FIW-Research Reports 2010/11 N° 07
September 2011

Structural Estimation of Gravity Models with Path-Dependent Market Entry

Peter Egger and Michael Pfaffermayr

Abstract

This paper develops a structural empirical general equilibrium model of aggregate bilateral trade with path dependence of country-pair level exporter status. Such path dependence is motivated through informational costs about serving a foreign market for first-time entry of (firms in) an export market versus continued export services to that market. We embed the theoretical model into a structural dynamic stochastic econometric model of bilateral selection into import markets and apply it to a data-set of aggregate bilateral exports among 120 countries over the period 1995-2004. In particular, we disentangle the role of changes in trade costs, in labor endowments, and in total factor productivity for trade, bilateral market entry, numbers of firms active, and welfare. Dynamic gains from trade differ significantly from static ones, and path-dependence in market entry cushions effects of impulses in fundamental variables that are detrimental to bilateral trade.

Keywords: Bilateral trade flows; Gravity equation; Dynamic random effects model; Sample selection

JEL-codes: F10; F12; F17

The FIW-Research Reports 2010/11 present the results of six thematic work packages “The financial and economic crisis of 2007-2010 and the European economy“, “Modelling the Effects of Trade Policy and the Transmission Mechanisms of the Economic Crisis on the Austrian Economy“, “The Gravity Equation“, “Macroeconomic Aspects of European Integration“, “Effects of International Integration on Income Distribution“ and “New Energy Policy and Security of Gas Supply“, that were announced by the Austrian Federal Ministry of Economics, Family and Youth (BMWFJ) within the framework of the “Research Centre International Economics” (FIW) in January 2010.

Structural Estimation of Gravity Models with Path-Dependent Market Entry

Peter Egger* and Michael Pfaffermayr***

Abstract

This paper develops a structural empirical general equilibrium model of aggregate bilateral trade with path dependence of country-pair level exporter status. Such path dependence is motivated through informational costs about serving a foreign market for first-time entry of (firms in) an export market versus continued export services to that market. We embed the theoretical model into a structural dynamic stochastic econometric model of bilateral selection into import markets and apply it to a data-set of aggregate bilateral exports among 120 countries over the period 1995-2004. In particular, we disentangle the role of changes in trade costs, in labor endowments, and in total factor productivity for trade, bilateral market entry, numbers of firms active, and welfare. Dynamic gains from trade differ significantly from static ones, and path-dependence in market entry cushions effects of impulses in fundamental variables that are detrimental to bilateral trade.

Keywords: Bilateral trade flows; Gravity equation; Dynamic random effects model; Sample selection

JEL codes: F10; F12; F17

---

*Department of Management, Technology, and Economics, ETH Zurich, WEH E6, Weinbergstrasse 35, 8092 Zurich, Switzerland. Co-affiliations: CEPR, CESifo, GEP, and WIFO.

**Department of Economics, University of Innsbruck, Universitaetsstrasse 15, 6020 Innsbruck, Austria. Co-affiliations: CESifo and WIFO.

Acknowldgements: The authors gratefully acknowledge funding from FIW. Comments at seminar presentations at the Universities of Basel, Freiburg, Göttingen, Lausanne, Mannheim, Pablo de Olavide at Sevilla, Syracuse, and Tübingen are gratefully acknowledged. Furthermore, we received useful comments at the annual meeting of the European Trade Study Group, the annual CESifo Global Economy Area Conference, the PEGGED Workshop at Villars, and the annual meeting of the Royal Economic Society. Especially, we thank Costas Arkolakis, Harald Badinger, Richard Baldwin, Jffrey Bergstrand, Bernd Fitzenberger, Piyusha Matreja, Marcelo Olarreaga, and Raymond Riezman for numerous useful comments on earlier drafts.

1 Introduction

Whether two countries trade with each other in a given year or not – often referred to as the *extensive country margin of bilateral trade* – can be explained with great success by their past export status. For a cross section of the major 120 countries in terms of their GDP over the time period 1995-2004, Table 1 suggests that 66% of the country-pairs display positive bilateral exports when they did so 3 years prior to that, 20% have zero exports when they did not have any exports 3 years prior to that, and 13% change their activity within 3 years on average. Moreover, 52% of the country-pairs have positive bilateral exports in 2004 and they did so in 1995, 20% report zero exports in 2004 and they did not have any exports in 1995, and 28% change their activity between 1995 and 2004. This evidence suggests that there is a strong role for persistence or path dependence to play both unconditional and, as we will show, conditional on exogenous determinants for the extensive margin of trade.

This paper delivers a structural empirical model which is capable of analyzing both the *extensive* and the *intensive margin* of aggregate bilateral goods trade with a path-dependent extensive margin of trade (e.g., due to learning of firms about fixed market entry costs) in general equilibrium. In particular, the work by Evenett and Venables (2002), Albornoz, Calvo Pardo, Corcos, and Ornelas (2010), and others points to such path dependence at the extensive margin of trade. The model we propose is based on a dynamic model for bilateral selection into export markets and a demand equation for bilateral goods exports which are interrelated through the deterministic and stochastic components of the data-generating process. This model fully respects general equilibrium constraints at both margins of trade and, unlike earlier work, pursues an iterated estimation of a general-equilibrium-consistent panel data model with dynamic selection into export markets.

– Table 1 –

By virtue of the chosen approach, the paper stands on the shoulders of previous research on structural modeling of bilateral trade flows. With the seminal papers of Eaton and Kortum (2002), Anderson and van Wincoop (2003), and Helpman, Melitz, and Rubinstein (2008), it became possible to infer empirically comparative static effects of determinants of bilateral trade flows which are consistent with general equilibrium, taking into account repercussions of changes of exogenous drivers of trade on endogenous product and, eventually, factor prices. Beyond earlier work, the structural models of Eaton and Kortum (2002) and Helpman, Melitz, and Rubinstein (2008) can explain zero trade flows and, hence, deliver answers to the question as to which extent trade responds to changes in fundamental variables through the *extensive* versus the *intensive* margins of bilateral trade.[1]

A key feature of the aforementioned general equilibrium models is that they are designed for empirical cross-section analysis. Hence, they do not distinguish between short-run and long-run responses of outcome to changes in fundamental variables. In principal, it is of course possible with such models to simply index endogenous and exogenous variables by time and analyze empirically a series of cross sections. Yet, there is no salient role for history to play in the sense that, conditional on the contemporaneous exogenous variables, those cross sections would be independent of each other. Hence, such theoretical work suggests that the analysis of panel data on

[1]This paper is mostly concerned with path dependence in the entry of markets at the aggregate bilateral level. Hence, it is only loosely related to recent work on the (static) determinants and effects of growth of product variety in new trade theory models along the lines of Broda and Weinstein (2006) and Feenstra and Kee (2008).

bilateral trade matrices can be performed for each period separately without any loss of insight.

In line with recent structural empirical work on aggregate bilateral trade flows, we model nominal bilateral goods trade as a function of an exporting country's supply potential, an importing country's demand potential, and trade barriers. As in Melitz (2003), Chaney (2005), or Helpman, Melitz, and Rubinstein (2008), the latter contain elements which are tied to the quantity of goods shipped (*variable trade costs*) and ones that entail fixed export market access costs (*fixed trade costs*). Apart from contemporaneous fundamentals, we allow the extensive margin of bilateral trade to depend on bilateral export status prior to a given point in time. For instance, this is consistent with firms' entering a market to generate information about that market as a public good which is available to suppliers from the same origin to that market in subsequent periods. This leads to a dynamic model of export market selection which is stochastically related to export demand at the intensive country margin.

We formulate a deterministic and a stochastic version of that model and apply it to data on bilateral aggregate exports of the aforementioned 120 countries in three-year intervals between 1995 and 2004. Our goal is to identify the main drivers of world trade for that period, which in the context of the model are (fixed and variable) trade costs, labor endowments, and productivity.[2] In particular, we shed light on the short-run and the long-run

[2]In a different context, Baier and Bergstrand (2001) have asked a related question in a non-structural model with tariffs, non-tariff trade costs, and GDP growth as the main drivers of trade in a static model. They found that 67% of total growth of trade flows for 16 OECD countries over 1958-1960 and 1986-1988 could be explained by GDP growth, 26% by tariff reductions, and 8% by changes in non-tariff trade costs. Hence, the lion's share is attributed to GDP growth, the latter being exogenous there but endogenous in general equilibrium models of trade and itself a function of tariffs and trade costs among other factors (such as total factor productivity and factor endowments). More recently, Anderson and Baier (2010) focus on comparative static effects of the main drivers of trade

responses – and, hence, of path-dependence – of trade in general equilibrium to the changes of these fundamentals. We do so in a fully nonlinear model. Our findings suggest that the average three-year change in (fixed and variable) trade costs – a reduction thereof – per country-pair between 1995 and 2004 triggered positive short-run and long-run effects on nominal bilateral exports. Increases in labor endowments and total factor productivity raised bilateral exports even more strongly in both the short run and the long run.

The remainder of the paper is organized as follows. The next section formulates a parsimonious endowment model with path-dependent export market entry. While we chose a model which is closest to Krugman's (1979), such a framework could easily be cast in the context of theoretical models à la Anderson (1979), Eaton and Kortum (2002), or Helpman, Melitz, and Rubinstein (2008). Section 3 embeds this model in a stochastic framework for dynamic selection into export markets and aggregate export demand. Also, that section provides details about the implementation of such a model for parameter estimation and counterfactual analysis. Section 4 describes features of the data-set of 120 countries and three-year intervals for 1995-2004 we apply this model to, and it summarizes estimation results. Section 5 describes the findings about the short-run (three-year) and long-run (thirteen-year) effects of changes in drivers of trade flows as observed over the period 1995-2004. The last section concludes with a summary of the most important findings.

---

in a static general equilibrium model with positive trade flows only.

2 An aggregate gravity model with path-dependent market entry

Consider a world with J countries indexed by $j = 1, ..., J$ and consumers with a love for variety for goods consumption in a single sector à la Dixit and Stiglitz (1977). It will be useful to introduce a time index and set out that model for two periods, say t and $t-1$. It suffices to focus mostly on the exposition of the model for period t, but, as will become clear below, the equilibrium in t will depend on export status (of firms) of country i with j in period $t-1$. Let us assume that all varieties in country i and period t are produced by using one factor of production, labor, at unit input costs of $w_{it}a_{it}$, where w_{it} denotes the wage rate and a_{it} the corresponding input coefficient (inverse labor or total factor productivity). Then, monopolistic competition and non-segmentation of consumer markets by firms implies mark-up pricing with mill price[3]

$$p_{it} = \tfrac{\sigma}{\sigma-1} w_{it} a_{it}, \qquad (1)$$

where $\sigma > 1$ is the (time-invariant) elasticity of substitution between varieties. An important consequence of the assumption of homogeneous technologies within countries is that, through (1), all firms in country i – of which there is a mass n_{it} in period t – behave in the same way so that we can write

[3]Notice that the chosen approach follows closely Krugman's (1979) and Redding and Venables' (2004) framework. Alternatively, one could allow for heterogeneous firms by assuming a fixed distribution of total factor productivity as in Melitz (2003) or Helpman, Melitz, and Rubinstein (2008). The latter approach would support comparative static results for trade costs which run through an additional channel, namely adjustment of the export market-specific lower cutoff level of productivity of active producers. While the latter may be important to consider for an analysis at the level of firms or individual sectors (see Das, Roberts, and Tybout, 2007; Kee and Krishna, 2008; Cherkashin, Demidova, Kee, and Krishna, 2009; for examples), selection-induced productivity effects tend to be negligible in estimated general equilibrium models at the aggregate (country) level (see Egger, Larch, Staub, and Winkelmann, 2011). Therefore, we suppress the less parsimonious outline for a model with heterogeneous firms, here.

utility-maximizing demand in j for an i-borne variety in period t, c_{ijt}, and the price index for the consumer basket in j and year t, P_{jt}, respectively, as

$$c_{ijt} = \frac{p_{ijt}^{-\sigma}}{P_{jt}^{1-\sigma}} Y_{jt}, \quad P_{jt}^{1-\sigma} = \sum_{i=1}^{J} n_{it} p_{ijt}^{1-\sigma} V_{ijt}, \tag{2}$$

where $p_{ijt} \geq p_{it}$ is the consumer price per unit of c_{ijt}, Y_{jt} is income (GDP) in country j in that period, and V_{ijt} is an indicator variable which is unity, if i-borne varieties are sold at market j in t and zero otherwise.

Each variety is assumed to be internationally tradable, but importing is subject to variable transportation costs. With variable iceberg-type trade costs for shipping goods from i to j in period t of $\tau_{ijt} - 1 \geq 0$, $p_{ijt} = p_{it}\tau_{ijt}$. We will assume below that τ_{ijt} also includes tariffs. However, there is no need to disentangle iceberg from policy trade costs in τ_{ijt} at this point. Notice that p_{ijt} applies to exports which are measured inclusive of cost, insurance, and freight. Moreover, we follow Melitz (2003) and Helpman, Melitz, and Rubinstein (2008) in assuming that a firm's profits are additively separable into export-market-specific profits. Accessing a particular export market j for i-borne firms in period t is associated with fixed sunk costs (incurred in the first year of entry of that export market) plus fixed period-specific costs.

Suppose i-borne firms did not deliver goods to market j in period $t-1$ but they start doing so in period t. Let us denote the sum of set-up and maintenance fixed costs per i-borne firm for serving market j for the first time in t by $w_{it} f_{ijt}$, where f_{ijt} measures the units of labor used for set-up and maintenance.[4] To capture path dependence through, e.g., generation

[4]To avoid complicated dynamics at the firm level which are not observable in aggregate data for multiple countries, we assume that each firm lives one period only (see Cherkashin, Demidova, Kee, and Krishna, 2009, for a similar assumption). However, there is a dynamic process of aggregate market entry in each period accruing to new firms' inheritance of public knowledge about exports markets from previous periods by previous exporters.

of information about a market as a public good for firms from the same exporting country, in a very parsimonious way, assume that prior exporting (in $t-1$) of any i-borne firms to that market results in proportionately lower fixed costs of $w_{it} f_{ijt} e^{-\delta}$ with $\delta \geq 0$ (see Hausmann and Rodrik, 2003, for an early argument along those lines). Then, fixed costs of i-borne firms from serving market j in year t may be written as $w_{it} f_{ijt} e^{-\delta V_{ij,t-1}}$, where $V_{ij,t-1} = 1$ if market j had been served by i-borne firms in the previous period and zero else. Most importantly, the presence of $e^{-\delta V_{ij,t-1}}$ in the fixed costs entails state-dependence in export status at the country-pair level.[5]

In equilibrium, for positive exports bilateral shipments per variety, x_{ijt}, equal τ_{ijt} times bilateral demand per variety, c_{ijt}. Then, per-firm shipments x_{ijt} gross of cost, insurance, and freight (cif) and the value of aggregate bilateral exports gross of cif, X_{ijt}, are determined as

$$
\begin{aligned}
x_{ijt} &= \tau_{ijt} c_{ijt} = \frac{p_{it}^{-\sigma} \tau_{ijt}^{1-\sigma}}{P_{jt}^{1-\sigma}} Y_{jt} V_{ijt}, && (3) \\
X_{ijt} &\equiv n_{it} p_{it} x_{ijt} = n_{it} \left(\frac{p_{it}^{1-\sigma} \tau_{ijt}^{1-\sigma}}{P_{jt}^{1-\sigma}} \right) Y_{jt} V_{ijt}. && (4)
\end{aligned}
$$

Denote the aggregate endowment with labor of country i at time t by L_{it}. Assuming full employment, the labor constraint reads

$$
L_{it} = n_{it} \sum_{j=1}^{J} V_{ijt} \left(a_{it} x_{ijt} + e^{-\delta V_{ij,t-1}} f_{ijt} \right), \qquad (5)
$$

where $\sum_{j=1}^{J} V_{ijt} \left(a_{it} x_{ijt} \right) = a_{it} \sum_{j=1}^{J} V_{ijt} x_{ijt}$ is the amount of labor used for

[5] It is straightforward to allow for a more flexible cost function with a more general pattern of path dependence such as $w_{it} f_{ijt} e^{-\sum_{d=1}^{D} \delta_d V_{ij,t-d}}$. However, in the application below there is too much multicollinearity across the $V_{ij,t-d}$ so that identification of the individual parameters δ_d is only possible with $D = 1$. Hence, we abstain from overburdening the model unnecessarily with notation.

production, and $\sum_{j=1}^{J}\left(V_{ijt}e^{-\delta V_{ij,t-1}}f_{ijt}\right)$ is the amount of labor used for set-up of business contacts in $\sum_{j=1}^{J}V_{ijt}\leq J$ markets.

