THE IMPACT OF TRADE AGREEMENTS ON CONSUMER WELFARE - EVIDENCE FROM THE EUROPEAN UNION'S COMMON EXTERNAL TRADE POLICY

Abstract. This paper estimates the consumer welfare impact of the new generation of trade agreements implemented by the European Union between 1993 and 2013. We decompose the overall effect into contributions of changes in prices, quality and variety. Estimating trade elasticities for narrow product categories of EU imports, we infer quality from data on imported values and volumes. For the EU as a whole, we find that trade agreements increased quality by 7% on average but did not affect prices or variety. This translates into a cumulative reduction in the consumer price index of 0.24% over our sample period. We also find a high degree of impact heterogeneity across EU countries, trading partners and the type of trade agreement, with high-income EU countries seeing much stronger quality increases and larger overall consumer benefits.

JEL Codes: F1, F5, F6, L6.

Keywords: Trade agreement, Quality, FTA, Elasticity, Variety, Unit Values.

Giuseppe Berlingieri — ESSEC & CEP

Holger Breinlich — Nottingham, CEP & CEPR

Swati Dhingra — LSE, CEP & CEPR

This Version: August 3, 2017

Acknowledgments. We thank Claudio Michelacci, Gianmarco Ottaviano, participants at BIS, Hertfordshire, HMT, LSE, Manchester and Nottingham, and four anonymous referees for helpful comments. This paper was formely titled "The Impact of Trade Agreements on Consumer Welfare".

Contacts: giuseppe.berlingieri@essec.edu; holger.breinlich@nottingham.ac.uk; s.dhingra@lse.ac.uk.

1. Introduction

A central tenet of trade theory is that lowering trade barriers increases welfare. Trade agreements between countries lower trade barriers on imported goods and provide welfare gains to consumers from increases in variety, access to better-quality products and lower prices. While a large literature estimates the overall gains from trade, less is known about the effects of specific trade agreements and the channels through which they increase welfare. We estimate the impact of trade agreements on consumer welfare through the channels of increased variety, better quality and lower prices, using the new generation of agreements implemented by the European Union (EU12) between 1993 and 2013.[1]

The paper makes two main contributions. The first contribution is to decompose the welfare impact of the EU's trade agreements on EU consumers into gains from new varieties, better quality and lower prices to explain which channel dominates the realization of gains from trade agreements. This is important because quality is not directly observable in standard trade data, and unit values (obtained from dividing revenues by quantities) confound increases in true prices and quality. If quality increases after the implementation of a trade agreement, interpreting unit values as true prices would imply an increase in the import price index. If unit values are instead interpreted as pure quality then we would over-estimate the impact of trade agreements on quality. Trade agreements are likely to have both price and quality effects, so that their welfare impact would be under-estimated when quality rises and prices fall.[2]

To disentangle these different welfare effects, we build on recent work that measures quality as a residual from data on import quantities and prices (e.g., Khandelwal et al. 2013). Building on this literature, we make the standard assumption that consumers have constant elasticities of substitution across differentiated varieties of products. This restricts the pro-competitive effects of trade, but enables us to use bilateral trade data to infer quality based on these elasticities, which we estimate extending Feenstra (1994) and Broda and Weinstein (2006) to include an improved treatment of measurement error in observed prices.

Having inferred quality, we estimate the impact of trade agreements on prices, quality and variety using a difference-in-difference strategy that exploits the timing and geographic variation in the EU's trade policy. We find that for the EU as a whole access to better-quality products is the primary contributor to welfare gains from trade agreements. Joining a free trade agreement

[1]Here and in the following, we use the term "European Union" (or EU12 or simply EU) to refer to the twelve member states prior to the 1995 enlargement (Belgium, Luxembourg, Germany, France, Italy, the Netherlands, the United Kingdom, Ireland, Denmark, Greece, Portugal and Spain). This geographic focus was chosen to keep the set of countries for our analysis constant over the sample period.

[2]While price and quality changes are substantively similar on the demand side of standard trade models, the decomposition of welfare gains into prices, quality and variety is of interest for supply reasons because firms might differ in their ability to reduce prices, upgrade quality and expand variety (Dhingra 2013).

increases the quality of products from the trade partner by 7% on average. This finding is robust to controlling for other determinants of quality such as per-capita income and we find no evidence of a potential endogeneity of the EU's trade agreements with respect to pre-trends in prices or quality.

Our overall results hide a substantial amount of treatment effect heterogeneity across EU countries, trading partners and types of trade agreements. For example, high-income EU countries (the United Kingdom, the Netherlands, Ireland and Belgium/Luxembourg) saw much stronger increases in quality than other EU countries. Indeed, for the group of low-income EU countries (Greece, Spain and Portugal), the impact of trade agreements worked almost exclusively through a reduction in prices rather than increased quality.

The second contribution of the paper is to show by how much changes in the import price index affect the consumer price index (CPI). For the EU12 as a whole, the quality-enhancing effect of trade agreements translates into a cumulative 0.24% CPI reduction between 1993-2013. Around 55% of this change results from the direct effect of higher-quality imports of final goods. The remaining 45% is due to reductions in the quality-adjusted price of imports of intermediate inputs which lowers the prices of domestic goods. While our focus is on the gains arising from imported goods throughout, we also provide tentative evidence that these gains were not partially offset by a reduction in the number of domestic varieties. Looking across EU countries, we find that CPI reductions are strongest for high-income countries (0.41%) and only 0.12%-0.13% for the remaining EU12 member states.

Our paper is related to the growing literature that estimates the gains from openness (Broda and Weinstein 2006; Feenstra and Weinstein 2010; Feenstra and Romalis 2012; Hallak and Schott 2011; Amiti and Khandelwal 2013; Demir 2011). In their classic paper, Broda and Weinstein (2006) estimate the gains to the US through the channel of import variety. Blonigen and Soderbery (2010) address the welfare gains from variety in the auto industry which has detailed data on varieties. Feenstra and Weinstein (2010) examine both variety and lower prices from import competition. Feenstra and Romalis (2012) and Hallak and Schott (2011) make methodological contributions to this literature and provide quality inferences. Hummels and Klenow (2005) examine the extent to which differences in trade volumes are explained by differences in quality across countries. They find that within product categories, richer countries export more units at higher prices to a given market, consistent with producing higher quality. Baldwin and Harrigan (2011) show that export zeros and export unit values are increasing in distance and decreasing in importer size, and that this fact can be explained by a quality-augmented version of a workhorse trade model. Johnson, 2012 estimates a heterogeneous firms trade model and finds that prices are increasing in the difficulty of entering a given destination market in most sectors.

The focus of these studies is on gains from openness and on how variety, quality or prices vary by trade partner characteristics such as the physical distance between them. We instead examine the changes in variety, quality and prices resulting from trade agreements. Focusing on trade agreements has several practical advantages. First, trade agreements are under the direct control of policy makers, unlike more general measures of openness such as trade-to-GDP ratios or simply changes over time. Understanding the effects of such agreements is thus more relevant from a policy perspective than understanding the gains from openness more generally. Second, bilateral trade agreements have become the preferred form of trade liberalization by the EU and other developed countries over the past decades and understanding their effects on consumers is of first-order importance for the design of trade policy. Finally, we can regress price, quality and variety measures on time- and origin-varying trade agreement variables, allowing us to disentangle the impact of these agreements from other determinants using standard difference-in-difference techniques.

Our paper also contributes to the literature on the effects of trade agreements (e.g., Baier and Bergstrand, 2007, 2009; Egger et al., 2008; Head and Mayer, 2014). These papers show that such agreements usually lead to a strong expansion in trade flows between partners. We adapt empirical techniques from this literature to identify the impact of trade agreements on prices, quality and variety. This enables us to compute implied consumer price index changes which is arguably of more direct relevance to overall welfare gains than the underlying trade flow changes.[3]

Finally, our results also mirror the research by Sheu (2014) who uses rich product-level data on the Indian computer printer industry and finds that quality was the primary source of welfare gains during the trade liberalization period of 1996-2005. We look at the entire range of tradable products which enables us to compute implied counterfactual changes in the economy-wide consumer price index. We also link price index changes more closely to specific trade policy changes using difference-in-differences regression techniques.

The remainder of the paper is organized as follows. Section 2 provides a model of consumer welfare that enables quality inferences and a decomposition of the consumer price index into

[3]A few studies have also attempted decompositions of the total effect of trade agreements on trade. For example, following Hummels and Klenow (2005), Baier et al. (2014) decompose trade flows into an extensive margin of variety (the weighted count of products of all products exported to a particular destination country) and an intensive margin of revenues from continuing varieties (the market share of these products in the total imports of the destination country). They find that both margins of trade are affected by trade agreements, and that the intensive margin elasticity of trade flows to trade agreements membership is larger than the extensive margin elasticity.

prices, quality and variety. Section 3 provides an empirical model to infer quality and estimate the impact of trade agreements on each channel. Section 4 discusses the data and its sources. Section 5 presents our empirical results and conducts a number of robustness checks. We also present the decomposition of the import and consumer price index and derive price index changes due to FTAs for the EU for the period 1993-2013. Section 6 concludes.

2. A Model of Consumer Welfare

In this section, we build on Broda and Weinstein (2006) to specify consumer demand. We derive the aggregate price index and the decomposition of price indices into prices, quality and variety. We use a partial equilibrium framework and initially focus on the direct effect of trade agreements on final good imports. The advantage of this approach is that it requires relatively few assumptions, apart from the specification of a demand system and the usual identifying assumptions of difference-in-difference estimates. However, in the second part of this section we also look at indirect supply-side impacts working through the cost-reducing effect of cheaper and better intermediate imports.

2.1. **Consumer Demand.** The representative consumer's welfare is defined over goods from S sectors and takes the following CES form:[4]

$$U = \left(\sum_{s=1}^{S} U_s^{\frac{\beta-1}{\beta}}\right)^{\frac{\beta}{\beta-1}}.$$

In our data, sectors are defined according to a version of the ISIC classification (13 sectors in total, see Section 5.4). Welfare in each sector is derived from composite imported and domestic goods as follows:

$$U_s = \left(M_s^{(\gamma_s-1)/\gamma_s} + D_s^{(\gamma_s-1)/\gamma_s}\right)^{\frac{\gamma_s}{\gamma_s-1}}. \tag{1}$$

We are interested in examining how consumer welfare changes over time with trade policy. The real income of a representative consumer at time t is $U_t = Y_t/P_t$ where Y_t is the representative consumer's total income and P_t is the price index associated with utility function U_t. With constant consumer income, the welfare change is equal to the economy-wide change in this price index:

$$P_t/P_{t-1} = \Pi_{s=1}^{S}\left(P_{s,t}/P_{s,t-1}\right)^{\omega_{st}}, \tag{2}$$

[4]Our choice of preference structure is driven by data availability and comparability to the existing literature. Since we have data on consumer expenditure and import shares, a nested CES structure is the most flexible specification possible. In an earlier version of the paper, we used a more restrictive Cobb-Douglas structure which yielded similar results.

where the ω_{st} are the Sato-Vartia weights associated with the CES price index.[5] The change in the sectoral price indices ($P_{s,t}$) in turn is given by:

$$\frac{P_{s,t}}{P_{s,t-1}} = \left(\frac{P_{sM,t}}{P_{sM,t-1}}\right)^{\omega_{sMt}} \left(\frac{P_{sD,t}}{P_{sD,t-1}}\right)^{\omega_{sDt}}, \tag{3}$$

where $P_{sM,t}$ and $P_{sD,t}$ are the prices of imported and domestic goods, and ω_{sMt} and ω_{sDt} the Sato-Vartia weights associated with (1).

Trade policy changes affect sectoral price indices through two channels. Directly, by changing the price, quality or variety of imported goods; and indirectly by affecting the production cost of domestically produced goods via changes in imported intermediate inputs. We discuss both channels in turn.

2.2. **Import Price Index Change.** The imported composite good M_s is a CES aggregate of products within sector s,

$$M_s = \left(\sum\nolimits_{h \in H_s} m_h^{\frac{\delta_s - 1}{\delta_s}}\right)^{\frac{\delta_s}{\delta_s - 1}},$$

where a product h is defined as a unique 6-digit code of the Harmonized System (HS) product classification and H_s is the set of HS products associated with sector s. HS 6-digit is the most disaggregated level at which data on import values and quantities are available in our data. Each product h is made up of varieties (indexed by z) shipped from different origin countries (indexed by o):

$$m_h^{(\sigma_h - 1)/\sigma_h} = \sum\nolimits_{o=1}^{O} \sum\nolimits_{z=1}^{n_h^o} \left(q_h^o(z)\, x_h^o(z)\right)^{(\sigma_h - 1)/\sigma_h}, \quad \sigma_h > 1,$$

where $q_h^o(z)$ denotes the quality of product h's variety z imported from origin o, $x_h^o(z)$ are the units of quantity consumed of that variety, and σ_h is the elasticity of substitution between varieties. The number of varieties imported from a given origin country is denoted by n_h^o. Let $p_h^o(z)$ denote the price of variety z of imported HS 6-digit product h. Then the price index for the imported composite product m_h is:

$$p_h = \left[\sum\nolimits_{o=1}^{O} \sum\nolimits_{z=1}^{n_h^o} \left(p_h^o(z) / q_h^o(z)\right)^{1-\sigma_h}\right]^{\frac{1}{1-\sigma_h}},$$

and the change in the import price index for sector s is

$$P_{sM,t}/P_{sM,t-1} = \Pi_{h \in H_s} \left(p_{h,t}/p_{h,t-1}\right)^{\omega_{ht}}, \tag{4}$$

where ω_{ht} are the Sato-Vartia weights associated with the HS products of sector s.

[5]That is, $\omega_{st} = \frac{s_{s,t} - s_{s,t-1}}{\ln s_{s,t} - \ln s_{s,t-1}} / \left(\sum_{i \in S} \frac{s_{i,t} - s_{i,t-1}}{\ln s_{i,t} - \ln s_{i,t-1}}\right)$, with $s_{s,t}$ being sector s's share in total expenditure. See Sato (1976) for details.

Because we have defined products h at the most disaggregated level of data available, we do not observe prices and quantities of individual varieties imported from country o. To make progress, we make the standard assumption that varieties are identical within a product-origin combination:

$$p_h = \left[\sum\nolimits_{o=1}^{O} n_h^o \left(p_h^o/q_h^o\right)^{1-\sigma_h}\right]^{\frac{1}{1-\sigma_h}}, \tag{5}$$

where n_h^o is the number of "hidden varieties" of product h imported from country o.

2.3. **Decomposition of the Import Price Index.** We would like to know the channels – prices, quality or variety – through which trade agreements change the import price index. To determine the sources of welfare gains, we use equation (5) to obtain an exact price index. Diewert (1976) defines an exact price index over a constant set of varieties as the ratio of price indices across time periods t and $t-1$. To use results from the price index literature, we define the quality- and hidden-variety adjusted price of the HS6 variety from origin o as $p_h^{o,adj} \equiv p_h^o/\left((n_h^o)^{1/(\sigma_h-1)}\, q_h^o\right)$. For a constant set of origin countries, the exact price index for product h is given by (see Feenstra, 1994, p.158):

$$p_{h,t}/p_{h,t-1} = \Pi_{o=1}^{O}\left(p_{h,t}^{o,adj}/p_{h,t-1}^{o,adj}\right)^{\omega_{ht}^o},$$

where ω_{ht}^o are the Sato-Vartia weights associated with each variety. If the set of varieties (i.e., origin countries) changes between periods, Feenstra (1994) shows that the exact price index can be rewritten as:

$$p_{h,t}/p_{h,t-1} = \left(\lambda_{h,t}/\lambda_{h,t-1}\right)^{1/(\sigma_h-1)}\,\Pi_{o\in I}\left(p_{h,t}^{o,adj}/p_{h,t-1}^{o,adj}\right)^{\omega_{ht}^o}, \tag{6}$$

where

$$\lambda_{hr} = \left(\sum\nolimits_{o\in I} p_{h,r}^{o,adj} x_{h,r}^o\right) / \left(\sum\nolimits_{o'\in I_r} p_{h,r}^{o',adj} x_{h,r}^{o'}\right) \quad r = t-1, t \tag{7}$$

and I denotes the set of varieties present in both periods, t and $t-1$ (and I_r is the set of varieties present in period r). That is, λ_{ht} is the share of continuing varieties in all varieties in period t, and $1-\lambda_{ht}$ is the share of new varieties. Similarly, $1-\lambda_{ht-1}$ is the share of varieties exiting between $t-1$ and t (expressed as a share of varieties available in $t-1$).

