1-\sigma} + \Lambda_t^h$. It augments the stationary component in (6) with a legacy term, Λ_t^h, which reflects the contribution of any survivors that would not enter under current conditions. This requires that we compare the current period cutoff c_{Vt}^U to a set of historical cutoffs extant in each of the previous s time periods since that last stationary equilibrium denoted by c_{V0}^U. We focus on two simple cases. The case in the first line below is the baseline: if the cutoff is no lower than c_{V0}^U, then there is no legacy. The second line reflects any case where conditions have worsened, $c_{Vt}^U < c_{V0}^U$, but may have partially improved. In the latter case there is a fraction β^t of legacy firms that still survive.

$$\Lambda_t^h = \begin{cases} 0 & \text{if } c_{Vt}^U \geq \max\left\{c_{V,t-s}^U\right\}_{s=1}^{s=t} \\ \beta^t\left(\tilde{\tau}_{Vt}\right)^{1-\sigma}\left[\int_{c_{Vt}^U}^{c_{V0}^U} c_v^{1-\sigma} dG\left(c\right)\right] & \text{if } c_{Vt}^U = \max\left\{c_{V,t-s}^U\right\}_{s=1}^{s=t-1} \text{ and } c_{Vt}^U < c_{V0}^U \end{cases}$$

We can then augment the approximation in (7) to reflect this. Using a first-order approximation around the cutoff and $\tilde{\tau}$ in the last stationary state $\theta = \left\{c_0^U, \tilde{\tau}_0\right\}$ we obtain

$$\ln \frac{P_{Vt}}{P\left(\theta\right)} \approx \left(1 - \beta^t\right)\left(\ln \frac{\tilde{\tau}_{Vt}}{\tilde{\tau}_0} + \hat{k}\left(\theta\right)\ln \frac{c_{Vt}^U}{c_0^U}\right) \tag{13}$$

A.2 Data

Our empirical analysis employs four main datasets:

(i) Monthly import value and weight data between June 2014 and June 2016 at the 8 digits of the Combined Nomenclature (CN) from Eurostat,

(ii) Daily contract price data from PredictIt.org betting on the June 2016 Brexit referendum outcome,

(iii) Most Favored Nation tariffs (MFN) at the HS6 level of the E.U., Australia, Canada, Japan and the U.S. from TRAINS, and

(iv) pound sterling exchange rate monthly averages from the IMF.

Most of these data are described in the main text. Below we provide details on how we construct price indices and measures of quality competition and sunk costs of exporting by industry.

A.2.1 U.K. Import Data

We employ monthly import value and weight data from Eurostat to compute the CES price indices at the HS4 level in equation 9, reproduced below for readability.

$$\Delta \ln P_{xV,t} = \sum_{\Omega_{xV,t}^{cont}} w_{xv,t} \ln \frac{p_{xv,t}}{p_{xv,t-12}} + \left(\sigma - 1\right)^{-1} \ln \frac{\psi_{xV,t}}{\psi_{xV,t-12}},$$

This expression includes three empirical measures we construct using this data:

1. The CN-exporter specific prices $p_{xv,t}$ and $p_{xv,t-12}$.

2. The CN-exporter specific Sato-Vartia weights $w_{xv,t}$.

3. The HS4-exporter specific value share of continuing CN-exporter varieties $\psi_{xV,t}$ and $\psi_{xV,t-12}$.

To construct prices, we employ the observed U.K. import value and weight, i.e. $p_{xv,t} = \frac{\text{Import Value}_{xv,t}}{\text{Import Weight}_{xv,t}}$. We employ weight and not quantities because the unit measure for quantities is not defined for most of the CN codes. Weights are measured in tons.

Sato-Vartia weights are defined as in Broda and Weinstein (2006):

$$w_{xv,t} \equiv \frac{(s_{xv,t} - s_{xv,t-12})/(\ln s_{xv,t} - \ln s_{xv,t-12})}{\sum_{xv \in \Omega_{xV,t}^{cont}}(s_{xv,t} - s_{xv,t-12})/(\ln s_{xv,t} - \ln s_{xv,t-12})}$$

where $s_{xv,t}$ is the market share of xv within xV at t. We compute shares as $s_{xv,t} \equiv \frac{\text{Import Value}_{xv,t}}{\sum_{xv \in \Omega_{xV,t}^{cont}} \text{Import Value}_{xv,t}}$. The set $\Omega_{xV,t}^{cont}$ contains all the continuing varieties observed at t and $t-12$, where t is months. This implies that the set of continuing varieties can vary over t within the sample.

Finally, we compute the share of continuing varieties at the HS4-exporter for t and $t-12$ level as follows:

$$\psi_{xV,t} \equiv \frac{\sum_{xv \in \Omega_{xV,t}^{cont}} \text{Import Value}_{xv,t}}{\sum_{xv \in \Omega_{xV,t}} \text{Import Value}_{xv,t}}$$

$$\psi_{xV,t-12} \equiv \frac{\sum_{xv \in \Omega_{xV,t-12}^{cont}} \text{Import Value}_{xv,t-12}}{\sum_{xv \in \Omega_{xV,t-12}} \text{Import Value}_{xv,t-12}}$$

where the sums are over the continuing varieties in the numerator and over the entire set of observed exporter-CN varieties in the denominator.

Given that our sample period is August 2015 to June 2016, we employ data from this period for variables indexed by t and from August 2015 to June 2015 for variables indexed by $t1 - 12$.