Market j-specific profits of i-borne firms in period t are given by

$$\pi_{ijt}=\tfrac{w_{it}a_{it}x_{ijt}}{\sigma-1}-w_{it}f_{ijt}e^{-\delta V_{ij,t-1}}, \tag{6}$$

where $V_{ij,t-1}=1$ if $i=j$; i.e., we assume that f_{iit} is small enough to ensure that active firms always serve consumers at least in the country they produce in at any period t.[6]

Non-negative profits in (6) for exports per firm from i to j in period t suggests that positive exports at free market entry require $x_{ijt}\geq x^{*}_{ijt}\equiv\frac{f_{ijt}e^{-\delta V_{ij,t-1}}}{a_{it}}(\sigma-1)$. Hence, i-borne firms will only start exporting to j in t, if $\tau_{ijt}c_{ijt}\geq\frac{f_{ijt}}{a_{it}}(\sigma-1)$ and, in case of prior exports between i and j, they will only continue exporting, if $\tau_{ijt}c_{ijt}\geq\frac{f_{ijt}e^{-\delta}}{a_{it}}(\sigma-1)$. No matter of whether they start or continue exporting in t, at free entry an i-borne firm's exports to j in t are determined by x^{*}_{ijt}. In equilibrium, usage of x^{*}_{ijt} in (5) determines the number of firms active in country i at time t as

$$n_{it}=\frac{L_{it}}{\sigma\sum_{j=1}^{J}V_{ijt}e^{-\delta V_{ij,t-1}}f_{ijt}}. \tag{7}$$

Since market j is only served in t by i-borne firms if this is profitable, we may introduce a latent variable V^{*}_{ijt} that reflects aggregate potentially

[6]We also assume throughout that the costs of entering foreign countries are low enough so that it pays off for firms to export somewhere abroad and to consumers in every country to import some varieties, in line with empirical stylized facts. When we estimate the model to the data, this outcome arises endogenously, consistent with those facts.

realizable profits of firms in i for serving consumers in j in period t as in (6):

$$V_{ijt}^{*} = \frac{w_{it}a_{it}x_{ijt}}{(\sigma-1)w_{it}f_{ijt}e^{-\delta V_{ij,t-1}}} \geq 1, \quad \text{or} \tag{8}$$

$$\breve{V}_{ijt}^{*} \equiv \frac{V_{ijt}^{*}}{V_{iit}^{*}} = \frac{\tau_{ijt}^{1-\sigma}}{\tau_{iit}^{1-\sigma}} \frac{Y_{jt}P_{jt}^{\sigma-1}}{Y_{it}P_{it}^{\sigma-1}} \frac{e^{-\delta} f_{iit}}{e^{-\delta V_{ij,t-1}} f_{ijt}} \geq 1. \tag{9}$$

Since $V_{iit}^{*} \geq 1$ by both assumption and observation (consumption from domestic producers is generally positive at the aggregate level), both V_{ijt}^{*} and $\breve{V}_{ijt}^{*}$ generate the same indicator variable V_{ijt} according to

$$V_{ijt} = \begin{cases} 1 & \text{if} \quad \ln \breve{V}_{ijt}^{*} \geq 0 \\ 0 & \text{else.} \end{cases} \tag{10}$$

In general equilibrium, total sales to all markets gross of ad-valorem tariffs charged by importers (referred to as including cost, insurance and, freight; cif) add up to GDP, Y_{it}, plus tariff revenues earned by i minus tariff revenues collected abroad from i's exports, T_{it}, so that

$$\breve{Y}_{it} \equiv Y_{it} + T_{it} = \sum_{h=1}^{J} X_{iht} = n_{it}p_{it}^{1-\sigma} \sum_{h=1}^{J} \left[V_{iht} \left(\frac{\tau_{iht}}{P_{ht}} \right)^{1-\sigma} Y_{ht} \right]$$

or, after defining $Y_t \equiv \sum_{h=1}^{J} Y_{ht}$, $\theta_{it} \equiv Y_{it}/Y_t$, $\breve{\theta}_{it} \equiv \breve{Y}_{it}/Y_t = \theta_{it}\breve{Y}_{it}/Y_{it} = \theta_{it} + T_{it}/Y_t$, and $\Pi_{it}^{1-\sigma} \equiv \sum_{h=1}^{J} V_{iht} \left(\frac{\tau_{iht}}{P_{ht}} \right)^{1-\sigma} \theta_{ht}$, similar to Anderson and van Wincoop (2003) and Anderson (2010), we obtain

$$\breve{Y}_{it} = n_{it}p_{it}^{1-\sigma} Y_t \Pi_{it}^{1-\sigma} \Rightarrow n_{it}p_{it}^{1-\sigma} = \breve{\theta}_{it}\Pi_{it}^{\sigma-1}. \tag{11}$$

The latter expressions illustrate that the adopted version of a Dixit and Stiglitz (1977) or Krugman (1979) model is isomorphic to the one of Anderson and van Wincoop (2003). Replacing $n_{it}p_{it}^{1-\sigma}$ by the expression in (11) and Y_{jt}

by $Y_t\theta_{jt}$ in (4) and recalling the definition of $P_{jt}^{1-\sigma}$ from (2), the generalized system of trade resistance equations à la Anderson and van Wincoop (2003) with possible zero trade flows and tariffs is then given by

$$\Pi_{it}^{1-\sigma} = \sum_{h=1}^{J} V_{iht}\tau_{iht}^{1-\sigma}P_{ht}^{\sigma-1}\theta_{ht}, \quad P_{jt}^{1-\sigma} = \sum_{h=1}^{J} V_{hjt}\tau_{hjt}^{1-\sigma}\Pi_{ht}^{\sigma-1}\breve{\theta}_h. \tag{12}$$

Defining $\mu_{it} \equiv \breve{\theta}_{it}\Pi_{it}^{\sigma-1}$ and $m_{jt} \equiv \theta_{jt}P_{jt}^{\sigma-1}$, we can rewrite aggregate nominal exports at cif from i to j in t as

$$X_{ijt} = Y_t\tau_{ijt}^{1-\sigma}V_{ijt}\mu_{it}m_{jt}, \quad \text{with} \tag{13}$$

$$\breve{\theta}_{it} = \mu_{it}\sum_{h=1}^{J} V_{iht}\tau_{iht}^{1-\sigma}m_{ht}, \quad \theta_{jt} = m_{jt}\sum_{h=1}^{J} V_{hjt}\tau_{hjt}^{1-\sigma}\mu_{ht}. \tag{14}$$

A key assumption in the paper is that firms consider the role of path dependence for market entry, but they do not look forward and equate the stream of future operating profits to the one of total (per-period and subsequently sunk entry) fixed costs when deciding about the timing of entry. In a separate paper, we analyze a model of the latter kind in general equilibrium. It turns out that when conditioning on observed fundamental variables, under certain assumptions, the estimation part of the problem is not much different from the problem with path dependence: while past export status exhibits a constant effect δ on the latent process determining the extensive margin here, it has a drift of the form $\delta \cdot t$, where t represents a time trend. However, counterfactual analysis is computationally extremely demanding with forward-looking managers and there are so many conceptual problems involved that this issue calls for a separate paper focusing on counterfactual analysis rather than estimation.

3 From theory to an empirical model: Implementation and estimation

To derive an econometric specification of the above gravity model with panel data, we need to specify the stochastic processes that arise from measurement error about or random shocks on exports. Finally, we ought to comment on some issues with the implementation of the model.

3.1 Adding disturbances

Let us take logs of the gravity equation in (13) and add a log-additive stochastic term $u_{X,ijt}$ to obtain

$$\ln X_{ijt} = \begin{cases} \ln Y_t + \ln \tau_{ijt}^{1-\sigma} + \ln m_{it} + \ln \mu_{jt} + u_{X,ijt} & \text{if } V_{ijt} = 1 \\ \text{unobserved} & \text{if } V_{ijt} = 0 \end{cases}, \tag{15}$$

where $u_{X,ijt}$ is the stochastic disturbance term. The trade resistance terms $\ln \mu_{it}$ and $\ln m_{jt}$ are determined as implicit solutions to the system of $2J$ equations (14) in $2J$ unknowns μ_{it} and m_{jt} for each period t following from the requirement of multilaterally balanced trade for each economy.

The unobserved latent variable for the propensity to export from i to j in year t based on (9) is log-transformed, augmented additively by the stochastic term $u_{V,ijt}$, and it is taken into account that $\frac{Y_{jt}P_{jt}^{\sigma-1}}{Y_{it}P_{it}^{\sigma-1}} = \frac{m_{jt}}{m_{it}}$, so that the expression can be written as

$$\ln \widetilde{V}_{ijt}^{*} = \ln \frac{\tau_{ijt}^{1-\sigma}}{\tau_{iit}^{1-\sigma}} + \ln \frac{m_{jt}}{m_{it}} + \delta V_{ij,-1} + \ln \frac{f_{iit}}{f_{ijt}} + u_{V,ijt}, \quad \text{with} \tag{16}$$

$$V_{ijt} = 1[\ln \widetilde{V}_{ijt}^{*} \geq 0]. \tag{17}$$

We will talk about the assumptions regarding $u_{X,ijt}$ and $u_{V,ijt}$ in the next subsection. With respect to variable trade costs and fixed export market access costs, our specification follows the literature (see, e.g., Helpman, Melitz, and Rubinstein, 2008) assuming

$$\ln \tau_{ijt}^{1-\sigma} = \sum_{k=1}^{K} \alpha_k \ln \zeta_{k,ijt}, \quad \ln f_{ijt} = \sum_{l=1}^{L} \beta_l \ln \chi_{l,ijt}, \tag{18}$$

where $\zeta_{k,ijt}$ and $\chi_{l,ijt}$ are variables related to variable and fixed trade costs, respectively. In practice, K may equal L and all factors determining $\ln \tau_{ijt}^{1-\sigma}$ may also affect $\ln f_{ijt}$. As long as the parameters α_k differ from the respective β_l, $\ln \tau_{ijt}^{1-\sigma}$ may still differ from $\ln f_{ijt}$, even if $\ln \zeta_{k,ijt} = \ln \chi_{l,ijt}$ for $k = l$. It may be desirable for identification to include at least one other element $\ln \chi_{l,ijt}$ beyond the ones of $\ln \zeta_{k,ijt}$ in small samples, but in large samples as ours, there is no need for the fundamentals behind $\ln \tau_{ijt}^{1-\sigma}$ and $\ln f_{ijt}$ to differ at all.

Obviously, even in the absence of zero trade flows (i.e., $V_{ijt} = 1$ for all ijt) and at known σ, Y_{it}, $\tau_{ijt}^{1-\sigma}$, the system in equation (14) could only be solved numerically.[7] Notice that we fully respect cross-equation restrictions of parameters in the empirical models (15)-(17).

3.2 Stochastic process and estimation

The actual implementation of the above model rests upon equations (15)-(17). Notice that export status at the country-pair level, V_{ijt}, is observed at any point in time t, but the underlying latent processes $\ln \widetilde{V}_{ijt}^{*}$ or $\ln \widehat{\widetilde{V}}_{ijt}^{*}$ are

[7]Baier and Bergstrand (2009) derived a linear approximation of the system of multilateral trade resistance terms (in the chosen notation Π_i and P_j) which is based on the first step of a Gauss-Newton iteration of the solution to the system of trade resistance equations (14). In Egger and Pfaffermayr (2010), we generalize this procedure to the case with some zero trade flows. However, we illustrate that this approximation does not work well due to discontinuities in the objective function.

not. The latter latent variables measure the net log benefits from exporting at all from i to j at time t. Hence, V_{ijt} measures and $\widetilde{V}^*_{ijt}$ determines what we may refer to as the *extensive margin* of exports at the aggregate country-pair level. The variable $\ln X_{ijt}$ is only observed if $\ln \widetilde{V}^*_{ijt} > 0$ and operating profits earned in country j are large enough to cover the fixed exporting (or export market access) costs.

The disturbances $u_{V,ijt}$ and $u_{X,ijt}$ in the models of $\widetilde{V}^*_{ijt}$ in (16) and $\ln X_{ijt}$ in (15), respectively, are specified as

$$u_{V,ijt} = \eta_{V,ij} + \lambda_{V0} V_{ij,0} + \varepsilon_{V,ijt} \quad (19)$$

$$u_{X,ijt} = \eta_{X,ij} + \varepsilon_{X,ijt}, \quad (20)$$

where $\eta_{V,ij}$ and $\eta_{X,ij}$ are time-invariant, country-pair-specific effects that are assumed to be uncorrelated with the other determinants of $\widetilde{V}^*_{ijt}$ (including $V_{ij,0}$) and of $\ln X_{ijt}$, respectively. $\eta_{V,ij}$ and $\eta_{X,ij}$ are identically and independently distributed normal random effects which may be correlated with each other, and λ_{V0} captures the (time-invariant) initial conditions, which are included to acknowledge the market entry dynamics introduced before. Moreover, $\varepsilon_{V,ijt}$ and $\varepsilon_{X,ijt}$ are identically and independently distributed normal disturbances which may be correlated with each other but are independent of $\eta_{V,ij}$ and $\eta_{X,ij}$ and the other determinants of $\widetilde{V}^*_{ijt}$ (including $V_{ij,0}$) and of $\ln X_{ijt}$.[8]

[8] In principal, it would be possible to allow not only $u_{V,ijt}$ (as we do) but even $\eta_{V,ij}$ to be correlated with some of the determinants of $\widetilde{V}^*_{ijt}$ and $\eta_{X,ij}$ with some of the determinants of $\ln X_{ijt}$. For instance, one could follow the so-called Mundlak-Chamberlain-Wooldridge device and include means of all determinants of $\widetilde{V}^*_{ijt}$ and $\ln X_{ijt}$ in the respective equations across time in addition to the original variables in the model. However, as this requires enough time variation in the explanatory variables, that approach is infeasible with numerous time-invariant covariates (such as bilateral distance or common borders, etc.) whose coefficient estimates are vital to the counterfactual analysis of the model. Accordingly, we have to resort to the somewhat stronger assumption of $\eta_{V,ij}$ and $\eta_{X,ij}$ as well as $\varepsilon_{V,ijt}$ and

Regarding the distribution of the disturbances, we assume specifically that $(\eta_{V,ij}, \eta_{X,ij}) \sim i.i.d.N(\mathbf{0}, \mathbf{V}_\eta)$ and $(\varepsilon_{V,ijt}, \varepsilon_{X,ijt}) \sim i.i.d.N(\mathbf{0}, \mathbf{V}_\varepsilon)$, where

$$\mathbf{V}_\eta = \begin{bmatrix} \sigma^2_{V,\eta} & \rho_\eta \sigma_{V,\eta}\sigma_{X,\eta} \\ \rho_\eta \sigma_{V,\eta}\sigma_{X,\eta} & \sigma^2_{X,\eta} \end{bmatrix}, \quad \mathbf{V}_\varepsilon = \begin{bmatrix} 1 & \rho_\varepsilon \sigma_{X,\varepsilon} \\ \rho_\varepsilon \sigma_{X,\varepsilon} & \sigma^2_{X,\varepsilon} \end{bmatrix}.$$

Since the variance of $\varepsilon_{V,ijt}$, the remainder disturbances in the extensive margin model, is not identified, we normalize it to unity without loss of generality (see the upper left cell of $\mathbf{V}_\varepsilon$). In that model, $\rho_\eta \neq 0$ and/or $\rho_\varepsilon \neq 0$ implies selection into export status, so that the stochastic process may be termed a generalized random effects sample selection model which allows for path-dependent aggregate bilateral export status.