Substituting for $p_h^{o,adj} = p_h^o/\left((n_h^o)^{1/(\sigma_h-1)}\, q_h^o\right)$ in equation (6), the exact price index can be decomposed into its components of variety, prices, quality, and hidden variety as follows:

$$(8) \qquad \frac{p_{h,t}}{p_{h,t-1}} = \underbrace{\left(\frac{\lambda_{ht}}{\lambda_{ht-1}}\right)^{\frac{1}{\sigma_h - 1}}}_{\text{Variety}} \times \underbrace{\Pi_{o\in I}\left(\frac{n^o_{ht}}{n^o_{ht-1}}\right)^{-\frac{\omega^o_{ht}}{\sigma_h - 1}}}_{\text{Hidden Variety}} \times \underbrace{\Pi_{o\in I}\left(\frac{p^o_{ht}}{p^o_{ht-1}}\right)^{\omega^o_{ht}}}_{\text{Prices}} \times \underbrace{\Pi_{o\in I}\left(\frac{q^o_{ht}}{q^o_{ht-1}}\right)^{-\omega^o_{ht}}}_{\text{Quality}}.$$

Equation (8) enables us to decompose the changes in the price indices arising from each of the sources of variety, prices and quality.[6]

Ideally, we would need data on prices, quality and the number of varieties within each HS 6-digit product (n^o_h), as well as on the degree of substitutability across varieties σ_h to implement this price index decomposition. Unfortunately, this is infeasible with standard trade data due to three empirical challenges.

First, as discussed the number of hidden varieties n^o_h is unobservable. While we observe the total value and volume of trade for each product-origin, we do not know how many different varieties n^o_h are imported within each product-origin. To overcome this problem, we will use proxies from the literature to control for hidden varieties in Section 3.

Second, the elasticities of substitution σ_h determine the importance of different varieties in the price index but are not directly observed. We will estimate the elasticities for each product using the structure of the CES demand and invoking assumptions on the supply side.

Finally, quality data is notoriously inadequate for welfare analysis because physical measures of quality are typically only available for a narrowly defined range of products that constitute a small share of consumer budgets. We will therefore use the CES demand structure to make quality inferences for all imported products. We will then estimate the impact of trade agreements on the inferred quality, prices and variety estimates to determine the sources of gains from trade agreements.

Once we have computed price index changes at the HS 6-digit product level, we can aggregate up to the sector-level import price index change ($P_{sM,t}/P_{sM,t-1}$) using Sato-Vartia shares computed from trade data.

2.4. **Domestic Price Index Change.** The sectoral price index change (3) is also determined by changes in the domestic price index for sector s, $P_{sD,t}/P_{sD,t-1}$. By analogy with the import side, we define a CES aggregator over domestic varieties, with the following associated price index:

[6] Note that the variety term captures the overall welfare contribution of new and exiting varieties, including their prices and quality. For example, if the total number of imported varieties remains constant but exiting varieties are replaced by higher-quality entering ones, continuing varieties lose market share and the variety term is smaller than unity. We thank an anonymous referee for pointing this out.

$$P_{sD,t} = \left(\sum_{z=1}^{Z_s} (p_s(z)/q_s(z))^{1-\delta_D} \right)^{\frac{1}{1-\delta_D}}.$$

Here, $q_s(z)$ and $p_s(z)$ denote the quality and price of domestic variety z within sector s, respectively, and δ_D is the elasticity of substitution between domestic varieties. We do not have information about the prices or quality of individual domestic firms, nor would it be possible to convincingly link these to individual trade agreements.[7] To make progress, and by analogy to our assumption about producers of foreign varieties, we assume that domestic firms are identical. We further assume that product quality and the number of producers is constant over time.[8] With these assumptions, the domestic component of the sectoral price index change is simply the price change of the representative domestic firm:

$$\frac{P_{sD,t}}{P_{sD,t-1}} = \prod_{z=1}^{Z_s} \left(\frac{p_{s,t}(z)/q_{s,t}(z)}{p_{s,t-1}(z)/q_{s,t-1}(z)} \right)^{\omega_z} = \frac{p_{sD,t}}{p_{sD,t-1}}.$$

The next step is to specify a production function for domestic firms. Following Blaum et al. (2015) we assume that a domestic firm z in sector s uses the production technology $y_{sz} = \phi_z l(z)^{\alpha_s} i(z)^{1-\alpha_s}$, where ϕ is total factor productivity (TFP), l is a primary factor (e.g., labor) and i are intermediates. With perfect pass-through from production costs to prices, this technology implies price changes of:[9]

$$\frac{p_{sD,t}}{p_{sD,t-1}} = \frac{\phi_{s,t}}{\phi_{s,t-1}} \left(\frac{w_t}{w_{t-1}} \right)^{\alpha_s} \left(\frac{P_{sI,t}}{P_{sI,t-1}} \right)^{1-\alpha_s}.$$

Thus, trade agreements can impact domestic prices through changes in TFP (ϕ), factor prices (w) and the cost of intermediate inputs (P_I). Given the difficulty of convincingly identifying productivity and factor price effects, we focus on intermediate inputs in the following.[10] By

[7]The consumer price data underlying the construction of official CPI data is not well-suited for our purposes for at least two reasons. First, it is usually less disaggregated than import price data derived from trade statistics. For example, UK CPI data are collected for around 800 items of which only part are for manufactured goods. By contrast, the HS system has information for over 4,000 products. Second, none of the available data distinguishes between domestically produced and imported products, let alone between products imported from different origin countries. Taken together, these limitations imply that it will be difficult to detect the impact of the individual trade agreements in our sample, most of which only accounted for a small share of total EU imports (and an even smaller share of total consumption). Furthermore, our identification strategy will rely on comparing prices of imports from different origins with and without trade agreements with the EU, which would not be possible with CPI data (or other sources such as barcode scanner data).

[8]We provide evidence for the latter assumption in Section 5.2.

[9]We discuss the implications of the constant markups assumption in more detail in Section 5.2.

[10]Again, the key identification challenge is that adequately capturing the effects of modern trade agreements requires the use of dummy variables which do not have a sectoral dimension (see Section 3.1). Hence, the usual approach in the literature to use cross-sectional variation in tariff reductions is ruled out here. In addition, the free trade agreements we consider account each for only a small share of EU12 imports and an even smaller share of domestic absorption, making the identification of effects on productivity and factor prices all but impossible. By contrast, our difference-in-difference identification strategy will rely on the much finer cross-origin variation in the trade data.

analogy to our preference structure for final goods, we assume that intermediates can be sourced domestically (i_D) or imported (i_M) according to:

$$i_s = \left((i_{sD})^{\frac{\varepsilon_s - 1}{\varepsilon_s}} + (i_{sM})^{\frac{\varepsilon_s - 1}{\varepsilon_s}} \right)^{\frac{\varepsilon_s}{\varepsilon_s - 1}}.$$

As in Blaum et al. (2015) and Caliendo and Parro (2015), we assume a roundabout production structure where firms use a sector-specific domestic input that is produced using the output of all other firms in the economy:

$$i_{sD} = \prod_{j=1}^{S} Y_{js}^{\varphi_{js}} \text{ and } Y_{js} = \left(\sum_{z=1}^{Z_j} y_{jzs}^{\frac{\varepsilon_j - 1}{\varepsilon_j}} \right)^{\frac{\varepsilon_j}{\varepsilon_j - 1}} = (Z_j)^{\frac{\varepsilon_j}{\varepsilon_j - 1}} y_{jzs}, \tag{9}$$

where φ_{js} is the value share of inputs that sector s sources from sector j, and where we have again assumed identical firms in each sector. Similar to imported final goods, the imported intermediate input bundle has two layers. At the upper level, it consists of several HS 6-digit products (h):

$$i_{sM} = \prod\nolimits_{h \in H_s} i_{hM}^{\varphi_{sh}}.$$

Second, we have different origin countries (o) within each HS6-digit code:

$$i_{hM} = \left(\sum\nolimits_{o=1}^{O} \sum\nolimits_{z=1}^{n_{hM}^o} \left(q_{hM}^o(z)\, i_{hM}^o(z) \right)^{(\varepsilon_{hM} - 1)/\varepsilon_{hM}} \right)^{\frac{\varepsilon_{hM}}{\varepsilon_{hM} - 1}},$$

where n_h^o is the number of imported varieties within each destination. Note that this structure is an exact analogue to imports of final good with the exception that the elasticities ε_{hM} now measure production substitutability within HS codes. As we show in Appendix A, this means that the exact price index of the product-h component of the imported intermediate bundle is again given by an expression analoguous to (8). Hence, we will be able to use the same approach to estimate the impact of trade agreements on the inferred quality, prices and variety of both final and intermediate goods.

The corresponding price index change for the imported intermediate input bundle is obtained by aggregating across the HS products used as inputs in sector s:

$$\frac{P_{sIM,t}}{P_{sIM,t-1}} = \prod\nolimits_{h \in H_s} \left(\frac{p_{hM,t}}{p_{hM,t-1}} \right)^{\varphi_{sh}},$$

Thus, the price change of intermediates is:

$$\frac{P_{sI,t}}{P_{sI,t-1}} = \left(\frac{P_{sIM,t}}{P_{sIM,t-1}} \right)^{\omega_{sIMt}} \left(\frac{P_{sID,t}}{P_{sID,t-1}} \right)^{\omega_{sIDt}}, \tag{10}$$

where ω_{sIMt} and ω_{sIDt} are the Sato-Vartia weights of imported and domestic intermediates.

Assuming that TFP and factor prices are not affected by trade agreements, we have that the change in domestic prices in sector s is

$$\frac{P_{sD,t}}{P_{sD,t-1}} = \frac{p_{sD,t}}{p_{sD,t-1}} = \left(\frac{P_{sI,t}}{P_{sI,t-1}}\right)^{1-\alpha_s}. \tag{11}$$

Note that the initial trade-agreement-induced reduction in input prices in sector s affects firms in all sectors of the economy through input-output linkages, further reducing input prices and hence sectoral prices, $p_{sD,t}$. To solve for the total change in $p_{sD,t}$, we use the exact price index associated with the production function of the domestic input bundle (9):

$$\frac{P_{sID,t}}{P_{sID,t-1}} = \prod_{j=1}^{S}\left(\frac{p_{jD,t}}{p_{jD,t-1}}\right)^{\varphi_{js}}. \tag{12}$$

Equations (10), (11) and (12) together imply the following system of S equations (one per sector):

$$\begin{aligned}
\ln\left(\frac{P_{sID,t}}{P_{sID,t-1}}\right) &= \ln\left(\prod_{j=1}^{S}\left(\left(\left(\frac{P_{jIM,t}}{P_{jIM,t-1}}\right)^{\omega_{jIMt}}\left(\frac{P_{jID,t}}{P_{jID,t-1}}\right)^{\omega_{jIDt}}\right)^{1-\alpha_j}\right)^{\varphi_{js}}\right) \\
&= \sum_{j=1}^{S}\omega_{jIMt}\left(1-\alpha_j\right)\varphi_{js}\ln\left(\frac{P_{jIM,t}}{P_{jIM,t-1}}\right)+\sum_{j=1}^{S}\omega_{jIDt}\left(1-\alpha_j\right)\varphi_{js}\ln\left(\frac{P_{jID,t}}{P_{jID,t-1}}\right). \qquad (13)
\end{aligned}$$

Once we have estimated the direct FTA-induced change in $P_{jIM,t}/P_{jIM,t-1}$, we can solve the system for $P_{sID,t}/P_{sID,t-1}$ to get the total effect on intermediate domestic prices. Written in matrix notation, the system (13) becomes:

$$\overrightarrow{\Delta P_{ID}} = \Omega_M\overrightarrow{\Delta P_{IM}} + \Omega_D\overrightarrow{\Delta P_{ID}},$$

where $\overrightarrow{\Delta X}$ are $S\times 1$ column vectors of (logs of) ideal price indices and Ω_M and Ω_D are $S\times S$ matrices of modified I-O input coefficients (i.e., the $\omega_{jI\{D,M\}t}\left(1-\alpha_j\right)\varphi_{js}$ from equation 13). Solving this system by matrix inversion yields:

$$\overrightarrow{\Delta P_{ID}} = \left(I-\Omega_D\right)^{-1}\Omega_M\overrightarrow{\Delta P_{IM}}.$$

With the solution for all $P_{sID,t}/P_{sID,t-1}$ in hand, we can compute the change in input prices and the resulting change in domestic prices using (10) and (11). The final step is to use trade and expenditure data to compute the relative shares of domestic and imported goods in the sectoral price indices (3) and to aggregate up the economy-wide consumer price index (2). We will implement this aggregation procedure in Section 5.4 but we first return to the estimation of price, quality and variety effects of trade agreements at the HS 6-digit level.

3. Quantifying the Impact of Trade Agreements

The decomposition in equation (8) is central to calculating the impact of trade agreements on consumer prices. As discussed, this equation applies regardless of whether the underlying HS 6-digit product is used for final consumption or as an intermediate input.[11] In this section, we describe how we estimate the effect of trade agreements on the individual components of (8).

We start with a straightforward estimating equation to determine the impact of trade agreements on prices. We then move towards inferring quality and estimating the impact of trade agreements on quality and variety. Our empirical methodology consists of four steps. First, we estimate the elasticities of substitution σ_h for each HS 6-digit product, based on Feenstra (1994) and Broda and Weinstein (2006). Then we infer quality and variety using the estimated elasticities and different proxies to control for hidden variety. Next, we estimate the impact of trade agreements on each of the components of (8). Once we have these estimates, we can replace $\lambda_{ht}/\lambda_{ht-1}$, p^o_{ht}/p^o_{ht-1} and q^o_{ht}/q^o_{ht-1} with predicted FTA-induced changes for the origin countries affected by the trade agreement.[12] This gives us a counterfactual change in the exact price index for each imported HS 6-digit product, which serves as the basic building block for the price index aggregation just described.

3.1. The Price Impact of TAs.

We model the impact of trade agreements on (non-quality- and non-hidden-variety-adjusted) prices as follows:

$$\ln p^o_{ht} = (\alpha_p)^o_h + (\alpha_p)_t + \beta_p \times FTA_{ot} + \mu^o_{ht},$$

where $(\alpha_p)^o_h$ and $(\alpha_p)_t$ are product-origin and time fixed effects. The dependent variable is the unit value obtained directly from trade data by dividing sales of a product-origin-time observation by its quantity. FTA_{ot} is a dummy which is one for origin countries o that have a free trade agreement with the EU at time t, and β_p measures the FTA's price impact. Following the literature on trade agreements (e.g., Baier and Bergstrand, 2007), we use a FTA_{ot} dummy rather than tariff changes as a regressor. The new generation of "deep" trade agreements contain

[11]The only difference for intermediates is that the elasticity parameters capture substitability in production rather than consumption. However, the elasticity estimation procedure (and hence the estimate) is the same irrespective of the product's use (see below).

[12]As will become clear, proxies for hidden varieties available from standard datasets are not well suited for estimating the impact of FTAs on n^o_{ht}. In practice, we will thus assume that n^o_{ht} does not change after the entry into force of an FTA. We return to this point below and discuss the implications for our estimated overall price index changes.

important measures such as the elimination of non-tariff barriers through harmonization of regulations on cross-border trade and investment that are harder to define at the level of a product. Tariff reductions or selective regulation measures are unlikely to capture these policy changes and would underestimate the impact of trade agreements on consumer welfare.

Note that while the choice of FTA partners by the EU is clearly not random, the use of product-origin fixed effects eliminates or at least reduces the most obvious problems arising from reverse causality and omitted variables bias by controlling for time-invariant trade partner characteristics. Given the absence of suitable instruments, the use of fixed effects is indeed the preferred approach in the trade agreement literature (e.g., Baier and Bergstrand, 2007). Nevertheless, if the EU systematically signed agreements with countries whose export prices are on a downward trajectory compared to other countries, the common-trend assumption underlying the standard difference-in-differences approach would be violated. The new generation of trade agreements negotiated by the EU are less likely to suffer from this problem, because they were motivated by regional geopolitics (such as the enlargement of the EU) or were primarily with smaller countries to catch up with the agreements negotiated by the United States. For example, Pelkmans and Brenton (1999) argue that economic factors are only one of the many which propel the EU towards such agreements. Brenton and Manchin (2003) go further to say that economic integration agreements are "at the forefront of EU policy towards developing countries and neighboring countries in Europe, including the countries of South-East Europe." DeBardeleben (2009) summarizes that the new agreements with Central and Eastern European countries cemented new governance linkages with some of Russia's closest neighbors. Through guarantees based on the NATO charter and integration into the fabric of Europe's social, political and economic life, the enlargement process largely removed these countries as 'zones of contention'between the West and Russia. The new trade agreements with these countries therefore reflected a policy objective of stabilizing post-Cold War Europe (Woolcock, 2007).[13]

Likewise, trade agreements with non-member countries such as Mexico reflected a "catch-up" game by the EU to keep pace with the trade agreements signed by the US. Baccini (2010) finds that a trade agreement with the EU is more likely for a country that had a trade agreement with the US in the previous year. The study argues that the EU reacted to trade agreements signed by the US with developing countries to avoid losing trade with the joining country or to establish its own regulatory standard in the international system. Woolcock (2007) documents that EU's FTA negotiations with Mexico, Mercosur, Central America, ASEAN and South Korea

[13]We will mainly focus the association agreements between the EU and the 2004 and 2007 accession countries that already came into force in the 1990s (see Section 4 below for details).

followed FTAs negotiated or envisaged by the US (e.g., CAFTA, US–Singapore, US–Thailand and US–Malaysia, and US–Korea agreements) and to a lesser extent Japan. This suggests that concern for future trade diversion rather than potential trade creation motivated many of the agreements signed by the EU.