About 13% of our sample can be identified as extreme outliers by using an interquartile range rule (i.e. observations that are either higher $o_H = q_3 + IQR * 1.5$ or lower than $o_L = q_1 - IQR * 1.5$, where q_3 is the third quartile, q_1 is the first quartile, and IQR is the interquartile range. In our baseline results, we employ a less strict definition of outliers and trim the top and bottom 2.5% of values. We present robustness to alternative trimming criteria.

A.3 Identification of Quality Competition

The goal of this section is to explain the methodology employed to classify industries into either quality or price competition. We observe average unit values at the 8-digit level by exporter-destination. These averages reflect the prices of firms with heterogeneous productivity and thus their composition. Therefore, for any given industry the set of exporting firms to a farther/tougher destination will be more productive. The empirical question we must address is whether this selection translates into lower observed average prices, as in Melitz (2003), or higher prices due to selection on quality, as in Baldwin and Harrigan (2011) and Johnson (2012).

A.3.1 Econometric Approach

We follow Baldwin and Ito (2012) to classify industries into either price or quality competition. For each exporter-CN pair $\{x, v\}$, e.g. German (x) widget (v) exports, we run separate log-linear regressions for unit value, which vary over time and destination, $p_{i,t}^{x,v}$

$$\log p_{i,t}^{x,v} = \beta_1^{x,v} \log d_i^x + \beta_2^{x,v} \log GDP_{i,t} + \beta_3^{x,v} \log GDPpc_{i,t} + \delta_t^{x,v} + \epsilon_{i,t}^{x,v} \quad \text{if } n_{i,t}^{x,v} \geq 20$$

We include exporter-CN-time fixed effects, $\delta_t^{x,v}$, and control for the destination i aggregate and per capita GDP. Conditional on these controls, a destination at higher distance, d, is costlier to access. Our null hypothesis is the negative relationship predicted by the baseline model with price competition, so we can't

reject it when $\beta_1^{x,v} < 0$. We interpret a positive and significant distance coefficient, $\beta_1^{x,v} > 0$, as evidence of quality competition.

A.3.2 Data

To construct unit values, we employ Eurostat 2012-2016 annual export values and weight from Germany, France and the Netherlands to the rest of the E.U. countries (excluding the U.K.) plus other OECD countries as well as Brazil, Russia, India and China. We select these three exporters because they are the top three E.U. suppliers to the U.K. We use distance, GDP and GDP per capita data from CEPII.

A.3.3 Results

In Table A4, we present the distribution of estimates for each country as in Baldwin and Ito (2012). In the last row, we include the overlap between the three, meaning the characteristic indicated at the top holds for the three countries.

We find a positive coefficient for the three countries in 1897 8-digit CN codes, suggesting quality competition. This is 20% of the total number of 8-digit products exported by the three countries. This overlap is almost identical to what Baldwin and Ito (2012) find at this level of aggregation for European countries if we consider the number of estimates we collect (7204). Table A4 mirrors their Table 5 in terms of using the top three exporters. In practice, this only implies using the Netherlands instead of the U.K.

A.3.4 Aggregation to the Industry Level

We follow three steps to translate the quality indicators into an HS4 measure for the degree of quality competition. First, we classify an 8-digit product v to have quality competition if $\beta^{x,v} > 0$ for all $x = $ Germany, France, and Netherlands.[32] Second, we aggregate this indicator to the HS4 level by weighting it using v shares in U.K. import from the E.U. within each HS4. We employ U.K. imports between August 2014 and June 2015 to compute these weights (the same base period as our price indices).[33] Finally, we split the baseline sample into below and above the median according to the HS4-level indicators. We also create an additional sample for the top quartile in terms of this measure. Figure A3 shows the prevalence of quality competition over HS4 percentiles.

A.4 Estimation Equation (11)

We use a first-order log approximation of (6) around a common point given by the cost cutoff and marginal trade costs not specific to the firm, $\left(c^U, \tilde{\tau}\right)$. This yields the equation (11) from section 3.3

$$\Delta \ln P\left(c_{Vt}^U, \tilde{\tau}_{Vt}\right) = \hat{k}\Delta \ln c_{Vt}^U + \Delta \ln \tilde{\tau}_{Vt} + u_{Vt}, \tag{14}$$
$$= \hat{k}\tilde{\beta} \cdot \left[\left(\omega_V^{BR} - \omega_V^{EU}\right)\Delta m_t\right] + \alpha_V + \alpha_t + e_{Vt}$$

To derive this equation we start with the equation for the cutoff

$$c_{Vt}^U = \underbrace{\left[\frac{a_{Vt}\tilde{\sigma}}{(1-\beta)K}\right]^{\frac{1}{\sigma-1}}}_{\equiv c_{Vt}^D} \times \underbrace{\left[1 + \frac{\beta\gamma\left(\bar{\omega}_{Vt} - 1\right)}{1 - \beta\left(1-\gamma\right)}\right]^{\frac{1}{\sigma-1}}}_{\equiv U_{Vt}}.$$

---

[32]We only use estimates from regressions with more than 20 degrees of freedom (89% of all exporter-CN estimations).

[33]As a robustness we also used an alternative weight: the U.K. import shares of $\{v, x\}$ within its HS4 imports from the respective exporter x.