For the sake of simplicity of the notation, let us collect the determinants of the indicator function V_{ijt} (the *extensive margin* of aggregate bilateral exports) and of continuous $\ln X_{ijt}$ (the *intensive margin* of aggregate bilateral exports) for observation ijt into the following vectors

$$\begin{aligned} \mathbf{w}_{V,ijt} &= \left[\ln \frac{\zeta_{1,ijt}}{\zeta_{1,iit}}, ..., \ln \frac{\zeta_{K,ijt}}{\zeta_{K,iit}}, \ln \frac{m_{jt}}{m_{it}}, V_{ij,t-1}, \ln \chi_{1,ijt}, ..., \ln \chi_{L,ijt}, V_{ij,0}, 1\right] \\ \mathbf{w}_{X,ijt} &= [\zeta_{1,ijt}, ..., \ln \zeta_{K,ijt}, \ln \mu_{it}, \ln m_{jt}, Y_t, 1], \end{aligned}$$

where $V_{ij,0}$ is included by following Wooldridge (2005) in $\mathbf{w}_{V,ijt}$ to model the initial condition of the dynamic process for the extensive margin (selection into export markets), and a constant is included at the end of both $\mathbf{w}_{V,ijt}$ and $\mathbf{w}_{X,ijt}$ for proper centering of the data. Taking into account the parametrization in (18), the parameter vectors corresponding to $\mathbf{w}_{V,ijt}$ and $\mathbf{w}_{X,ijt}$,

$\varepsilon_{X,ijt}$ to be generally uncorrelated with other determinants of the extensive and the intensive margin of exports. Moreover, the findings of Baier and Bergstrand (2007) suggest that, e.g., the endogeneity of trade regionalism is much less an issue in panel data models than in cross-section models.

respectively, are

$$\beta_V = [\alpha_1, ..., \alpha_K, 1, \delta, \beta_1, ..., \beta_L, \lambda_{V0}, \beta_0] \tag{21}$$

$$\beta_X = [\alpha_1, ..., \alpha_K, 1, 1, 1, \alpha_0], \tag{22}$$

where β_0 and α_0 are the coefficients of the constants in the two models. Notice that, for counterfactual analysis, the coefficients on $\ln \frac{\zeta_{1,ijt}}{\zeta_{1,iit}}, ..., \ln \frac{\zeta_{K,ijt}}{\zeta_{K,iit}}$ in the specification of the latent process (16) underlying the extensive margin of aggregate bilateral trade have to equal the ones on $\zeta_{1,ijt}, ..., \ln \zeta_{K,ijt}$ in the specification of the intensive margin of exports (15). Moreover, general-equilibrium-consistent counterfactual analysis requires that the coefficients on $\frac{m_{jt}}{m_{it}}$ in (16) as well as the ones on $\ln \mu_{it}$, $\ln m_{jt}$, and Y_t in (15) are unity each.

Then, we can write the models to be estimated as follows:

$$\begin{aligned} V_{ijt} &= 1[\ln \widetilde{V}^*_{ijt} = \mathbf{w}_{V,ijt}\beta_V + \eta_{Vij} + \varepsilon_{V,ijt} > 0] \qquad (23) \\ &= 1[A_{ijt} + \eta_{Vij} + \varepsilon_{V,ijt} > 0] \\ \ln X_{ijt} &= \mathbf{w}_{X,ijt}\beta_X + \eta_{Xij} + \varepsilon_{X,ijt} \qquad (24) \\ &= B_{ijt} + \eta_{Xij} + \varepsilon_{X,ijt}. \end{aligned}$$

Recently, Raymond, Mohnen, Palm and Schim van der Loeff (2007, 2010) analyzed such models which allow to test and correct for sample selection with a dynamic process.[9] Following Wooldridge (2005) and Raymond, Mohnen, Palm, and Schim van der Loeff (2007, 2010), we specify the likelihood of

[9]In contrast to sample selection models for panel data as, e.g., in Wooldridge (1995), this model permits accounting for state dependence in the selection equation for the extensive margin of exports. Unlike previously applied selection models for structural gravity equations, this model is applicable with panel data and allows entertaining the time variation in trade with path dependence at the extensive margin.

country-pair ij, starting in $t = 1$ conditional on the regressors in $\mathbf{w}_{V,ijt}$ (including the initial conditions) and $\mathbf{w}_{X,ijt}$ and integrate out the country-pair-specific random effects $\eta_{V,ij}$ and $\eta_{X,ij}$ as

$$\begin{aligned} \mathcal{L}_{ij} &= \int_{-\infty}^{\infty}\int_{-\infty}^{\infty} \Pi_{t=1}^{T} L_{ijt}\phi(\eta_{V,ij},\eta_{X,ij})d\eta_{V,ij}d\eta_{X,ij}, \qquad (25) \\ L_{ijt} &= \Pi_{t=1}^{T}\Big\{\Phi\left(-A_{ijt}-\eta_{V,ij}\right)^{1-V_{ijt}}\left[\tfrac{1}{\sigma_{X,\varepsilon}}\phi\left(\tfrac{\ln X_{ijt}-B_{ijt}-\eta_{X,ij}}{\sigma_{X,\varepsilon}}\right)\times \right. \\ &\left.\Phi\left(\tfrac{A_{it}+\eta_{V,ij}+\frac{\rho_\varepsilon}{\sigma_{X,\varepsilon}}\left(\ln X_{ijt}-B_{ijt}-\eta_{Xij}\right)}{\sqrt{1-\rho_\varepsilon}}\right)\right]^{V_{ijt}}\Big\}, \qquad (26) \end{aligned}$$

where $\phi(\eta_{V,ij},\eta_{X,ij})$ denotes the density of the bivariate normal of the random country-pair effects as defined above, and $\Phi(\cdot)$ and $\phi(\cdot)$ in the expression for L_{ijt} denote the cumulative distribution function and the density, respectively, of the univariate normal distribution.

The likelihood in (25)-(26) can be numerically maximized to estimate the model parameters – namely the elements in $\mathbf{w}_{V,ijt}$ and $\mathbf{w}_{X,ijt}$ as well as those in $\mathbf{V}_\eta$ and $\mathbf{V}_\varepsilon$ – using a two-step Gauss-Hermite quadrature for integrating out the random country-pair effects (see Appendix 1 for details). For this, one chooses a (not too large) number of sample points. The procedure is computationally demanding, since, with a bivariate normal, the number of sample points implies a number of evaluation points of that number squared. We use 49 evaluation points of the Hermite polynomial and a weight for each of them to approximate the density of the bivariate normal distribution in the likelihood function (see Appendix 1 for further details).[10]

Since (5) for observation ijt depends on $\ln \mu_{it}$, $\ln m_{it}$, and $\ln m_{jt}$ which themselves depend on the estimated model parameter estimates, we pursue an iterative approach to parameter estimation and solving for $\ln \mu_{it}$, $\ln m_{it}$,

[10]Hence, with seven sample points and a bivariate normal, there are 49 points at which the likelihood has to be evaluated iteratively.

and, $\ln m_{jt}$ for all ijt. Hence, at each iteration step of the likelihood optimization, the multilateral resistance terms are solved iteratively. More precisely, we use starting values of θ_{it}, $\ln \mu_{it}$, $\ln m_{it}$ for all it and jt in Step 1 and optimize (25) to obtain estimates of the elements of β_V and β_X as well as those of $\mathbf{V}_\eta$ and $\mathbf{V}_\varepsilon$. Then, we solve for all $\ln \mu_{it}$ and $\ln m_{it}$ from the $2JT$ equations in (14) through a Newton procedure in Step 2. With the new values for all $\ln \mu_{it}$ and $\ln m_{it}$ at hand, one obtains new values of the latent variable $\ln \widetilde{V}^*_{ijt}$, etc. We iterate Steps 1 and 2 until convergence to obtain theory-consistent parameter estimates from maximum likelihood estimation. With the chosen grid of 49 evaluation points (based on seven sample points) with a bivariate normal for the stochastic process, parameter estimation of a random effects model cum dynamic sample selection and endogenous multilateral resistance terms takes roughly two days on a modern multi-core computer for a data-set as large as ours.[11]

Overall, the model accounts for three types of instantaneous effects of increasing trade costs on bilateral trade flows similar to Eaton and Kortum (2002), Melitz (2003), Chaney (2008), or Helpman, Melitz, and Rubinstein (2008). First, there is a direct effect due to the adjustment at the intensive margin as in (24) through higher (variable) trade costs on consumer prices in the destination country. Second, higher (variable as well as fixed) trade costs, eventually, may lead to zero bilateral trade flows as captured by the extensive margin relationship in (23). Finally, these direct consequences of higher trade costs at the extensive and intensive margins cause multilateral effects through trade by virtue of the price index effects captured by (14).[12]

[11] As is demonstrated in Egger and Pfaffermayr (2010), the gain from estimating a linearly approximated model à la Baier and Bergstrand (2009) is only marginal and comes at the cost of a potentially high approximation bias of benchmark and counterfactual predicted outcome variables.

[12] As said before, by focusing on homogeneous firms within countries, we rule out effects

In contrast to previous structural empirical work on bilateral trade flows, our model generates dynamic effects of changes in trade barriers through dynamic adjustment at the extensive margin of aggregate bilateral trade. In our empirical analysis, we aim at fleshing out the instantaneous versus the long-run effects of changes in country size versus trade costs on the extensive and intensive margin of trade and, taking general equilibrium feedback effects and implied parameter constraints in the model fully into account for both estimation and counterfactual analysis.

4 Data and estimation results

4.1 Data

The panel data-set employed in this paper is based on three-year averages of bilateral exports among 120 countries in five periods (see Appendix 2 for a list of countries by continent): 1992 ($t = 0$), 1995 ($t = 1$), 1998 ($t = 2$), 2001 ($t = 3$), 2004 ($t = 4$). We use three-year intervals so as to keep the number of time periods, T, small enough, since maximum likelihood estimation of the stochastic model is computationally quite demanding. Both X_{ijt} and V_{ijt} are based on nominal aggregate bilateral export flows in current US dollars as published in the United Nations' COMTRADE database. Figures on exporter and importer nominal GDP in current US dollars for the respective years come from the World Bank's World Development Indicators.

Furthermore, we employ three types of trade barriers. First of all, we use average (trade-weighted) applied bilateral tariffs information about which is

of higher trade costs on average productivity of firms exporting from a given country to a specific destination country. However, previous evidence suggests that this effect is of minor importance in aggregate data (see Egger and Larch, 2010; Egger, Larch, Staub, and Winkelmann, 2011).

available from the World Bank's WITS Database. Since the source data on weighted tariffs exhibit a large number of missing values, we interpolated and imputed missing tariff data using exogenous predictors (see Appendix 3 for details). Since such a procedure (and even trade weighting alone) leads to measurement error, we follow Wansbeek and Meijer (2000, p. 29) by constructing indicator variables so as to capture quantiles of the distribution of tariffs. Using a rough approximation of the distribution of measurement error-prone tariff data, e.g., from trade weighting or imputation, through discrete variables is a valid alternative to instrumental variables estimation to reduce measurement error (see Wansbeek and Meijer, 2000). More specifically, we generate five indicator variables, which are associated with quintiles of the imputed tariff levels. We use zero tariff rates (as charged within deep preferential trade agreements such as customs unions or free trade areas) as the base which fully captures preferential trade agreement membership. In this way we are able to obtain a maximum coverage of countries and time periods, which is a prerequisite for both sample selection model estimation and solution for endogenous terms μ_{it} and m_{jt} in (14) consistent with world trade general equilibrium. Second, we use trade cost measures which are related to geographical distance between countries from the Centre d'Études Prospectives et d'Informations Internationales' Geographical Database. In particular, we use bilateral distance (in kilometers) between economies' economic centers and an indicator reflecting common land borders between countries from that source. Third, we employ measures of cultural distance in terms of a common official language indicator variable, past colonial relationship, and a common colonizer indicator variable from that source.

Denote average applied bilateral tariff levels charged by country j on varieties from i in year t in quintile $\kappa = 2, ..., 5$ by $1 \geq b_\kappa - 1 \geq 0$. Average applied bilateral tariff levels in percent are $100(b_\kappa - 1)$, and they amount to 2.96%, 7.07%, 11.62%, and 21.42%, in the second, third, fourth, and fifth quintile, respectively, of the distribution in the average year t. This information is important for interpretation of the parameter estimates. We choose a notation so that $\zeta_2, ..., \zeta_5$ (e.g., in Table 4 below) correspond to quintile indicators for the second to the fifth quintile. Given that tariffs in the lowest quintile are captured by $b_1 = 1$, the estimated coefficients $\widehat{\alpha}_2, ..., \widehat{\alpha}_5$ on the indicators $\zeta_2, ..., \zeta_5$ can be interpreted as follows: $\widehat{\alpha}_\kappa = -\widehat{\sigma} \ln b_\kappa$ for $\kappa = 2, ..., 5$ so that $\widehat{\sigma} = -\frac{\widehat{\alpha}_\kappa}{\ln b_\kappa}$. Hence, the model principally permits estimation of σ.[13]

– Table 2 –

Table 2 summarizes features of the data on nominal exports in logs GDP, and the geographical (bilateral distance in logs and a non-contiguity binary indicator), cultural (binary indicator variables on no common language, no past colonial relationships between exporter and importer, and the two countries not having had a common colonizer), and political trade barriers (quintiles for tariff rates).[14] While the bloc on the left-hand side of Table 2 provides information on average levels of these variables over the information period and their standard deviation, the bloc on the right-hand side provides average three-year changes for the time-variant subset of variables (i.e., except for the geographical and cultural indicators).

[13] However, since there are several levels of κ, the estimates for σ may differ across them if they are not restricted to be the same. In general, there are various ways of estimating σ which eventually will give different point estimates. See Eaton and Kortum (2002) for a similar finding in a static Ricardian model of bilateral trade where what we refer to as an estimate of σ corresponds to an estimate of comparative advantage.

[14] We use binary indicators on non-contiguity, absence of a common language, etc., so that the parameters on these binary elements of $\ln \tau_{ijt}^{1-\sigma}$ and $\ln f_{ijt}$ always measure the role of higher barriers associated with an absence of the respective trade facilitation through contiguity, common language, etc., on the extensive and intensive country margins of exports.