Finally, it is worth noting that our results are less likely to suffer from political economy concerns because the EU negotiates trade agreements on behalf of its member countries' governments, implying there are fewer concerns that country-specific factors are driving FTA partner choice. In the light of these arguments, we believe that endogeneity bias is less likely in the context of EU FTAs. Below, we also provide additional evidence that the FTAs in our sample are not correlated with pre-trends in quality or prices.

We estimate β_p through the following first-differenced regression:

$$\ln p^o_{ht} - \ln p^o_{ht-r} = \left[(\alpha_p)_t - (\alpha_p)_{t-r}\right] + \beta_p \times (FTA_{ot} - FTA_{ot-r}) + \left(\mu^o_{ht} - \mu^o_{ht-r}\right). \tag{14}$$

Once we have estimates for β_p, we can compute a counterfactual price change due to the FTA as $\hat{p}^o_{ht}/\hat{p}^o_{ht-r} = \exp\left(\hat{\beta}_p \times (FTA_{ot} - FTA_{ot-r})\right)$. Replacing p^o_{ht}/p^o_{ht-1} with $\hat{p}^o_{ht}/\hat{p}^o_{ht-1}$ in (8) gives us the counterfactual impact of trade agreements on the price of HS product h from origin o at time t. Note that it does not matter whether product h is imported as an intermediate input or for final consumption – equation (14) applies regardless.

3.2. **Inferring Quality.** We now discuss how to estimate the quality impact of trade agreements. Under CES import demand, the total import value of HS 6-digit product h from origin o at time t is

$$X^o_{ht} = n^o_{ht} \left(p^o_{ht}/q^o_{ht}\right)^{1-\sigma_h} p^{\sigma_h - 1}_{ht} E_{ht},$$

where E_{ht} is total expenditure on HS-product h. Taking logs and defining $\alpha_{ht} \equiv \ln p^{\sigma_h-1}_{ht} E_{ht}$,

$$\ln X^o_{ht} = \alpha_{ht} + (1 - \sigma_h) \ln p^o_{ht} + \ln n^o_{ht} - (1 - \sigma_h) \ln q^o_{ht}.$$

Bilateral trade data give trade values X^o_{ht} and unit values p^o_{ht}, but not direct measures of quality. In principle, we could infer quality by regressing trade values on prices as follows:

$$\ln X^o_{ht} = \alpha_{ht} + (1 - \sigma_h) \ln p^o_{ht} + \epsilon^o_{ht},$$

where α_{ht} are product-time fixed effects and $\epsilon^o_{ht} \equiv \ln n^o_{ht} - (1 - \sigma_h) \ln q^o_{ht}$ is an error term capturing quality. But this approach would give biased estimates of σ_h because prices are likely to be correlated with quality so that there is an endogeneity problem.

Following Khandelwal et al. (2013), we can infer quality (up to a constant) if we have estimates for σ_h and proxies for n^o_{ht} through the following relationship:

$$\ln X^o_{ht} - (1-\hat{\sigma}_h)\ln p^o_{ht} = \alpha_{ht} + \ln \hat{n}^o_{ht} - (1-\hat{\sigma}_h)\ln q^o_{ht}. \tag{15}$$

To obtain estimates of σ_h, we build on Feenstra (1994) and Broda and Weinstein (2006) but extend their methodology to include an improved treatment of measurement error in our price proxy (unit values), which also affects the weighting matrix of the WLS estimator. Appendix B contains a detailed exposition of our estimation strategy. Our key generalization is to account for the measurement error that is likely to be present in observed prices of the variety which is used to normalize the price changes in the elasticity estimation. In practice, the Feenstra-Broda-Weinstein methodology and our approach yield very similar estimates. The median of $\hat{\sigma}_h$ across our HS 6-digit estimates is 3.20 compared to 3.16 using the FBW approach.[14] Below, we also show that our results are not affected by using the FBW estimates instead.

To capture hidden varieties, we use an approach similar to Amiti and Khandelwal (2013). Hidden varieties are defined as a function of the GDP of the partner country, $n^o_{ht} = (GDP^o_t)^{\beta_g}$ where g refers to a group of products within the same HS 3-digit code. This draws on standard trade models (Krugman, 1980) that predict that the number of varieties produced is increasing in a country's size. Note that we add flexibility by allowing the relationship between varieties and GDP to vary across broad product categories, although imposing a uniform coefficient does not change the following results.

As a robustness check, we will also proxy for n^o_{ht} more directly using the number of firms in a given country, year and sector. This information is available from UNIDO for 4-digit ISIC manufacturing industries across countries, which we map into our HS codes. The hidden variety term is now modelled as $n^o_{ht} = (N^o_{it})^{\eta}$ for products within a more aggregate UNIDO industry i. The underlying assumption here is that the number of unobservable varieties varies across industries (but not across products) for each origin country. The downside of this approach is that it reduces the number of available observations by close to 50%.

Note that while (15) applies to final products, it is straightforward to show that an exact analogue holds for intermediate products. The only difference is again that the final consumption good elasticity σ_h is replaced by the intermediate input elasticity ε_{hM} (see Appendix A). Furthermore, our FBW-type estimation procedure and the resulting elasticity estimate are the same irrespective of an imported product's use. Hence, for a given HS product, the exact

[14]We have elasticity estimates for approximately 4,000 HS6 products. Our value for $\hat{\sigma}_h$ compares to median estimates by Broda and Weinstein (2006, Table IV) for $\hat{\sigma}_h$ of 3.1-3.7 for the TSUSA/HTS classification (ca. 11,000-14,000 products depending on the year in question) and 2.7-2.8 for SITC-5-digit (1,500-2,700 products).

same quality estimation procedure applies and identical quality estimates result independent of whether the product is used for final consumption or intermediate use. This will prove useful below as we will argue that for our purposes, it is not possible to neatly divide HS products into final and intermediate use categories (see Section 5.4).

3.3. **The Quality Impact of FTAs.** Having estimated elasticities of substitution, we can use the two proxies of hidden varieties to infer quality. Defining $\ln Y^o_{ht} \equiv \ln X^o_{ht} - (1-\hat{\sigma}_h)\ln p^o_{ht}$ from equation (15), the first proxy for hidden variety implies the following regression:

$$\ln Y^o_{ht} = \alpha_{ht} + \beta_g \ln GDP^o_t + \varepsilon^o_{ht},$$

where $\varepsilon^o_{ht} \equiv -(1-\hat{\sigma}_h)\ln q^o_{ht}$. We regress $\ln Y^o_{ht}$ on product-time FE and GDPs of the origin countries to obtain quality inferences from the residual and the estimated elasticities of substitution. Using the second proxy for hidden variety, the implied regression for quality inferences is:

$$\ln Y^o_{ht} = \alpha_{ht} + \eta \ln N^o_{it} + \varepsilon^o_{ht}.$$

Note that in both cases, quality is estimated as a residual. Holding price and the number of hidden varieties fixed, any remaining variation in imports of product h from origin o is attributed to quality. In line with the existing literature (e.g., Khandelwal, 2010), we choose this indirect approach because more direct information on quality-related attributes is not available for the wide range of product and origin countries in our sample.

Having inferred quality, we use the same approach as for prices to model the impact of trade agreements on quality:

$$\ln \hat{q}^o_{ht} = (\alpha_q)^o_h + (\alpha_q)_t + \beta_q \times FTA_{ot} + \mu^o_{ht},$$

where $(\alpha_q)^o_h$ and $(\alpha_q)_t$ are product-origin and time fixed effects. We again estimate β_q through a difference-in-differences regression:

$$\ln \hat{q}^o_{ht} - \ln \hat{q}^o_{ht-r} = \left[(\alpha_q)_t - (\alpha_q)_{t-r}\right] + \beta_q \times (FTA_{ot} - FTA_{ot-r}) + (\mu^o_{ht} - \mu^o_{hs}). \tag{16}$$

Once we have estimates for β_q, we can compute a counterfactual quality change due to the FTA as $\hat{q}^o_{ht}/\hat{q}^o_{ht-r} = \exp\left(\hat{\beta}_q \times (FTA_{ot} - FTA_{ot-r})\right)$. Replacing q^o_{ht}/q^o_{ht-1} with $\hat{q}^o_{ht}/\hat{q}^o_{ht-1}$ in (8) gives us the counterfactual impact of FTAs on the quality component of the product-specific price index.

3.4. **The Variety Impact of FTAs.** The contribution of changes in the number of available varieties to the exact HS-level price index is given by $\lambda_{ht}/\lambda_{ht-1}$ from equation (8). Intuitively, a lower value for λ_{ht} implies a larger share of new varieties and hence a reduction in the price index between periods $t-1$ and t. By contrast, a lower value of λ_{ht-1} signals a higher share of disappearing varieties and is associated with a larger increase in the exact price index. Using expression (7), we can directly compute $\lambda_{ht}/\lambda_{ht-r}$ using our trade data and estimates for σ_h and n_h^o obtained from the previous steps, where r denotes the lag used in the computation.[15] To estimate the impact of FTAs on $\lambda_{ht}/\lambda_{ht-r}$, we estimate

$$\ln\left(\lambda_{ht}/\lambda_{ht-r}\right) = \alpha_t + \beta_v \times \left(FTA_{ot} - FTA_{ot-r}\right) + \mu_{ht}^o, \tag{17}$$

where we use the same lag r as for our previous difference-in-differences regressions. Note that $\lambda_{ht}/\lambda_{ht-r}$ already measures a change over r periods so that differencing is not required here. For comparability with our price and quality FTA-impact estimates, we estimate (17) on the same set of observations. While $\lambda_{ht}/\lambda_{ht-r}$ does not vary by origin country, different products are imported from different sets of countries so that there is variation in the FTA-regressor across both products and time.

Once we have obtained estimates of β_v, we compute the FTA-induced contribution of variety growth to the exact HS-level price index as $\widehat{\lambda_{ht}/\lambda_{ht-r}} = \exp\left(\hat{\beta}_v \times \left(FTA_{ot} - FTA_{ot-r}\right)\right)$.

4. Data Sources and Descriptive Statistics

We apply the methods just described to bilateral trade data for the European Union (EU) which has negotiated a large number of trade agreements over the past two decades. We examine the overall impact of EU FTAs implemented during our sample period 1993-2013. We use the term "European Union" (or EU12 or simply EU) to refer to the twelve member states prior to the 1995 enlargement (Belgium, Luxembourg, Germany, France, Italy, the Netherlands, the United Kingdom, Ireland, Denmark, Greece, Portugal and Spain). This geographic focus was chosen to keep the set of countries for our analysis constant over the sample period.

We need data on the value and the quantity of EU imports for each product, information about the timing of FTAs, and data for our proxies for hidden varieties (GDP and the number of firms). We obtain origin-specific EU import data at the 6-digit HS level for the period 1993-2013 from the United Nation's COMTRADE database, accessed through the World Bank's WITS interface. These trade data are classified according to the version of the Harmonized System that was in force at the time of reporting. To achieve comparability over time, we map all data into the 6-digit level of the HS0 (1988/1992) version of the Harmonized System,

[15] Again, if product h is used as an intermediate input, the interpretation of the elasticity parameter changes, capturing substitutability between inputs from different origins.

using concordances provided by WITS. Data on trade agreements and their implementation dates are available from the European Commission.[16] The information on expenditure and import shares as well as the input-output structure come from the OECD Inter-Country Input-Output database (2017 edition), which provide data for the 1995-2011 period.[17] For the baseline exercise, the tables are aggregated at the EU12 level, while for the aggregation results of the heterogeneity analysis we will also use individual country data as well as tables for three country groups. The input-output structure for the initial and final years of our sample is assumed to be the same of the first and final year available in the database, respectively 1995 and 2011. Finally, GDP data are from the World Bank Development Indicators and the number of firms per sector from the UNIDO Industrial Statistical Database (INDSTAT4).[18]

Table 1 presents a list of all countries with which the EU signed FTAs between 1993 and 2013. The first group consists of countries which became member states in 2004 and 2007, respectively. These countries all signed association agreements with the EU several years prior to accession and we use the respective dates for our FTA dummy. In robustness checks below, we also control for EU membership to see whether the later EU accession had effects over and above the trade liberalization measures implemented as part of the association agreements.

The EU also negotiated a number of additional trade agreements over our sample period. Within Europe, these include a customs union with Turkey, and FTAs or Association Agreements with the Faroe Islands, Macedonia, Croatia, Albania, Bosnia and Herzegovina, Montenegro and Serbia. In the Mediterranean, these include agreements with Israel, Algeria, Egypt, Jordan, Lebanon, Morocco and Tunisia. Further afield, the EU also implemented FTAs with Mexico, South Africa, Chile, Korea, Peru, Columbia, Costa Rica, El Salvador, Guatemala, Honduras, Nicaragua and Panama.[19]

To illustrate the trade volumes covered by the FTAs analyzed here, Figure 1 shows the value of EU12 imports by groups of origin countries. In 2012, the single most important source country for the EU12 was China which accounted for \$293bn or 12.5% of total imports from non-EU12 countries, followed by the United States (\$253bn, 10.8%), Russia (\$149bn, 6.4%) and

[16]See http://ec.europa.eu/enterprise/policies/international/facilitating-trade/free-trade/index_en.htm#h2-1.

[17]These data contain information on the expenditure shares needed for the construction of the Sato-Vartia weights ω_{st}, ω_{sMt}, ω_{sDt}, as well as the modified matrices of input-output coefficients Ω_M and Ω_D, which also include ω_{sIMt} and ω_{sIDt}, the Sato-Vartia weights of imported and domestic intermediates.

[18]To improve time and country coverage we use both the ISIC Rev. 3 and the ISIC Rev. 4 versions of the database, and we transform all the data in ISIC Rev. 3 using a concordance table available from the authors on request.

[19]Croatia is a special case as it only became a EU member at the very end of our sample period (in 2013). In this paper, we group Croatia with the "non-EU" agreements. We do not include the 1995 accession countries (Austria, Finland and Sweden) in either FTA group because they already had FTAs in place with the EU at the beginning of our sample period and because we would only have two pre-accession years of data to estimate accessions effects.

Switzerland ($115bn, 4.9%). This compares to imports of $279bn (11.9%) for the 2004/2007 EU accession countries, and imports of $256bn (10.9%) for the non-EU FTA partners.

5. Empirical Results

We start with our baselines estimates for the impact of the EU's trade agreements on the prices, quality and variety of imports. Section 5.2 examines the robustness of our results and Section 5.3 shows how the estimates vary across different countries, product groups and trade agreements. Finally, Section 5.4 presents our price index aggregation and estimates the total CPI effect of the EU's trade agreements.

5.1. **Baseline Results.** Table 2 presents results for our baseline specifications (14), (16) and (17). We choose a five-year lag for these initial estimations and regress the five-year log change in prices, quality and variety on the five-year change in the FTA dummy as well as year fixed effects. The coefficient on the FTA dummy thus tells us whether imports from origin countries with which the EU has implemented a free trade agreement within the last five years saw more pronounced changes in the dependent variable (e.g., prices) than imports from other countries. As discussed above, the same estimation equations apply regardless of whether a product is imported for final or intermediate input use and hence we include all HS products in the regressions underlying Table 2.[20]

Columns (1) and (4) show that FTAs had essentially no effect on import prices and variety - the corresponding coefficients are statistically insignificant and close to zero. However, FTAs did increase the quality of imported goods by around 7% (column 2) and this effect is highly statistically significant. We obtain very similar results if we use the Broda-Weinstein elasticities instead of our own estimates (column 3). Column (5) combines the price and quality estimates by showing that the net effect of FTAs on quality-adjusted prices is also negative and significant.[21]

These results highlight the importance of taking quality into account. A naive approach which simply regresses unit values on indicators of trade liberalization (such as our FTA dummy) would erroneously conclude that trade agreements have no impact on consumers. At least for the case of the FTAs implemented by the EU12, the entire effect works through changes in quality.

[20]See Section 5.4 for why a division of HS products into final and intermediate use categories is problematic for our purposes. We will allow for different types of heterogeneity in Section 5.3, however, including across broad ISIC categories.

[21]Column (5) also adjusts prices for hidden varieties, i.e., $p_h^{o,adj} = p_h^o / \left((n_h^o)^{1/(\sigma_h - 1)} \, q_h^o \right)$. Not adjusting for hidden varieties yields very similar estimates.