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We take logs, difference the cutoff equation, and then simplify as follows

$$\Delta \ln c_{Vt}^U = \frac{1}{\sigma - 1} \left(\Delta \ln a_{Vt} + \Delta \ln \left[1 + \frac{\beta \gamma (\bar{\omega}_{Vt} - 1)}{1 - \beta (1 - \gamma)} \right] \right)$$

$$\approx \frac{1}{\sigma - 1} \left(\Delta \ln a_{Vt} + \frac{\beta \gamma}{1 - \beta (1 - \gamma)} \Delta (\bar{\omega}_{Vt} - 1) \right)$$

$$= \frac{1}{\sigma - 1} \left(\Delta \ln a_{Vt} + \frac{\beta \gamma}{1 - \beta (1 - \gamma)} \left(\omega_V^{BR} - \omega_V^{EU} \right) \Delta m_t \right) \quad (15)$$

The second line follows by using the approximate $\ln(1 + x) \approx x$. The third lines replaces $\Delta \bar{\omega}_{Vt}$ with the tail risk differential from section 2.2, equation (5).

We then replace $\Delta \ln c_{Vt}^U$ in equation (14) with (15) and simplify as follows

$$\Delta \ln P \left(c_{Vt}^U, \tilde{\tau}_{Vt} \right) \approx \Delta \ln \tilde{\tau}_{Vt} + \frac{\hat{k}}{\sigma - 1} \left(\Delta \ln a_{Vt} + \frac{\beta \gamma}{1 - \beta (1 - \gamma)} \left(\omega_V^{BR} - \omega_V^{EU} \right) \Delta m_t \right) + u_{Vt}$$

$$\approx \Delta \ln \tau_{Vt} / (N_V \tilde{\sigma})^{1/(\sigma-1)} - \frac{\hat{k} \sigma}{\sigma - 1} \Delta \ln \tau_{Vt} + \frac{\hat{k}}{\sigma - 1} \Delta \ln D_t + \hat{k} \tilde{\beta} \times \left(\omega_V^{BR} - \omega_V^{EU} \right) \Delta m_t + u_{Vt}$$

$$\approx \hat{k} \tilde{\beta} m_0 \times \left(\omega_V^{BR} - \omega_V^{EU} \right) \sum_{l=0,\dots,L} r_l^m (mbv_{t-l} - mbv_0) + \alpha_V + \alpha_t + e_{Vt}$$

$$= \hat{k} \tilde{\beta}_0 \times \left(\eta^{BR} - \eta^{EU} \right) \left(1 - a_{Vt}^{MFN} / a_{Vt}^{EU} \right) m_0 \sum_{l=0,\dots,L} r_l^m mbv_{t-l} + \tilde{\alpha}_V + \alpha_t + e_{Vt}$$

$$= \sum_l \mathcal{E}_{t-l} \left[mbv_{t-l} \left(a_{Vt}^{MFN} / a_{Vt}^{EU} - 1 \right) \right] + \tilde{\alpha}_V + \alpha_t + e_{Vt} \quad (16)$$

From line 1 to 2, we collect terms after replacing $\tilde{\tau}_{Vt}$, a_{Vt}, and $\tilde{\beta}$ by their definitions. In line 3, we replace Δm_t by the approximation (10) from section 3.2 and collect terms. The unobserved industry time fixed effects in line 3 are given by $\alpha_V + \alpha_t + e_{Vt} \equiv \left(1 - \frac{\hat{k}\sigma}{\sigma-1} \right) \Delta \ln \tau_{Vt} - (\sigma - 1)^{-1} \Delta \ln N_V + \frac{\hat{k}}{\sigma-1} \Delta \ln D_t + u_{Vt}$ plus an error term reflecting the approximation of Δm_t. We then replace $\omega_V^x = \eta^x \left[a_{Vt}^{MFN} / a_{Vt}^{EU} - 1 \right]$ for $x = \{BR, EU\}$, factor out -1, and collect terms to obtain line 4. We collect terms in the time-invariant common pre-referendum level of the Brexit uncertainty proxied by mbv_0 and redefine the industry fixed effects as $\tilde{\alpha}_V \equiv \alpha_V - \hat{k} \tilde{\beta} m_0 \times \left(\omega_V^{BR} - \omega_V^{EU} \right) \sum_{l=0,\dots,L} r_l^m mbv_0$.

The final line collects terms into the identified coefficients $\mathcal{E}_{t-l} \equiv \left[\hat{k} \tilde{\beta} \times \left(\eta^{BR} - \eta^{EU} \right) m_0 r_l^m \right]$. We assume the change in applied tariffs or any other industry-specific barriers is negligible, i.e. $\Delta \tau_{Vt} \approx 0$. Even if there were some variation, most should be absorbed by the industry-exporter and exporter-time effects in our baseline. Changes in the unobserved mass of exporters by industry, $\Delta \ln N_V$, and time-varying U.K. expenditure $\Delta \ln D_t$ also appear in the fixed effects. Any constant time trend for N_V and D_t is already differenced out as is seasonal variation because we always compare to the same month in the previous year.