According to Table 2, about 21% of the observations fall into the lowest and as many into the highest quintile of the tariff distribution (zero tariffs),[15] while about 20%, 19%, and 19% of the observations fall into the second, third, and the fourth quintile, respectively. The allocation of observations across quintiles is not exactly identical to 20% due to characteristics of the distribution of tariffs. In the average three-year period, more than 4% of the observations enter the lowest quintile of tariffs (from wherever) and slightly more than 1% enter the second quintile. Anyone of the upper three quintiles looses observations in the average three-year period between 1992 and 2004. The majority of observations does neither have a common land border or a common language, nore a common colonizer in the past. More than 60% of the country-pairs did have positive exports in 1989. In the average period, about 72% of the country-pairs had positive bilateral exports and about 70% of the country-pairs saw positive bilateral exports three years earlier.

In terms of the notation in the previous section, we have up to $K = L = 10$ elements $\alpha_k \ln \zeta_{k,ijt}$ for $k = 1, ..., 10$ in $\ln \tau_{ijt}^{1-\sigma}$ and $\beta_l \ln \chi_{l,ijt}$ for $l = 1, ..., 10$ in $\ln f_{ijt}$, namely the aforementioned tariff, geographical, and cultural barriers which determine $\ln \tau_{ijt}^{1-\sigma}$ and $\ln f_{ijt}$, respectively. Recall that we impose the restriction that the estimate of $\ln \tau_{ijt}^{1-\sigma}$ is identical between the extensive ($\ln \widetilde{V}_{ijt}^{*}$) and intensive margin equations ($\ln X_{ijt}$), but the inclusion of $\ln f_{ijt}$ along with $\ln \tau_{ijt}^{1-\sigma}$ in the extensive country margin model allows for identification of the parameters β_l apart from α_k.

[15] Exports of about 24% of the observations in the sample happen within a preferential trade agreement.

4.2 Estimation results

In this subsection, we summarize the estimation results of dynamic selection models for the fully non-linear model as introduced in the previous sections. In any case, the parameters have to be estimated iteratively, since the multilateral resistance terms in (14) depend on the endogenous V_{ijt}.

Table 3 summarizes parameter estimates and their standard errors for three models (labeled A to C) each. In a vertical dimension, the table exhibits two blocs, where the one at the top refers to the extensive country margin as in equation (23), and the one at the bottom refers to the intensive country margin as in equation (24). All models are based on the fully nonlinear model involving implicit solutions to (14) at every step of the maximum likelihood estimation.

Models A and B allow the stochastic terms to be correlated across the extensive and intensive margin models. While Model A assumes bivariate normality so that dependence can be captured by an inverse Mills' ratio as outlined in Section 3.2, Model B is a semi-parametric counterpart. The latter model replaces the inverse Mills' ratio in the outcome equation by a third-order polynomial of the linear prediction of the extensive margin model (i.e., of $\ln \widehat{\widetilde{V}}^*_{ijt}$). Conditional on the polynomial function, the stochastic terms between the two equations are assumed independent. We suppress the coefficients of the polynomial function but note that they are jointly significant at one percent with the data at hand. Helpman, Melitz, and Rubinstein (2008) interpret such a model as to control for both endogenous selection into export markets and firm heterogeneity within countries (in our case, average productivity of firms in i that serve market j in year t).[16] Unlike

[16]This interpretation involves many more assumptions than homogeneous firm models do. For instance, one has to specify the distribution function for firm productivity and the boundaries of the support region of possible productivity draws (inter alia, one

in Helpman, Melitz, and Rubinstein (2008), the (here, dynamic) selection equation and the outcome equation have to be estimated simultaneously rather than in two steps. Hence, maximum likelihood estimation has to be carried out iteratively until convergence, since the predictions of the control function change as the parameters of the models change.

Model C assumes that there is no endogenous selection into the extensive margin of trade and condition on the indicator V_{ijt} in the intensive margin outcome model as an exogenous variable. Hence, V_{ijt} in the intensive margin equation for $\ln X_{ijt}$ is not treated as a Bernoulli response variable to $\widetilde{V}^*_{ijt}$, unlike in Models A and B. Accordingly, the parameters of the latent process $\widetilde{V}^*_{ijt}$ are not estimated in these models but the multilateral resistance terms in (14) are solved by conditioning on the observed contemporaneous bilateral export status V_{ijt}. We consider Model A to be the preferred reference case, while the other models are inferior due to assumptions made for counterfactual analysis (Model B) or endogeneity bias (Models C).

Due to the parameter restrictions imposed, the estimates of α_k are identical for all determinants of $\ln \tau_{ijt}^{1-\sigma}$ in either equation within a model. However, we assume that the same variables affect $\ln \tau_{ijt}^{1-\sigma}$ and $\ln f_{ijt}$ so that $K = L$

needs to take a stance whether this support region is the same across countries or not; for instance, identical potential productivity support across all countries is assumed in Helpman, Melitz, and Rubinstein, 2008). Moreover, in static models, comparative static effects tend to be very similar between homogeneous and heterogeneous firm models when assuming identical distribution functions and possible productivity support regions across countries (see Egger, Larch, Staub, and Winkelmann, 2011). Therefore, we use Model A for further reference. That model assumes as an approximation that firms only differ in terms of productivity across countries but are identical within economies. As Helpman, Melitz, and Rubinstein (2008) admit, controlling for firm heterogeneity links productivity of the average firm in country i which serves consumers in j (here, in year t) to the propensity to export for the marginal firm. It does not serve to control for the productivity of the average firm active in market i, i.e., the inverse of what we dubbed a_{it}. It is the latter, which we are primarily interested in, and Helpman, Melitz, and Rubinstein (2008) admit that a_{it} is proportional to producer prices and, hence, implicitly taken care of in estimation anyway.

and $\ln \zeta_{k,ijt} = \ln \chi_{l,ijt}$ for all $k = l$, but α_k may differ from β_l. For the sake of brevity, we therefore always report parameter estimates for $\alpha_{k,ijt} + \beta_{k,ijt}$ in the extensive margin models, since they refer to the same fundamental trade cost variables. Moreover, only the extensive margin equation includes (endogenous) $V_{ij,t-1}$ and $V_{ij,0}$ and, hence, delivers parameter estimates for δ and λ_{V0}, respectively.[17]

– Table 3 –

For the selection equations, we assess the goodness of fit by Matthew's correlation coefficient (MCC). This correlation coefficient is based on a cross tabulation of V_{ijt} and $\widehat{V}_{ijt}$ and it is related to the χ^2-statistic for a 2×2 contingency table by $|MCC| = \sqrt{\frac{\chi^2}{N}}$, where N is the number of observations. For log positive export flows at the intensive margin, we measure the goodness of fit by the correlation between the observed and predicted values. Table 3 shows that for the former, we obtain an MCC of 0.520 and 0.593. With respect to the latter, the fit is quite similar across all three estimated models, amounting to 0.744, 0.743 and 0.740 for Models A-C, respectively.[18]

[17] As said before, it is sometimes argued that exclusion restrictions have to be made for identification of endogenous selection. Yet, this is only an issue in small samples and would be irrelevant here. However, in our case the common colonial relationship dummy may be excluded from the extensive margin model for stochastic reasons and the exclusion of $V_{ij,t-1}$ and $V_{ij,0}$ from the outcome equation is dictated by the model in Section 2.

[18] We would like to emphasize that the results for Models A and B are quite similar and even those for Model A and C compare closely. For instance, the correlation coefficient of $\ln \hat{\mu}_{it} + \ln \hat{m}_{jt}$ between Models A and B 0.982 and the one for Models A and C to 0.996. The high corresponding correlation coefficients suggest that the estimated multilateral resistance terms are quite similar across estimated models. The same holds true for estimated $\ln X_{ijt}$ at $V_{ijt} = 1$ where the correlation coefficient between Models A and B amounts to 0.977 and the one between Models A and C amounts to 0.997. The correlation coefficient for the predicted V_{ijt} between Models A and B amounts to 0.932. V_{ijt} is taken as given in Model C and we know from Table 3 that the correlation coefficients between observed and predicted V_{ijt} in Models A and B amount to 0.520 and 0.593, respectively. Obviously, a disadvantage of Model C is that counterfactual experiments may not display an impact of changes in fundamentals on V_{ijt}, since the latter is fixed to the observed value which is inconsistent with general equilibrium.

The estimation results in Table 3 suggest the following conclusions. First, the positive and highly significant coefficient of previous exporting clearly points to the importance of dynamics and path dependence at the extensive margin of bilateral exports. We estimate the impact of knowledge-creation through first market entry for subsequent exporters to that market at a fixed-cost reduction of about $100 \cdot e^{0.431} - 100 \simeq 53.87\%$. Hence, dynamic market entry plays a role *beyond* contemporaneous (or conditional on) fundamentals so that static model results would be misleading. The parameter estimates in the semiparametric selection Model B are comparable to their parametric counterparts in Model A. Finally, the point estimates and standard errors on ρ_η and ρ_ε – i.e., correlation of the disturbances between the processes of $\widetilde{V}^*_{ijt}$ and $\ln X_{ijt}$ – suggests that contemporaneous export status V_{ijt} should not be treated as exogenous (as in Model C) but as a Bernoulli response variable (as in the other models).

Regarding the role of variable trade costs for the extensive and the intensive margin, we find that all elements of $\ln \tau_{ijt}^{1-\sigma}$ display negative parameters (α_k) which are highly significantly different from zero in Model A. Hence, variable trade barriers of any kind specified deter both the probability to export at all for country-pairs and, at $V_{ijt} = 1$, the volume of bilateral exports.

5 Counterfactual analysis

5.1 Preliminaries

With (5)-(7) and (13)-(14), we can now conduct a counterfactual analysis of changes in the variables underlying $\tau_{ijt}^{1-\sigma}$ and f_{ijt} as well as of changes in factor endowments, L_{it}, and the inverse of total factor productivity, a_{it}. For this, note that the level of a_{it} is hard to measure. However, defining real output as $\Upsilon_{it} = n_{it}\overline{y}_{it}$, with $\overline{y}_{it} \equiv \sum_j^J x_{ijt}$, and aggregate tariff income of country i in year t as Ξ_{it}, and using these terms in the definition of nominal GDP, $Y_{it} = \frac{\sigma}{\sigma-1} w_{it} a_{it} \Upsilon_{it} = w_{it} L_{it} + \Xi_{it}$, we obtain

$$a_{it} = \frac{\sigma - 1}{\sigma} \frac{L_{it}}{\Upsilon_{it}} \left(\frac{Y_{it}}{Y_{it} - \Xi_{it}} \right).$$

Now, the ratio of counterfactual to baseline inverse total factor productivity is

$$\frac{a_{it}^c}{a_{it}} = \frac{\Upsilon_{it}/L_{it}}{\Upsilon_{it}^c/L_{it}^c} \frac{1 - \Xi_{it}/Y_{it}}{1 - \Xi_{it}^c/Y_{it}^c}.$$

Hence, while the level of a_{it} is hard to measure, we can measure, for instance, the change of a_{it} over time, $\frac{a_{i,t+1}}{a_{it}}$, by the inverse change in real output per worker, $\left(\frac{\Upsilon_{it}/L_{it}}{\Upsilon_{i,t+1}/L_{i,t+1}} \right)$ (using GDP at constant producer prices) from period t to $t + 1$, together with the change of the trade-weighted ad-valorem tariff factor, $\frac{1-\Xi_{it}/Y_{it}}{1-\Xi_{it}^c/Y_{it}^c}$.[19]

Using $P_{it} \equiv m_{it}^{\frac{1}{\sigma-1}} \theta_{it}^{\frac{1}{1-\sigma}}$, we can define a measure of the equivalent variation as a measure of welfare change in percent (this is the change of real GDP in terms of consumer prices, Y_{it}/P_{it}, in percent) as

$$EV_{it} \equiv 100 \cdot \left(\frac{Y_{it}^c/P_{it}^c}{Y_{it}/P_{it}} - 1 \right).$$

[19] In their model of the determinants of export variety, Feenstra and Kee (2008) allow total factor productivity to be determined endogenously in a non-linear systems estimation approach. While we do not consider heterogeneous firms or responses of total factor productivity to endogenous variables, this would be principally possible also with our general equilibrium model. One could even allow tariff indicators to be endogenous and analyze a system of equations where only geographical (distance and absence of a common land border) and cultural trade barriers (absence of a common language, of a past colonial relationship, or of a common colonizer) along with factor endowments L_{it} would be exogenous. However this would push the importance of the adopted structural assumptions quite far, and we resort to stronger assumptions about exogeneity at the advantage of simplicity of an already complicated structural empirical general equilibrium model with path dependence at the extensive margin.

In general, we calculate changes between baseline and counterfactual equilibria based on the estimates of Model A for each experiment.

5.2 Design of experiments

Recall that, by design of our data-set, $t = 0$ corresponds to the initial year of 1992, while $t = 1, ..., 4$ correspond to $1995, 1998, 2001, 2004$. Hence, $V_{ij,t-1}$ refers to three years prior to the one referred to by t. For the analysis of the role of fundamental variables to the model on outcome, we will compute equilibria which are based on $\tau_{ijt}^{1-\sigma}$, f_{ijt}, L_{it}, a_{it}, and $V_{ij,t-1}$ as observed or estimated from data used for estimation. In general, we will use model predictions based on such values and parameters for the observation period, namely the years $1995, 1998, 2001, 2004$, as benchmark equilibrium values. Using estimated parameters from the data and assuming an elasticity of substitution of $\sigma = 5.74$,[20] we then consider four counterfactual equilibria for all countries and country-pairs for those years. To some extent, such an analysis is related to an impulse-response analysis in empirical macroeconomics. The four experiments considered are the following.

Freezing bilateral tariffs: For this experiment, we change tariff-related trade costs as captured by the indicator variables for quintiles of tariffs, which Models A-C are based upon, in 1998, 2001, and 2004 so as to eliminate the experienced tariff change since 1995 from the data. This

[20] We have derived all impulse-response results for three alternative levels of the elasticity of substitution, namely $\sigma \in \{5; 7; 10\}$. Since $\widehat{\alpha}_\kappa < 0$ and log ad-valorem tariff factors $\ln b_\kappa > 0$ for all quintiles $\kappa = 2, ..., 5$, our results suggest that $\widehat{\sigma} > 1$ throughout, which is consistent with the corresponding model assumption. However, there is variation about $\widehat{\sigma}$ across $\kappa = 2, ..., 5$, as expected, and the corresponding point estimates are in the range of $\widehat{\sigma} \in [4.36, ..., 6.79]$. Since the number of observations in each of the upper four quintiles is about the same, the average value of $\widehat{\sigma} \simeq 5.74$. The latter seems plausible against the background of previous work at the aggregate level of bilateral trade (see Anderson and van Wincoop, 2003; or Bergstrand, Egger, and Larch, 2009). Hence, our estimates are broadly in line with the assumption of $\sigma = 5$.

leads to counterfactual levels of $\tau_{ijt}^{c1-\sigma}$ and f_{ijt}^c which in turn lead to counterfactual export status $V_{ij,t-1}^c$ for $t-1=2,3$, but it leaves L_{it} and a_{it} for every year t as observed in the data.

Freezing past export status: For this experiment, we keep $V_{ij,t-1}$ constant at its level in 1995. Hence, outcome may change only in response to contemporaneous changes in $\tau_{ijt}^{1-\sigma}$, f_{ijt}, L_{it}, and a_{it} as observed in the data between 1995 and 2004 for any ijt, but these changes may not stimulate dynamic adjustment at the extensive country margin of trade.

Freezing labor endowments: For this experiment, we set L_{it}^c in each year after 1995 to the level L_{i1}, which corresponds to 1995. Inter alia, this leads to changes at the extensive margin so that $V_{ij,t-1}^c \neq V_{ij,t-1}$ in $t-1=2,3$. All other variables such as $\tau_{ijt}^{1-\sigma}$, f_{ijt}, and a_{it} are as observed for any ijt.