Our findings are consistent with a number of recent papers that also document quality-upgrading in response to trade liberalization or other changes in trade costs. Using plant-product-level data for the post-NAFTA period 1994-2004, Iacovone and Javorcik (2010) show that Mexican firms entering the U.S. market upgrade product quality in preparation for exporting because of the higher demand for quality by U.S. consumers. Similarly, Verhoogen (2008) shows that Mexican exporters upgraded quality in response to the currency devaluation of the late-1994 peso crisis. Given that the EU12's level of per-capita income is high relative to almost all of its recent trade agreements partners, quality upgrading for the EU market seems a plausible mechanism underlying our results. Below, we also present additional evidence that quality increases were particularly pronounced for the group of EU countries with the highest per-capita income.

There is also evidence that lower import barriers can lead to quality-upgrading by domestic firms by improving access to imported intermediate inputs (Fan et al., 2015; Bas and Strauss-Kahn, 2015). In view of the fact that most of the EU's agreements had a reciprocal element, reduced import barriers in the exporting countries could be a complementary explanation for our findings.[22]

Finally, if trade barrier reductions take the form of an elimination of quotas, liberalization can also lead to quality-downgrading as firms no longer face quantity restrictions encouraging the export of high-value products (Harrigan and Barrows, 2009). Clearly, this explanation is at odds with our finding of quality-upgrading, consistent with the fact that quota elimination was at best a small part of the EU's trade agreements.

5.2. **Robustness.** Tables 3-10 examine the robustness of our baseline results. In Table 3, we control for changes in per-capita GDP of the origin country of EU imports. Per-capita GDP has been shown to be an important correlate of the quality of exported goods (e.g., Schott, 2004; Feenstra and Romalis, 2012) and is thus a potential origin-time varying omitted variable. At the same time, however, there is strong evidence that productivity and per-capita GDP are themselves positively influenced by increased export opportunities (e.g., Trefler, 2004). So the estimates in Table 3 are best interpreted as eliminating the indirect effect of FTAs on quality working through changes in per-capita GDP. In any case, the coefficient estimates are almost identical to our baseline estimates from Table 2. As expected, per-capita GDP is also positively correlated with quality and prices.

[22]Unfortunately, it is not possible to provide a more direct test in our setting given the difficulties of measuring the degree of reciprocity. (As discussed, non-tariff barriers are the key component of the FTAs in our sample, so that the total effects are best captured by binary variables.)

Table 4 examines the importance of functional form. Column (1) uses two-year instead of five-year differences. As before, the effect of the FTAs on prices and variety is insignificant. The impact on quality is still statistically significant although the coefficient magnitude decreases by around 50%. This is consistent with the idea that quality upgrading might take time and will not materialise fully and immediately after the FTA has come into force.

Column (2) estimates a fixed-effects regression where the log of quality and price is regressed on product-origin country and year fixed effects. That is, instead of taking differences over time we eliminate the product-origin effects $(\alpha_p)_h^o$ through a within-transformation. The results for quality are again similar to our baseline. In contrast, FTAs are now estimated to have a significantly positive impact on prices. We cannot use the fixed effects approach for variety because the variety component of the price index change is inherently defined in first differences (equation 17).

In columns (3) and (4) we use a single two-year and five-year difference for each FTA, respectively. Specifically, we estimate the change in quality, prices and variety from one year prior to the implementation year of the FTA to either one or four years after the implementation year (i.e., t-1 to t+1 and t-1 to t+4). We would expect to find slightly larger effects compared to our baseline estimates because on average we now evaluate the effect of the FTA after a longer time period.[23] This is indeed the case, with the estimated impact on quality in column (4) being almost 50% larger than in the baseline specification.

One potential criticism of our baseline difference-in-differences approach is that, for a given product, we include all origin countries with which the EU has not implemented a free trade agreement in our control group. In Table 5, we instead select the relevant control group for each FTA through a simple matching procedure. We first use a logit regression to predict the likelihood that an FTA between the EU and a given origin country was implemented between t-5 and t (i.e., $FTA_{ot} - FTA_{ot-5}$) as a function of bilateral distance, a contiguity dummy and initial GDP of the origin country. We use the resulting propensity score estimates to construct control groups for each FTA partner using radius matching and imposing an exact match on year and product code.[24] Table 5 shows results for different values of the imposed matching radius ($r = 0.05$, $r = 0.1$ and $r = 0.2$). We again find a positive and significant impact of FTAs

[23]For example, if an FTA took place in year t, our baseline five-year difference estimates represent an average of the effects from t-5 to t, t-4 to t+1, t-3 to t+2, t-2 to t+3 and t-1 to t+4.

[24]That is, for each FTA implemented by the EU, we first select a group of control countries with a similar propensity score to the FTA partner country (where similarity is defined by the matching radius). For each product and year, we then calculate the difference in the five-year log difference of the variable in question (e.g., quality) between the partner country and the average of the control group. The average treatment effect on the treated (ATT) for the FTA in question is the average across all relevant product-year differences, and the treatment effect reported in Table 5 is the average of the FTA-specific ATTs. Standard errors are obtained via cluster-bootstrapping, where clusters are at the partner-time level for comparison with the clustering of standard errors in our baseline specifications.

on the quality of imported goods. FTAs now also seem to have a positive impact on prices although the coefficient estimate is again much smaller in magnitude than that for quality.[25]

In Table 6 we examine the sensitivity of our results to different ways of proxying for hidden varieties ($\hat{n}^o_{ht}$ in equation 15). In column 1, we do not control for hidden varieties at all. In column 2, we proxy $\hat{n}^o_{ht}$ by origin fixed effects, in column 3 by origin-by-1-digit-HS fixed effects and in column 4 by origin-by-3-digit-HS fixed effects. Column 5 uses the number-of-firms proxy described in Section 3. That is, the hidden variety term is now modelled as $n^o_{ht} = (N^o_{it})^\eta$ where N^o_{it} is the number of firms from origin country o at time t within the more aggregate UNIDO industry i to which HS product h belongs, and η is a parameter estimated in our first-stage quality regression.[26] The results in Table 6 show that our estimates of the impact of FTAs on quality are not particularly sensitive to the way we control for hidden varieties. The estimated FTA coefficients are either similar in magnitude or slightly larger than in our baseline regression.

In Table 7 we include an EU membership dummy in our first-difference regression ($EU_t - EU_{t-5}$). Most of the 2004/2007 accession countries had signed association agreements with the EU in the mid-1990s which had already liberalized bilateral trade well before the actual EU accession. The coefficient on the EU dummy will thus tell us whether EU accession had additional price and quality effects over and above the earlier trade liberalization. One reason we might expect such effects is a reduction in uncertainty over future access to European markets which were eliminated for these countries by fully joining the EU (see Handley and Limão, 2012). The results in Table 7 show that EU accession did indeed have an additional positive and significant effect, increasing quality by about 5% in addition to the 7% increase brought about by FTAs. In contrast to the FTAs however, EU accession also led to a significant price increase of about 12%. Importantly, including the EU accession dummy leaves the coefficient on the FTA dummy basically unchanged for all three dependent variables.[27]

Table 8 examines the relative role of tariff and non-tariff barriers in explaining our results. As discussed, one advantage of using an FTA dummy variable is that it also captures non-tariff barriers (NTBs) which are otherwise hard to measure because of the lack of reliable data. However, tariff data are available for around 70% of the observations in our sample and in Table 8, we include origin-specific EU import tariffs at the HS 6-digit level as an additional regressor.[28] The results show that tariff reductions had no effect on quality, and the coefficient of our FTA

[25]Note that the variety measure does not vary by origin country so that we cannot implement our matching procedure for this variable.

[26]To allocate trade flows to industries, we use a mapping from HS0 to ISIC Rev. 3 available from the World Bank's WITS website.

[27]One explanation for this is that the *EU* and the *FTA* dummies are highly correlated in levels (correlation coefficient 0.52) but uncorrelated in changes (correlation coefficient -0.07).

[28]We use effectively applied tariffs from the TRAINS database which are the lower of the EU's MFN tariff for a given HS6-digit product and any preferential tariff granted (e.g., as part of a free trade agreement).

dummy barely changes when we control for tariffs. We conclude that the quality increases we documented earlier are most likely due to FTA-induced changes in NTBs or additional effects such a reduction in uncertainty.

In Table 9, we revisit the issue of the potential endogeneity of our FTA dummy. While we have argued above that endogeneity is less likely in the EU context, a concern with our identification strategy is that the EU could have systematically signed free trade agreements with countries whose product quality was on an upward trajectory, violating the common trend assumption underlying our difference-in-differences specification. In Table 9, we report results of a placebo check where we regress our differenced price, quality and variety measures on an indicator for whether the EU implemented a trade agreement with a given origin country within the next five years. If the EU systematically signed agreements with countries with underlying pre-trends in our variables of interest, the coefficient on this lead indicator should be statistically significant. As the results in Table 9 show, however, this is not the case for our sample of trade agreements, suggesting that pre-trends are unlikely to explain our earlier results.[29]

In our next robustness check, we investigate whether the EU's trade agreements led to a loss of domestic varieties. There is no data on domestically produced varieties at the HS 6-digit level, so we cannot use the same approach as in Section 2. Instead, we follow Hsieh et al. (2016) in associating varieties with plants or establishments, and measure the number of varieties as the number of establishments in a given sector.[30] We also require a new identification strategy to link trade agreements to changes in our domestic variety proxy. We construct an industry-specific exposure indicator which measures the share in total EU absorption (i.e., production plus imports minus exports) accounted for by origin countries with which the EU has a trade agreement. This indicator is defined as $E_{ts} = \sum_o FTA_{ot} \times ABS_{os}$, where FTA_{ot} is a binary indicator taking the value one if the EU had a trade agreement with country o at time t, and ABS_{os} is the share of country o in total EU absorption in industry s in 1993.[31] This

[29]In an unreported robustness check, we have also computed the trade flow increases implied by our quality-adjusted prices and the relevant price elasticities. For the median HS product in our data ($\sigma_h = 3.2$), we obtain a predicted trade flow increase of 16%. This magnitude seems to be plausible given a median five-year change in actual trade flows in our sample of 26% and a median change of 43% for the five-year periods around trade agreements.

[30]Hsieh et al. (2016) show theoretically that welfare changes arising from changes in the set of establishments serving a country can be decomposed into changes in the number of establishments and their average productivity. An exact measure of the total change requires separate data on the sales of continuing firms which is not available in our context. Thus, the approach we outline in the following is best thought of as an approximation unless we are in the special case where entering, exiting and continuing establishments have the same productivity.

[31]Production data and the count of establishments are from the UNIDO database and sectors are defined at the 4-digit level of the ISIC Rev. 3 classification (145 sectors). We use of a time-invariant absorption share to reduce problems arising from the potential endogeneity of our absorption measure.

indicator captures the importance of the countries with which the EU signs trade agreements. For example, if the EU signs a new agreement with an important producer, E_{ts} will increase by more than if the agreement is signed with a country that only accounts for a very small share of total absorption in industry s. Next, we regress the log of the count of establishments in the EU in industry s on our exposure variable, controlling for sector and year fixed effects. If import competition triggered by new trade agreements led to a reduction in domestically produced varieties, we would expect a negative and significant coefficient on E_{ts}. As Table 10 shows, our estimated coefficient is indeed negative once we control for sector fixed effects, but far from conventional levels of statistical significance. We conclude that there is no prima fascia evidence that trade agreements led to a reduction in domestic varieties.[32]

Finally, we discuss the role of the assumption of constant markups in our setting. We have assumed a CES constant-markup setting and focused on import prices because the available consumer price information is less detailed than the fine product categorization of trade data and does not allow for a convincing identification of trade agreement effects due to the absence of cross-origin variation.[33] In terms of welfare changes, this means that our estimates should be interpreted as the lower bound of the potential gains to final consumers, as long as import competition lowers the markups charged by firms to consumers. We would ideally like to estimate how much consumers gain from reductions in markups charged by firms when import competition intensifies. Unfortunately, this is not feasible in our setting because of the same data constraints that arise in the literature pioneered by Broda and Weinstein (2006) which uses trade data. Additional firm-level market shares are needed to determine the markups charged to consumers.[34] However, a large empirical literature finds that increased competition from integration lowers markups charged in the importing country. Dhingra and Morrow (2012) show that the CES demand setting, where markups do not fall with import competition, provides a lower bound for the gains from integration in models of imperfect competition. As we work with a CES demand setting where markups do not vary, our results can be interpreted as a lower bound of the potential gains from trade to consumers.[35]

[32]By contrast, Hsieh et al. find that the Canada-US Free Trade Agreement (CUSFTA) led to negative net welfare effects from establishment entry and exit. This difference seems plausible given that CUSFTA represented a much more significant shock for Canadian firms given the size and proximity of the US.

[33]See footnote 7 for details.

[34]See Blaum et al. (2015), Feenstra and Weinstein (2010) or Bernard and Dhingra (2015) for different approaches that rely on firm-level or other data not available for the broad range of countries in our sample.

[35]Throughout, we have focused on simple c.i.f. import values, rather than c.i.f. values inclusive of tariffs, as the basis of our analysis. This is because tariff data are missing for a large part of our sample and because tariff revenue might of course benefit consumers in other ways. If we base our unit value and quality estimates on c.i.f. prices inclusive of tariffs, we unsurprisingly find a slightly larger role for price reductions (many agreements contain a tariff reduction component) and a correspondingly smaller quality effect. However, this does not affect our overall CPI effects presented below, nor the earlier headline finding that only quality matters for the EU12

5.3. **Heterogeneity Analysis.** Our headline results in Table 2 present an average treatment effect across products, exporting countries and EU12 member states. We now investigate how treatment effects vary across these groups. The main challenge faced in this context is that the data and the corresponding regression results become increasingly noisy as we move to finer levels of disaggregation, especially for some of the smaller countries and trading partners. Disaggregation also creates issues of spurious significance when analysing dozens or even hundreds of coefficient estimates, as statistically significant effects will appear by virtue of Type I errors alone. We use a two-pronged approach to address these issues. We first present results for different groupings of importing countries, trading partners and products which we believe are of interest and are justifiable from an ex-ante perspective. We then use a version of our baseline regression sample where we split up total EU imports into imports by each of the twelve member states. We run our baseline equations on these importer-exporter-product-year specific data and interact the trade agreement indicator with a set of explanatory variables such as the income level of the importer, bilateral distance or initial import shares. This allows us to highlight broad trends in terms of heterogeneity while imposing enough structure to significantly reduce problems of spurious significance.[36]

Panel A of Table 11 shows results for regressions where we split our sample into three groups of EU12 countries. We do so according to per-capita income levels in 2005 which also broadly corresponds to a split along geographic lines: Portugal, Greece and Spain (low income, “South”); Italy, France, Germany and Denmark (middle income, “Central”); and Belgium/Luxembourg, the United Kingdom, Ireland and the Netherlands (high income, “North”).[37]

There is a substantial amount of heterogeneity across these three groups. In particular, the group of the four richest EU member states saw a much larger increase in the quality of imported products - around 13% as opposed to a statistically insignificant 0.3% for Portugal, Greece and Spain. By contrast, the latter group saw a stronger and statistically highly significant decline in import prices close to 6%. One potential explanation for this heterogeneity is that demand for quality is higher in high income countries (e.g., Hallak, 2010), so that the incentives for quality upgrading by exporters discussed above might have been stronger.

---

as a whole (the price effects in Table 2 are still statistically insignificant, and the quality effects are significantly positive).

[36]We have also used the same disaggregated dataset to perform a large number of sample splits, generating treatment effects for each exporter-FTA partner-ISIC combination in the data. We then regressed the resulting set of coefficient estimates on the same set of explanatory variables. Results were qualitatively very similar to the interaction regressions presented below (available from the authors on request).

[37]We use GDP per capita on a PPP basis for 2005 from the World Bank Development Indicators for this ranking. Belgium and Luxembourg report trade data jointly, and we use a population-weighted mean of Belgium and Luxembourg GDP per capita.

In Panel B, we perform similar sample splits by income level for the EU's trade agreement partner countries. We use the World Bank's income classification to allocate partner countries to three groups: low and lower middle income, upper-middle income, and high income. Results again reveal substantial differences across income groups. We find large and statistically significant increases in product quality only for the two lower income groups but not for high income FTA partners. Overall, agreements with low and lower middle income countries seem most beneficial to EU consumers in terms of price, quality and variety effects. For upper-middle income partners, strong increases in quality are partially offset by higher prices. For high income FTA partners, we did not find any significant effects.[38]

Panel C splits partner countries by type of agreement and geographic location. We distinguish future EU members (the 2004/2007 accession countries), the Stabilization and Association Agreements with the Balkan countries (the ex-Yugoslav republics and Albania), the Euro-Med Assocation Agreements with the EU's meditarranean neighbors and non-regional free trade agreements with Korea, South Africa and South and Central American countries. We find the strongest quality effects for future EU members and the Balkan countries, followed by the non-regional agreements. While we observe no statistically significant quality increase for the Euro Med countries, prices as measured by unit values did fall by around 7% after the implementation of trade agreements with the EU. These results demonstrate that the correlation of quality increases and partner country income levels observed in panel B is complex and also reflects differences across types of agreements and the geographic location of the EU's partner countries.