Figure A1: Brexit Referendum Contract and Pound/Euro Exchange Rate
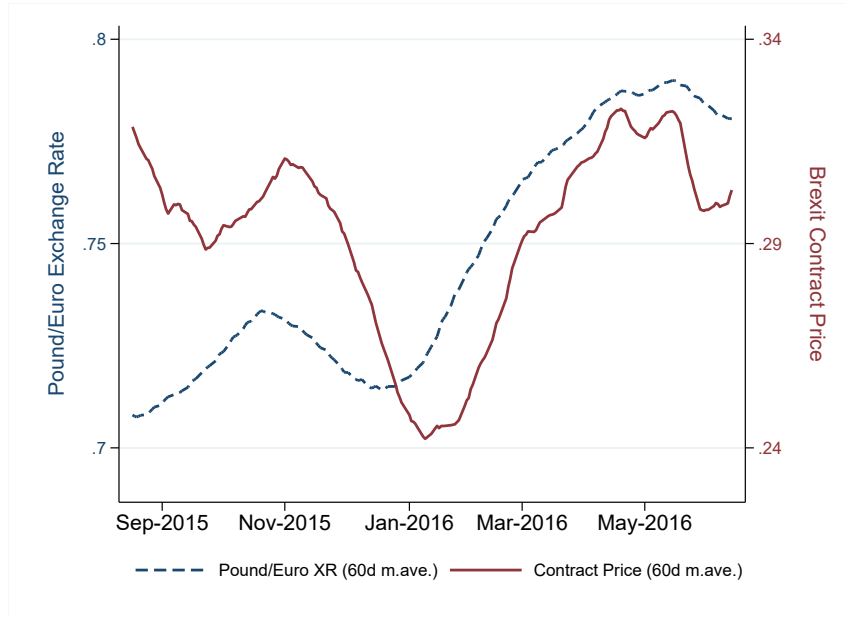
Notes: Average monthly contract price used in the baseline sample and pound/euro exchange rate.

Figure A2: MFN and Pound Depreciation Risk Distribution.
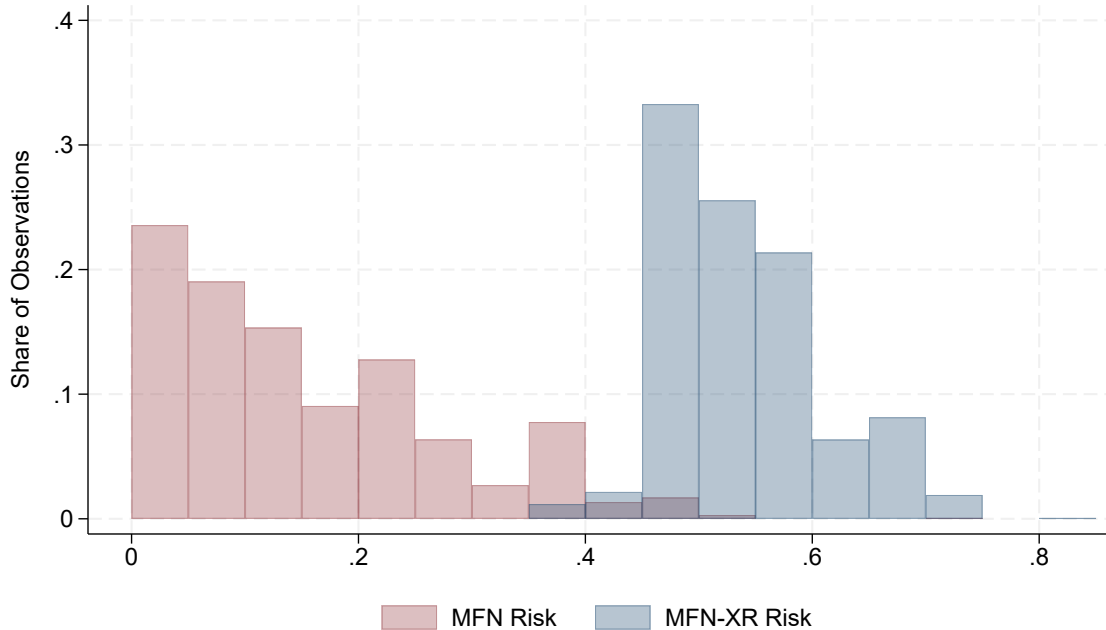
Notes: MFN risk calculated as $1 - (\tau_V^{MFN})^{-\sigma}$. MFN and pound depreciation risk, XR-MFN risk, calculated as $1 - (\tau_V^{MFN} \hat{e}_x)^{-\sigma}$, where $\hat{e} \equiv \frac{\hat{e}_{x,2017}}{e_{x,sp}}$, and $e_{x,sp}$ is the average pound to exporter x's currency exchange rate over the sample period (August 2015-June 2016.

Figure A3: Distribution of Quality Competition Measure
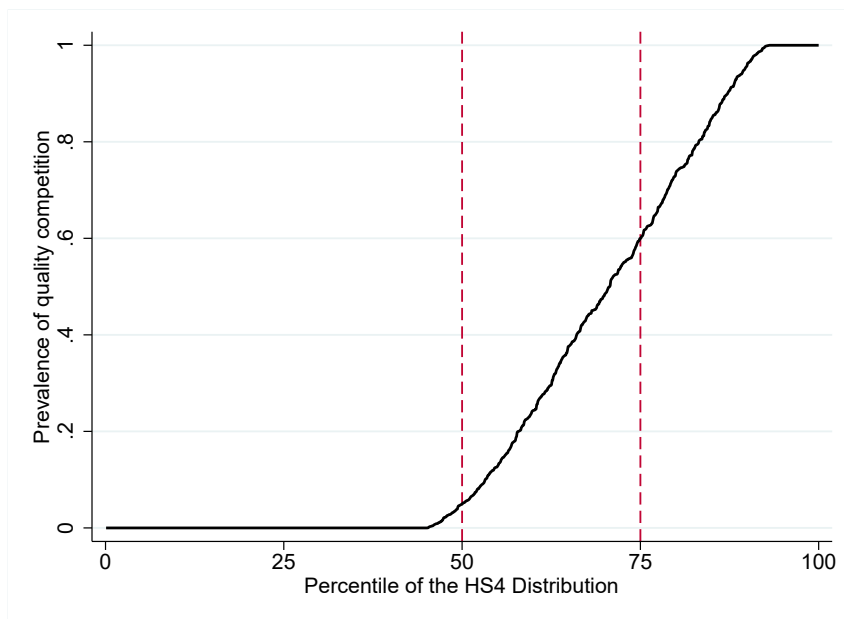
Table A1: MFN Risk Price Effects-- outlier robustness