Freezing total factor productivity: For this experiment, we set a_{it}^c to its level of 1995 in every year after 1995 but let $\tau_{ijt}^{1-\sigma}$, f_{ijt}, and L_{it} change as observed in the data for any ijt. Again, this will lead to a change in aggregate bilateral export status so that also $V_{ij,t-1}^c \neq V_{ij,t-1}$ in $t-1=2,3$.

Then, for each experiment we calculate counterfactual bilateral export flows (X_{ijt}^c), GDP (Y_{it}^c), price terms μ_{it}^c and m_{jt}^c, endogenous export status (V_{ijt}^c), and equivalent variation (EV_{it}^c) as described in Appendix 4.

A comparison of the four counterfactuals analyzed with the benchmark equilibrium for 2004 addresses the role of observed changes in all fundamental variables involved in our model.

– Tables 4a and 4b –

Table 4a summarizes average differences between 2004 and 1995 of $\widehat{\tau^{1-\sigma}}_{ijt}$ and $\widehat{f}_{ijt}$ based on parameter estimates and data. All changes are expressed in percent of the corresponding levels in 2004. In particular $\widehat{\tau_{ijt}^{1-\sigma}}$ and $\widehat{f}_{ijt}$ reflect weighted changes of tariffs according to the associated tariff quintiles country-pairs belong in. However, total bilateral fixed costs are composed not only of $\widehat{f}_{ijt}$ but also of $e^{-\widehat{\delta}\widehat{V}_{ij,t-1}}$ so that it is useful to report changes of $\widehat{\tau_{ijt}^{1-\sigma}}$ and $\widehat{f}_{ijt}$ along with ones of $e^{-\widehat{\delta}\widehat{V}_{ij,t-1}}$ and $e^{-\widehat{\delta}\widehat{V}_{ij,t-1}}\widehat{f}_{ijt}$ in Table 4a. In general, we have grouped countries into four blocs – EFTA members as of 2004,[21] EU members as of 2004,[22] NAFTA members,[23] and a Rest of the World which consists of the remaining 89 countries our estimates are based upon (see Appendix 2 for a detailed list). We report changes for average country-pairs within and across blocs of countries and, underneath those figures, standard deviations. This is done to illustrate that there is much variation both within and across blocs of countries in tariff-related impulses in variable and fixed trade costs. There are entries in the diagonal elements because these blocks consist of multiple countries.[24]

Table 4a points to relatively large differences in trade and fixed costs between 2004 and 1995. At first glance, it seems surprising that these changes

[21] *European Free Trade Agreement:* Iceland, Norway, and Switzerland.

[22] *European Union:* Austria, Belgium, Cyprus, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Latvia, Lithuania, Malta, Netherlands, Luxembourg, Poland, Portugal, Slovak Republic, Slovenia, Spain, Sweden, and United Kingdom.

[23] *North American Free Trade Agreement:* Canada, Mexico, and United States.

[24] In principal, one could even analyze changes within countries. However, $V_{ii,t-1} = 1$ for all $ii, t-1$, and tariffs do not change for intranational trade. That does not mean that there are no intranational responses to changes in foreign tariffs. Changes in intranational trade in response to bilateral variable or fixed trade costs are indirect responses to changes abroad. When comparing counterfactual with benchmark equilibria for nominal trade, we will report intranational and international responses of countries' outcomes to changes in fundamentals separately.

are not smaller for intra-EU relationships than for other blocks. However, we should bear in mind that we define the EU as of 2004 so that the figures in Table 4a account for the extensive liberalizations between the ten entrants to the Union in 2004 and the tariff liberalizations between the 15 incumbents and those entrants even prior to the Union's enlargement (see Egger and Larch, 2010). Negative figures in the upper left panel of Table 4a indicate that $\widehat{\tau_{ijt}^{1-\sigma}}$ was lower and, hence, trade costs were higher in 1995 than in 2004. For many cells in that panel, the corresponding differences were in the double digits in terms of 2004 levels of $\widehat{\tau_{ijt}^{1-\sigma}}$. In general, the variation of differences in $\widehat{\tau_{ijt}^{1-\sigma}}$ within the cells of the panel appears as big as the one across block-wise averages.

The upper right panel of Table 4a suggests that fixed costs $\widehat{f}_{ijt}$ of exporting to a new market were higher in 1995 than in 2004 within and between most country blocks on average. The average difference was smaller than the one in $\widehat{\tau_{ijt}^{1-\sigma}}$, though. However, the variance in that comparison is bigger within most cells than the one of averages across cells. Again, there was a relatively big change for intra-EU25 relationships over the observation period. The lower right panel of Table 4a suggests that average fixed bilateral market entry costs declined more extensively due to dynamic market entry than they would have without it. To see this, compare the cells in the panel at the lower right with the corresponding ones at the upper right of Table 4a.

Table 4b summarizes average changes in L_{it} and a_{it} for country blocs. Since both L_{it} and a_{it} are unilateral, there is no need for a bloc-by-bloc decomposition of the corresponding changes, unlike in Table 4a. Within all country blocks, the labor force was smaller in 1995 than in 2004 and more labor input was required to produce one unit of output on average prior to 2004. Estimated changes in technology between the reference years were obviously much larger than ones in labor endowments over the considered decade. As in Table 4a, the standard deviation of changes is meant to provide information about the variation of differences between benchmark and counterfactual values – i.e., of impulses within blocs at the end of the observation period.

Since there is large variation in the impulses across countries, country-pairs, and blocs thereof, we will also report averages and standard deviations of outcome responses across blocs below. The latter will not only vary through the heterogeneity of changes but also through the heterogeneity of levels of these variables in 1995.

Of course, since the model at stake is highly nonlinear, the total change in predicted outcome is not linearly separable into the ones contributed by the four fundamental variables. However, the magnitudes of the responses of outcome to the associated changes will shed light on the relative importance of these changes, given the magnitudes of observed or estimated changes in $\widehat{\tau_{ijt}^{1-\sigma}}$ and $e^{-\widehat{\delta}\widehat{V}_{ij,t-1}}$, L_{it}, a_{it}, and inertia at the extensive country margin of trade, $\widehat{V}_{ij,t-1}$.

5.3 Counterfactual versus benchmark equilibria

In Tables 5-8, we summarize the associated responses to the accumulated shocks described in Tables 4a and 4b on outcomes of interest. In particular, such outcomes are nominal bilateral flows at the extensive as well as the intensive country margins, X_{ijt} (see Table 5),[25] nominal trade flows at the

[25] At the level of country-pairs, X_{ijt} may be zero in the benchmark equilibrium, the counterfactual equilibrium, in neither situation or under either circumstance. We avoid loosing observations which entailed a change at the extensive country-pair margin of trade by aggregating total exports up by country block and then computing percentage changes after aggregating. We do not report the standard deviation of changes within blocks in

intensive country-pair margin only, i.e., X_{ijt} at given $X_{ijt} > 0$ in both the benchmark and the counterfactual equilibria (see Table 6), endogenous export status as a binary measure of the extensive country-pair margin of trade only, V_{ijt} (see Table 7), intra-national nominal sales, X_{iit} (see Table 8), the equivalent variation, EV_{it} (see Table 8), and the number of firms active, n_{it} (see Table 8). We do so for an elasticity of substitution of $\sigma = 5.74$, which is consistent with the data. Akin to the change in $\tau_{ijt}^{1-\sigma}$ and f_{ijt}, L_{it}, and a_{it}, all counterfactual equilibria are expressed as percentage changes relative to the benchmark equilibrium as of 2004. However, notice that outcomes in 2004 are informed by and depend on the history since 1995 (and the one before that). For the counterfactual equilibrium in 2004, it matters not only that but also when changes in fundamentals and associated responses happen.

– Tables 5-8 –

The figures in Table 5 suggest that, across the board, technological change and the changes in trade and fixed costs together were the most important drivers of aggregate nominal bilateral exports. There were significant impulses in trade and fixed costs and, consequently, sizable responses of intra-EU and EU-EFTA trade. Those accrued to the liberalization of tariffs with the entrants to the EU via-à-vis the EU incumbents as well as the EFTA countries and also other blocks. Notice that we compare counterfactuals with higher trade and fixed costs to benchmark ones with lower such costs. Hence, the upper left panel of Table 5 should be interpreted as to illustrate that the experienced reduction in tariffs led to large positive responses of nominal block-wise trade until 2004. The table suggests that technological progress was relatively more important than trade and fixed cost changes

that case to avoid dropping zeros.

that on average. This is not to say that (ad-valorem and fixed) trade costs are less important than technology as such. It rather means that the blunt tariff liberalization impulse in an already relatively liberalized world as of 1995 did not cause further strong responses of trade flows. In other words, the lion's share of tariff liberalizations occurred at times prior to our sample period – when information about applied tariffs was even scarcer than from 1995 onwards. Technological improvements appear to have had a much stronger impact at given liberalization than tariff changes did.[26]

Similar to Table 5, Table 6 reports responses of predicted changes in trade flows between 1995 and 2004 in a counterfactual configuration of fundamental variables and the one in a benchmark configuration over the observation period. Yet, in contrast to Table 5, we now focus on changes at the intensive country margin so that $V_{ijt} = 1$ is required in both the benchmark and the counterfactual equilibrium before aggregating trade flows and computing changes thereof per block. This allows us to report not only average block-wise changes but also the standard deviation in responses across the country-pairs behind each cell of the four panels in Table 6. Similar to changes in total aggregate trade in Table 5, where V_{ijt} could have been zero in either the counterfactual or the benchmark equilibrium, the estimated change in total factor productivity between 1995 and 2004 appears to have had the largest effect on nominal trade among the considered experiments. On average, trade among previously trading economies within and across country blocks would have grown by more than 20 percent less during the observation period, if technology had stayed constant after 1995 (see Table 6). The corresponding effect was particularly large within the EU25 block where a large fraction

[26] It may well be that trade itself was an important carrier to technological change, which lies beyond the possibilities of inference with the structure imposed here (see Feenstra and Kee, 2008, for evidence along those lines).

of both the exporters and the importers experienced dramatic productivity improvements.

The average country-pair in the sample would have seen a growth of trade at the intensive country margin which would have been slower by more than 12 percent without the growth of the labor force over the same time span (see Table 6). Not surprisingly, that effect was largest among the 90 ROW countries in the sample, where L_{it} changed the most according to Table 4b. The change in trade costs and fixed costs together exhibited an impact on the growth of nominal trade flows at the intensive margin which was not much less important than the one of changing labor force since 1995. The average country-pair would have seen a growth in trade which would have been almost 6 percent slower than in the benchmark situation between 1995 and 2004. Both the impulses and the responses were largest for EU-ROW, EFTA-ROW, and NAFTA-ROW trade. However, even the response of intra-EU25 trade was quite sizable. According to the lower right panel in Table 6, a lack of dynamic adjustment at the extensive country-pair margin and its associated impact on fixed costs alone would have led to 3.7 percent less of growth at the intensive country margin of trade than in the benchmark equilibrium. For that effect, it matters where and when changes in economic fundamentals such as trade and fixed costs, labor endowments, or technology occur.

A comparison of Tables 5 and 6 suggests where most of the changes at the extensive margin occurred. Table 7 considers this difference explicitly by focusing on changes in V_{ijt} rather than the difference between Tables 5 and 6 as such. The latter is interesting, but it is even more so in combination with Table 7, since changes at the extensive country margin may be composed of minor changes in the number of trading relationships but big jumps in values of trade or vice versa.[27]

Table 7 suggests, as expected, that a change in neither of the fundamental factors considered had a big impact on the extensive country margin of trade within or across EFTA, EU, and NAFTA. This is not surprising, since most of the countries maintained bilateral trade relations in the benchmark equilibria, and even the sizable impulses in fundamentals were not big enough to change that much. However, some of the impulses were as big in the ROW bloc, within and with which most of the existing zero trade flows in aggregate bilateral trade matrices occurred in the data. The extensive binary country margin of trade increased on average by about 6 percentage points in the average three-year interval between 1995 and 2004, according to Table 2. This change was mostly due to changes of relationships with or of the ROW. According to the upper left panel of the table, that change was mainly induced by a reduction in trade and fixed market entry costs. Growth of the labor force had a qualitatively similar but quantitatively less important effect. If anything, technology growth and market entry dynamics cushioned the stimulus of trade and fixed costs as well as population growth on the extensive country margin of trade with the ROW.[28]

Table 8 summarizes responses to the impulses in Tables 4a and 4b of three further outcomes: intra-national sales (or "trade", X_{iit}) for the average country in a block;[29] the change in the number of firms active;[30] and the

[27]This is also the case for other extensive margins of trade such as the extensive product margin analyzed in Feenstra and Kee (2008) or Kehoe and Ruhl (2009).

[28]Recall that, due to the nonlinearity of the model, we may not simply add up the changes in the four panels to arrive at the total predicted change of V_{ijt}. The proposed structural model predicts both average levels and average changes well. This suggests that changes in trade and fixed costs interact with changes in the labor force in a way so that the joint impact is significantly larger than the sum of te individual impacts.

[29]Since domestic sales are never zero, it suffices to consider X_{iit} without any distinction of the extensive and the intensive country margin, there.

[30]The latter is an aggregate, single-sector counterpart to the multi-sector analysis in Broda and Weinstein (2006), Feenstra and Kee (2008), or Kehoe and Ruhl (2009). How-

equivalent variation, which corresponds to the differences in the change of real GDP between an unobserved counterfactual configuration of economic fundamentals and the predicted benchmark.

The first panel in Table 8 pertains to nominal intra-national goods trade which is interesting to compare either to the diagonal or the row sum entries of the respective impulse-specific panels in Tables 5 or 6. Not surprisingly, changes in trade and fixed costs had a smaller impact on intra-national than on international trade. The reason is that such changes have direct effects only on transactions with foreign consumers. Hence, all the effects on nominal intra-national sales are indirect in scope. Notice also that responses differ qualitatively in our model or others allowing for changes at the extensive country margin of trade as captured by V_{ijt} (see Helpman, Melitz, and Rubinstein, 2008) from standard Krugman-type models. The reason is that trade costs affect real aggregate output, real output of the average firm, as well as the number of firms active, here. This is fundamentally different from models in the vein of Krugman (1979), Anderson (1979), or Anderson and van Wincoop (2003) where aggregate real output is independent of (ad valorem) trade costs. Partly for reasons of real effects of trade costs, average block-wise responses in intra-national sales tend to have the same sign as the responses of international sales (exports) in Tables 5 and 6. Moreover, effects on intranational sales are qualitatively similar to the ones on international trade, since effects on wages or mill prices affect either outcome in a similar way.

Had trade costs and fixed costs not declined, or the labor force, or technology, or the number of markets served per country not increased since 1995, then firm numbers would have grown slower (or even have declined)

ever, unlike there it is influenced by a dynamic process about fixed costs and market entry at the country margin consistent with general equilibrium in the proposed model.

in response.[31] In other words, the decline in trade and fixed costs and the expansion of the labor force, technological progress, and aggregate export market entry have each in isolation contributed to faster growth of the number of firms in the average country. This can be seen from the second panel in Table 8. Among those, market entry dynamics and the growth of the labor force were of the most and of about equal importance for the average country and country block. In particular, export market entry at the extensive country margin and the associated depression of fixed market service costs contributed to much growth of firm numbers in both the ROW as well as the EU as of 2004. The stimuli on firm numbers were biggest in the ROW, NAFTA, and EFTA. Trade and fixed cost reductions exhibited their biggest impact on the growth of firm numbers in the EU and ROW. Technological progress was of minor importance for the entry of firms or export market entry but was obviously more important for output per firm (firm size).[32]

The last panel of results in Table 8 suggests that reductions in trade and fixed costs since 1995 together had a relatively small welfare-enhancing impact, irrespective of which country block we look at. However, there was a big variation about that magnitude among the 90 countries in the ROW. The figures in the other columns suggest that dynamic adjustment at the extensive country margin, growth of the labor force and, in particular, technology improvements entailed much bigger stimuli for the growth of trade than trade and fixed costs. Had technological progress taken place in isolation, the average economy would have grown slower by more than 18 percent

[31] This is true on average with the model, estimates, and data at hand. Yet, the reduction in trade and fixed costs alone even triggered negative effects on intranational sales in the EU.