Table 12 disaggregates our results by product groups. We use 13 sectors, which for non-service industries corresponds to a level of aggregation slightly above the two-digit level of the ISIC classification. We map the HS codes underlying our baseline estimates into these sectors and re-estimate our baseline specification for each ISIC subsample separately.[39] For all sectors, quality increases were either positive and significant or insignificantly different from zero. Overall, six out of thirteen sectors saw statistically significant FTA-induced increases in the quality of imported goods, with the strongest effects observed for machinery and motor vehicles, followed by textile products and wood, paper and printing. Estimates for prices and variety are again much smaller and stastically and economically insignificant for the vast majority of sectors. One exception is "Machinery and equipment n.e.c." where we estimate a statistically significant price decline of around 9%.

[38]We caution that this last group only contains two partner countries, Israel and Korea; the agreement with Korea also only came into force in 2011, so that the resulting effects might not yet be fully present in our data.

[39]The mapping from HS0 codes to ISIC Rev. 3 comes from the World Bank's WITS website. Note that only a small number of HS codes map into service industries, so we present one aggregate estimate for the entire service sector (ISIC 40_99). (There are not enough HS codes to obtain further breakdowns within services.) We will return to the issue of direct import effects in services in our aggregation exercise below.

We next turn to the interaction regressions using the fully disaggregated sample. That is, we now model the impact of trade agreements on prices and quality as

$$\ln y_{ht}^{od} = (\alpha_p)_h^{od} + (\alpha_p)_t + \beta_{y0} \times FTA_{ot} + \sum_{j=1}^{J} \beta_{yj} \times FTA_{ot} \times C_j + \sum_{j=1}^{J} \gamma_{yj} C_j + \mu_{ht}^{od},$$

where $y \in \{p, q\}$ denotes price or quality, and the C_j are a set of time-invariant product, exporter and importer specific variables, such as bilateral distance, the type of trade agreement or the initial market share of exporting country o in total imports by importing (EU12) country d in HS product h (see Table 13 for the full list). As discussed, we no longer use aggregate EU12 imports, hence the new subscript d and the use of exporter-importer-product fixed effects, $(\alpha_p)_h^{od}$. As before, we eliminate all time-invariant variables by taking first differences over time:

$$\ln \Delta_5 y_{ht}^{od} = \Delta_5 \alpha_{Pt} + \beta_{y0} \times \Delta_5 FTA_{ot} + \sum_{j=1}^{J} \beta_{yj} \times C_j \times \Delta_5 FTA_{ot} + \Delta_5 \mu_{ht}^{od}, \tag{18}$$

where Δ_5 denotes five-year differences and $\Delta_5 \alpha_{Pt} = (\alpha_p)_t - (\alpha_p)_{t-5}$ enters the specification as a set of year fixed effects.

Table 13 presents results for the estimation of (18) using prices and quality as the dependent variable, respectively.[40] In columns (1) and (4), we use interaction terms similar to the categorical variables from Tables 11 and 12: dummy variables for the three income groups of partner and EU12 countries and dummy variables for the different categories of trade agreements. Columns (2) and (5) instead interact a number of alternative explanatory variables with the first-differenced FTA dummy. First, we use two standard gravity equation variables – bilateral distance between the FTA partner and the EU country in question and a dummy for whether they share a common official language. Second, we control for the initial, pre-agreement trade share of the partner country in the total imports of an EU12 country as a measure of the importance of a given trade agreement. Third, we include the time elapsed since the implementation of an FTA to control for the fact that some of the agreements in our sample started less than five years before the end of our sample period, so that their effect might not be fully captured by our first-difference regressions. Finally, we also use a measure of bilateral differences in revealed comparative advantage as a rough measure of potential gains

[40] Note that our variety measure does not vary by origin country, so that we cannot estimate interaction regressions for this part of the price decomposition.

from specialization.[41] Columns (3) and (6) combine both sets of regressors. In all regressions, we also include fixed effects for the 13 ISIC groups as part of the interaction variables (C_j).

The results for our categorical indicators are broadly similar to before. Higher EU importer income levels are associated with significantly stronger quality effects, agreements with Euromed countries reduce prices and FTAs with future EU members and Balkan countries increase quality.[42] However, we no longer find that agreements with lower income countries raise quality by more than those with high income countries. Columns 2-3 show that higher initial market shares, having an official language in common and larger bilateral differences in revealed comparative advantage are all correlated with a stronger price-reducing effect of FTAs, although the RCA effect disappears once we control for our categorical variables in column 3. RCA differences also correlate positively with quality increases, whereas common language is consistently associated with lower quality effects.[43]

5.4. **Aggregation.** We now use our estimates to implement the decomposition of the economy-wide exact price index outlined in Section 2. For this, we require expenditure and import shares for different commodity groups (to compute ω_{st}, ω_{sMt}, and ω_{sDt}), as well as intermediate input shares for the calculation of the indirect domestic price effect (Ω_M and Ω_D, which also include ω_{sIMt} and ω_{sIDt}, the Sato-Vartia weights of imported and domestic intermediates). We use information about these shares from the OECD Inter-Country Input-Output database for the countries in our sample, aggregated to the EU12 level. As before, we work at a level of aggregation slightly above the ISIC 2-digit level (13 sectors). Having allocated HS 6-digit imported products to the different ISIC sectors, we then use input-output data to compute sectoral linkages and aggregate up to the economy wide CPI.

In principle, our expenditure and input shares can be combined with any of the coefficient estimates presented in the previous sections. Here, we focus on four sets of estimates. First, we use our aggregate coefficients from Table 2, which do not vary across exporters, importers or product groups. That is, we compute our predicted price, quality and variety changes ($\hat{q}^o_{ht}/\hat{q}^o_{ht-r}$, $\hat{p}^o_{ht}/\hat{p}_{ht-r}$ and $\widehat{\lambda_{ht}/\lambda_{ht-r}}$) using the same aggregate coefficients across all expenditure groups s. Second, we allow coefficients to vary by ISIC groups (see Table 12 for details of these estimates). Third, we again impose coefficient homogeneity across expenditure groups

[41] The revealed comparative advantage of country i in product h is measured as $RCA_{ih} = (X_{ih}/X_i) / (X_{wh}/X_w)$, where X_{ih} denotes exports of h by country i, X_i are total exports of i, X_{wh} are world exports of h and X_w are total world exports. Our bilateral RCA difference measure for countries i and j is defined as $rcadiff_{ij} = \sum_h |RCA_{ih} - RCA_{jh}|$ and hence captures differences in the export structure of the two countries.

[42] These coefficient estimates are relative to the excluded category which are observations for low income EU countries, low and lower middle income partner countries, non regional FTAs and ISIC group 01_05.

[43] We have also experimented with other explanatory variables such as variables from the World Governance Indicators and the World Bank's Doing Business database. These regressors were statistically insignificant throughout.

but allow estimates to vary across the three EU income groups defined earlier. This allows us to compute aggregate price effects for each of these groups. Finally, we use the interaction regression estimates from Table 13 to see how robust our aggregate results are in the presence of treatment effect heterogeneity.

For each set of coefficient estimates, we decompose the price index changes for each HS code according to equation (8), restated here using five-year lags for comparability with our estimates:

$$\frac{p_{ht}}{p_{ht-5}} = \underbrace{\left(\frac{\lambda_{ht}}{\lambda_{ht-5}}\right)^{\frac{1}{\sigma_h-1}}}_{\text{Variety}} \times \underbrace{\Pi_{o\in I}\left(\frac{n^o_{ht}}{n^o_{ht-5}}\right)^{-\frac{\omega^o_{ht}}{\sigma_h-1}}}_{\text{Hidden Variety}} \times \underbrace{\Pi_{o\in I}\left(\frac{p^o_{ht}}{p^o_{ht-5}}\right)^{\omega^o_{ht}}}_{\text{Prices}} \times \underbrace{\Pi_{o\in I}\left(\frac{q^o_{ht}}{q^o_{ht-5}}\right)^{-\omega^o_{ht}}}_{\text{Quality}}.$$

We compute price, quality and variety changes due to the FTA as follows:

$$\begin{aligned} \hat{p}^o_{ht}/\hat{p}^o_{ht-5} &= \exp\left(\hat{\beta}_p \times (FTA_{ot} - FTA_{ot-5})\right) \\ \hat{q}^o_{ht}/\hat{q}^o_{ht-5} &= \exp\left(\hat{\beta}_q \times (FTA_{ot} - FTA_{ot-5})\right) \\ \widehat{\lambda_{ht}/\lambda_{ht-5}} &= \exp\left(\hat{\beta}_v \times (FTA_{ot} - FTA_{ot-5})\right) \end{aligned}$$

where the $\hat{\beta}$ are either the aggregate estimates from Table 2, the ISIC-level estimates from Table 12, the country group level estimates from Table 11 or the interaction regression coefficients from Table 13.[44] We use origin Sato-Vartia weights ω^o_{ht} and the elasticity of substitution estimates $\hat{\sigma}_h$ to calculate the HS 6-digit level price index changes (p_{ht}/p_{h-5}) and its price, quality and variety components.[45]

The first row of Table 14 reports summary statistics for the total HS-level price index effects and its components across HS products and years for our set of aggregate estimates from Table 2.[46] Note that while the coefficients $\hat{\beta}$ do not vary across products, estimated HS-level price effects are also influenced by the distribution of shares across origin countries which does vary

[44]For the interaction regressions, we predict $\hat{y}^{od}_{ht}/\hat{y}^{od}_{ht-5} = \exp\left(\hat{\beta}_{y0} \times \Delta_5 FTA_{ot} + \sum_{j=1}^{J} \hat{\beta}_{yj} \times C_j \times \Delta_5 FTA_{ot}\right)$ where y denotes price or quality. As discussed, we cannot run interaction regressions for variety. Instead, we run the baseline variety regression (17) separately for each EU12 country and compute importer-specific predictions $\widehat{\lambda_{hdt}/\lambda_{hdt-5}}$ based on the resulting coefficient estimates ($\hat{\beta}_{vd}$).

[45]As discussed, we do not try to estimate an FTA-induced change in hidden varieties because our proxies are too crude to expect reliable results. A large body of work on firm-level export market entry after trade liberalization suggests that the number of hidden varieties should be positively affected by FTAs. As such, our counterfactual price index change is likely to be an underestimate.

[46]Because we are using overlapping five-year differences in our baseline estimation, we can calculate counterfactual FTA impacts for every year in our sample period. Table 14 reports sample statistics across HS products and years.

by HS code. The FTA impact on the mean HS product was to reduce p_{ht}/p_{ht-5} by -0.35%. This is entirely explained by increases in quality, with the impact of the price and variety components being negligible.

The next aggregation step is to use the shares of each HS product in the total imports accounted for by a given ISIC sector to compute the import price indices for final goods and intermediate inputs. Note that we use the same set of HS codes for each of these indices. This is because in practice, it is not possibly to neatly classify HS codes as either final goods or intermediate inputs - the same product can be used for both purposes.[47] However, our input-output tables provide information on the share of each ISIC industry allocated to final and intermediate input use. We use these shares to weight domestic and imported components in final and intermediate sector-level price indices. Thus, the relative contribution of each HS-level price change to intermediate and final good price changes is implicitly determined by the use shares of the corresponding ISIC sector.

Row 2 of Table 14 shows that, when we use the set of aggregate coefficient estimates, FTAs reduced the final good import price index by an average of -0.42% across ISIC groups and years. FTA-induced changes in domestic prices working through intermediate input linkages come in somewhat smaller at -0.06% (row 3). The overall ISIC-level price change is a weighted average between these two price changes. Because the import share is usually substantially smaller than the share of domestically produced goods ($\omega_{sM} < \omega_{sD}$ in equation 3), the overall ISIC-level average price change is -0.14%, much closer to the domestic price change than the import price change (row 4).

The final step is to use expenditure shares to aggregate across ISIC price indices to arrive at an economy wide consumer price index change. Row 5 in Table 14 reports a counterfactual FTA effect over our entire sample period. This is done by chaining the predicted five-year changes over the periods 1993-1998, 1998-2003, 2003-2008 and 2008-2013. We also compute bootstrapped 95% confidence intervals for the aggregate price effects by resampling 200 times from the original regression sample underlying the $\hat{\beta}$ coefficient estimates. For our aggregate estimates (Table 14), we find that the cumulative effect of FTAs over the period 1993-2013 was to lower the aggregate EU12 consumer price index by 0.24%, with a 95% confidence interval of $[-0.19\%, -0.28\%]$. Put differently, in the absence of FTAs, EU quality- and variety-adjusted consumer prices would be higher by around a quarter of a percent. While this is not a large effect, it still amounts to substantial savings for EU consumers of around €24 billion per year

[47] As a consequence, standard attempts at mapping HS products often lead to contradictory results. For example, the UN's Broad Economic Categories will yield a different set of final consumption goods than when using concordances from products to consumption classifications such as COICOP (see OECD, 2001).

given that total consumer expenditure in the EU12 was approximately €10 trillion in 2013.[48] A substantial part of this change is due to imported intermediate inputs. If we switch off this channel by setting the domestic price index change to 0%, the overall CPI effect declines to −0.13% (see row 6).

One advantage of using our aggregate estimates is that we can easily verify how our different robustness checks from Section 5.2 translate into aggregate CPI effects. In fact, our baseline results from Table 2 are at the lower end of the various quality coefficients presented in Tables 3-8, so that the overall CPI effect reported above is best seen as a conservative estimate. For instance, if we instead use estimates based on the Broda-Weinstein elasticities (Column 3 of Table 2), the overall impact increases to −0.30% (Row 7 of Table 14); and if we use estimates based on the number-of-firms proxy for hidden varieties (Column 5 of Table 6), the overall impact doubles to −0.49%.[49]

Table 15 summarizes the overall CPI effect for the aggregation exercises based on heterogeneous coefficient estimates. Panel (1) shows that using ISIC-level estimates instead of aggregate estimates slightly increases the overall impact to −0.27% or €27 billion per year (−0.18% if we switch off the intermediate input channel). Looking at the country group estimates (Panel 2), consumers in all three country groups benefitted from the EU's trade agreements but at −0.41% the CPI reductions for high-income country consumers were three times larger than those of consumers in the middle- and low-income groups. For both the ISIC and country-level aggregations, the aggregate effects are precisely estimated and we can always reject the null of no overall price effect. Finally, Panel (3) shows that allowing for more coefficient heterogeneity

[48] Since we estimate the cumulative effect, the consumer savings are smaller in years before 2013, but potentially higher for all years after, given the lag in the quality effect and the fact that some FTAs were signed towards the end of the period. We also note that our predicted changes in the economy-wide consumer price index will in general be different from CPI figures published by national statistical agencies. This is because of different aggregation schemes and, more importantly, because published CPI figures often do not adequately control for changes in quality and almost never for changes in the number available varieties. As such, our results should be interpreted as the true cost of living changes brought about by FTAs which existing CPIs would like to measure but are currently unable to do.

[49] Note that we are also using our aggregate estimates to infer price, quality and variety changes for the service sector (ISIC 40_99). While there is a small set of HS codes that maps into ISIC 40_99, a better interpretation of this approach is that we are assuming that the FTA impact on services trade (which is not directly measurable due to the absence of unit values for services imports) is the same as on goods trade. Given the nature of the new generation of "deep" trade agreements and their focus on non-tariff barriers, this seems a legitimate albeit crude way of capturing the overall FTA effect. In any case, switching off all direct service sector effects by setting the relevant import price changes to zero does not change the overall CPI effect by much (−0.20% instead of −0.24%). This is because the import share of services is only around 6-7%, limiting the direct impact of imports. (Note that service industries will still benefit from the indirect effect working through intermediate goods even if there are no direct import price effects.)

increases the mean estimated CPI effect to −0.47% (−0.30% without intermediates).[50] However, the noise in the underlying coefficient estimates also increases substantially, leading to a widening of the 95% confidence interval to [−0.17%, −0.76%]. Hence, we cannot reject the hypothesis that the aggregate effect is equal to −0.24% (the CPI reduction predicted using our baseline estimates).[51]

6. Conclusion

This paper examines the consumer welfare effects of the new generation of trade agreements implemented by the European Union between 1993 and 2013. We find that for the EU12 as a whole, these agreements increased welfare primarily by raising the quality of imported products from partner countries. There are important differences across EU countries, trading partners and the type of trade agreement, however. For example, high-income EU countries (the United Kingdom, the Netherlands, Ireland and Belgium/Luxembourg) saw a much stronger increase in quality than other EU countries. Indeed, for the group of low-income EU countries (Greece, Spain and Portugal), the impact of trade agreements worked almost exclusively through a reduction in prices rather than through increased quality.