Yearly Changes in Ideal Price Index by month (ln) 8/15-6/16

	1	2	3	4	5
Outlier definition:	No outliers	Top and bottom 1% obs.	Top and bottom 2.5% obs.	Top and bottom 5% obs.	Extreme outliers
Pr(Brexit) × MFN Risk	1.163	1.432	0.923	0.785	0.793
	(0.467)	(0.373)	(0.296)	(0.257)	(0.286)
N	75,834	74,479	72,448	69,009	71,792

Notes: Includes all industries with export sunk costs as defined in the text. The baseline sample is in column 3 (2.5%) and replicates Table 2 column 4 so all notes in that Table apply here as well. The others provide baseline estimates with different samples excluding no outliers (column 1) or alternatives based on top and bottom percentages. Extreme outliers defined as above Q3 + 3 x IQR and below Q1 - 3 x IQR, where Q1 and Q3 are the first and third quartiles and IQR is the inter-quartile range. We instrument the risk variables by their respective median HS4-specific risk across four large countries (Australia, Canada, Japan and US). All estimations include exporter-HS4 and exporter-month fixed effects.

Table A2: MFN Risk Price Effects-- alternative aggregation levels

Yearly Changes in Ideal Price Index by month (ln) 8/15-6/16

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Aggregation Level	HS2-exporter			HS4-exporter			HS6-exporter		
	Ideal Price Index	Continuing Varieties	Variety Term	Ideal Price Index	Continuing Varieties	Variety Term	Ideal Price Index	Continuing Varieties	Variety Term
Pr(Brexit) × MFN Risk	1.521	1.200	0.321	0.923	0.790	0.134	0.533	0.472	0.0614
	(0.535)	(0.473)	(0.216)	(0.296)	(0.281)	(0.0859)	(0.222)	(0.221)	(0.0292)
N	14,046	14,046	14,046	72,448	72,448	72,448	199,452	199,452	199,452
IV weak ID KP F-stat	30.77	30.77	30.77	296.1	296.1	296.1	1036	1036	1036

Notes: Dependent variable yearly changes in ideal price index and its decomposition defined at the level indicated at the column headers and month for EU exporters to the UK (2015 membership). We instrument the risk variables by their respective median HS4-specific risk across four large countries (Australia, Canada, Japan and US). Coefficients report the sum of current and two monthly lags. Robust standard errors clustered at the exporter-HS4 level below estimates. All estimations include exporter-HS4 and exporter-month fixed effects.

Table A3: MFN Risk Effect Decomposition into Import Values, Price, and Quantity

Yearly Changes by month (ln) 8/15-6/16

Dependent Variable (ln):	1 Import Value	2 Price Index	3 Quantity Index
Pr(Brexit) × MFN Risk	-0.955	0.923	-1.879
	(0.579)	(0.296)	(0.653)
N	72,448	72,448	72,448

Notes: Dependent variable yearly changes in dependent varialbe defined at the exporter-HS4-month level for EU exporters to the UK (2015 membership). Import values are aggregated by exporter-HS4-month. Quantity index is computed by subtracting the ideal price index from import values in natural logs. We instrument the risk variables by their respective median HS4-specific risk across four large countries (Australia, Canada, Japan and US). Coefficients report the sum of current and two monthly lags. Robust standard errors clustered at the exporter-HS4 level below estimates. All estimations include exporter-HS4 and exporter-month fixed effects.

Table A4: Estimated Distance Coefficients at the 8-digit Combined Nomenclature Level by Country

	Total number of 8-digit products exported	Number of estimated 8-digit coefficients	Number of 8-digit lines where distance is significant at 5%	Number with positive coefficient (quality competition)
Germany	9482	8642	5700	5497
France	9427	7804	4323	4066
Netherlands	9372	7669	3962	3530
Overlap	*9291*	*7204*	*2061*	*1897*

B On-line Quantification Appendix

The partial equilibrium price effects do not account for second-round impacts from higher overall price index as E.U. prices increase. For exporters that continue exporting to the U.K. market, even after an adverse shock to MFN tariff levels, market conditions will not deteriorate to the full extent implied by higher tariffs. Here provide the assumptions and derivations required to compute bounds on an attenuation factor for the partial effect in our baseline estimation.

