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How confident can we be of CGE-based assessments of Free Trade Agreements? ☆

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Abstract

Computable General Equilibrium models, widely used for the analysis of Free Trade Agreements, are often criticized for having poor econometric foundations. This paper improves the linkage between econometric estimates of key parameters and their usage in CGE analysis in order to better evaluate the likely outcome of a Free Trade Area of the Americas (FTAA). Our econometric work focuses on estimation of a particular parameter, the elasticity of substitution among imports from different countries, which we show to be central to our evaluation of the normative impacts of the FTAA. We match the data in the econometric exercise to the policy experiment at hand, and employ both point estimates and the associated standard errors in our FTAA analysis which takes explicit account of the degree of uncertainty in the underlying parameters. In particular, we sample from the distribution of parameter values given by our econometric estimates in order to generate a distribution of model results, from which we can construct confidence intervals. We find that imports increase in all regions of the world as a result of the FTAA, and this outcome is robust to variation in the trade elasticities. Nine of the thirteen FTAA regions experience a welfare gain in which we are more than 95% confident. We conclude that there is great potential for combining econometric work with CGE-based policy analysis in order to produce a richer set of results that are likely to prove more satisfying to the sophisticated policy maker.

JEL classification: C15; C68; F11; F15

Keywords: Trade elasticities; General equilibrium; Systematic sensitivity analysis; Free Trade Agreements

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> "The econometric cat was set among the pigeons when a second government-commissioned modelling study on the FTA was finally released... The second reason for the contradictory results is differing assumptions about an arcane economic relationship known as Armington elasticity." (Australian Financial Review, 2003)

1. Introduction

With the proliferation of Free Trade Agreements (FTAs) over the past decade, demand for quantitative analysis of their likely impacts has surged. The main quantitative tool for performing such analysis is Computable General Equilibrium (CGE) modeling. Yet these models have been widely criticized for performing poorly (Kehoe, 2002) and having weak econometric foundations (McKitrick and Ross, 1998; Jorgenson, 1984). FTA results have been shown to be particularly sensitive to assumptions on the price elasticity of export demand (henceforth, the trade elasticity). As will be demonstrated in Section 2, small trade elasticities generate large terms of trade effects by reducing the responsiveness of export demand. On the other hand, small trade elasticities reduce the likelihood of trade diversion, as import sourcing becomes less sensitive to relative prices. Of course, large trade elasticities lead to the opposite results. Critics are understandably wary of results being heavily influenced by the authors' choice of trade elasticities. Indeed, the sensitivity of welfare results to the choice of trade elasticities has even surfaced in the popular press as witnessed in the opening quotation to this paper.[1]

Where do these trade elasticities come from? CGE modelers typically draw the elasticities from econometric work that uses time series price variation to identify an elasticity of substitution between domestic goods and composite imports (Alaouze, 1977; Alaouze et al., 1977; Stern et al., 1976; Gallaway et al., 2003). This approach has three problems: the use of point estimates as "truth", the downward bias in the magnitude of the point estimates created by problems in the estimation technique, and a mis-match between the data sample and source of variation in the econometric exercise and the policy experiment explored in the CGE exercise.

Consider the first problem. CGE modelers typically take point estimates drawn from the econometric literature, while ignoring the precision of these estimates. As we will make clear below, the confidence one has in various CGE conclusions depends critically on the size of the confidence interval around parameter estimates. Standard "robustness checks" such as systematically raising or lowering the substitution parameters fail to properly address this problem because they ignore information about which parameters are known with precision and which are highly uncertain.

A second problem with most existing studies derives from the use of import price series to identify home vs. foreign substitution. This approach tends to systematically understate the true elasticity because these estimates take price variation as exogenous when estimating the import demand functions, while ignoring quality variation. When quality is high, import demand and prices will be jointly high. This biases estimated elasticities toward zero. A related point is that the fixed-weight import price series used by most authors are theoretically inappropriate for estimating the elasticities of interest. CGE modelers generally examine a nested utility structure, with domestic production substituting for a CES composite import bundle. The appropriate price series is then the corresponding CES price index among foreign varieties. Constructing such an index requires knowledge of the elasticity of substitution among foreign varieties (see below). By

[1] In this article, two studies of the Australia-USA FTA are discussed, one which reports a gain, and one which reports a loss. Differing assumptions about the benefits of services liberalization was the first reason identified, while the second difference for the contradictory results was identified as the assumptions about the Armington elasticities.

using a fixed-weight import price series, previous estimates place too much weight on high foreign prices, and too small a weight on low foreign prices. In other words, they overstate the degree of price variation that exists, relative to a CES price index. Reconciling small trade volume movements with large import price series movements requires a small elasticity of substitution. This problem, along with unmeasured quality variation, helps explain why typical estimated elasticities are relatively small.

The third problem with the existing literature is that estimates taken from other researchers' studies typically employ different levels of aggregation, and exploit different sources of price variation, from what policy modelers have in mind. Employment of elasticities in experiments ill-matched to their original estimation can be problematic. For example, estimates may be calculated at a higher or lower level of aggregation than the level used in the policy analysis. Estimating substitutability across sources for paddy rice gives one a quite different answer than estimates that look at agriculture as a whole. In addition, when analyzing Free Trade Agreements, the principle policy experiment is a change in relative prices among foreign suppliers caused by lowering tariffs within the FTA. Understanding the substitution this will induce across those suppliers is critical to gauging the FTA's effects. Using home vs. foreign commodity elasticities rather than elasticities of substitution among imports supplied from different countries may be quite misleading. Moreover, these "sourcing" elasticities are critical for constructing composite import price series to appropriately estimate home vs. foreign substitutability.

In summary, the history of estimating the substitution elasticities governing trade flows in CGE models has been checkered at best. Yet these parameters are central to the welfare results of such studies. Clearly there is a need for improved econometric estimation of these trade elasticities that is well-integrated into the CGE modeling framework. This paper provides such estimation and integration, and has several significant merits. First, we choose our experiment carefully. Our CGE analysis focuses on the prospective Free Trade Agreement of the Americas (FTAA) which has periodically moved to the forefront of the negotiating agendas for countries in the Western Hemisphere. It is potentially one of the most important FTAs currently "in play".[2] It also fits nicely with the source data used to estimate the trade elasticities, which are largely based on world-wide imports into North and South America.

Our assessment is done in a perfectly competitive, comparative static setting in order to emphasize the role of the trade elasticities in determining the conventional gains/losses from such an FTA.[3] As highlighted by the quotation at the start of this paper, this type of model is still widely used by government agencies for the evaluation of such agreements. In fact, the GTAP model (Hertel, 1997) which we employ in this paper is actively used in dozens of public research institutions around the world. Extensions to incorporate imperfect competition are straightforward, but involve the introduction of additional parameters (markups, extent of unexploited scale

[2] Since the inception of this work, movement on the FTAA has come to a standstill. The summit of the Americas, hosted by Argentina in the spring of 2006, saw agricultural exporters in South America demanding reforms of US agricultural subsidieis. The US, on the other hand, maintained that discussion of such reforms belongs in the global talks under the auspices of the WTO. With the Summer, 2006, breakdown of the WTO negotiations, largely due to differences of opinion over agricultural subsidies, a return to the FTAA negotiating table seems unlikely in the near future.