Using expenditure shares of EU consumers, we also compute the aggregate consumer price index effects implied by our estimates. Our baseline results suggest that the trade agreements implemented by the EU lowered the CPI by 0.24%, saving EU consumers about €24 billion per year. Of this overall effect, we attribute around 55% to the direct effect on the prices and quality of imported products. The remaining part is due lower domestic prices brought about by a reduction in the cost of imported intermediate inputs. Looking across EU country groups we find that high-income member states saw substantially stronger quality increases and hence larger overall consumer benefits.

[50]The sectoral shares and input-output linkages are constructed at the relevant level of aggregation, the three country groups for Panel 2 and individual countries for Panel 3. In the latter exercise we then aggregate from the country results to the EU level using countries' shares in aggregate EU12 expenditure in 2013. Using other forms of aggregation, such as time-varying Cobb-Douglas or Sato-Vartia shares, leads to virtually identical results.

[51]In unreported results, we have also examined the distributional effects of the EU's FTA by matching the ISIC-level price changes from Table 15 (Panel 1) to data on consumer expenditure by income deciles. Our findings show that the FTA impact was moderately regressive in the sense that richer households saw slightly stronger CPI reductions than poorer households. This is due to the fact that richer households tend to have higher expenditure shares for products with higher estimated quality-adjusted prices falls (e.g., motor vehicles).

Appendix A. Additional Results for Intermediate Inputs

A.1. **Derivation of Equation** (8) **for Intermediate Inputs.** We start by restating the expression for the HS 6-digit level intermediate input bundle:

$$i_{hM}=\left(\sum_{o=1}^{O}\sum_{z=1}^{n_{hM}^{o}}\left(q_{hM}^{o}\left(z\right)i_{hM}^{o}\left(z\right)\right)^{\left(\varepsilon_{hM}-1\right)/\varepsilon_{hM}}\right)^{\frac{\varepsilon_{hM}}{\varepsilon_{hM}-1}}.$$

The associated price index is:

$$\left[\sum_{o=1}^{O}\sum_{z=1}^{n_{hM}^{o}}\left(p_{hM}^{o}\left(z\right)/q_{hM}^{o}\left(z\right)\right)^{1-\varepsilon_{hM}}\right]^{\frac{1}{1-\varepsilon_{hM}}},\quad \varepsilon_{hM}>1.$$

By analogy to the demand side, we assume that varieties are identical within a product-origin combination:

$$p_{hM}=\left[\sum_{o=1}^{O}n_{hM}^{o}\left(p_{hM}^{o}/q_{hM}^{o}\right)^{1-\varepsilon_{hM}}\right]^{\frac{1}{1-\varepsilon_{hM}}},$$

where n_{hM}^{o} is the number of "hidden varieties" of product h imported from country o.

To use results from the price index literature, we define the quality- and hidden-variety adjusted price of the HS6 variety from origin o as $p_{hM}^{o,adj}\equiv p_{hM}^{o}/\left(\left(n_{hM}^{o}\right)^{1/\left(\varepsilon_{hM}-1\right)}q_{hM}^{o}\right)$. For a constant set of origin countries, the price index ratio over time for product h is given by:

$$p_{hM,t}/p_{hM,t-1}=\Pi_{o=1}^{O}\left(p_{hM,t}^{o,adj}/p_{hM,t-1}^{o,adj}\right)^{\omega_{hMt}^{o}},$$

where ω_{hMt}^{o} are the Sato-Vartia weights associated with each variety. If the set of varieties (i.e., origin countries) changes between periods, the exact price index can be rewritten as:

$$p_{hM,t}/p_{hM,t-1}=\left(\lambda_{hM,t}/\lambda_{hM,t-1}\right)^{1/\left(\varepsilon_{hM}-1\right)}\Pi_{o\in I}\left(p_{hM,t}^{o,adj}/p_{hM,t-1}^{o,adj}\right)^{\omega_{hMt}^{o}},\tag{A.1}$$

where

$$\lambda_{hMr}=\left(\sum_{o\in I}p_{hM,r}^{o,adj}x_{hM,r}^{o}\right)/\left(\sum_{o'\in I_r}p_{hM,r}^{o',adj}x_{hM,r}^{o'}\right)\quad r=t-1,t$$

and I denotes the set of varieties present in both periods, t and $t-1$ (and I_r is the set of varieties present in period r). Substituting back for $p_{hM}^{o,adj}$ into (A.1), the exact price index can be decomposed into its components of variety, prices, quality, and hidden variety as follows:

(A.2)

$$\frac{p_{hM,t}}{p_{hM,t-1}}=\underbrace{\left(\frac{\lambda_{hMt}}{\lambda_{hMt-1}}\right)^{\frac{1}{\varepsilon_{hM}-1}}}_{\text{Variety}}\times\underbrace{\Pi_{o\in I}\left(\frac{n_{hMt}^{o}}{n_{hMt-1}^{o}}\right)^{-\frac{\omega_{hMt}^{o}}{\varepsilon_{hM}-1}}}_{\text{Hidden Variety}}\times\underbrace{\Pi_{o\in I}\left(\frac{p_{hMt}^{o}}{p_{hMt-1}^{o}}\right)^{\omega_{hMt}^{o}}}_{\text{Prices}}\times\underbrace{\Pi_{o\in I}\left(\frac{q_{hMt}^{o}}{q_{hMt-1}^{o}}\right)^{-\omega_{hMt}^{o}}}_{\text{Quality}}.$$

This is of course the exact analogue to equation (8) for final goods, with the difference that the elasticity parameters ε_{hM} now measure production substitutibility within HS codes. Crucially, however, these parameters can be estimated using the same data and the same Feenstra-Broda-Weinstein-type procedure described in Appendix B. Thus, for a given elasticity estimate for product h, the only difference is one of interpretation. If we believe that an HS product is an intermediate input, our elasticity estimate is a proxy for ε_{hM}; if we believe an HS product is a final good, our elasticity estimate is a proxy for σ_h. As we have argued earlier, however, it is impossible to neatly divide HS products into final and intermediate products given that the same product often has both uses. Hence, we perform the above decomposition only once per HS code and instead rely on information on the share of each upper-level ISIC industry allocated to final and intermediate input use from input-output tables (see Section 5.3). This means that the relative contribution of each HS-level price change to intermediate and final goods price changes is implicitly determined by the use shares of the corresponding ISIC sector, allowing us to avoid a sharp binary classification of an HS code as either a final or intermediate product.

A.2. **Import Demand Equation for Quality Estimation.** Recall that we defined the HS 6-digit level intermediate input bundle as:

$$i_{hM} = \left(\sum\nolimits_{o=1}^{O} \sum\nolimits_{z=1}^{n_{hM}^{o}} \left(q_{hM}^{o}\left(z\right) i_{hM}^{o}\left(z\right)\right)^{\left(\varepsilon_{hM}-1\right)/\varepsilon_{hM}}\right)^{\frac{\varepsilon_{hM}}{\varepsilon_{hM}-1}}.$$

If varieties are identical within a product-origin combination, this simplifies to:

$$i_{hM} = \left(\sum\nolimits_{o=1}^{O} n_{hM}^{o}\left(q_{hM}^{o} i_{hM}^{o}\right)^{\left(\varepsilon_{hM}-1\right)/\varepsilon_{hM}}\right)^{\frac{\varepsilon_{hM}}{\varepsilon_{hM}-1}}.$$

Hence, import demand given expenditure on h and the associated CES price index is given by:

$$X_{hM}^{o} = n_{hM}^{o}\left(p_{hM}^{o}/q_{hM}^{o}\right)^{1-\varepsilon_{hM}} p_{hM}^{\varepsilon_{hM}-1} E_{hM},$$

where E_{hM} denotes total expenditure on imported intermediate input h, and p_{hM} is the CES price index. This is the exact analogue of the import demand equation in Section 3.2 of the paper, with the demand side elasticity parameter σ_h replaced by the supply side elasticity ε_{hM}. Hence, all the remaining steps in Section 3.2 carry through for both final and intermediate products.

Appendix B. Estimating Elasticities

The estimation strategy follows Feenstra (1994). The import demand equation for each variety of product h can be expressed in terms of shares and changes over time:

$$\Delta \ln s^o_{ht} = \varphi_{ht} - (\sigma_h - 1)\Delta \ln p^o_{ht} + \Delta\epsilon^o_{ht}, \tag{B.3}$$

where: $\varphi_{ht} = (\sigma_h - 1)\ln p_{ht}$ and $\Delta\epsilon^o_{ht}$ is treated as an unobservable random variable, reflecting changes in the number of varieties and quality.

The dependent variable $\Delta \ln s^o_{ht}$ and the regressor $\Delta \ln p^o_{ht}$ might be correlated with the error term due to the simultaneous determination of import prices and quantities. So equation (B.3) cannot be directly estimated and some assumptions on the supply side of the economy have to be made. Simultaneity bias is corrected by allowing the supply of variety o to vary with the amount of exports, the export supply equation is defined as follows:

$$\Delta \ln p^o_{ht} = \psi_{ht} + \frac{\omega_h}{1+\omega_h}\Delta \ln s^o_{ht} + \Delta\delta^o_{ht}, \tag{B.4}$$

where: $\psi_{ht} = \omega_h \Delta \ln w_{sMht}/(1+\omega_h)$ (w_{sMht} is total expenditures on product h), ω_h is the inverse supply elasticity (assumed to be the same across countries) and $\Delta\delta^o_{ht}$

The identification strategy relies on the following assumption:

$$E(\Delta\epsilon^o_{ht}\Delta\delta^o_{ht}) = 0. \tag{B.5}$$

This implies that shocks to demand and supply at the variety level are uncorrelated. It is convenient to eliminate φ_{ht} and ψ_{ht} by choosing a reference country k and differencing demand and supply equations, denoted in (B.3) and (B.4), relative to country k.

$$\Delta^k \ln s^o_{ht} = -(\sigma_h - 1)\Delta^k \ln p^o_{ht} + \Delta^k \epsilon^o_{ht} \tag{B.6}$$

$$\Delta^k \ln p^o_{ht} = \frac{\rho_h}{(\sigma_h - 1)(1-\rho_h)}\Delta^k \ln s^o_{ht} + \Delta^k \delta^o_{ht} \tag{B.7}$$

where $\Delta^k x^o_{ht} = \Delta x^o_{ht} - \Delta x^k_{ht}$, $\rho_h = \omega_h\left(\sigma_h - 1\right)/\left(1 + \omega_h\sigma_h\right)$ and satisfies $0 \leq \rho_h \leq \left(\sigma_h - 1\right)/\sigma_h < 1$. In order to take advantage of the identification strategy equation (B.6) and (B.7) are then multiplied together to obtain:

$$(\Delta^k \ln p^o_{ht})^2 = \theta_1(\Delta^k \ln s^o_{ht})^2 + \theta_2(\Delta^k \ln p^o_{ht}\Delta^k \ln s^o_{ht}) + u^o_{ht}, \tag{B.8}$$

where $\sigma_h = f(\theta_1, \theta_2)$ as shown in the following Proposition.

Proposition 1. *So long as $\theta_1 > 0$, then σ_h and ρ_h are defined as follows:*

$$\rho_h = \frac{1}{2} + \left(\frac{1}{4} - \frac{1}{4 + (\theta_2^2/\theta_1)}\right)^{1/2} \quad \textit{if } \theta_2 > 0$$

$$\rho_h = \frac{1}{2} - \left(\frac{1}{4} - \frac{1}{4 + (\theta_2^2/\theta_1)}\right)^{1/2} \quad \textit{if } \theta_2 < 0$$

$$\sigma_h = 1 + \left(\frac{2\rho_h - 1}{1 - \rho_h}\right)\frac{1}{\theta_2} \quad \textit{in both cases}$$

If $\theta_1 < 0$, but $\theta_1 > -\theta_2^2/4$, it is still possible to obtain a value for σ_h exceeding unity but $\rho \notin [0, 1]$.

It is still not possible to consistently estimate equation (B.8) because prices and expenditure shares are correlated with the error term. Nevertheless, it is possible to obtain a consistent estimator for the thetas and hence for the elasticity of substitution by averaging (B.8) over time. The estimation is still possible because σ_h and the supply elasticity are assumed to be constant over the varieties of the same product; the former due to the CES demand structure, the latter for the particular form of the supply curve, whose elasticity is assumed to be equal across all supplying countries.

Hence, taking the sample means of the variables, equation (B.8) can be rewritten as:

$$\text{(B.9)} \qquad \overline{(\Delta^k \ln p_{ht}^o)^2} = \theta_1 \overline{(\Delta^k \ln s_{ht}^o)^2} + \theta_2 \overline{(\Delta^k \ln p_{ht}^o \Delta^k \ln s_{ht}^o)} + \overline{u_{ht}^o}.$$

From the assumption that underlines the identification strategy, $E\left[\overline{u_{ht}^o}\right] = 0$.[52] This implies that the expectation of the error term in (B.9) converges to zero and the equation can be consistently estimated. Let $\widehat{\theta}_1$ and $\widehat{\theta}_2$ denote the estimates of θ_1 and θ_2 obtained by running weighted least squares (WLS) on (B.9), it turns out that these are equivalent to the Hansen's (1982) GMM estimator defined as follows:

$$\text{(B.10)} \qquad \widehat{\beta}_{GMM} = \arg\min_{\theta \in \Theta} \overline{u_{ht}^o}(\beta)' W \overline{u_{ht}^o}(\beta),$$

where $\beta = \binom{\sigma_h}{\rho_h}$, and W is a positive definite weighting matrix. In order to get a consistent estimate for σ_h, first $\widehat{\theta}_1$ and $\widehat{\theta}_2$ are obtained by running WLS on (B.9), then σ_h is computed using Proposition 1. Whenever an infeasible value for σ_h is obtained (<1), a constrained numerical minimization of equation (B.10) is performed using the Nelder and Mead's (1965) simplex algorithm.[53]

[52]Supply and demand error terms are assumed to be independent, see equation (B.5). A further condition to get identification requires to have some difference in the relative variances of the two error terms. See Feenstra (1994).

[53]The variables are constrained as follows: $1 < \sigma_h \leq 135.5$ and $0 \leq \rho_h < 1$. Once a solution is obtained the non-linear condition $\rho_h \leq (\sigma_h - 1)/\sigma_h$ is checked. If the condition is not satisfied, it implies that a bigger value of σ_h would be needed. In the rare event that this happens (on average less than 0.1% of all estimates), σ_h

A direct measure of prices is not available, so p^o_{ht} is calculated as a unit value. This implies that prices are measured with some error. In order to mitigate this problem, Feenstra (1994) suggests adding a constant to equation (B.9) in order to capture the variance of the measurement error. Broda and Weinstein (2006) refine the method by imposing some structure to the form of the error, which also affects the strategy for the form of the weighting matrix in the WLS estimator.

Let p^{oi}_{ht} be the price of a particular product of variety o of product h; so that the trade value $p^o_{ht}x^o_{ht} = \sum_i p^{oi}_{ht}$ because the quantity of each product, x^{oi}_{ht}, always equals one (i.e.: in case of more items of the same product, the same price is added several times). They assume that product prices are measured with an i.i.d. error such that $p^{oi}_{ht} = \widetilde{p}^{oi}_{ht}\zeta^{oi}_{ht}$ where $\widetilde{p}^{oi}_{ht}$ is the true price and p^{oi}_{ht} is the measured price. In this case the error has mean zero and:

$$\begin{aligned} var(\ln \zeta^{oi}_{ht}) &= \sigma^2 \\ cov\left(\ln \zeta^{oi}_{ht}, \ln \zeta^{o'j}_{hs}\right) &= 0 \ \ \forall\ o \neq o',\ t \neq s,\ i \neq j \end{aligned}$$

By assuming that the log of the geometric mean price of a variety is approximately equal to the log of the arithmetic mean, it is possible to compute the variance of $\ln p^o_{ht}$ as follows:

$$\begin{aligned} \sigma^2_{\ln p^o_{ht}} \equiv var\left[\ln\left(\frac{\sum_i p^{oi}_{ht}}{x^o_{ht}}\right)\right] &\approx var\left[\ln\left(\left(\prod_i p^{oi}_{ht}\right)^{1/x^o_{ht}}\right)\right] = \\ = \frac{1}{(x^o_{ht})^2} var\left(\sum_i \left(\ln \widetilde{p}^{oi}_{ht} + \ln \zeta^{oi}_{ht}\right)\right) &= \frac{1}{(x^o_{ht})^2} x^o_{ht}\sigma^2 = \frac{1}{x^o_{ht}}\sigma^2. \end{aligned}$$

Thanks to the assumptions on structure of the error, we obtain:

$$\begin{aligned} E(\Delta^k \ln p^o_{ht})^2 &= E\left((\ln p^o_{ht} - \ln p^o_{ht-1}) - (\ln p^k_{ht} - \ln p^k_{ht-1})\right)^2 = \\ &= \delta^2_{ho_kt} + \sigma^2\left(\frac{1}{x^o_{ht}} + \frac{1}{x^o_{ht-1}} + \frac{1}{x^k_{ht}} + \frac{1}{x^k_{ht-1}}\right), \end{aligned}$$

where $\delta^2_{ho_kt}$ is the variance of the true price differences over time and with respect to variety k. Averaging this across all periods:

$$E\frac{1}{T}\sum_t(\Delta^k \ln p^o_{ht})^2 = \frac{1}{T}\sum_t \delta^2_{ho_kt} + \sigma^2\frac{1}{T}\sum_t\left(\frac{1}{x^o_{ht}} + \frac{1}{x^o_{ht-1}} + \frac{1}{x^k_{ht}} + \frac{1}{x^k_{ht-1}}\right).$$

is assumed to be equal to 140. All results are very robust to these assumptions, the choice of a max value of sigma equal to 140 has no impact on the final result.