To obtain bounds on this attenuation, we assume that in the event of a hard Brexit and return to MFN tariffs that the state is absorbing such that the tail risk is augmented by a new term $g_t \equiv (1-\beta)\sum_{T=0}^{\infty}\beta^T \left(\frac{P_T}{P_t}\right)^{\sigma-1} > 1$, as follows: $\omega_{Vt} = \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma} \times g_t$. This new term reflects the transition path to a higher overall price index for continuing firms after moving to higher tariffs. As such, the cutoff expression now depends on demand conditions, a tail risk term, and the g factor, i.e. $c_{xVt}^{U} = c\left(a_t, \tilde{\omega}_{xVt}, g\right)$.

As in the main text, we approximate the price index around a common stationary point: $\theta = \left\{\tilde{\tau}, c^U\right\}$,

$$\ln \frac{P\left(c_{xVt}^{U}, \tilde{\tau}_{xVt}\right)}{P\left(\theta\right)} = \Delta_0 \ln \tilde{\tau}_{xVt} + \hat{k}\left(\Delta_0 \ln c_{xVt}^{U}\right) + u_{xVt} \tag{17}$$

and then take the difference between two stationary states for the same xV to obtain

$$\Delta_t \ln P\left(c_{xVt}^{U}, \tilde{\tau}_{xVt}\right) = \Delta_t \ln \tilde{\tau}_{xVt} + \hat{k}\left(\Delta_t \ln c_{xVt}^{U}\right) + \Delta_t u_{xVt}$$

We then replace the cutoff condition with the new augmented term and simplify as follows:

$$
\begin{aligned}
\Delta_t \ln P\left(c_{xVt}^{U}, \tilde{\tau}_{xVt}\right) &= \Delta_t \ln \tilde{\tau}_{xVt} + \frac{\hat{k}}{\sigma-1}\Delta_t \ln a_{xt} + \hat{k}\tilde{\beta}\left(\tilde{\omega}_{xV}^{BR} - \tilde{\omega}_{x}^{EU}\right)\Delta_t m_t + \hat{k}\Delta_t u_{xV}^{c} + \Delta_t u_{xVt} \\
&= \alpha_{xt,xV} + \hat{k}\tilde{\beta} \times \left(\tilde{\omega}_{xV}^{BR} - \tilde{\omega}_{x}^{EU}\right)\Delta_t m_t + \tilde{u}_{xVt} \\
&= \alpha_{xt,xV} + \hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)\left[g\left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma} - 1\right]\Delta_t m_t + \tilde{u}_{xVt}
\end{aligned}
$$

The error term and the exporter-industry-time effect are given by $\tilde{u}_{xVt} \equiv \Delta_t u_{xVt} + \hat{k}\Delta_t u_{xV}^{c}$ and $\alpha_{xt,xV} \equiv \Delta_t \ln \tilde{\tau}_{xVt} + \frac{\hat{k}}{\sigma-1}\Delta_t \ln a_{xt}$. We can further simply use this equation to show that the estimated baseline coefficients include general equilibrium adjustment factor, as follows:

$$
\begin{aligned}
\Delta_t \ln P\left(c_{xVt}^{U}, \tilde{\tau}_{xVt}\right) &= \alpha_{xt,xV} + \hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)g\left[\left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma} - \frac{1}{g}\right]\Delta_t m_t + \tilde{u}_{xVt} \tag{18} \\
&= \alpha_{xt,xV} + \hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)g\left[\left(\left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma} - 1\right)\right]\Delta_t m_t \\
&\quad + \hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)\left(1 - \frac{1}{g}\right)\Delta_t m_t + \tilde{u}_{xVt} \\
&= \alpha_{xt,xV} + \hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)gm_0\left[\left(1 - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right)\Delta_t \ln m_t\right] \\
&\quad + \hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)\left(1 - \frac{1}{g}\right)m_0\Delta_t \ln m_t + \tilde{u}_{xVt} \\
&= \tilde{\alpha}_{xt,xV} + \left\{\hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)gm_0\right\}\left[\left(1 - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right)\Delta_t \ln m_t\right] + \tilde{u}_{xVt}.
\end{aligned}
$$

The first term above is $\tilde{\alpha}_{xt,xV} \equiv \alpha_{xt,xV} + \hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)\left(1 - \frac{1}{g}\right)m_0\Delta_t\ln m_t$ is absorbed by the fixed effects in our baseline estimation. After we substitute equation (10), it is straightforward to see the terms in $\{\cdot\}$ on the last line are the same as the elasticity $\mathcal{E}$ in our baseline estimation equation (16) augmented by g.

B.1 Assumption and Derivations for Quantification

We make the following assumptions:

1. The expenditure on E.U. imports is Cobb-Douglas with share $\mu = 0.07$ relative to all other U.K. expenditure.

2. The productivity distribution is Pareto with shape parameter k such that $\hat{k} = \left(1 - \frac{k}{\sigma-1}\right)$

3. The elasticity of substitution is $\sigma = 4$

B.2 Correction for GE feedback of aggregate prices into import prices.