[3] It is clear that this traditional, comparative static analysis is likely to understate the impacts of lowering trade barriers on trade flows and economic welfare. For example, imports have been shown to have beneficial, "procompetitive effects" on domestic markups (Levinsohn, 1993; Ianchovichina et al., 2000). Increased aggregate exports have also been associated with aggregate productivity gains as more productive firms expand their share of total sales (Bernard and Jensen, 2001). Finally, researchers have identified beneficial effects from increased foreign direct investment (Hallward-Driemeier et al., 2002), which is a likely by-product of an FTAA.

economies) as well as structural assumptions (entry/no-entry, nature of inter-firm rivalry) that introduce further uncertainty.

Since our focus is on the effects of a *preferential* FTA we estimate elasticities of substitution across multiple foreign supply sources by exploiting cross-sectional variation in delivered prices across many importer-exporter pairs. This technique fixes exporter supply characteristics, including factory gate prices, at a point in time. Delivered prices still vary across country pairs due to bilateral variation in ad-valorem trade costs (freight and tariffs). We exploit this variation in delivered prices to trace out import demand curves for each commodity and identify the elasticity of substitution.

We do not employ time series variation in prices. Exporter price series exhibit a high degree of multicolinearity, and in any case, would be subject to unmeasured quality variation as described previously. Similarly, cross-sectional tariff variation by itself would be of limited use since by their very nature, Most Favored Nation (MFN) tariffs are non-discriminatory, affecting all suppliers in the same way. Instead we employ a unique data set drawing on not only tariffs, but also bilateral transportation costs for goods traded internationally (Hummels, 1999). Transportation costs vary in cross-section much more widely than do tariffs, allowing more precise estimation of the trade elasticities that are central to CGE analysis of FTAs. We have highly disaggregated commodity trade flow data, and are therefore able to provide estimates that precisely match the commodity aggregation scheme employed in the subsequent CGE model. We follow the GTAP aggregation scheme which includes 40 merchandise commodities covering food products, natural resources and manufactures (Dimaranan and McDougall, 2002). Using this approach, we are able to estimate trade elasticities for all merchandise commodities that are significantly different form zero at the 95% confidence level.

Rather than producing point estimates of the resulting welfare effects, we report confidence intervals instead. These are based on repeated solution of the model, drawing from a distribution of trade elasticity estimates constructed based on the econometrically estimated standard errors. There is now a long history of CGE studies based on Systematic Sensitivity Analysis (SSA: Pagan and Shannon, 1987; Wigle, 1991; Harrison and Vinod, 1992; Harrison et al., 1993;) However, with the notable exception of Abdelkhalek and Dufour (1998), who estimate a two-parameter, one sector, CGE model for Morocco, all of these studies have taken their parameter distributions “from the literature”.[4] None of the disaggregated studies has been accompanied by an econometric study in which the key parameters and their distributions are estimated using data samples and variation that closely match the policy experiment considered in the CGE analysis.

For this paper, we use the Gaussian Quadrature (GQ) approach to SSA, which has proven to be the most efficient and unbiased approach to systematically assessing the sensitivity of model results to parametric uncertainty (DeVuyst and Preckel, 1997; Arndt, 1996). We find that many of the results are qualitatively robust to uncertainty in the trade elasticities. In those cases where our findings are not robust, we explore the source of underlying uncertainty. In this way, the paper addresses the fundamental question: How Robust Are CGE Analyses of Free Trade Agreements?

2. Explaining welfare changes: the role of trade elasticities

Due to the centrality of the trade elasticities to our argument, we begin by specifying the nested CES import demands. Expenditure on each composite commodity i in region s, E_{is}, is determined in general equilibrium by a combination of demand for the composite commodity in private

[4] Harrison et al. (1993) allude to the estimation of some parameters, but do not present these in their paper.

consumption, public consumption, investment demand and intermediate input demand. Therefore, for purposes of partial equilibrium estimation, we treat this expenditure level as exogenous, and focus on changes in its composition. The composite commodities are modeled as being a constant elasticity of substitution (CES) function of domestic and imported goods (1), and, at the second level of this preference structure, imports from different countries are combined in a CES function (2):

$$Q_{is} = \left[\beta_{\mathrm{Dis}} QD_{is}^{\frac{\varphi_i - 1}{\varphi_i}} + \beta_{\mathrm{Mis}} QM_{is}^{\frac{\varphi_i - 1}{\varphi_i}}\right]^{\frac{\varphi_i}{\varphi_i - 1}} \tag{1}$$

$$QM_{is} = \left[\sum_{r=1}^{R} b_{irs} \times QMS_{irs}^{\frac{\sigma_i - 1}{\sigma_i}}\right]^{\frac{\sigma_i}{\sigma_i - 1}} \tag{2}$$

Here, the index s denotes the importing region, Q_{is} is utility of consuming composite commodity i in this region, while QD_{is} is utility from domestically produced i, and QM_{is} is utility from composite imports (obtained from Eq. (2)). The parameters β_{Dis} and β_{Mis} represent commodity-specific preference weights on domestic versus imported goods, and φ_i is the elasticity of substitution between domestic and imported sources of good i in region s. We assume that this elasticity is equal across importing regions. In a similar fashion, the composite demand for imports (2) is a CES function of bilateral imports of i, sourced from different exporting regions r: QMS_{irs}. In Eq. (2), b_{irs} is a preference weight and σ_i is the elasticity of substitution among imports from different exporters. Again, we assume that this elasticity is identical across regions.

To determine aggregate demand for imports of commodity i into region s, the importing region maximizes Eq. (1), conditional on aggregate commodity expenditure for each commodity E_{is}, giving rise to the following import expenditure equation:

$$I_{is} = \frac{\beta_{\mathrm{Mis}}^{\varphi_{is}}}{PM_{is}^{\varphi_{is} - 1}} \times \frac{1}{P_{is}^{1 - \varphi_{is}}} E_{is} \tag{3}$$

where the composite commodity price index is given by:

$$P_{is} = [\beta_{\mathrm{Dis}}^{\varphi_{is}} PD_{is}^{1 - \varphi_{is}} + \beta_{\mathrm{Mis}}^{\varphi_{is}} PM_{is}^{1 - \varphi_{is}}]^{1/1 - \varphi_{is}}. \tag{4}$$

The optimal sourcing of imports from different exporters is obtained by maximizing (2), conditional on composite import spending I_{is}, given the import prices from different sources, PMS_{irs}:

$$QMS_{irs} = b_{irs}^{\sigma_i} I_{is} PMS_{irs}^{-1} \left[\frac{PMS_{irs}}{PM_{is}}\right]^{1 - \sigma_i} = b_{irs}^{\sigma_i} QM_{is} \left[\frac{PMS_{irs}}{PM_{is}}\right]^{-\sigma_i} \tag{5}$$

The price index over the imported commodities is given by:

$$PM_{is} = \left[\sum_{r=1}^{R} b_{irs}^{\sigma_i} (PMS_{irs})^{1 - \sigma_i}\right]^{1/1 - \sigma_i} \tag{6}$$

Changes in welfare in response to an FTA may be decomposed using the method of Huff and Hertel (1996), who provide an analytical decomposition of the Equivalent Variation (EV) for the representative household in region s.[5] It is similar in spirit to that of Baldwin and Venables (1995); however, unlike the latter decomposition, it allows for non-homothetic preferences, domestic taxes and subsidies, and, most importantly, it assumes products are differentiated by origin (Armington, 1969). This decomposition is also implemented numerically to decompose non-local welfare changes. For the sake of brevity, we focus on the case where there are no export taxes, and domestic taxes are applied only to consumption and production. (This assumption will be relaxed in the empirical analysis below.) As we will show, the elasticity of substitution is a key parameter determining not only the trade volume responses to the FTAA, but also their welfare implications.