This implies that the equation (B.9) should be modified by adding the following error adjustment term to the right hand side:

$$err\_adj = \widehat{\theta}_3 \frac{1}{T} \sum_t \left(\frac{1}{x^o_{ht}} + \frac{1}{x^o_{ht-1}} + \frac{1}{x^k_{ht}} + \frac{1}{x^k_{ht-1}} \right), \tag{B.11}$$

where $\widehat{\theta}_3 = \widehat{\sigma^2}$ is a parameter to be estimated. This equation generalises Broda and Weinstein's (2006) approach, which which does not take into account the first difference with respect to the reference country k. Since prices of variety k might also be measured with error, our estimates are likely to be more robust to measurement error if the choice of the reference country is accounted for.

Broda and Weinstein (2006) use a similar line of reasoning for the weighting matrix of their WLS estimator. Heteroskedasticity is likely to be present because if prices are measured with error, so are their sample variances. They correct for this heteroskedasticity by assuming that the sample variances are inversely related to the quantity of goods (used to calculate unit values) and to the number of periods. We follow the same strategy, and define the variance in order to account for the first differencing with respect to the reference country k. Hence the weights are given by:

$$weight = T^{3/2} \left(\frac{1}{x^o_{ht}} + \frac{1}{x^o_{ht-1}} + \frac{1}{x^k_{ht}} + \frac{1}{x^k_{ht-1}} \right)^{-1/2}. \tag{B.12}$$

The elasticities will be estimated both following the Broda and Weinstein's (2006) baseline estimation strategy and using the modified error adjustment term and weighting scheme.

References

Amiti, Mary and Amit K. Khandelwal, "Import competition and quality upgrading," *Review of Economics and Statistics*, 2013, *95* (2), 476–490.

Baccini, Leonardo, "Explaining formation and design of EU trade agreements: The role of transparency and flexibility," *European Union Politics*, 2010, *11* (2), 195–217.

Baier, Scott L. and Jeffrey H. Bergstrand, "Do free trade agreements actually increase members' international trade?," *Journal of international Economics*, 2007, *71* (1), 72–95.

______ and ______, "Estimating the effects of free trade agreements on international trade flows using matching econometrics," *Journal of international Economics*, 2009, *77* (1), 63–76.

______, ______, and Michael Feng, "Economic integration agreements and the margins of international trade," *Journal of International Economics*, 2014, *93* (2), 339–350.

Baldwin, Richard and James Harrigan, "Zeros, quality, and space: Trade theory and trade evidence," *American Economic Journal: Microeconomics*, 2011, *3* (2), 60–88.

Bas, Maria and Vanessa Strauss-Kahn, "Input-trade liberalization, export prices and quality upgrading," *Journal of International Economics*, 2015, *95* (2), 250–262.

Bernard, Andrew B. and Swati Dhingra, "Contracting and the Division of the Gains from Trade," Technical Report, Working paper 2015.

Blaum, Joaquin, Claire LeLarge, and Michael Peters, "The Gains from Input Trade in Firm-Based Models of Importing," *NBER Working Paper*, aug 2015, *21504.*

Blonigen, Bruce A. and Anson Soderbery, "Measuring the benefits of foreign product variety with an accurate variety set," *Journal of International Economics*, 2010, *82* (2), 168–180.

Brenton, Paul and Miriam Manchin, "Making EU trade agreements work: the role of rules of origin," *The World Economy*, 2003, *26* (5), 755–769.

Broda, Christian and David E. Weinstein, "Globalization and the Gains from Variety*," *Quarterly Journal of Economics*, 2006, *121* (2), 541–585.

Caliendo, Lorenzo and Fernando Parro, "Estimates of the Trade and Welfare Effects of NAFTA," *Review of Economic Studies*, 2015, *82* (1), 1–44.

DeBardeleben, Joan, "The impact of EU enlargement on the EU-Russian relationship," *A Resurgent Russia and the West: The European Union, NATO, and Beyond. Dordrecht, The Netherlands: Republic of Letters Publishing*, 2009, pp. 93–112.

Demir, F. Banu, *Trading tasks and quality*, Department of Economics, University of Oxford, 2011.

Dhingra, Swati, "Trading away wide brands for cheap brands," *The American Economic Review*, 2013, *103* (6), 2554–2584.

——— **and John M. Morrow**, "Monopolistic Competition and Optimum Product Diversity under Firm Heterogeneity," *Working Paper*, nov 2012.

Diewert, W. Erwin, "Exact and superlative index numbers," *Journal of econometrics*, 1976, *4* (2), 115–145.

Egger, Hartmut, Peter Egger, and David Greenaway, "The trade structure effects of endogenous regional trade agreements," *Journal of international Economics*, 2008, *74* (2), 278–298.

Fan, Haichao, Yao A. Li, and Stephen R. Yeaple, "Trade liberalization, quality, and export prices," *Review of Economics and Statistics*, 2015, *97* (5), 1033–1051.

Feenstra, Robert C., "New product varieties and the measurement of international prices," *American Economic Review*, 1994, *84*, 157.

——— **and David E. Weinstein**, "Globalization, markups, and the US price level," *NBER Working Paper*, 2010.

——— **and John Romalis**, "International prices and endogenous quality," Technical Report, National Bureau of Economic Research 2012.

Hallak, Juan C., "A product-quality view of the linder hypothesis," *The Review of Economics and Statistics*, 2010, *92* (3), 453–466.

——— **and Peter K. Schott**, "Estimating Cross-Country Differences in Product Quality*," *The Quarterly journal of economics*, 2011, *126* (1), 417–474.

Handley, Kyle and Nuno Limão, "Trade and Investment under Policy Uncertainty: Theory and Firm Evidence," NBER Working Papers 17790, National Bureau of Economic Research, Inc Jan 2012.

Hansen, Lars Peter, "Large Sample Properties of Generalized Method of Moments Estimators," *Econometrica*, 1982, *50* (4), 1029–1054.

Harrigan, James and Geoffrey Barrows, "Testing the theory of trade policy: Evidence from the Abrupt End of the Multifiber Arrangement," *The Review of Economics and Statistics*, 2009, *91* (2), 282–294.

Head, Keith and Thierry Mayer, "Gravity Equations: Workhorse, Toolkit, and Cookbook," *Handbook of International Economics*, 2014, *4*, 131.

Hsieh, Chang-Tai, Nicholas Li, Ralph Ossa, and Mu-Jeung Yang, "Accounting for the New Gains from Trade Liberalization," Technical Report, National Bureau of Economic Research 2016.

Hummels, David and Peter J. Klenow, "The Variety and Quality of a Nation's Exports," *American Economic Review*, 2005, *95* (3), 704–23.

Iacovone, Leonardo and Beata S. Javorcik, "Multi-Product Exporters: Product Churning, Uncertainty and Export Discoveries," *The Economic Journal*, 2010, *120* (544), 481–499.

Johnson, Robert C., "Trade and prices with heterogeneous firms," *Journal of International Economics*, 2012, *86* (1), 43–56.

Khandelwal, Amit, "The long and short (of) quality ladders," *The Review of Economic Studies*, 2010, *77* (4), 1450–1476.

Khandelwal, Amit K., Peter K. Schott, and Shang-Jin Wei, "Trade Liberalization and Embedded Institutional Reform: Evidence from Chinese Exporters," *The American Economic Review*, 2013, *103* (6), 2169–2195.

Krugman, Paul R., "Scale Economies, Product Differentiation, and the Pattern of Trade," *American Economic Review*, 1980, *70* (5), 950–959.

Nelder, John A. and Roger Mead, "A Simplex Method for Function Minimization," *The Computer Journal*, 1965, *7* (4), 308–313.

OECD, *COICOP-CPC and CPC-COICOP correspondence tables*, Paris: Organisation for Economic Co-operation and Development, 2001.

Pelkmans, Jacques and Paul Brenton, "Bilateral trade agreements with the EU: Driving forces and effects," in "Multilateralism and Regionalism in the Post-Uruguay Round Era," Springer, 1999, pp. 87–121.

Sato, Kazuo, "The ideal log-change index number," *The Review of Economics and Statistics*, 1976, pp. 223–228.

Schott, Peter K., "Across-product Versus Within-product Specialization in International Trade," *The Quarterly Journal of Economics*, 2004, *119* (2), 647–678.

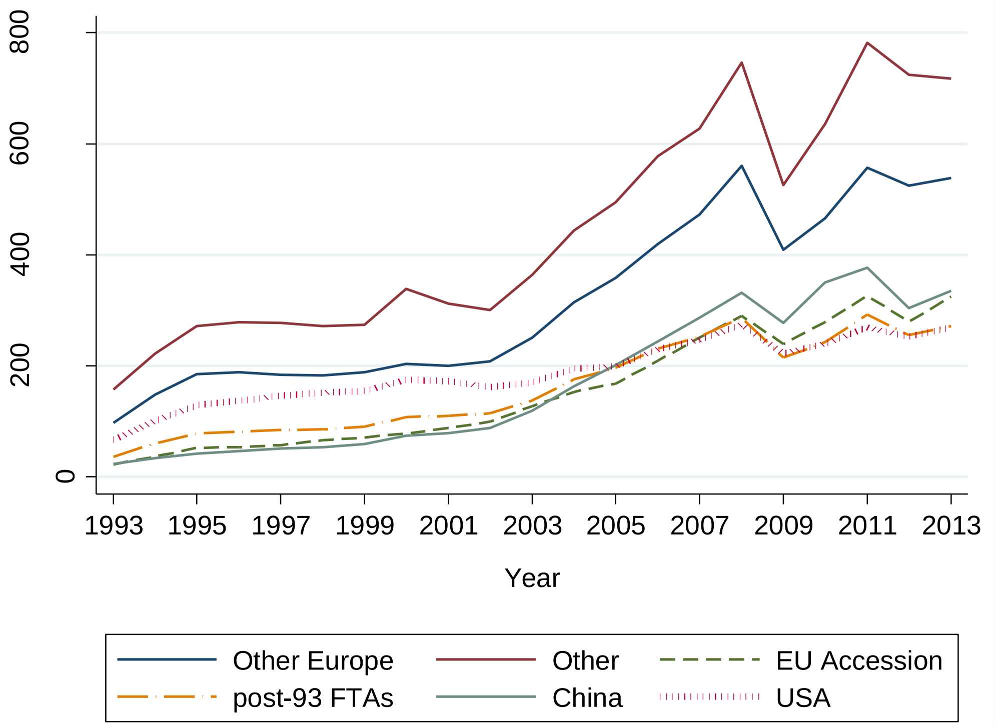
Sheu, Gloria, "Price, quality, and variety: Measuring the gains from trade in differentiated products," *American Economic Journal: Applied Economics*, 2014, *6* (4), 66–89.

Trefler, Daniel, "The Long and Short of the Canada-US Free Trade Agreement," *The American Economic Review*, 2004, *94* (4), 870–895.

Verhoogen, Eric A., "Trade, Quality Upgrading, and Wage Inequality in the Mexican Manufacturing Sector*," *The Quarterly Journal of Economics*, 2008, *123* (2), 489–530.

Woolcock, Stephen, "European Union policy towards free trade agreements," *EUPE, Working Brussels Papers, Working Brussels Papers-3/2007. Disponible sur: www. ecipe. org/pdf/EWP-3-2007.pdf*, 2007, *3/2007.*

Figure 1: EU12 Imports by Origin, 1993-2013



Notes: Figures shows EU12 imports by group of origin countries. "EU Accession" are the 2004/2007 accession countries, "Other Europe" are Russia, Switzerland, Norway, Ukraine, Belarus, Austria, Sweden, Finland and Iceland, "post-93 FTAs" are the non-accession countries with which the EU has signed FTAs after 1993 (also see Table 1) and "Other" are all other countries.

Table 1: List of EU Free Trade Agreements, 1993-2013

Partner country	Year of Entry into Force	EU Accession Date	Partner country	Year of Entry into Force	EU Accession Date
Cyprus	1973	2004	Israel	2000	--
Malta	1971	2004	Algeria	2006	--
Czech Republic	1995	2004	Egypt	2004	--
Slovakia	1995	2004	Jordan	2002	--
Estonia	1998	2004	Lebanon	2003	--
Latvia	1998	2004	Morocco	2000	--
Lithuania	1998	2004	Tunisia	1998	--
Hungary	1994	2004	Mexico	2000	--
Poland	1994	2004	South Africa	2000	--
Slovenia	1995	2004	Chile	2003	--
Bulgaria	1995	2007	South Korea	2011	--
Romania	1995	2007	Peru	2013	--
Turkey	1996	--	Columbia	2013	--
Faroe Islands	1997	--	Costa Rica	2013	--
Macedonia	2004	--	El Salvador	2013	--
Croatia	2005	2013	Guatemala	2013	--
Albania	2009	--	Honduras	2013	--
Bosnia and Herzegovina	2009	--	Nicaragua	2013	--
Montenegro	2010	--	Panama	2013	--
Serbia	2010	--			

Table 2: The Impact of FTAs on Prices, Quality and Variety: EU12, 1993-2013

	(1)	(2)	(3)	(4)	(5)
	$\ln(price_{oht})$ - $\ln(price_{oht-5})$	$\ln(quality_{oht})$ - $\ln(quality_{oht-5})$	$\ln(quality_{oht})$ - $\ln(quality_{oht-5})$	$\ln(variety_{ht})$	$\ln(adj.price_{oht})$ - $\ln(adj.\ price_{oht-5})$
FTA_{ot}- FTA_{ot-5}	0.0002	0.0674***	0.0884***	0.00098	-0.0666***
	(0.0127)	(0.0152)	(0.0193)	(0.00124)	(0.0185)
N	1,613,652	1,613,652	1,613,652	1,613,652	1,613,652
R^2	0.009	0.001	0.001	0.011	0.004
Year FE, α_t	Yes	Yes	Yes	Yes	Yes
Elasticity Estimates	--	Section 3	BW Method	--	Section 3

Notes: Table shows results from OLS regressions (robust standard errors in brackets, clustered at the origin country-year level). Quality in columns (2) and (3) is estimated using the elasticity estimates from Section 3 and estimates based on Broda and Weinstein's (2008) methodology, respectively; both quality measures are adjusted for hidden varieties. Variety in column (4) refers to $\lambda_{jt}/ \lambda_{jt-1}$ from Section 2. Column (5) uses quality- and hidden-variety adjusted prices. FTA_t - FTA_{t-5} takes a value of 1 when there is an FTA in force between the EU12 and the trade partner in period t but not in period t-5, and 0 otherwise. The RHS contains year fixed effects throughout. ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively.

Table 3: GDP Controls

	(1)	(2)	(3)
	$\ln(price_{oht})$-$\ln(price_{oht-5})$	$\ln(quality_{oht})$- $\ln(quality_{oht-5})$	$\ln(variety_{ht})$
FTA_{ot}- FTA_{ot-5}	-0.0048	0.0644***	0.0009
	(0.0119)	(0.0145)	(0.0013)
$\ln(pcGDP_{ot})$ – $\ln(pcGDP_{ot-5})$	0.1784***	0.1065***	0.0022**
	(0.0152)	(0.0191)	(0.0011)
N	1,613,388	1,613,388	1,613,388
R^2	0.010	0.001	0.011
Year FE, α_t	Yes	Yes	Yes

Notes: Table shows results from OLS regressions (robust standard errors in brackets, clustered at the origin country-year level). See Table 2 and text for details. ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively.