We start with equation (18) and derive the change in $a(m)$ due to aggregate price changes

$$
\begin{aligned}
E\ln\frac{P_{xV}(m,.)}{P_{xV}(m',.)} &= -\hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)gm_0 E\left[\frac{1}{g} - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right]\ln\frac{m}{m'} + E\left(\alpha_x(m) - \alpha_x(m')\right) \\
&= -\hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)gm_0 E\left[\frac{1}{g} - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right]\ln\frac{m}{m'} + \hat{k}E\left[\ln\frac{P(m)}{P(m')}\right] \\
&= -\hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)gm_0 E\left[\frac{1}{g} - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right]\ln\frac{m}{m'} + \hat{k}\mu E\left[\sum_{xV} w_{xV}\ln\frac{P_{xV}(m,.)}{P_{xV}(m',.)}\right]
\end{aligned}
$$
(19)

where in second line we use $E\left(\alpha_x(m) - \alpha_x(m')\right) = \frac{\hat{k}}{\sigma-1}E\left[\ln\frac{a(m)}{a(m')}\right] = \hat{k}E\left[\ln\frac{P(m)}{P(m')}\right]$. We keep all other prices constant, therefore in the third line we employ $E\ln\frac{P(m)}{P(m')} = \mu E\left[\sum_{xV} w_{xV}\ln\frac{P_{xV}(m,.)}{P_{xV}(m',.)}\right]$. This assumes a CD aggregator between E.U. imports and all other U.K. expenditure with share μ. We also use the Sato-Vartia weights w_{xV} described in appendix A.2 to aggregate over exporter-industry price indices.

We now need an expression for $\sum_{xV} w_{xV}\ln\frac{P_{xV}(m,.)}{P_{xV}(m',.)}$, which we obtain by aggregating the left and right sides equation (18) using the weights w_{xV}.

$$
\sum_{xV} w_{xV}\ln\frac{P_{xV}(m,.)}{P_{xV}(m',.)} = -\hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)gm_0 E\left[\frac{1}{g} - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right]\ln\frac{m}{m'} + \hat{k}\mu\sum_{xV} w_{xV}\ln\frac{P_{xV}(m,.)}{P_{xV}(m',.)} + \sum_{xV} w_{xV}\tilde{u}_{xVt}
$$

$$
\left(\sum_{xV} w_{xV}\ln\frac{P_{xV}(m,.)}{P_{xV}(m',.)}\right)\left(1 - \hat{k}\mu\right) = -\hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)gm_0 E\left[\frac{1}{g} - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right]\ln\frac{m}{m'} + \sum_{xV} w_{xV}\tilde{u}_{xVt}
$$

$$
\sum_{xV} w_{xV}\ln\frac{P_{xV}(m,.)}{P_{xV}(m',.)} = -\frac{1}{\left(1 - \hat{k}\mu\right)}\hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)gm_0 \sum_{xV} w_{xV}\left[\frac{1}{g} - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right]\ln\frac{m}{m'} + \frac{\sum_{xV} w_{xV}\tilde{u}_{xVt}}{1 - \hat{k}\mu}
$$

$$
\hat{k}\mu\sum_{xV} w_{xV}\ln\frac{P_{xV}(m,.)}{P_{xV}(m',.)} = -\frac{\hat{k}\mu}{\left(1 - \hat{k}\mu\right)}\hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right)gm_0 \sum_{xV} w_{xV}\left[\frac{1}{g} - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right]\ln\frac{m}{m'} + \frac{\hat{k}\mu}{\left(1 - \hat{k}\mu\right)}\sum_{xV} w_{xV}\tilde{u}_{xVt}
$$

We replace this expression in (19) and simplify

$$E \ln \frac{P_{xV}(m,.)}{P_{xV}(m',.)} = -\hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right) gm_0 E \left[\frac{1}{g} - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right] \ln \frac{m}{m'}$$

$$- \frac{\hat{k}\mu}{\left(1 - \hat{k}\mu\right)} \hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right) gm_0 \sum_{xV} w_{xV} \left[\frac{1}{g} - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right] \ln \frac{m}{m'} + \frac{\hat{k}\mu}{\left(1 - \hat{k}\mu\right)} \sum_{xV} w_{xV} \tilde{u}_{xVt}$$

$$= -\left(1 + \frac{\hat{k}\mu}{\left(1 - \hat{k}\mu\right)}\right) \hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right) gm_0 E \left[\frac{1}{g} - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right] \ln \frac{m}{m'}$$

$$= -\left(\frac{1}{\left(1 - \hat{k}\mu\right)}\right) \hat{k}\tilde{\beta} \times \left(\eta^{BR} - \eta^{EU}\right) gm_0 E \left[\frac{1}{g} - \left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right] \ln \frac{m}{m'}. \tag{20}$$

In the second line we assume the weighted average of the error term is zero, $\frac{\hat{k}\mu}{(1-\hat{k}\mu)} \sum_{xV} w_{xV} \tilde{u}_{xVt} = 0$, and collect terms. The third line simplifies the leading term in $(\cdot)$ to obtain a correction factor $\frac{1}{(1-\hat{k}\mu)}$ in GE, which we will bound below.

B.3 Deriving bounds for the adjustment factor g

We now show that under reasonable assumptions we can also provide bounds on the adjustment factor g.

We start by considering how large the steady-state price increase would be if tariffs would increase to MFN rates with certainty. This requires a parametric assumption that the productivity distribution is Pareto with shape parameter k. We then obtain an expression for the upper bound on g of

$$\max g \mid_{\omega^{EU}=1} < \max \left(\frac{P\left(\tau^{mfn}\right)}{P\left(\tau^{EU}\right)}\right)^{\sigma-1} = \left(\exp\left(\ln \frac{P\left(\tau^{mfn}\right)}{P\left(\tau^{EU}\right)}\right)\right)^{\sigma-1} = \left(\exp\left(\mu \frac{\left(\frac{\sigma k}{\sigma-1} - 1\right)\frac{1}{\sigma-1}}{1 - \mu\left(1 - \frac{k}{\sigma-1}\right)} \ln \tau^{mfn}\right)\right)^{\sigma-1}. \tag{21}$$

This expression for the upper bound is obtained through the following steps.