The EV decomposition is given by Eq. (7) where the subscript i is indexed over the traded commodities, r denotes exporter region and s refers to the importing region. ψ_s is a scaling factor which is normalized to one initially, but changes as a function of the marginal cost of utility in the presence of non-homothetic preferences (McDougall, 2002).

$$\mathrm{EV}_s = (\psi_s)\left\{\begin{array}{l} \sum_{i=1}^{N}\sum_{r=1}^{R}(\tau_{\mathrm{Mirs}}PCIF_{irs}dQMS_{irs}) \\ +\sum_{i=1}^{N}(\tau_{\mathrm{CDis}}PD_{is}dQD_{is}) \\ +\sum_{i=1}^{N}(\tau_{\mathrm{CMis}}PM_{is}dQM_{is}) \\ +\sum_{i=1}^{N}(\tau_{\mathrm{Ois}}PD_{is}dQO_{is}) \\ +\sum_{i=1}^{N}\sum_{r=1}^{R}(QMS_{irs}dPFOB_{irs}) \\ -\sum_{i=1}^{N}\sum_{r=1}^{R}(QMS_{irs}dPCIF_{irs}) \end{array}\right\} \tag{7}$$

The first four summations on the right-hand side of Eq. (7) measure the changes in efficiency of resource utilization in region s. These involve the interaction of tax/subsidy distortions with the change in associated quantities. Consider what happens when we eliminate the bilateral tariff on imports of commodity i from one of the FTAA partner countries. The relevant term appears in the first summation:

$$\mathrm{EV}(\tau_{Mirs}) = \psi_s(\tau_{\mathrm{Mirs}}PCIF_{irs}dQMS_{irs}) \tag{8}$$

[5] The Huff-Hertel EV decomposition is obtained by starting with the equation for regional income as a function of endowment income, plus taxes less subsidies. Into this equation, they substitute the general equilibrium conditions for zero profits, price linkages, and market clearing for tradables and non-tradables. They deflate income by deducting the appropriate price index from both sides of the equation, thereby obtaining Eq. (7).

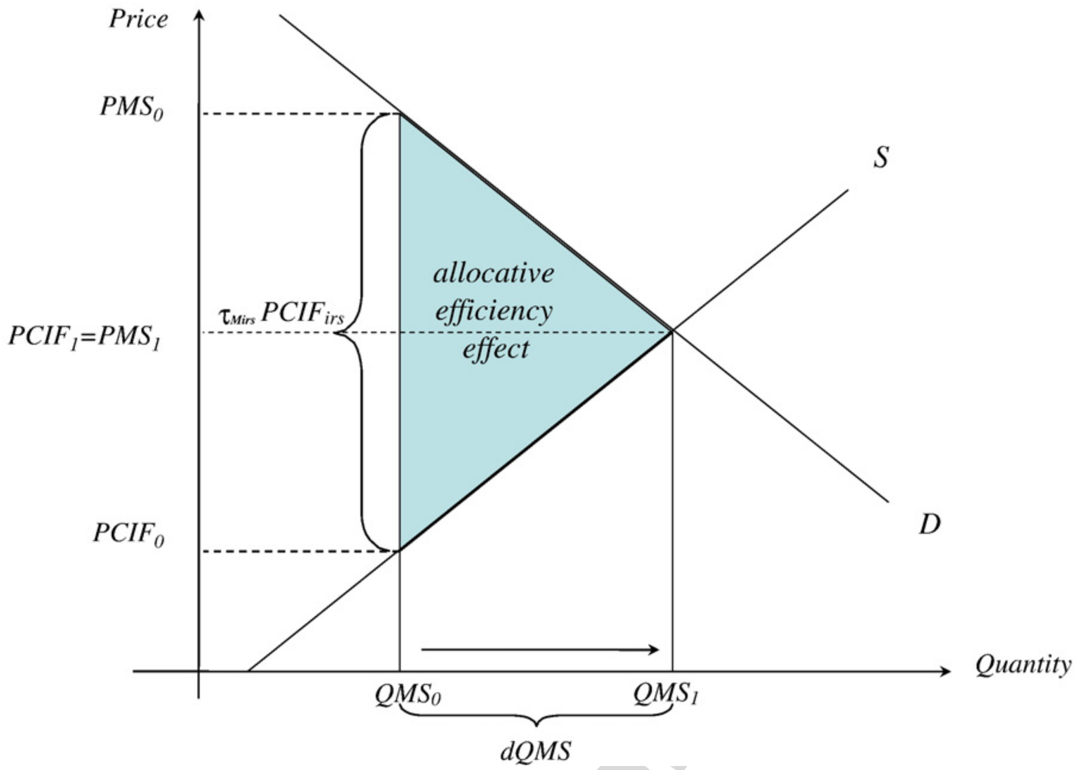

Fig. 1. Allocative efficiency gains from tariff elimination.

Here, (τ_{Mirs} $PCIF_{irs}$) is the per unit tariff revenue on imports of good i from r into s, associated with the *ad valorem* tariff rate τ_{Mirs}. This is multiplied by the change in the volume of imports of i from r into s: $dQMS_{irs}$. The "Harberger triangle" that we are measuring with this term may be seen in Fig. 1. In order to evaluate the area of this triangle as the tariff is eliminated, we must consider both the "base" ($\tau_{Mirs}PCIF_{irs}$) and the "height" ($dQMS_{irs}$).[6] By continually reevaluating the base of this triangle as the tariff is eliminated, we track the diminishing gap between *PCIF* and *PMS*. In this way, we are able to accurately measure its area, which is then added to the aggregate welfare measure.

In order to properly perform the numerical integration depicted in Fig. 1, the welfare decomposition equations must be solved in conjunction with the CGE model, using appropriate solution procedures. We use version 8 of the GEMPACK software suite (Harrison and Pearson, 2002) which is ideally suited to this problem, as it solves the non-linear CGE model using a linearized version of the behavioral equations, coupled with updating equations that link the change, in this case $dQMS_{irs}$, with the levels variables, QMS_{irs}. Standard extrapolation techniques can be used to obtain arbitrarily accurate solutions to any well-posed non-linear problem (Harrison and Pearson, 1996).[7]

Note from Eq. (7) that, in addition to tariffs, we consider volume interactions with consumption taxes on household purchases of both domestic goods (τ_{CDis}) and imported goods

[6] For those accustomed to computing "Harberger triangles" as 1/2 base*height, it may appear that we need a 1/2 premultiplying the right hand side of Eq. (8). However, this is not required. Our numerical integration procedure continually re-evaluates the base of the "triangle". Indeed, it ensures the accurate determination of the area between the supply and demand curves, even when these are non-linear.

[7] For purposes of this paper, we require that 95% of the variables and levels variables are accurate to four digits. Another useful check is to compare EVs computed from Eq. (7) with that computed directly from the utility function. These match, to machine accuracy.

(τ_{CMis}). Taxes (or subsidies) on output also play a role. If $\tau_{\mathrm{Ois}}<0$, then the production of commodity i in region s is subsidized and an expansion of output ($dQO_{is}>0$) will contribute negatively to efficiency and hence to EV. The absence of terms associated with income, export and input taxes, is due to the fact that we are assuming these taxes are zero in this stylized example. (In the empirical section below, this assumption will be relaxed.)