Table 4: Different Functional Forms

	(1)	(2)	(3)	(4)
	Two-Year Differences, $\ln(y_{oht})-\ln(y_{oht-2})$	OLS with Product-Time Fixed Effects, $\ln(y_{oht})$	Single Two-Year Difference, $\ln(y_{oht})-\ln(y_{oht-2})$	Single Five-Year Difference, $\ln(y_{oht})-\ln(y_{oht-5})$
FTA price effect (SE)	0.000985	0.0524***	0.00209	0.0198
	(0.0131)	(0.0165)	(0.0173)	(0.0222)
FTA quality effect (SE)	0.0433***	0.0839***	0.0574***	0.107***
	-0.0136	-0.0213	(0.0168)	(0.0288)
FTA variety effect (SE)	-0.00125	--	-0.00233	-0.00490*
	(0.00111)		(0.00144)	(0.00251)
Time FE, α_t	Yes	Yes	Yes	Yes
Origin-product FE, α_{oh}	No	Yes	No	No
Definition FTA dummy	FTA_t- FTA_{t-2}	FTA_t	FTA_t- FTA_{t-2}	FTA_t- FTA_{t-5}

Notes: Table shows results for OLS regressions of the variable in the first column (y = price, quality, variety) on the FTA dummy in levels or changes. See the last row for the definition of the FTA dummy in each regression. The top row indicates the functional form of the dependent variable in each regression. Robust standard errors are reported in brackets below the coefficient estimates, clustered at the origin country-year level throughout. See Table 2 and text for details. ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively.

Table 5: Matched Control Group

	(1)	(2)	(3)
	Five-Year Overlapping Differences, $\ln(y_{oht})$-$\ln(y_{oht-5})$	Five-Year Overlapping Differences, $\ln(y_{oht})$-$\ln(y_{oht-5})$	Five-Year Overlapping Differences, $\ln(y_{oht})$-$\ln(y_{oht-5})$
FTA price effect (SE)	0.0352	0.0263	0.0248
	(0.0095)***	(0.0100)***	(0.0133)*
FTA quality effect (SE)	0.1037	0.1092	0.1363
	(0.0177)***	(0.0264)***	(0.0483)***
FTA variety effect (SE)	--	--	--
Definition FTA dummy	FTA_t- FTA_{t-5}	FTA_t- FTA_{t-5}	FTA_t- FTA_{t-5}
Matching Radius	0.20	0.10	0.05

Notes: Table shows propensity score matching estimates for the effect of FTAs on the variables in the first column (y = price, quality, variety). Propensity scores are obtained from a logit regression of changes in the FTA dummy, FTA_t - FTA_{t-5}, on standard gravity variables (distance, contiguity, GDP). We use radius matching to choose a control group, imposing that matches need to be from the same product-year bin as the treated observation. Columns (1) - (3) report estimates for different choices of matching radius. Standard errors are obtained via cluster-bootstrapping (clusters at the origin country-year level). See text for details. ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively.

Table 6: Controlling for Hidden Varieties, Alternative Approaches

	(1)	(2)	(3)	(4)	(5)
	$\ln(qual_{oht})$ - $\ln(qual_{oht-5})$	$\ln(qual_{oht})$ - $\ln(qual_{oht-5})$	$\ln(qual_{oht})$ - $\ln(qual_{oht-5})$	$\ln(qual_{oht})$ - $\ln(qual_{oht-5})$	$\ln(qual_{oht})$ - $\ln(qual_{oht-5})$
FTA_{ot}- FTA_{ot-5}	0.0774***	0.102***	0.0983***	0.101***	0.138*
	(0.0178)	(0.0197)	(0.0195)	(0.0192)	(0.0799)
N	1,613,652	1,613,652	1,613,652	1,613,652	782,930
R^2	0.001	0.002	0.001	0.001	0.005
Time FE, α_t	Yes	Yes	Yes	Yes	Yes
Proxy for Hidden Varieties	None	Origin FE	Origin x 1-digit-HS FE	Origin x 3-digit-HS FE	$\ln(\# \text{ firms}_{oht})$

Notes: Table shows results from OLS regressions of the quality measures in the top row on the first-differenced FTA dummy, FTA_t - FTA_{t-5} (robust standard errors in brackets, clustered at the origin country-year level). The last row indicates how the quality measures are adjusted for "hidden" varieties. See text and Table 2 for details. ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively.

Table 7: Controlling for EC Status

	(1)	(2)	(3)
	$\ln(price_{oht})$-$\ln(price_{oht-5})$	$\ln(quality_{oht})$-$\ln(quality_{oht-5})$	$\ln(variety_{ht})$
FTA_{ot} - FTA_{ot-5}	0.00882	0.0712***	0.000862
	(0.0129)	(0.0153)	(0.00124)
EU_{ot} - EU_{ot-5}	0.120***	0.0517**	-0.00166
	(0.0147)	(0.0233)	(0.00124)
N	1,613,652	1,613,652	1,613,652
R^2	0.009	0.001	0.011
Time FE, α_t	yes	yes	yes

Notes: Table shows results from OLS regressions of the measures in the top row on the first-differenced FTA dummy, FTA_t - FTA_{t-5}, and the first-differenced EU dummy, EU_t - EU_{t-5} (robust standard errors in brackets, clustered at the origin country-year level). See text and Table 2 for details. ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively.

Table 8: The Effect of Tariffs

	(1)	(2)	(3)
	$\ln(price_{oht})$- $\ln(price_{oht-5})$	$\ln(quality_{oht})$- $\ln(quality_{oht-5})$	$\ln(variety_{ht})$
FTA_{ot} - FTA_{ot-5}	0.0171	0.0774***	0.0006
	(0.0136)	(0.0148)	(0.0012)
$Tariff_{oht}$ – $Tariff_{oht-5}$	0.00112*	-0.00031	0.0005***
	(0.00066)	(0.00084	(0.0001)
N	1,177,343	1,177,343	1,177,343
R^2	0.008	0.001	0.012
Time FE, αt	yes	yes	yes

Notes: Table shows results from OLS regressions of the measures in the top row on a first-differenced FTA dummy and changes in import tariffs (robust standard errors in brackets, clustered at the origin country-year level). $Tariff_t$ - $Tariff_{t-5}$ is the percentage point change in effectively applied EU import tariffs at the HS 6-digit level between periods t-5 and t. See text and Table 2 for details. ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively.

Table 9: Pre-Trend Regressions

	(1)	(2)	(3)
	$\ln(price_{oht})$-$\ln(price_{oht-5})$	$\ln(quality_{oht})$-$\ln(quality_{oht-5})$	$\ln(variety_{ht})$
FTA_{ot+5}- FTA_{ot}	-0.00913	0.0101	0.00155
	(0.0138)	(0.0149)	(0.000986)
N	1,613,652	1,613,652	2,295,129
R^2	0.009	0.001	0.011
Time FE, αt	yes	yes	yes

Notes: Table shows results from OLS regressions of the measures in the top row on a lead indicator for free trade agreements (robust standard errors in brackets, clustered at the origin country-year level). FTA_{ot+5}- FTA_{ot} takes a value of one if origin country *o* implemented a free trade agreement within five years from period *t*, and zero otherwise. See text and Table 2 for details. ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively.

Table 10: Domestic Varieties

	(1)	(2)
	$\ln(variety_{st})$	$\ln(variety_{st})$
$Exposure_{st}$	4.844	-1.642
	(4.107)	(2.450)
N	2,268	2,268
R^2	0.051	0.955
Year Fixed Effects	yes	yes
Sector Fixed Effects	no	yes

Notes: Table shows results from OLS regressions of the number of varieties on a measure of exposure to import competition (robust standard errors in brackets, clustered at the 4-digit ISIC level). See text for details and variable definitions. ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively.

Table 11: Heterogeneity – EU and Partner Country Income Groups

	(1)	(2)	(3)
	$\ln(price_{oht}) - \ln(price_{oht-5})$	$\ln(quality_{oht}) - \ln(quality_{oht-5})$	$\ln(variety_{ht})$
Panel A: EU12 Income/Geographic Groups			
EU12 Low/South	-0.0556***	0.00303	0.00650**
	(0.00866)	(0.0127)	(0.00255)
EU12 Middle/Central	-0.00676	0.0276***	0.00254**
	(0.00515)	(0.00712)	(0.00120)
EU12 High/North	-0.0175**	0.131***	0.00372**
	(0.00755)	(0.0123)	(0.00188)
Panel B: Partner Countries by Income Groups			
Low and Lower-Middle Income FTA Partners	-0.0259	0.0917***	0.00349**
	(0.0180)	(0.0195)	(0.00165)
Upper-Middle Income FTA Partners	0.0563***	0.109***	-0.00181
	(0.0159)	(0.0239)	(0.00231)
High Income FTA Partners	-0.00550	0.00890	-0.00142
	(0.0218)	(0.0317)	(0.00156)
Panel C: Partner Countries by Agreement Type			
Future EU members	0.101***	0.193***	-0.00360
	(0.0137)	(0.0316)	(0.00288)
Balkans	0.0299	0.172***	0.0153***
	(0.0218)	(0.0386)	(0.00374)
Euro Med	-0.0709***	0.0325	0.00203
	(0.0184)	(0.0214)	(0.00178)
Non-Regional	0.0362**	0.0572***	0.0832***
	(0.0180)	(0.0202)	(0.0314)
Year FE, α_t	Yes	Yes	Yes

Notes: Table shows results from OLS regressions of the measures in the top row on the first-differenced FTA dummy and year fixed effects for the subsamples indicated in the first column (robust standard errors in brackets, clustered at the origin country-year level). See text and Table 2 for details.

Table 12: Heterogeneity – Results by ISIC Groups

ISIC	ISIC Group Name	Prices		Quality		Variety		No. Obs.
		Coeff.	SE	Coeff.	SE	Coeff.	SE	
01_05	Products of agriculture, forestry, fisheries and aquaculture	0.0129	(0.0174)	0.0040	(0.0294)	0.0045	(0.0035)	93456
10_14	Mining products	0.0172	(0.0416)	-0.0026	(0.0876)	0.0016	(0.0101)	21060
15_16	Food and Tobacco	0.0150	(0.0139)	-0.0284	(0.0245)	0.0067*	(0.0038)	126722
17_19	Textiles	0.0199	(0.0148)	0.1079***	(0.0272)	0.0020	(0.0015)	368126
20_22	Wood, paper, printing	0.0036	(0.0275)	0.0981***	(0.0286)	-0.0021	(0.0027)	86410
23_24	Coke and chemicals	0.0134	(0.0193)	-0.0254	(0.0363)	-0.0040*	(0.0024)	200120
25_26	Rubber and plastic	0.0128	(0.0191)	0.0655**	(0.0280)	-0.0033	(0.0020)	124954
27_28	Metal products	-0.0021	(0.0163)	0.0241	(0.0242)	0.0030	(0.0030)	191304
29	Machinery and equipment n.e.c.	-0.0911**	(0.0377)	0.1573**	(0.0710)	0.0002	(0.0024)	139626
30_33	Machinery, equipment, television, instruments	0.0398	(0.0248)	0.1942***	(0.0667)	-0.0015	(0.0011)	138363
34_35	Motor vehicles, transport equipment	-0.0059	(0.0521)	0.1265***	(0.0457)	0.0079	(0.0076)	43277
36_37	Furniture; other manufactured goods n.e.c.	-0.0117	(0.0324)	0.0116	(0.0360)	0.0053**	(0.0027)	47660
40_99	Services	-0.0496	(0.1096)	0.1044	(0.1961)	0.0207	(0.0146)	3103

Notes: Table shows coefficient estimates and standard errors from regressions of prices, quality and variety indicators on the first-differenced FTA indicator and year fixed effects carried out separately for each ISIC group listed in the first column (robust standard errors in brackets, clustered at the origin country-year level). See text and Table 2 for details. ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively.

Table 13: Heterogeneity – Interaction Regressions

	(1) dln(price)	(2) dln(price)	(3) dln(price)	(4) dln(quality)	(5) dln(quality)	(6) dln(quality)
dFTA	0.0202	0.220	-0.00274	0.0125	-0.355	-1.003***
	(0.0241)	(0.166)	(0.234)	(0.0311)	(0.259)	(0.292)
dFTAxHigh_EU	0.0235*		0.0188	0.0674***		0.0718***
	(0.0131)		(0.0134)	(0.0217)		(0.0204)
dFTAxMed_EU	0.0149		0.00996	0.0584***		0.0341*
	(0.00998)		(0.0105)	(0.0210)		(0.0202)
dFTAxH_Partner	0.0309		0.0550**	-0.0281		-0.0429
	(0.0248)		(0.0265)	(0.0314)		(0.0401)
dFTAxUM_Partner	0.0265		0.00682	-0.118***		-0.0869***
	(0.0265)		(0.0207)	(0.0325)		(0.0303)
dFTAxFutureEU	0.0407		0.0167	0.102***		0.150**
	(0.0283)		(0.0545)	(0.0342)		(0.0743)
dFTAxBalkans	0.0441		0.0268	0.0969***		0.179***
	(0.0388)		(0.0583)	(0.0372)		(0.0649)
dFTAxEuroMed	-0.0906***		-0.123***	-0.0601**		-0.0245
	(0.0249)		(0.0335)	(0.0294)		(0.0558)
dFTAxlog(distance)		5.76e-05	-0.00359		-0.0142	0.0348
		(0.0106)	(0.0202)		(0.0127)	(0.0291)
dFTAxCommon Language		-0.0578***	-0.0353**		-0.0608**	-0.0581*
		(0.0200)	(0.0173)		(0.0295)	(0.0318)
dFTAxlog(share)		-0.0128***	-0.0137***		0.0107	0.0117
		(0.00314)	(0.00311)		(0.00844)	(0.00898)
dFTAxlog(time)		0.0162	0.0475*		-0.0524**	-0.0403
		(0.0174)	(0.0263)		(0.0216)	(0.0273)
dFTAxlog(RCA Difference)		-0.0286**	-0.00974		0.0626***	0.0813***
		(0.0112)	(0.0106)		(0.0184)	(0.0161)
Observations	4,741,033	4,741,033	4,741,033	4,741,033	4,741,033	4,741,033
R-squared	0.021	0.021	0.022	0.001	0.001	0.001

Notes: Table shows results of OLS interaction regressions using exporter-importer-HS 6-digit level data. Each of the regressors shown in the first column is interacted with the FTA dummy. The regression also includes year fixed effects and interaction terms between the FTA dummy and ISIC group dummies. Robust standard errors are in brackets, clustered at the origin country-year level. The excluded interaction category in columns (1), (3), (4) and (6) is Low Income EU – Low, Lower Middle Partner, Non Regional FTAs, ISIC 01_05. See Table 2 and text for details. ***, ** and * denote statistical significance at the 1%, 5% and 10% level, respectively.

Table 14: Counterfactual Price Changes – Aggregate Estimates

Aggregation Level	Mean	SD	Min	Max
(1) HS 6d – Total	-0.35%	0.71%	-6.50%	0.30%
– Price	0.00%	0.00%	0.00%	0.02%
– Quality	-0.35%	0.72%	-6.52%	0.00%
– Variety	0.00%	0.01%	0.00%	0.31%
(2) ISIC import price indices	-0.42%	0.40%	-1.64%	-0.05%
(3) ISIC domestic price indices	-0.06%	0.05%	-0.17%	-0.01%
(4) ISIC indices	-0.14%	0.13%	-0.61%	-0.01%
(5) Cumulative Price Index Effect, 1993-2013 (95% CI)	-0.24% [-0.19%, -0.28%]			
(6) Cumulative Price Index Effect, 1993-2013, No Intermediates (95% CI)	-0.13% [-0.10%, -0.15%]			
(7) Cumulative Price Index Effect, 1993-2013 (95% CI), Int., BW Elasticities	-0.30% [-0.23%, -0.36%]			
(8) Cumulative Price Index Effect, 1993-2013 (95% CI), Int., Unido	-0.49% [-0.42%, -0.56%]			

Notes: Table shows descriptive statistics for FTA-induced five-year changes in the subcomponents of the consumer price index from equation (2.8). Rows 5-8 show the cumulative effect over the sample period on the aggregate price index for different scenarios and coefficient estimates. The confidence intervals of the cumulative effects are computed using bootstrapped standard errors. See text for details.

Table 15: Counterfactual Price Changes – Estimates Allowing for Heterogeneity

Type of Estimate Heterogeneity	Cumulative Price Index Effect, 1993-2013 (95% CI)	Cumulative Price Index Effect, 1993-2013, No Intermediates (95% CI)
(1) ISIC-Level Estimates (Table 12)	-0.27% [-0.07%, -0.47%]	-0.18% [-0.07%, -0.29%]
(2) EU Country Group Estimates (Table 11A)		
– Low (Greece, Spain, Portugal)	-0.12% [-0.05%,-0.18%]	-0.07% [-0.04%,-0.10%]
– Middle (Denmark, Germany, France, Italy)	-0.13% [-0.07%,-0.19%]	-0.08% [-0.05%,-0.11%]
– High (Ireland, UK, Netherlands)	-0.41% [-0.34%,-0.48%]	-0.27% [-0.23%,-0.31%]
(3) Interaction Regressions (Table 13)	-0.47% [-0.17%, -0.76%]	-0.30% [-0.13%, -0.47%]

Notes: Table shows the total cumulative effect on the consumer price index for the period 1993-2013. Figures in panels (1) and (3) are for the EU12 as a whole and figures in panel (2) for the country groups listed. All confidence intervals are computed using bootstrapped standard errors. See text for details.