[32] To see this, consider the relatively large effects of technological progress on welfare in Table 8 and nominal trade flows in Tables 5-6 and contrast them with its small impact on market entry in Table 7 and firm numbers in Table 8.

(or roughly 1.8 percent per annum) over the last decade. Dynamic adjustment at the extensive margin (which is associated with lower fixed costs) explains a welfare change of about one-tenth of that magnitude on average. The realized growth of the labor force appears to have been about one-third less important for trade than technological progress was. Notice that the welfare effects reported in Table 8 are accumulated effects which depend not only on the difference in fundamentals between 1995 and 2004 but also the spacing of the associated difference in time. Since responses take time to accumulate, simple inference about welfare effects in static models as suggested by Arkolakis, Costinot, and Rodriguez-Clare (2009) is not possible in a dynamic setting as this one.

5.4 Impulse-response functions for average welfare and the intensive margin of trade

In Tables 5-8, we summarized responses of outcome to shocks in fundamentals by considering only the year 2004 for the comparison of counterfactual and benchmark equilibria. In part, these responses consisted of accumulated contemporaneous responses and amplified effects through dynamic adjustment at the extensive margin. It is the purpose of this subsection to disentangle accumulated immediate (contemporaneous) responses from the amplification effect accruing to dynamic adjustment through path dependence at the extensive country margin of trade. By such an analysis, we aim at disentangling dynamic from static gains from trade.

– Figure 1 –

In Figure 1, we display changes in response to impulses on the four fundamental variables across the four years 1995, 1998, 2001, and 2004. For the sake of brevity, we consider responses of the average country or country-pair in a period and over time. In general, one source of a dynamic pattern in responses is time pattern of impulses, and the other one is sluggish adjustment of outcome, in particular, at the extensive margin of exports. We aim at disentangling the two by displaying the total response by a blue line and the immediate response without dynamics by a red line in the figure.

Figure 1 contains six panels: three of them pertain to a response in equivalent variation (at the top; compare to the bottom panel of Table 8) and three to the intensive country margin of nominal bilateral exports (at the bottom; compare to Table 6). In a horizontal dimension, we report responses to alternative impulses: keeping trade and fixed costs (left), labor endowments (center), and labor input coefficients (labor productivity; right) constant at their levels in 1995 for all countries and country-pairs. In all panels, we consider responses of outcome between 1995 to 2007 to changes in fundamentals between 1995 and 2004 (i.e., there is one period outside of the sample period).

The six panels suggest that dynamic adjustment at the extensive country margin dampens the detrimental effects of shocks for the average country (with equivalent variation) or country pair (with nominal exports at the intensive country margin). At first glance, this seems surprising, since we see that there is a positive impact of lagged dependent market entry on the probability of entering in any period. However, notice that the without-dynamics loci are based on equilibria which do not consider adjustments of $V_{ij,t-1}$ across time but enforce immediate adjustment through resource and other general equilibrium constraints. Hence, V_{ijt} changes but only due to contemporaneous impulses in economic fundamentals. Ceteris paribus, this reduces the propensity to enter a randomly drawn new market. However, a contemporaneous detrimental impulse of fundamentals on outcome is cushioned by sluggish adjustment of $V_{ij,t-1}$. Some markets would not be served in the absence of a fixed-cost-reducing effect of path-dependent $V_{ij,t-1}$. Therefore, negative shocks of fundamentals will always be moderated by the aggregate learning effect through $V_{ij,t-1}$ as an argument of bilateral time-specific fixed market entry costs.

Moreover, Figure 1 illustrates that the biggest marginal responses happened at the beginning and the end of the sample period. The results for 2007 relative to 2004 suggest that path dependence at the extensive country margin triggers dynamic effects on outcomes such as welfare and nominal trade.

6 Conclusions

This paper formulates a structural general equilibrium model which involves adjustment dynamics at the extensive country margin of aggregate bilateral trade. We postulate that fixed costs of aggregate export market entry depend on the earlier presence of exporters from the same country in that market. Reasons for that may be learning or other forms of information exchange, establishing a public good character for knowledge about bilateral market access. Unless there are negative shocks or adverse changes in fundamentals, firms would then always serve a market if knowledge existed about it.

Otherwise, the model is a large-numbers monopolistic competition version of the framework of Dixit and Stiglitz (1977) or Krugman (1979). All firms in a market are homogeneous, do not segment export markets with respect to pricing, use labor as the only factor input, and exhibit the same productivity. While it would be straightforward to allow firms to be heterogeneous with regard to their total factor productivity (e.g., as in Melitz, 2003), previous work suggests that aggregate quantitative analysis can safely ignore such heterogeneity. Firms exhibit variable and fixed costs of serving a market, and profits are linearly separable across countries. Hence, firms may decide to stay out of a market if the associated profits do not cover the fixed costs of doing so without inducing direct effects on their activity in other markets.

Structural estimation of that model rests upon two pillars: a dynamic panel data discrete choice process for the extensive country margin of exports which is coupled stochastically and in terms of parameter restrictions with a panel data model for the intensive country margin of exports; and a nonlinear process of for multilateral trade balance (through multilateral trade resistance) which depends on the endogenous extensive country margin of trade. We estimate parametric and semi-parametric bivariate dynamic sample selection versions of that model.

The results can be summarized as follows. First, there is clear evidence of dynamic adjustment at the extensive margin of trade conditional on observable fundamentals of bilateral trade flows as suggested by the theoretical model. Second, the structural model points to differences in the relevance of four alternative drivers of bilateral trade: trade costs and fixed market entry costs; labor endowments; labor productivity; and market entry dynamics. The data suggest that, after 1995, changes in labor input coefficients and labor endowments were (much) more important drivers of both the extensive and the intensive margin of trade than contemporaneous trade and fixed cost changes. Part of the reason of this result are bigger impulses in labor productivity and endowments relative to changes in tariffs. However, there is a lot of variation in the responses across countries and country-pairs which does not only accrue to the heterogeneity of impulses in the decade after 1995 but also to the heterogeneity of country size as well as trade costs and market entry costs.

The paper sheds light on sizable *dynamic gains from trade*. Without market entry dynamics – i.e., in the absence of dynamic gains to exporters from knowledge acquisition about foreign market entry – the model predicts that negative shocks to trade would induce larger time-specific and accumulated responses of levels of trade or welfare, irrespective of whether the impulse is on contemporaneous trade and fixed market entry costs, labor endowments, or labor productivity. At the extensive margin of bilateral aggregate exports, market entry dynamics (e.g., knowledge acquisition about foreign markets) were almost as important as rising productivity on average.

References

Abramovitz, M. and I. Stegun (1964), *Handbook of Mathematical Functions With Formulas, Graphs and Mathematical Tables*, National Bureau of Standard Applied Mathematics, Washington DC.

Albornoz, F., H.F. Calvo Pardo, G. Corcos, and E. Ornelas (2010), Sequential exporting, CEP Discussion Paper no. 974, London School of Economics.

Anderson, J.E. (1979), A theoretical foundation for the gravity equation, *American Economic Review* **69**(1), 106-16.

Anderson, J.E. (2010), The incidence of gravity, in S. Brakman and P. Bergeijk (eds.), *The Gravity Model in International Trade: Advances and Applications*, Cambridge University Press.

Anderson, J.E. and E. van Wincoop (2003), Gravity with gravitas: A solution to the border puzzle, *American Economic Review* **93**(1), 170-192.

Arkolakis, C., A. Costinot, and A. Rodriguez-Clare (2009), New trade models, same old gains?, NBER Working Paper No. 15628.

Baier, S.L. and J.H. Bergstrand (2001), The growth of world trade: Tariffs, transport costs, and income similarity, *Journal of International Economics* **53**(1), 1-27.

Baier, S.L. and J.H. Bergstrand (2007), Do free trade agreements actually increase members' international trade?, *Journal of International Economics* **71**(1), 72-95.

Baier, S.L. and J.H. Bergstrand (2009), Bonus vetus OLS: A simple method for approximating international trade-cost effects using the gravity equation, *Journal of International Economics* **77**(1), 77-85.

Bergstrand, J.H., P. Egger, and M. Larch (2008), Gravity redux: Structural estimation of gravity equations, elasticities of substitution, and economic welfare under asymmetric bilateral trade costs, Working Paper, University of Notre-Dame.

Broda, C. and D.E. Weinstein (2006), Globalization and the gains from variety, *Quarterly Journal of Economics* **121**(2), 541-85.

Chaney, T. (2008), Distorted gravity: The intensive and extensive margins of international trade, *American Economic Review* **98**(4), 1707-21.

Cherkashin, I., S. Demidova, H.L. Kee, and K. Krishna (2009), Firm heterogeneity and costly trade: A new estimation strategy and policy experiments, unpublished manuscript, Pennsylvania State University.

Das, M., M. Roberts and J. Tybout (2007), Market entry costs, producer heterogeneity and export dynamics, *Econometrica* **75**(3), 837-73.

Dixit, A.K. and J.E. Stiglitz (1977), Monopolistic competition and optimum product diversity, *American Economic Review* **67**(3), 297-308.

Eaton, J. and S. Kortum (2002), Technology, geography, and trade, *Econometrica* **70**(5), 1741-79.

Egger, P. and M. Larch (2010), An assessment of the Europe Agreements' effects on bilateral trade, GDP, and welfare, *European Economic Review*, forthcoming.

Egger, P., M. Larch, K. Staub, and R. Winkelmann (2011), The trade effects of endogenous preferential trade agreements, *American Economic Journal: Economic Policy*, forthcoming.

Egger, P. and M. Pfaffermayr (2010), Bonus vetus OLS? Issues with the linearization of empirical new trade theory models, *unpublished manuscript*.

Evenett, S. and A.J. Venables (2002), Export growth by developing economies: Market entry and bilateral trade, unpublished manuscript, University of St. Gallen.

Feenstra, R.C. (2002), Border effects and the gravity equation: Consistent methods for estimation, *Scottish Journal of Political Economy* **49**(5), 491-506.

Feenstra, R.C. (2004), *Advanced International Trade: Theory and Evidence*, Princeton University Press.

Feenstra, R.C. and H.L. Kee (2008), Export variety and country productivity: Estimating the monopolistic competition model with endogenous productivity, *Journal of International Economics* **74**(2), 500-18.

Fossati, S. (2009), Dating U.S. Business Cycles with Macro Factors, *unpublished manuscript,* University of Washington.

Götz, G. (1999), Monopolistic competition and the diffusion of new technology, *Rand Journal of Economics* **30**(4), 679-93.

Hausmann, R. and D. Rodrik (2003), Economic development as self-discovery, *Journal of Development Economics* **72**(2), 603-33.

Helpman, E., M. Melitz, and Y. Rubinstein (2008), Estimating trade flows: Trading partners and trading volumes, *Quarterly Journal of Economics* **123**(2), 441-87.

Kee, H.L. and K. Krishna (2008), Firm level heterogeneous productivity and demand shocks: Evidence from Bangladesh, *American Economic Review* **98**(2), 457-62.

Kehoe, T.J. and K.J. Ruhl (2009), How important is the new goods margin in international trade?, Staff Report 324, Federal Reserve Bank of Minneapolis.

Krugman, P.R. (1979), Increasing returns, monopolistic competition, and international trade, *Journal of International Economics* **9**(4), 469-79.

Matthews, B.W., (1975), Comparison of the predicted and observed secondary structure of T4 phage lysozyme. Biochimica et Biophysica Acta **405**, 442–451.

Raymond, W., P. Mohnen, F. Palm, and S. Schim van der Loeff (2007), The behavior of the maximum likelihood estimator of dynamic panel data sample selection models, CESifo Working Paper No. 1992, May 2007.

Raymond, W., P. Mohnen, F. Palm and S. Schim van der Loeff (2007b), Persistence of innovation in Dutch manufacturing: Is it spurious?, *Review of Economics and Statistics* **92**(3), 495-504.

Reinganum, J.E. (1981), Market structure and the diffusion of new technology, *Bell Journal of Economics* **12**(2), 618-24.

Wansbeek, T. and E. Meijer (2000), *Measurement Error and Latent Variables*, Elsevier.

Wooldridge, J.A. (1995), Selection corrections for panel data models under conditional mean independence assumptions, *Journal of Econometrics* **68**(1), 115-32.

Wooldridge J.A. (2005), Simple solutions to the initial conditions problem in dynamic, nonlinear panel data models with unobserved heterogeneity, *Journal of Applied Econometrics* **20**(1), 39-54.

Appendix 1. Details on the maximum likelihood estimation procedure

Following Raymond, Mohnen, Palm, and Schim van der Loeff (2007, 2010), the likelihood of country-pair ij at period t, starting in $t = 1$ and conditional on the regressors in $\mathbf{w}_{V,ijt}$ (including the initial conditions) and $\mathbf{w}_{X,ijt}$ is given by terms in (25)-(26). We integrate out the country-pair-specific random effects $\eta_{V,ij}$ and $\eta_{X,ij}$ using a two-step Gauss-Hermite quadrature, which is based on

$$\int_{-\infty}^{\infty} e^{-z^2} f(z)dz \approx \sum_{m=1}^{M} w_m f(a_m),$$

where, e^{-z^2} plays the role of the normal density and $f(z)$ is any continuous function of z. w_m and a_m are the weights and abscissas, respectively, as defined by the Hermite polynomial (see, e.g., Abramovitz and Stegun, 1964), where m indexes to the integration points of which there are M. Use the transformation of the random variables $z_{V,ij} = \frac{\eta_{V,ij}}{\sigma_{V,\eta}\sqrt{2(1-\rho_\eta^2)}}$ and $z_{X,ij} = \frac{\eta_{X,ij}}{\sigma_{X,\eta}\sqrt{2(1-\rho_\eta^2)}}$ with the likelihood weights w_p and w_m and corresponding abscissas a_p and a_m. Then, we can approximate the likelihood function as

$$\begin{aligned} L_{ijt} \approx\ & \frac{\sqrt{(1-\rho_\eta^2)}}{\pi} \sum_{p=1}^{M} w_p \Pi_{t=1}^{T} \left(\frac{1}{\sigma_{X,\varepsilon}} \Phi \left(\frac{\ln X_{ijt} - B_{ijt} - a_p \sigma_{X,\eta} \sqrt{2(1-\rho_\eta^2)}}{\sigma_{X,\varepsilon}} \right) \right)^{V_{ijt}} \\ & \times \sum_{m=1}^{M} w_m \left(e^{2\rho_\eta a_p a_m} \Pi_{t=1}^{T} \left(\Phi \left(-A_{ijt} + a_m \sigma_{V,\eta} \sqrt{2(1-\rho_\eta^2)} \right) \right)^{1-V_{ijt}} \right) \\ & \times \Phi \left(\frac{-A_{ijt} + a_m \sigma_{V,\eta} \sqrt{2(1-\rho_\eta^2)} + \frac{\rho_\varepsilon}{\sigma_{V,\varepsilon}} \left(\ln X_{ijt} - B_{ijt} - a_p \sigma_{X,\eta} \sqrt{2(1-\rho_\eta^2)} \right)}{\sqrt{(1-\rho_\eta^2)}} \right). \end{aligned}$$

Note that the double integral in (25) is then approximated by a weighted double summation over all abscissa points a_p and a_m.