The final two terms on the RHS of Eq. (7) refer to the terms of trade (TOT) effects for region s. These determine how the global efficiency gains are shared amongst regions. If region s's export-weighted FOB prices rise, relative to her import-weighted CIF prices, then the TOT will improve. Since one region's export prices (inclusive of international transport services exports) are another region's import prices, the improved TOT for region s translates into a TOT deterioration in the rest of the world (taken as a group).

In summary, each region's welfare gains can be decomposed into terms of trade and allocative efficiency components. The essence of the FTAA experiment involves eliminating the trade taxes within the block, i.e., $\tau_{Mik\ell}=0$, $\forall i$ and $\forall k, \ell \in FTAA$. This, in turn, induces a shift in the sourcing of imports, away from exporters outside the block and towards exporters within the block. As seen from (5), the extent of this shifting depends on the Armington elasticity of substitution, σ_i. Log differentiating (5) and (6), and converting to percent changes (lower case denotes the percentage change in the associated upper case levels variables), we obtain the following two equations describing import sourcing in region s:

$$qms_{irs} = qm_{is} - \sigma_i[pms_{irs} - pm_{is}] \tag{9}$$

$$pm_{is} = \sum_{r=i}^{R} \Theta_{irs} pms_{irs} \tag{10}$$

where $pms_{irs}=(tm_{irs}+pcif_{irs})$, and tm_{irs} is the percentage change in $(1+\tau_{Mirs})$. The coefficient Θ_{irs} is the share of total import expenditure on i in s, sourced from region r.

Now, if we assume, for the sake of exposition, that *there are no domestic taxes whatsoever*, and we convert the simple changes in Eq. (7) into percentage changes, thereupon substituting in Eq. (9), we obtain the following decomposition of the local change in welfare of region s:

$$\mathrm{EV}_s = (\psi_s/100)\left\{\begin{array}{l} \displaystyle\sum_{i=1}^{N}\sum_{r=1}^{R}(TR_{irs})(qm_{is}-\sigma_i[(tm_{irs}+pcif_{irs})-pm_{is}]) \\ \displaystyle +\sum_{i=1}^{N}\sum_{r=1}^{R}(PFOB_{isr}QMS_{isr})pfob_{isr} \\ \displaystyle -\sum_{i=1}^{N}\sum_{r=1}^{R}(PCIF_{irs}QMS_{irs})pcif_{irs} \end{array}\right\} \tag{11}$$

where $TR_{irs}=(\tau_{\mathrm{Mirs}}PCIF_{irs}QMS_{irs})$=tariff revenue on commodity i imported from r into s. Provided region s is small, relative to supplier markets, it will have little impact on import prices ($pcif_{irs}\cong 0$). However, even a small country can have an impact on its own export prices in this differentiated product framework, so $pfob_{isr}\neq 0$. The size of the export price changes will be determined by the export demand elasticity, which approaches the value $-\sigma_i$ for a country that is small in its export markets ($\Theta_{isr} \cong 0$) (recall Eqs. (9) and (10)). This decomposition makes clear why, in our econometric exercise, we focus so intently on σ_i, the elasticity of substitution among imports. The welfare consequences of a single, small economy's FTA measures will depend first and foremost on the value of σ_i. Large values of σ_i will cause the elements of the first term in (11) to become larger in absolute value, as the shift in import sourcing becomes more pronounced. On the other hand, large values for σ_i, which determine the export demand elasticity facing region s, serve to dampen the changes in export price.

It is the first term in Eq. (11) which determines whether or not net trade diversion or trade creation takes place in this FTA. If, for example, $\sigma_i = 0$, then the pattern of import sourcing will remain unchanged and the sole effect of lower tariffs will be to lower the cost of composite imports (Eq. (10)), thereby leading to an increased demand for imports by Eq. (9) ($qms_{irs} = qm_{is} > 0$), and the efficiency gain collapses to: $\psi_s qm_{is} \sum_{i=1}^{N} \sum_{r=1}^{R} (\tau_{Mirs} PCIF_{irs} QMS_{irs})$. This is the case of pure trade creation.

In practice, we expect the value of σ_i to be quite large for most products—a point confirmed by the econometric work reported below. This means that the substitution term in Eq. (9) will be dominant in the determination of qms_{irs}. In this situation, the key issue is how the changes in bilateral imports, qms_{irs}, correlate with bilateral tariff revenues. Clearly, if the preferential FTA eliminates tariffs on bilateral flows that are already lightly taxed (low tariff revenue), while leaving in place tariffs on flows that are heavily taxed ($TR_{irs} \gg 0$) then there will be potential for trade diversion, as the bilateral import changes will be negatively correlated with tariff revenue (i.e., $qms_{irs} > 0$ for $TR_{irs} \cong 0$ and $qms_{irs} < 0$ for $TR_{irs} \gg 0$). In the ensuing empirical analysis of the FTAA, we will examine this trade creation/diversion effect in all participating countries. We will also explore its sensitivity to the econometrically estimated uncertainty in σ_i, a topic to which we now turn.

3. Econometric specification and estimation of trade elasticities

Our econometric estimation focuses on the second level of the two-level, Armington structure. The key parameter is σ_i, the elasticity of substitution among imports from different sources for a given commodity.[8] As described in the introduction, many papers estimate this parameter by examining time series variation in import prices and quantities. This approach is highly problematic if there are unmeasured factors, such as quality, that shift both supply and demand. In contrast, we follow the approach in Hummels (1999), who identifies σ_i by exploiting cross-sectional variation in delivered prices. We condition on an exporter and commodity, which fixes supply characteristics (quantity and quality supplied and FOB prices) at a point in time. We then identify the elasticity of substitution from variation over importers in delivered prices which arises from bilateral variation in ad-valorem trade costs.

More formally, the "power" of the trade cost for an imported commodity, T_{irs}, equals one plus the *ad valorem* rates for freight and insurance, as well as the tariffs, which vary by commodity i, importer s, and exporter r: $T_{irs} = (1 + \tau_{Firs} + \tau_{Mirs}) = 1 + freight_{irs} + tariff_{irs}$. Therefore, the delivered cost of imports is given by $PMS_{irs} = T_{irs} PFOB_{ir}$, which varies across importers, for a given exporter and commodity, only to the extent that trade costs vary.

[8] It is also the case that the data set that we have available only covers imports, and therefore does not lend itself to estimation of the upper level nest.