(1) We use the definition of the price index in (6) for each exporter(x)-industry(V) in log differences, which depends on the aggregate price index through the cost cutoff. This yields:

$$\Delta \ln P_{V,x} = \Delta \ln (\tau_V) + \Delta \ln (D_t) + \left(1 - \frac{k}{\sigma-1}\right)\left[\frac{1}{\sigma-1}\Delta \ln (a_V) - \ln U_V\right]$$

$$= \left(1 - \frac{k}{\sigma-1}\right)(-\ln U_V) + \left(\frac{\sigma k}{\sigma-1} - 1\right)\frac{1}{\sigma-1}\Delta \ln \tau_V + \left(1 - \frac{k}{\sigma-1}\right)\Delta \ln \left(P\left(\varepsilon Y_t\right)^{\frac{1}{\sigma-1}}\right) \tag{22}$$

(2) We can solve for $\Delta \ln P_{V,x}$ when the aggregate price change is $\Delta \ln P = \mu \mathbb{E}_{xV}\left[\ln \frac{P_{xV}(m,.)}{P_{xV}(m',.)}\right]$ around an initial E.U. steady state $(\omega^{EU} = 1)$ to obtain

$$\mathbb{E}_{xV}\left[\Delta \ln P_{V,x}\right] = \frac{\left(\frac{\sigma k}{\sigma-1} - 1\right)\frac{1}{\sigma-1}}{1 - \mu\left(1 - \frac{k}{\sigma-1}\right)} \Delta \ln \tau_V$$

Using $\Delta \ln P = \mu \mathbb{E}_{xV}\left[\ln \frac{P_{xV}(m,.)}{P_{xV}(m',.)}\right]$, the aggregate price elasticity with respect to a tariff change is $\left[\mu \frac{\left(\frac{\sigma k}{\sigma-1} - 1\right)\frac{1}{\sigma-1}}{1 - \mu\left(1 - \frac{k}{\sigma-1}\right)}\right]$.

3

B.4 Bounds for the aggregate import price index change

We combine this aggregate price elasticity with the expression for the general equilibrium price adjustment in (20)

$$
\begin{aligned}
E\left[\ln \frac{P_{xV}(m,.)}{P_{xV}(m',.)}\right] &= \frac{1}{\left(1-\hat{k}\mu\right)}\left(-\hat{k}\tilde{\beta}\times g^{BR}m_0\right)E\left(\frac{1}{g}-\left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right)\ln\frac{m}{m'} \\
&= \left[\frac{1}{\left(1-\hat{k}\mu\right)}\right]\sum_l \mathcal{E}_{t-l}E\left(\left[\frac{1}{g}-1\right]+\mathbb{E}_{xV}\left[1-\left(\frac{\tau_{xV}^{MFN}}{\tau_{xV}^{EU}}\right)^{-\sigma}\right]\right)\Delta mbv
\end{aligned}
$$

As we describe in section 4, we use the estimated cross-elasticity of 0.923, the weighted change in the tail risk factor 0.148, and the full change in Δmbv with 33% Bregret to compute a partial effect of 11.2 lp. We can now adjust that impact by replacing those numbers in the expression above to obtain

$$
E\left[\ln\frac{P_{xV}(m,.)}{P_{xV}(m',.)}\right] = \left[\frac{1}{\left(1-\hat{k}\mu\right)}\right]\cdot 0.923\cdot E\left(\left[\frac{1}{g}-1\right]+0.148\right)(1.23\cdot 0.67) \tag{23}
$$

The lower bounds on the price index change can be obtained for reasonable ranges of the Pareto shape parameter $k\in(3,9)$, using E.U. import share of $\mu=0.07$, $\sigma=4$, and the bound for g from equation (21). The attenuation due to the GE correction factor $\frac{1}{(1-\hat{k}\mu)}$ is at most 0.88 for $k=9$ and only 0.94 for $k=6$. Likewise, the upper bound on g is 1.03 for $k=9$ and 1.01 to 1.02 for $k=4$ to 6. When these upper bounds are combined, we obtain lower bounds on the aggregate E.U. import price index in equation (23) increase of 7.9 lp for $k=9$ or 9.1 lp for $k=6$. We can then obtain an aggregate effect by multiplying this lower bound by the E.U. import share in total expenditure. We focus on $k=6$, which implies a lower bound on aggregate prices of 0.67 lp ($=0.07\times 9.1$)

We can provide more detail by considering IO tables that breakout the share of imports in total expenditure by households and firms on manufactured goods. The simple approach above assumes no input-output multiplier. This implies that any import price increases for firms accrue either to the owners, which have the same marginal utility of income as consumers, or are fully passed on to consumers. Using the 2014 U.K. national IO tables from the World Input-Output Data (Timmer et al., 2015), we compute that imports are 46% of expenditure of manufactured goods, of which 28% is firm expenditure and 17% is from households. E.U. imports are 52% of total U.K. imports in 2015, which implies the increase in the manufacturing price index is 2.2 log points ($=9.1\times 0.52\times 0.46$). Because manufactured goods are 24% of total firm and household expenditure, the effect on aggregate prices is 0.53 log points. As such, the effect is nearly the same as the simple overall E.U. import expenditure share approach above.