Appendix 2. List of included countries by continent

Africa (47 countries): Algeria, Angola, Benin, Botswana, Burkina Faso, Burundi, Cameroon, Cape Verde, Central African Republic, Chad, Comoros, Congo (Democratic Republic of), Congo (Republic of), Côte d'Ivoire, Djibouti, Egypt, Ethiopia, Gabon, Gambia, Ghana, Guinea, Guinea-Bissau, Kenya, Lesotho, Madagascar, Malawi, Mali, Mauritania, Mauritius, Morocco, Mozambique, Namibia, Niger, Nigeria, Rwanda, Senegal, Seychelles, Sierra, Leone, South Africa, Sudan, Swaziland, Tanzania, Togo (United Rep. of), Tunisia, Uganda, Zambia, Zimbabwe.

Americas (33 countries): Antigua and Barbuda, Argentina, Barbados, Belize, Bolivia, Brazil, Canada, Chile, Colombia, Costa Rica, Dominica, Dominican Republic, Ecuador, El Salvador, Grenada, Guatemala, Guyana, Haiti, Honduras, Jamaica, Mexico, Nicaragua, Panama, Paraguay, Peru, Saint Kitts and Nevis, Saint Lucia, Saint Vincent and the Grenadines, Suriname, Trinidad and Tobago, United States, Uruguay, Venezuela.

Asia (40 countries): Armenia, Azerbaijan, Bahrain, Bangladesh, Bhutan, Brunei Darussalam, Cambodia, China, Georgia, Hong Kong, India, Indonesia, Iran, Israel, Japan, Jordan, Kazakhstan, Korea, Kuwait, Kyrgyzstan, Lebanon, Malaysia, Maldives, Mongolia, Nepal, Oman, Pakistan, Philippines, Qatar, Russian Federation, Saudi Arabia, Singapore, Sri Lanka, Syrian Arab Republic, Tajikistan, Thailand, Turkmenistan, United Arab Emirates, Viet Nam, Yemen.

Europe (36 countries): Albania, Austria, Belarus, Belgium and Luxembourg, Bulgaria, Cyprus, Czech Republic, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Latvia, Lithuania, Luxembourg, Macedonia (former Yugoslav Rep. of), Malta, Moldova (Rep. of), Netherlands, Norway, Poland, Portugal, Romania, Slovak Republic, Slovenia, Spain, Sweden, Switzerland, Turkey, Ukraine, United Kingdom. **Pacific (9 countries):** Australia, Fiji, Kiribati, New Zealand, Papua New Guinea, Samoa, Solomon Islands, Tonga, Vanuatu.

Appendix 3. Imputation of tariffs and construction of tariff quintile indicators

Table 9 summarizes the parameter estimates of the (log-)linear econometric model which we used to impute bilateral log tariff factors, $\ln(1 + \text{tariff rate}_{ijt})$. The estimated model includes, inter alia, $\frac{1}{J}\sum_{i=1}^{J}\ln(1+\text{tariff rate}_{ijt})$ as a regressor. Usually, one would avoid doing so to prevent an endogeneity bias. However, since we are interested in imputation rather than causal analysis in Table 9, such a procedure is innocuous.

– Table 9 –

Notice that the models we employ in Table 3 are based on 57,120 observations for which we need tariff quintiles. Average (trade-weighted) applied bilateral tariff rates are non-missing for 42,537 observations. Hence, 14,583 (about one-quarter) of the bilateral log tariff factors have to be imputed. The regression in Table 9 is based on a larger number of data-points than the regressions in Table 3 are. This helps predicting tariff rates in the earlier years of the sample period the export models are based upon. Some of the imputed observations use data before and after missing data-points, most of them are informed by non-missing bilateral tariffs in later years of the sample period. The imputation models work relatively well with a within R^2 of almost 40%.

We the use the $57,120$ observations on partly imputed bilateral log tariff factors ($\ln(1+\text{tariff rate}_{ijt})$) and allot them into quintiles. Finally, we define five binary indicator variables capturing which quintile of $\ln(1+\text{tariff rate}_{ijt})$ exports from i to j in year t are associated with. The use of tariff quintiles rather than actually observed and imputed tariff rates helps reducing the measurement error (see Wansbeek and Meijer, 2000).

Appendix 4. Solving the fully nonlinear model in counterfactual equilibrium

Based on known (or estimated) parameters including σ, known counterfactual GDP shares θ_{it}^c and $\breve{\theta}_{it}^c$, and counterfactual trade barriers $(\tau_{it}^{1-\sigma})^c$ and f_{it}^c for each period, we may solve for counterfactual trade resistance terms from the system (14) by using

$$V_{ijt}^c = 1\left[\ln\left(\frac{\tau_{ijt}^{1-\sigma}}{\tau_{iit}^{1-\sigma}}\right)^c + \ln\frac{m_{jt}^c}{m_{it}^c} + \ln\frac{f_{iit}^c}{f_{ijt}^c} + \delta(V_{ij,t-1}^c - 1)\right], \qquad (27)$$

where $\tau_{ijt}^{1-\sigma}$ and f_{ijt}^c depend on the same variables capturing trade barriers by assumption. Of course, $\theta_{it}^c = Y_{it}^c/(\sum_{i=1}^J Y_{it}^c)$ is not observed, but Y_{it}^c it can be solved for by using

$$\begin{aligned} Y_{it}^c &= \frac{p_{it}^c}{p_{it}}\frac{\Upsilon_{it}^c}{\Upsilon_{it}}Y_{it} = \left(\frac{\mu_{it}^c/n_{it}^c}{\mu_{it}/n_{it}}\right)^{\frac{1}{1-\sigma}}\frac{\Upsilon_{it}^c}{\Upsilon_{it}}Y_i = \\ &= \left(\frac{\mu_{it}^c}{\mu_{it}}\right)^{\frac{1}{1-\sigma}}\left(\frac{L_{it}^c}{L_{it}}\right)^{\frac{\sigma}{\sigma-1}}\left(\frac{\sum_{j=1}^J V_{ijt}^c e^{-\delta V_{ij,t-1}^c} f_{ijt}^c}{\sum_{j=1}^J V_{ijt} e^{-\delta V_{ij,t-1}} f_{ijt}}\right)^{\frac{1}{1-\sigma}}\frac{a_{it}}{a_{it}^c}Y_{it}, \qquad (28)\end{aligned}$$

where we employed $\Upsilon_{it} \equiv n_{it}\overline{y}_{it}$ and $\overline{y}_{it} \equiv \sum_j^J x_{ijt}$ for the baseline scenario and an analogous definition for Υ_{it}^c. Moreover, we used $p_{it} = (\mu_{it}/n_{it})^{\frac{1}{1-\sigma}}$ from (11) and assume throughout that $f_{iit}^c = f_{iit}$. For estimation, replace estimates of V_{ijt} by ones of V_{ijt}^c from (27) and Y_{it} by Y_{it}^c from (28) in (14). In particular, use $\widehat{V}^c_{ijt} = 1[P(\ln \widehat{\widetilde{V}}^c_{ijt} > 0) > \frac{1}{TN(N-1)} \sum_{t=1}^{T} \sum_{j=1}^{J} \sum_{i\neq j} P(\ln \widehat{\widetilde{V}}^c_{ijt} > 0)]$ as an estimate for V^c_{ijt} in (27).

Notice that (14) and (27)-(28) have to be solved simultaneously (or iteratively until convergence), since, in counterfactual equilibrium, (14) depends on (27) and (28) both of which are a function of the multilateral resistance terms in (14).

The counterfactual analysis requires the prediction of the exporter status at the country pair level (V_{ijt}) for both the baseline scenario and the counterfactual. For constructing a predicted binary indicator $\hat{V}_{ijt}$ based on the continuous $\widehat{\widetilde{V}}^*_{ijt}$, we follow Fosatti (2009) and minimize a cost-weighted misclassification cost function in a grid search to obtain these predictions:

$$\begin{aligned} \widehat{V}_{ijt} &= 1 \text{ if } \Phi\left(\widehat{\widetilde{V}}^*_{ijt}\right) > c^*_t & (29) \\ c^*_t &= \arg\min_{c_t} \sum_{i=1}^{J} \sum_{j=1}^{J} (1-q)\, V_{ijt} \left(1 - \widehat{V}_{ijt}\right) + q\,(1 - V_{ijt})\, \widehat{V}_{ijt}, & (30) \end{aligned}$$

where the weights are given by $q \in [0.535, 0.585, 0.620., 0.623]$ for periods 1995, 1998, 2001, and 2004, respectively. These weights are chosen to minimize the difference in the share of predicted versus observed non-zero exports. Table 3 summarizes measures of goodness of fit for Models A-C.

Table 1 - Persistence of the extensive margin of bilateral exports in 120 countries
(Included years are 1995, 1998, 2001, 2004, and 1992 as the initial period)

Country-pairs with	Fraction	Country-pairs with	Fraction
Exports in t and t-3	0.66	Exports in 2004 and 1992	0.52
Exports neither in t nor in t-3	0.20	Exports neither in 2004 nor in 1992	0.20
Exports in t but not in t-3	0.10	Exports in 2004 but not in 1992	0.27
Exports in t-3 but not in t	0.03	Exports in 1992 but not in 2004	0.01

Table 2 - Descriptive statistics on log exports, on elements of $(1-\sigma)\ln \tau_{ijt}$ and $\ln f_{ijt}$
(120 countries; included years are 1992, 1995, 1998, 2001, and 2004)

	Levels		Three-year changes	
Variables	Mean	Std.dev.	Mean	Std.dev.
Log bilateral exports ($\ln X_{ijt}$)	1.132	3.826	0.301	1.243
Extensive margin of bilateral exports (V_{ijt})	0.715	0.451	0.064	0.357
Lagged extensive margin of bilateral exports ($V_{ij,t-1}$)	0.698	0.459	-	-
Initial condition for extensive margin of bilateral exports ($V_{ij,0}$)	0.529	0.499	-	-
GDP share (θ_{it})	0.008	0.031	0.000	0.003
Variables in $(1-\sigma)\ln \tau_{ijt}$ and $\ln f_{ijt}$				
Lowest quintile of bilateral tariffs (binary)	0.214	0.410	0.040	0.285
Second quintile of bilateral tariffs (binary)	0.204	0.403	0.012	0.381
Third quintile of bilateral tariffs (binary)	0.193	0.395	-0.012	0.414
Fourth quintile of bilateral tariffs (binary)	0.194	0.396	-0.003	0.416
Highest quintile of bilateral tariffs (binary)	0.194	0.396	-0.037	0.334
Log bilateral distance	3.792	1.510	-	-
Non-contiguity (binary)	0.972	0.165	-	-
No common language (binary)	0.843	0.364	-	-
No colonial relationship (binary)	0.897	0.304	-	-
No common colonizer (binary)	0.976	0.152	-	-

Table 3 - Regression results for specification variants of the extensive and intensive country margins of bilateral exports (120 countries; 1995-2004)

Determinants of bilateral exports			Model A	Model B	Model C
			Extensive margin		
Variables in ln f_{ijt}	Acronym	Param.	(Dep. is Bernoulli response variable V_{ijt})		(V_{ijt} exogenous)
Lowest quintile of bilateral tariffs (binary)	$\ln \chi_1 + \ln \zeta_1$	$\beta_1 + \alpha_1$	Basis	Basis	-
			Basis	Basis	-
Second quintile of bilateral tariffs (binary)	$\ln \chi_2 + \ln \zeta_2$	$\beta_2 + \alpha_2$	-0.225 ***	-0.273 ***	-
			0.056	0.056	-
Third quintile of bilateral tariffs (binary)	$\ln \chi_3 + \ln \zeta_3$	$\beta_3 + \alpha_3$	-0.396 ***	-0.486 ***	-
			0.059	0.059	-
Fourth quintile of bilateral tariffs (binary)	$\ln \chi_4 + \ln \zeta_4$	$\beta_4 + \alpha_4$	-0.671 ***	-0.753 ***	-
			0.060	0.060	-
Highest quintile of bilateral tariffs (binary)	$\ln \chi_5 + \ln \zeta_5$	$\beta_5 + \alpha_5$	-0.737 ***	-0.786 ***	-
			0.063	0.063	-
Log bilateral distance	$\ln \chi_6 + \ln \zeta_6$	$\beta_6 + \alpha_6$	-1.564 ***	-1.623 ***	-
			0.021	0.022	-
Non-contiguity (binary)	$\ln \chi_7 + \ln \zeta_7$	$\beta_7 + \alpha_7$	-3.297 ***	-4.283 ***	-
			0.575	1.128	-
No common language (binary)	$\ln \chi_8 + \ln \zeta_8$	$\beta_8 + \alpha_8$	-0.226 ***	-0.224 ***	-
			0.096	0.097	-
No common colonizer (binary)	$\ln \chi_9 + \ln \zeta_9$	$\beta_9 + \alpha_9$	-0.419 ***	-0.571 ***	-
			0.118	0.111	-
Lagged dependent variable	$V_{ij,t-1}$	δ	0.431 ***	0.390 ***	-
			0.027	0.027	-
Initial condition	$V_{ij,0}$	λ_{V0}	3.656 ***	6.483 ***	-
			0.068	0.092	-
Goodness of fit [a)]			0.520	0.593	
			Intensive margin		
Variables in $(1-\sigma) \ln \tau_{ijt}$	Acronym	Param.	(Dependent variable is $\ln X_{ijt}$)		
Lowest quintile of bilateral tariffs (binary)	$\ln \zeta_1$	α_1	Basis	Basis	Basis
			Basis	Basis	Basis
Second quintile of bilateral tariffs (binary)	$\ln \zeta_2$	α_2	-0.169 ***	0.002	-0.069 ***
			0.026	0.025	0.025
Third quintile of bilateral tariffs (binary)	$\ln \zeta_3$	α_3	-0.425 ***	-0.162 ***	-0.264 ***
			0.028	0.028	0.028
Fourth quintile of bilateral tariffs (binary)	$\ln \zeta_4$	α_4	-0.747 ***	-0.407 ***	-0.544 ***
			0.030	0.030	0.030
Highest quintile of bilateral tariffs (binary)	$\ln \zeta_5$	α_5	-0.849 ***	-0.454 ***	-0.602 ***
			0.032	0.032	0.032
Log bilateral distance	$\ln \zeta_6$	α_6	-1.125 ***	-0.963 ***	-1.186 ***
			0.013	0.018	0.018
Non-contiguity (binary)	$\ln \zeta_7$	α_7	-1.517 ***	-1.273 ***	-1.141 ***
			0.142	0.207	0.163
No common language (binary)	$\ln \zeta_8$	α_8	-0.304 ***	-0.059	-0.207 ***
			0.049	0.068	0.072
No colonial relationship (binary)	$\ln \zeta_9$	α_9	-2.456 ***	-2.315 ***	-2.593 ***
			0.087	0.169	0.180
No common colonizer (binary)	$\ln \zeta_{10}$	α_{10}	-0.674 ***	-0.537 ***	-0.563 ***
			0.063	0.084	0.090
$\sigma_{V,\eta}$		$\sigma_{V,\eta}$	2.982 ***	3.085 ***	-
$\sigma_{X,\eta}$		$\sigma_{X,\eta}$	2.601 ***	2.352 ***	2.514 ***
ρ_η		ρ_η	0.720 ***	-	-
$\sigma_{X,\varepsilon}$		$\sigma_{X,\varepsilon}$	1.135 ***	1.048 ***	1.050 ***
ρ_ε		ρ_ε	-0.187 ***	-	-
Goodness of fit [b)]			0.744	0.743	0.740

Notes: The estimation includes the data for 1995, 1989, 2001 and 2004 using 1992 as starting values. The total number of observations is 57,120 out of which exhibits 43,896 strictly positive trade flows. ***, **, and * indicate significance levels of 1, 5 and 10 percent, respectively. Three time dummies are included in $(1-\sigma) \ln \tau_{ijt}$, but not reported.