Because it is difficult to observe the quantity demanded, we multiply both sides of Eq. (5) by end-user prices, PMS_{irs} to get the amount of bilateral trade in value terms, V_{irs}:

$$V_{irs} = b_{irs}^{\sigma_i} I_{is} \left[\frac{PFOB_{ir} T_{irs}}{PM_{is}} \right]^{1-\sigma_i} \tag{12}$$

which, by taking natural logarithms results in Eq. (13):

$$\ln V_{irs} = \ln I_{is} + (1-\sigma_i)\ln PFOB_{ir} - (1-\sigma_i)\ln PM_{is} + (1-\sigma)_i \ln T_{irs} + \sigma_i \ln b_{irs} \tag{13}$$

It is commonly observed that countries with similar languages and cultures trade more with one another than would be predicted solely on the basis of observable trade costs. We model these effects as entering through bilateral variation in the commodity-specific preference parameters b_{irs}, treating them as a function of physical distance: Dist, similarity of language: Lang, and adjacency of the trading countries: Adj. In addition, the preference weights may reflect differences in quality b_{ir} that are specific to an exporter-commodity and commonly perceived by all importers. The specific functional relationship is as follows:

$$b_{irs} = b_{ir} \mathrm{Dist}_{rs}^{\delta_{1i}} e^{\delta_{2i} \mathrm{Lang}_{rs} + \delta_{3i} \mathrm{Adj}_{rs}} \tag{14}$$

Substituting Eq. (14) into Eq. (13), we obtain the following:

$$\begin{aligned} \ln V_{irs} = {} & \ln I_{is} + (1-\sigma_i)\ln PFOB_{ir} + (1-\sigma_i)\ln T_{irs} - (1-\sigma_i)\ln PM_{is} + \sigma_i \ln b_{ir} \\ & + \sigma_i \delta_{1i} \ln \mathrm{Dist}_{rs} + \sigma_i \delta_{2i} \mathrm{Lang}_{rs} + \sigma_i \delta_{3i} \mathrm{Adj}_{rs} \end{aligned} \tag{15}$$

To implement this equation, we include a vector of exporter-commodity intercepts in order to sweep out variation in supply characteristics (FOB prices, quality) at a point in time.[9] This makes all export sources comparable, excepting for differences in delivered prices arising from differences in freight costs and tariffs. Were we to estimate Eq. (15) separately for each importer and commodity, import expenditures I_{is} and importer's price index PM_{is} would enter the constant. However, this would offer too few observations to yield precise estimates. Instead, we stack import demands for multiple importers, allowing each to have its own intercept. That is, we include a vector of importer-commodity intercepts that sweep out importer expenditures and the importer's price index (which are both difficult to measure and endogenous). This gives rise to the final estimating equation:

$$\begin{aligned} \ln V_{irs} = {} & a_0 + a_{is} + a_{ir} + \beta_{0,i} \ln(1 + \mathrm{freight}_{irs} + \mathrm{tariff}_{irs}) + \beta_{1,i} \ln \mathrm{Dist}_{rs} + \beta_{2,i} \mathrm{Lang}_{rs} \\ & + \beta_{3,i} \mathrm{Adj}_{rs} + \varepsilon_{irs} \end{aligned} \tag{16}$$

where a_{is} and a_{ir} are vectors of importer-commodity and exporter-commodity intercepts. The parameter of interest, $\beta_{0,i} = 1-\sigma_i$, is identified from bilateral variation in trade costs. The parameter $\beta_{1,i} = \sigma_i \delta_{1i}$, where δ_{1i} is the price-equivalent preference shifter associated with distance, and similarly for $\beta_{2,i}$, $\beta_{3,i}$. We employ these preference shifters only as controls, and do not otherwise make use of the information in the δ_{Ki} parameters.

[9] A more general model might include the number of varieties supplied as a factor in the import demand equation. These too would be swept out by the fixed effects.

4. Data and elasticity estimates

The data used in estimation are taken from Hummels (1999). Given the emphasis in this study on the FTAA, it is appropriate that these data are a compilation of detailed customs information on imports into six FTAA countries (Argentina, Brazil, Chile, Paraguay, USA, and Uruguay) and one non-FTAA economy (New Zealand).[10] In order to estimate Eq. (16) we also require data on physical distance among the countries as well as comprehensive tariff data.[11] The final dataset contains 187,000 observations with the following variables: 5-digit SITC code of the commodity traded, *fob* and *cif* values for each trade flow, applied tariff rates, trade distance and two dummy variables to indicate common language or countries' adjacency. In addition, we dropped extreme observations where measured trade costs were either negative or greater than 4 times the *fob* value of the product.

At this point we face an interesting choice. We could aggregate the 5-digit, SITC trade flows and trade costs according to the 40 traded merchandise commodity groups used in our CGE model (Table 1). The advantage of this aggregation approach is that it exactly matches the data variation contained in the CGE exercise (i.e. a single value of trade for each bilateral pair in each of the 40 commodity groups). An alternative approach retains the variation across bilateral pairs and 5-digit level commodities within each of the 40 GTAP categories, constraining the elasticity of substitution to be equal within each broad sector. The main advantage of the pooling approach is that it provides greater within-sector variation in tariffs and transport costs which is useful for identifying the relevant substitution elasticities. We employ the pooling technique in order to yield more precise estimates.[12]

The results of Ordinary Least Squares estimation of Eq. (16) are presented in Table 1. Note that all of these estimates are positive and are significantly different from zero. Based on a simple t-test, each of the 40 estimated elasticities of substitution allows us to reject the hypothesis that the estimated elasticity is zero at the 95% confidence level.

Table 1 also contrasts these estimates with the original elasticities of substitution among imports from the version 5 GTAP database (Dimaranan and McDougall, 2002).[13] As noted previously, the GTAP parameters are widely used in the analysis of FTAs.[14] If we compute the simple average of the 40 estimates it is 7.0, which is somewhat larger than the average for the previous GTAP parameters (5.3). Although these two averages are fairly similar, there is much greater sectoral variation in the econometrically estimated elasticities. In fact, the most striking thing about the GTAP parameters is that they show no variability within broad sectors such as food and agriculture, and metal products. This is because the source studies were not conducted at a disaggregate level.[15]

[10] The choice of New Zealand as a non-FTAA economy is purely pragmatic. Data were readily available and transport costs to this country are generally quite high relative to the other importers, thus introducing additional variation in this key variable.

[11] For more information see the appendix in Hummels (1999).

[12] In the FTAA analysis section, we also discuss the results obtained from using the aggregation approach.

[13] The version 5 GTAP parameter file was taken from the SALTER project (Jomini et al., 1994). These trade elasticities are based on a synthesis of estimates from the literature and original econometric work for one country — New Zealand.

[14] For a sampling of these applications, visit the GTAP web: www.gtap.org).

[15] It is also striking to observe that the largest estimated elasticity of substitution is for natural gas (34.4) and the lowest is for other mineral products (1.8), yet for the GTAP model, both of these products are assigned the value of 5.6, corresponding to the generic estimate for natural resource products.

5. Application to the FTAA

5.1. Background

The recent growth of Free Trade Agreements in the Western Hemisphere began in the 1980s as Latin American countries initiated significant trade and economic policy reforms. Over the past two decades, more than forty regional or bilateral trade agreements have been implemented within the hemisphere. As the rapid growth of Free Trade Agreements

Table 1
Elasticities of substitution among imports from different sources

Sector	GTAP V5 Elasticity	Estimated elasticity	S.D.	Num. Obs.
Paddy rice	4.4	10.1*	4.0	26
Wheat	4.4	8.9*	4.2	32
Cereal grains nec	4.4	2.6*	1.1	131
Vegetables, fruit, nuts	4.4	3.7*	0.4	1199
Oil seeds	4.4	4.9*	0.8	239
Plant-based fibers	4.4	5.0*	2.4	71
Crops nec	4.4	6.5*	0.4	1796
Bovine cattle, sheep and goats, horses	5.6	4.0*	0.7	156
Animal products nec	5.6	2.6*	0.3	813
Wool, silk-worm cocoons	4.4	12.9*	2.7	76
Forestry	5.6	5.0*	0.7	529
Fishing	5.6	2.5*	0.6	527
Coal	5.6	6.1*	2.4	71
Oil	5.6	10.4*	3.8	56
Gas	5.6	34.4*	14.3	8
Minerals nec	5.6	1.8*	0.3	1584
Bovine meat products	4.4	7.7*	1.9	211
Meat products nec	4.4	8.8*	0.9	411
Vegetable oils and fats	4.4	6.6*	0.7	717
Dairy products	4.4	7.3*	0.8	547
Processed rice	4.4	5.2*	2.6	62
Sugar	4.4	5.4*	2.0	156
Food products nec	4.4	4.0*	0.1	6917
Beverages and tobacco products	6.2	2.3*	0.3	998
Textiles	4.4	7.5*	0.1	14,375
Wearing apparel	8.8	7.4*	0.2	9090
Leather products	8.8	8.1*	0.3	3457
Wood products	5.6	6.8*	0.2	4120
Paper products, publishing	3.6	5.9*	0.2	6597
Petroleum, coal products	3.8	4.2*	1.1	344
Chemical, rubber, plastic products	3.8	6.6*	0.1	61,603
Mineral products nec	5.6	5.8*	0.2	6240
Ferrous metals	5.6	5.9*	0.3	5524
Metals nec	5.6	8.4*	0.4	3194
Metal products	5.6	7.5*	0.2	9926
Motor vehicles and parts	10.4	5.6*	0.3	2238
Transport equipment nec	10.4	8.6*	0.4	1843
Electronic equipment	5.6	8.8*	0.2	8916
Machinery and equipment nec	5.6	8.1*	0.1	44,386
Manufactures nec	5.6	7.5*	0.2	7586

*Estimate significant at 95% confidence level.

continued, the idea for a Free Trade Area of the Americas arose as a logical next step in the economic integration of the hemisphere and a major impetus of the FTAA effort, launched in April 1998 in Santiago, Chile, has been to simplify the complex network of existing bilateral agreements in place in the region (FTAA Tri-Partite Commission, 2002, 2003; Diao and Somwaru, 2001; Harrison et al., 2002).

Given the focus on merchandise trade in the Americas, and in light of the large dimensions of the GTAP data base used as the empirical basis for our simulations, we begin by aggregating the sixty-six regions of the version 5 GTAP database to the seventeen regions shown in Table 2. Full v. 5 GTAP country detail is preserved in the Western Hemisphere, while composite regions are formed for the rest of the world, including: Asia–Oceania (ASOC), the fifteen-member European Union (EU15), rest of Europe (OEUR), and Middle East and Africa (MEAF). Services activities are aggregated into four broad categories to further reduce model size. Because of the complex set of existing preferential agreements on merchandise in trade in the Americas, a fair assessment of the changes that occur due to further liberalization under the FTAA umbrella requires accounting for these in our benchmark data. We do this by updating the database to account for the pre-existing applied tariff structure in the liberalizing regions as of 2001 using the MAcMap tariffs data base (Bouët et al., 2004).

5.2. Simulation and SSA

The experiment we consider for simulation is a stylized representation of the FTAA whereby we eliminate all tariffs on intra-regional merchandise trade. Our choice of experiment is driven by the ease of interpretation of results, as complete elimination of tariffs represents an upper bound on potential liberalization gains from FTAA implementation (Young and Huff, 1997). In addition, viewing the experiment in terms of total potential gains also alleviates from the analysis continuing uncertainty about the negotiated extent of tariff elimination.

Table 2
Percentage change in aggregate imports: Mean, 95% Confidence Interval, and Coefficient of Variation

Region	Confidence interval			Coefficient of variation	
	Lower	Mean	Upper	Total imports	Sector average
Canada	1.07	1.32	1.57	0.10	1.17
USA	2.07	2.19	2.31	0.03	0.10
Mexico	7.36	8.38	9.39	0.06	0.06
Central America	10.76	11.16	11.56	0.02	0.01
Colombia	9.22	10.37	11.51	0.06	0.22
Peru	7.93	8.41	8.89	0.03	0.16
Venezuela	4.69	4.81	4.93	0.01	0.18
Other Andean Pact	4.65	5.39	6.12	0.07	0.24
Argentina	4.13	4.26	4.39	0.02	0.12
Brazil	8.07	8.33	8.58	0.02	0.18
Chile	5.39	5.51	5.64	0.01	0.17
Uruguay	2.04	2.24	2.45	0.05	0.80
Other South America	5.68	5.98	6.27	0.03	0.01
Asia–Oceania	−0.32	−0.31	−0.31	−0.01	−0.11
European Union	−0.23	−0.23	−0.23	−0.01	2.01
Other Europe	−0.15	−0.15	−0.15	−0.01	−0.10
Mid-East and Africa	−0.29	−0.27	−0.25	−0.04	−0.10

As noted in the introduction, a primary aim of this paper is to present results from our experiment that account for the uncertainty in model parameters via systematic sensitivity analysis. The most common approach to generating such results is to employ a Monte Carlo procedure to repeatedly solve the model using a random vector of elasticities. As an alternative to Monte Carlo, we employ the Gaussian Quadrature (GQ) numerical integration technique developed by DeVuyst and Preckel (1997). These authors show that an approximating discrete distribution can be obtained based on known lower order moments of the model parameters, and that selectively solving the model based on the moments of this approximate distribution generates sensitivity results consistent with the Monte Carlo approach, with far fewer simulations required.

We must make some assumptions about the underlying parameter distributions to make robustness claims using the GQ procedure. We assume that all parameters altered in the model are independently and normally distributed. Beyond this, we also assume that the elasticity of substitution in the domestic-import substitution nest (recall φ_i in Eq. (1)) is tied to σ_i via the "rule of two" so that these elasticities vary together and by the same proportion in repeated solutions of the model.[16]

We choose to focus our results in the next section on variability in model results only with respect to the trade elasticities. This choice is logical when considering that all of our shocks involve the elimination of border import measures. However, in order to further support this claim, we conducted a supplementary analysis in which two separate sensitivity experiments are undertaken, thereupon comparing the results in terms of the predicted variability of the regional welfare (EV). Results are reported in Appendix Table A-4, available from the authors upon request. We find that the welfare impacts of the FTAA are far more sensitive to the trade elasticities than to variation in all of the remaining model parameters. Specifically, the coefficient of variation from systematic variation in the trade elasticities is, on average, more than twelve times as large as that resulting from comparable variation of the non-trade parameters in the model.[17]

6. Results

Since we use distributions, rather than point estimates, for the trade elasticities, our results also come in the form of distributions. Therefore, the most natural thing to look at is the mean value for each variable of interest, along with the associated standard deviation, or the coefficient of variation (standard deviation/mean). This information, accompanied by an assumption regarding the shape of the underlying distribution of endogenous variables (we assume normality, as with the parameters) allows formation of confidence intervals for welfare changes as well as other model results. These will be the focal point for our discussion.

[16] The "rule of two" links φ_i with the estimated value for σ_i as follows: $\sigma_i = 2\varphi_i$. This rule was first proposed by Jomini et al. (1994) and was retained in the GTAP parameter file. Recently this rule was tested by Liu, Arndt and Hertel (2004) in a back-casting exercise with a simplified version of the GTAP model. While those authors reject the validity of the GTAP trade elasticities, they fail to reject the rule of two, thereby lending additional support to this approach.

[17] Obviously we face a problem in specifying the degree of uncertainty in the non-trade parameters. In order to make the two experiments comparable, we take the trade weighted average of the coefficient of variation in the trade elasticities and use this to determine the proportional variation in other model parameters.