Table 4a - Average change in trade costs and fixed costs across 4 blocks (120 countries; 1995-2004)
(All changes are in percent. Figures are based on data excluding intranational relations.)

	Changes in trade costs $\tau_{ijt}^{1-\sigma}$ in percent				
	Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	-12.5	-10.3	-8.0	-7.1	-8.2
	2.5	1.3	0.8	4.9	4.9
EFTA	-12.5	-10.3	-8.0	-7.1	-8.3
	2.5	1.4	0.8	4.9	4.9
NAFTA	-12.5	-10.3	-8.0	-7.1	-8.3
	2.5	1.4	0.9	4.9	4.9
ROW	-12.5	-10.3	-8.0	-7.1	-8.3
	2.5	1.3	0.8	4.9	4.9
Total	-12.5	-10.3	-8.0	-7.1	-8.3
	2.5	1.3	0.8	4.9	4.9

	Changes in fixed costs f_{ijt} in percent				
	Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	2.8	1.3	2.0	1.3	1.6
	4.7	6.3	0.8	5.7	5.5
EFTA	2.4	0.6	2.0	1.1	1.4
	0.5	0.2	0.8	4.7	4.1
NAFTA	2.4	0.6	2.0	0.9	1.2
	0.5	0.2	0.8	3.2	2.8
ROW	3.3	1.4	2.4	1.7	2.1
	7.1	6.6	4.9	6.5	6.6
Total	3.2	1.3	2.3	1.6	1.9
	6.5	6.4	4.3	6.2	6.3

	Changes in $\exp(-\delta V_{ij,t-1})$ in percent				
	Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	0.4	0.7	0.0	0.1	0.2
	4.6	6.4	0.0	5.7	5.4
EFTA	0.0	0.0	0.0	0.1	0.1
	0.0	0.0	0.0	5.5	4.8
NAFTA	0.0	0.0	0.0	-0.3	-0.2
	0.0	0.0	0.0	3.0	2.6
ROW	0.9	0.8	0.4	0.6	0.6
	6.9	6.5	4.6	6.4	6.5
Total	0.8	0.8	0.3	0.5	0.5
	6.3	6.3	4.0	6.2	6.2

	Changes in fixed costs $\exp(-\delta V_{ij,t-1})f_{ijt}$ in percent				
	Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	3.4	2.5	2.0	1.8	2.1
	11.9	16.1	0.8	13.3	13.0
EFTA	2.4	0.6	2.0	1.8	1.9
	0.5	0.2	0.8	13.0	11.3
NAFTA	2.4	0.6	2.0	0.8	1.1
	0.5	0.2	0.8	5.1	4.5
ROW	4.7	2.6	3.0	2.8	3.2
	18.0	16.6	12.2	16.0	16.4
Total	4.3	2.5	2.7	2.5	2.9
	16.5	16.2	10.6	15.3	15.5

Table 4b - Average change in labor endowment and input coefficients in four blocks (120 countries; 1995-2004)
(All changes are in percent.)

	Changes in percent	
Exporting block	Labor (L_{it})	Input coeff. (a_{it})
EU	-2.3	35.9
	5.2	21.4
EFTA	-6.5	22.3
	1.8	9.9
NAFTA	-9.8	22.8
	1.1	1.2
ROW	-13.6	22.7
	9.0	23.2
Total	-11.0	25.4
	9.3	22.9

Table 5 - Average responses of overall aggregate nominal bilateral trade (X_{ijt}) at the extensive and intensive margins in alternative counterfactual experiments across four country blocks (120 countries; 1995-2004)
(All changes are in percent. Figures are based on data excluding intranational relations.)

	Changes in trade and fixed costs				
	Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	-10.1	-9.1	-7.5	-10.5	-9.3
EFTA	-10.5	-9.5	-6.2	-11.8	-9.5
NAFTA	-11.5	-11.5	-8.0	-13.2	-11.1
ROW	-5.6	-4.6	-0.3	-6.0	-4.1
Total	-9.5	-8.7	-5.5	-10.4	-8.5

	Changes in labor endowments				
	Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	-3.8	-5.0	-9.0	-4.1	-5.5
EFTA	-4.5	-6.6	-9.3	-4.3	-6.2
NAFTA	-7.5	-8.0	-10.2	-9.5	-8.8
ROW	-7.5	-7.6	-15.2	-7.9	-9.6
Total	-5.8	-6.8	-10.9	-6.5	-7.5

	Changes in input coeffcients				
	Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	-24.3	-17.8	-24.1	-23.5	-22.4
EFTA	-16.6	-19.2	-24.7	-19.1	-19.9
NAFTA	-18.7	-15.6	-21.0	-13.1	-17.1
ROW	-29.4	-29.4	-30.2	-27.8	-29.2
Total	-22.3	-20.5	-25.0	-20.9	-22.2

	Changes in past export status (dynamics)				
	Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	-0.8	-1.1	-1.1	-3.6	-1.7
EFTA	-1.2	-1.4	-1.2	-4.3	-2.0
NAFTA	-1.2	-1.3	-0.9	-2.0	-1.4
ROW	-3.0	-3.4	-1.6	-4.7	-3.2
Total	-1.6	-1.8	-1.2	-3.7	-2.1

Table 6 - Average responses of aggregate nominal bilateral trade (X_{ijt}) at the intensive margin in alternative counterfactual experiments across four country blocks (120 countries; 1995-2004)
(All changes are in percent. Figures are based on data excluding intranational relations.)

	Changes in trade and fixed costs				
	Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	-10.5	-10.0	-6.9	-11.7	-11.3
	2.6	2.7	2.3	4.5	4.2
EFTA	-10.7	-10.1	-7.0	-11.8	-11.4
	1.7	2.2	1.0	4.1	3.7
NAFTA	-11.4	-10.9	-7.8	-12.7	-12.3
	1.7	1.8	1.2	4.0	3.6
ROW	-5.6	-5.1	-2.0	-6.7	-6.3
	4.7	4.9	4.8	6.1	5.9
Total	-6.8	-6.4	-3.2	-8.1	-7.6
	4.8	4.9	4.8	6.2	5.9

	Changes in labor endowments				
	Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	-2.5	-5.2	-7.5	-6.9	-5.9
	4.9	3.6	3.4	6.3	6.1
EFTA	-4.5	-7.3	-9.5	-8.7	-7.8
	3.6	1.2	1.3	5.3	5.2
NAFTA	-6.2	-8.7	-10.9	-10.2	-9.4
	3.4	1.3	1.0	5.1	4.9
ROW	-11.6	-13.9	-16.0	-15.5	-14.6
	7.4	6.7	6.5	7.9	7.9
Total	-9.5	-11.9	-14.0	-13.3	-12.5
	7.8	7.0	6.8	8.3	8.3

	Changes in input coeffcients				
	Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	-27.5	-23.9	-27.7	-20.4	-22.2
	8.1	7.7	7.1	10.4	10.2
EFTA	-24.2	-20.3	-24.2	-16.8	-18.7
	6.2	5.2	4.8	8.7	8.7
NAFTA	-21.1	-17.2	-21.3	-13.8	-15.5
	4.4	2.2	1.2	7.1	7.2
ROW	-25.8	-22.1	-26.1	-18.9	-20.7
	10.3	9.9	9.3	12.3	12.1
Total	-26.0	-22.3	-26.3	-19.0	-20.8
	9.7	9.4	8.8	11.8	11.6

	Changes in past export status (dynamics)				
	Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	-1.7	-1.5	-1.2	-3.1	-2.7
	1.8	1.2	1.2	2.4	2.3
EFTA	-1.7	-1.5	-1.2	-3.2	-2.8
	1.4	0.3	0.3	2.1	2.1
NAFTA	-1.4	-1.3	-1.0	-2.8	-2.5
	1.3	0.3	0.2	2.0	1.9
ROW	-3.0	-2.9	-2.6	-4.5	-4.0
	2.6	2.3	2.3	3.1	3.0
Total	-2.7	-2.5	-2.2	-4.1	-3.7
	2.5	2.2	2.1	3.0	2.9

Table 7 - Summary of average responses of the extensive binary country margin of aggregate nominal bilateral trade (V_{ijt}) in alternative counterfactual experiments across four country blocks (120 countries; 1995-2004)
(All changes are in percentage points. Figures are based on data excluding intranational relations.)

	Changes in trade and fixed costs						Changes in labor endowments				
	Importing block						Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total	Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	-0.2	0.0	0.0	-0.9	-0.7	EU	0.4	0.0	0.0	-0.1	0.0
	4.3	0.0	0.0	11.8	10.4		6.0	0.0	0.0	11.4	10.3
EFTA	0.0	0.0	0.0	-0.4	-0.3	EFTA	0.0	0.0	0.0	0.7	0.6
	0.0	0.0	0.0	13.6	11.8		0.0	0.0	0.0	12.2	10.6
NAFTA	0.0	0.0	0.0	0.0	0.0	NAFTA	0.0	0.0	0.0	0.4	0.3
	0.0	0.0	0.0	0.0	0.0		0.0	0.0	0.0	6.1	5.3
ROW	-0.8	-1.5	0.0	-1.1	-1.1	ROW	0.0	-0.4	0.0	-0.3	-0.2
	9.6	12.1	0.0	11.8	11.2		6.5	6.1	0.0	10.7	9.8
Total	-0.7	-1.1	0.0	-1.1	-1.0	Total	0.1	-0.3	0.0	-0.2	-0.1
	8.6	10.5	0.0	11.7	10.9		6.2	5.3	0.0	10.8	9.8

	Changes in input coeffcients						Changes in past export status (dynamics)				
	Importing block						Importing block				
Exporting block	EU	EFTA	NAFTA	ROW	Total	Exporting block	EU	EFTA	NAFTA	ROW	Total
EU	0.4	0.0	0.0	1.3	1.1	EU	0.9	0.0	0.0	1.7	1.4
	6.0	0.0	0.0	11.3	10.2		9.5	0.0	0.0	14.2	13.0
EFTA	0.0	0.0	0.0	2.2	1.7	EFTA	0.0	0.0	0.0	2.2	1.7
	0.0	0.0	0.0	17.1	14.9		0.0	0.0	0.0	17.1	14.9
NAFTA	0.0	0.0	0.0	0.7	0.6	NAFTA	0.0	0.0	0.0	1.1	0.8
	0.0	0.0	0.0	8.6	7.5		0.0	0.0	0.0	10.5	9.1
ROW	-0.9	-1.1	0.0	-0.2	-0.4	ROW	0.0	-0.4	0.0	-0.1	-0.1
	9.6	10.5	0.0	12.4	11.7		7.8	6.1	0.0	10.7	9.9
Total	-0.6	-0.8	0.0	0.2	0.0	Total	0.1	-0.3	0.0	0.3	0.3
	8.8	9.1	0.0	12.3	11.4		7.9	5.3	0.0	11.7	10.8

Table 8 - Average change in intra-national trade, equivalent variation (EV_{it}), and the number of firms (n_{it}) in four country blocks (120 countries; 1995-2004)
(All changes are in percent.)

	Change in intra-national sales X_{iit} in percent			
	Response to changes in percent in			
Exporting block	Trade and fixed costs (τ_{ijt}, f_{ijt})	Labor (L_{it})	Input coeff. (a_{it})	Dynamics (V_{ijt})
EU	1.7	-2.0	-20.9	-2.1
	2.7	5.7	12.8	2.9
EFTA	0.1	-4.9	-14.6	-1.7
	0.7	3.2	8.8	1.3
NAFTA	0.0	-7.6	-15.0	-1.1
	0.5	4.3	8.1	0.5
ROW	-0.1	-9.8	-12.1	-4.2
	2.0	10.4	13.0	4.5
Total	0.2	-8.4	-13.6	-3.7
	2.2	10.0	13.2	4.3

	Change in n_{it} in percent			
	Response to changes in percent in			
Exporting block	Trade and fixed costs (τ_{ijt}, f_{ijt})	Labor (L_{it})	Input coeff. (a_{it})	Dynamics (V_{ijt})
EU	0.1	-1.8	-1.5	-3.1
	3.4	7.2	2.7	8.6
EFTA	-0.7	-6.9	-2.7	-1.6
	2.2	3.0	1.6	1.1
NAFTA	-1.1	-10.0	-0.5	0.0
	0.4	2.0	1.1	1.1
ROW	0.3	-13.3	0.4	-12.0
	3.5	9.8	3.4	13.8
Total	0.2	-10.8	-0.1	-9.6
	3.4	10.2	3.3	13.2

	EV_{it} in percent			
	Response to changes in percent in			
Exporting block	Trade and fixed costs (τ_{ijt}, f_{ijt})	Labor (L_{it})	Input coeff. (a_{it})	Dynamics (V_{ijt})
EU	1.1	-2.3	-26.8	-0.7
	2.1	7.6	10.9	2.6
EFTA	0.8	-8.4	-19.5	-0.2
	2.1	2.9	6.9	0.5
NAFTA	0.9	-13.0	-20.8	0.2
	1.8	2.1	2.6	0.5
ROW	-0.3	-16.5	-15.7	-3.3
	2.9	11.9	15.9	4.2
Total	0.0	-13.4	-18.1	-2.6
	2.8	12.3	15.3	4.0

Table 9 - Imputation of missing bilateral ad-valorem tariff rates

Dependent variable is log(1+tariff rate$_{ijt}$)	Coef./Std.err.	
Importer and time specific average of log(1+tariff rate$_{ijt}$): ъ$_{jt}$	-2.567	***
	0.200	
ъ$_{jt}$ where i and j are neigbours	0.123	***
	0.029	
$\log(Y_{jt}/L_{jt})$	0.025	*
	0.015	
$\log(Y_{jt})$	-0.009	
	0.015	
$\log(Y_{it}/L_{it})$	-0.002	
	0.005	
$\log(Y_{it})$	0.007	
	0.006	
$\lvert\log(Y_{jt})-\log(Y_{it})\rvert$	0.002	***
	0.000	
$\lvert\log(Y_{it}/L_{it})-\log(Y_{jt}/L_{jt})\rvert$	0.003	***
	0.000	
$\log(Y_{jt}/L_{jt})$ x ъ$_{jt}$	-0.019	
	0.013	
$\log(Y_{jt})$ x ъ$_{jt}$	-0.048	***
	0.009	
$\log(Y_{jt}/L_{jt})$ x ъ$_{jt}$	-0.205	***
	0.006	
$\log(Y_{it})$ x ъ$_{jt}$	0.071	***
	0.004	
Overall trend	-0.002	***
R^2	0.393	
Observations	84282	

Notes: Constant, exporter fixed effects, importer fixed effects and importer specific time trends are not reported. Missing observations have been replaced by predictions from this regression. In a second step 5 dummies are generated that take the value one if the log(1+tariff rate$_{ijt}$) belongs to quintile k, k=1,...,5. Lastly, all countries which never report exports or receive imports are skipped both as exporter and importer. This leaves us with a balanced panel of 120 exporter and importer countries.

Figure 1 - Effects of fundamentals on the equivalent variation and nominal bilateral trade with and without path dependence (dynamics) at the extensive country margin of aggregate bilateral exports