The first question that comes to mind when undertaking the FTAA analysis with these newly estimated trade elasticities is whether or not they yield markedly different results when compared to those obtained using the literature-based parameter set of the v.5 GTAP parameter file. Before proceeding with our formal analysis of the FTAA results, we seek to briefly address this question. As noted previously, the average of the newly estimated trade elasticities does not differ greatly when compared to the simple average of the elasticities adopted from the literature. However, within broad sectors of the economy there is considerable variation across the two sets of parameters. And recall that net trade diversion depends critically on the size of these elasticities, as they interact with differential tariffs and trade flows (recall Eq. (11)). So sectoral variation could make a difference in the welfare outcome.

Therefore, we begin with a simple comparison of the two sets of regional welfare estimates. These estimates are reported, along with other results which we will soon discuss, in Table 5. The first column of entries are the result of a deterministic simulation using the version 5 GTAP elasticities of substitution and the mean welfare result from the SSA analysis is reported in the column with header EV, mean. Our significance criterion is that the GTAP V5 welfare result does not lie within the confidence interval of the welfare result from the simulation with the trade elasticities estimated in this paper. A comparison of these two outcomes reveals that the welfare results are significantly different from the GTAP V5 point estimate in all cases except for the regions Venezuela, Andean Pact, Uruguay, Other South America, Other Europe, and Middle East/Africa.

Now let us turn to an in-depth analysis of the FTAA using the full SSA analysis based on the econometrically estimated trade elasticities. Our approach to analyzing the results in this paper will be to investigate the elements of Eq. (7) individually, thereafter examining their combined impact on welfare. We begin with the tariff-related efficiency effect. Since this is driven by changes in import volume, let us first consider what happens to imports. Table 2 reports the mean percentage change in regional import volume as a result of the FTAA experiment. Aggregate import volume increases in all FTAA regions, while falling in the non-FTAA regions. Furthermore, 95% confidence intervals, constructed based on the assumption of normality, show that we can be confident in all of these increases. The largest increases are for Colombia and Other Central America. Our 95% confidence intervals for these two regions do not overlap with that of Peru, which shows the third largest increase in total imports. These large increases in imports may be directly attributed to the relatively larger tariff rates for these countries. Most of the aggregate import volume changes are between +4% and +9%, with some exceptions. The US and Canada, which already enjoy free trade with one another, show a smaller increase in imports. Also, there is a very low import volume increase for Uruguay. This can be attributed to the relative loss of preferential access that occurs under FTAA when Uruguay's partners in MERCOSUR liberalize with other regions in the Americas.

One interesting question that arises in the context of our analysis is whether there is greater certainty about more aggregate variables than about disaggregate variables produced by this model. The last two columns in Table 2 address this issue in the case of import volumes. First we report the coefficient of variation (the ratio of the standard deviation to the mean) for the change in national imports that is reported in column 2 of Table 2. When this ratio is small, we can infer a relatively higher level of confidence in the result. Note that it is quite small for Chile and Venezuela (0.01), whereas it reaches its maximum value in the case of the rest of the Andean Pact (0.07). The final column in Table 2 reports the *average* coefficient of variation for the percentage change in national imports, *at the sector level*. This gives an indication of the average degree of precision for the more disaggregate results. As can be seen by comparing these two entries for

Table 3
Components of allocative efficiency effects due to FTAA ($US million): Mean and Coefficient of Variation (CV)

Region	Imports		Consumption		Production		Other
	Mean	CV	Mean	CV	Mean	CV	Mean
Canada	760.60	0.12	246.40	0.11	40.78	0.21	−212.06
USA	545.25	0.06	−0.01	−0.23	−74.44	−0.09	−495.42
Mexico	975.35	0.19	2.50	0.29	174.80	0.23	103.37
Central America	322.05	0.05	217.79	0.05	−10.58	−0.15	−8.56
Colombia	508.55	0.10	15.38	0.08	4.22	0.23	98.54
Peru	9.35	0.15	23.68	0.08	−2.69	−0.38	123.03
Venezuela	−10.50	−0.09	2.83	0.08	0.00	−0.13	1.44
Other Andean Pact	17.26	0.17	11.48	0.13	−5.12	−0.39	22.83
Argentina	32.61	0.08	*−0.19*	*−0.80*	−0.23	−0.02	−29.45
Brazil	148.88	0.04	431.82	0.02	−10.62	−0.17	788.52
Chile	−15.69	−0.06	33.40	0.03	−5.46	−0.04	9.40
Uruguay	1.57	0.36	3.88	0.10	*0.22*	*0.73*	−4.16
Other South America	*−0.60*	*−1.04*	7.64	0.19	−1.56	−0.03	−2.72
Asia–Oceania	−252.95	−0.07	−34.53	−0.03	46.53	0.09	−362.48
European Union	−226.71	−0.06	−120.41	−0.05	60.52	0.03	−407.48
Other Europe	−49.62	−0.04	−15.67	−0.04	6.17	0.08	−42.99
Mid-East and Africa	−66.17	−0.07	−10.27	−0.10	1.91	0.17	−14.03

Note: Results in italics indicate that the result can not be signed at the 95% confidence level, i.e. the confidence interval includes zero.

each country, the more disaggregate results are less precise. This is intuitive in that we often expect a certain degree of offset at the aggregate level (when one sector's imports are low, another's may be high).

Next, turn to the allocative efficiency effects associated with the import volume changes. These are reported in the first column of Table 3, which gives both the mean and the coefficient of variation (CV) associated with the tariff-related allocative efficiency component of Eq. (7). Recall from Eq. (11) that that this can be re-written as the (welfare-scaled) summation of the tariff revenue-weighted import quantity changes. From the mean values of this variable, we see that there *is* net welfare creation from trade for 10 of the 13 FTAA partners. In the cases of Venezuela, Chile and the Other South America, this import efficiency term is negative, despite the fact that aggregate import volume rises (recall Table 2). This is due to the fact that the *welfare* contribution of the trade volume change depends on the interaction between tariff rates and trade flows. As tariffs are eliminated on intra-FTAA flows, the associated welfare weight is also eliminated. If the tariffs on extra-FTAA imports are large, and if the associated FTAA-driven decline in the trade volume is also large, these negative numbers can dominate the overall welfare effect, leading to net trade diversion.

Table 4 explores this trade diversion phenomenon in detail for the case of machinery and equipment imports into Chile. Chile is notable for its uniform tariff structure (8% across all sources/products in 2001). This is efficient in that it promotes the sourcing of imports from the least cost supplies of any given product, as well as discouraging substitution across import categories in response to differential tariffs by product. Of course, there remains the distortion of import/domestic choices, as with any tariff regime. We focus the discussion here on other machinery and equipment imports as this contributes the largest share of the aggregate efficiency loss in column one of Table 3.

Table 4
Analysis of trade diversion in Chile: Machinery and equipment

	Welfare change ($US mill.)($EV_s$)	Volume change ($US mill.) ($dQMS_{irs}$)	Initial tariff rate (τ^0_{Mirs})	Updated tariff rate (τ^1_{Mirs})	Price initial ($PCIF^0_{irs}$)	Price updated ($PCIF^1_{irs}$)
Canada	1.00	25.54	0.08	0.00	0.93	0.93
USA	18.20	450.01	0.08	0.00	0.93	0.93
Mexico	2.70	71.77	0.08	0.00	0.93	0.93
Central America	0.01	0.28	0.08	0.00	0.93	0.92
Colombia	0.30	8.12	0.08	0.00	0.99	0.96
Peru	0.04	1.19	0.08	0.00	0.81	0.97
Venezuela	0.07	1.82	0.08	0.00	0.93	0.81
Other Andean Pact	0.01	0.42	0.08	0.00	0.81	0.91
Argentina	1.28	33.31	0.08	0.00	0.93	0.81
Brazil	4.54	106.97	0.08	0.00	0.96	0.92
Chile	0.00	0.00	0.00	0.00	1.00	0.96
Uruguay	0.02	0.58	0.08	0.00	0.93	1.00
Other South America	0.00	0.11	0.08	0.00	0.94	0.93
Asia–Oceania	−16.14	−189.46	0.08	0.08	0.94	0.93
European Union	−23.41	−284.15	0.08	0.08	0.93	0.92
Other Europe	−1.19	−14.50	0.08	0.08	0.93	0.92
Mid-East and Africa	−0.80	−9.72	0.08	0.08	0.93	0.93
Total	−13.356	202.303	n.a.	n.a.	n.a.	n.a.

The individual columns in Table 4 correspond to components of Eq. (8). The first entry reports the non-linear solution value for $\mathrm{EV_s}$ (τ_{Mirs}) for i=*ome*, s=Chile and r=all regions.[18] The driving force behind each of these entries is the underlying change in bilateral import volume, $dQMS_{irs}$ reported in column two of the table. This is measured in $US 1997, where one unit of the good is the amount that could be purchased in the source country for $1 million in the initial (i.e., pre-FTAA) equilibrium. Since these goods are differentiated products, the sum of these trade volumes is not particularly meaningful. But if we did perform this summation, we would find that this crude estimate of import volume showed an *increase* of $202 million, with the rises in intra-FTAA imports of machinery and equipment into Chile more than offsetting the declines in extra-FTAA imports. This increased import volume stands in sharp contrast to the negative total welfare change.

The difference between the simple volume summation and the welfare change derives from the bilateral weights applied to these volume changes: $\tau_{\mathrm{Mirs}}PCIF_{irs}$. In this regard, it is instructive to consider both the initial (0) and ending (1) values for the tariff rate and import price. These are also reported in Table 4. Note that the proportional $PCIF_{irs}$ changes (typically less than 10%) are an order of magnitude smaller than the changes in intra-regional tariffs (which are equal to −100%), so we focus our attention on the latter. While the initial reductions in the tariff on intra-FTAA imports (e.g., from 0.08 to 0.075) bring fairly large welfare gains, (recall Fig. 1), the final reductions (e.g., from 0.005 to 0.000) bring almost nothing. Yet, the final reduction in τ_{Mirs} continues to lead to substantial displacement of extra-FTAA imports (recall Eq. (9)). Given the absence of any cuts to extra-FTAA tariffs, these volume reductions come to dominate the welfare

[18] Note that the non-linear solution to Eq. (7) requires that we incorporate it individually into the model's solution. This permits us to capture the interaction between changes in the levels of τM_{irs} and $PCIF_{irs}$ on the one hand, and the volume changes, $dQMS_{irs}$, on the other.

story. This is why the welfare loss due to reduced imports of machinery and equipment from the EU is nearly twice as large as the gain due to increased imports from USA, even though the absolute value of the trade volume change with respect to USA is nearly twice that of the EU.

Recall from our earlier analysis (e.g., Eq. (11)) that the elasticity of substitution among imports, by source, is a critical determinant of the allocative efficiency effect associated with tariff changes. Yet these elasticities are uncertain, and we have characterized this uncertainty in our systematic sensitivity analysis. So it is of some interest to explore the relationship between uncertainty in the trade elasticities and uncertainty in these welfare contribution terms themselves. We examine this issue statistically for the welfare changes associated with FTAA flows in the model. Consider the welfare term in Eq. (8), $\mathrm{EV_s}$ (τ_{Mirs}). In a typical CGE analysis, there is but one value of this term for each commodity i, exporter r, and importer s. However, in our approach, we have a value for this term for each solution of the model, every time with a different set of trade elasticities. To better understand the standard error in $\mathrm{EV_s}$ (τ_{Mirs}) across model solutions, we regress the standard error in this variable on the depth of the associated bilateral tariff cut,τ_{rs}, which is the same across model solutions, and an interaction between the depth of the tariff cut and $SE_{\sigma i}$, the standard error of the substitution elasticity among imports by source, for commodity i. We include the depth of the tariff cut as an explanatory variable to control for differences in relative dispersion in the welfare contribution variable, because larger tariff cuts will increase variability of the welfare variable for a constant standard error for the elasticity. Estimates are reported in Eq. (17),[19] along with the associated *T*-statistics (in parentheses):

$$SE_{WQIMirs} = \underset{(6.464)}{0.020}\,\tau_{irs} + \underset{(6.936)}{0.023}\,\tau_{irs}SE_{\sigma i} \tag{17}$$

The OLS coefficients in Eq. (17) are significant and positive indicating that both the depth of the tariff reduction and the interaction of the tariff with the variability of the elasticity are important in explaining variability in the allocative welfare effects. The positive relationship between the two standard errors is as we would expect, since we hypothesize that uncertainty in the model parameter should be carried over to the allocative welfare component as demonstrated in Section 2. The coefficient levels indicate that we predict a change of 0.043 = (0.020 + 0.023) in the standard error of the welfare contribution variable when the tariff cut is increased by one percent and the standard error of the substitution elasticity is simultaneously increased by one. The mean standard error for the allocative welfare effect is 0.163, so this predicted change represents about twenty-five percent of the mean for the dependent variable.

We now return to Table 3 to discuss the remaining elements of the efficiency story. For Chile, the next most important efficiency change relates to the consumption taxes,[20] which apply equally to consumption of imports and domestic goods. Thus, the increased consumption of imported goods boosts overall consumption and results in positive contributions to aggregate welfare in all of the FTAA regions where adequate consumption taxation data are available.[21] Production taxes

[19] The regression is based on 5337 non-zero bilateral trade flows in the FTAA region. The R2 for the regression is 0.040, indicating fairly low explanatory power as we would expect given the large number of omitted variables (in the regression model) that affect the welfare term in the GE model. Since our primary concern is to characterize the relationship between uncertainty in the welfare term and the trade elasticity we feel the regression model is serviceable as it controls for the dominant effect of size of the tariff cut, and shows the significant positive relationship between the two measures of uncertainty.

[20] This includes the Chilean value-added taxes which are most naturally modeled as a consumption tax.

[21] The negative number for Argentina results from the apparent exemption of imported oil products from consumption taxation. This appears to be an error in the original data base for that country.

(subsidies) also play a big role in the welfare decomposition in some regions. In USA, the expansion of subsidized grains output leads to a negative welfare contribution, whereas in Mexico, the expansion of taxed manufacturing activity at the expense of untaxed fuel and agriculture improves efficiency.

The final column in Table 3 reports the combined efficiency impacts of intermediate import taxes, primary factor taxes (subsidies), and export taxes (subsidies). (These were suppressed in Eq. (7) for the sake of brevity.) The negative efficiency contribution in the US derives from land and capital subsidies for program crops and dairy export subsidies, whereas agricultural export taxes in Brazil play a key role in the positive welfare contribution in that country.

Next, turn to Table 5, which reports the aggregate welfare effects, by country, decomposed into their efficiency and terms of trade components. The aggregate efficiency effect is simply the summation of the results reported in Table 3. This decomposition permits us to explain the efficiency loss in the USA and Venezuela. In Venezuela, this is caused by net trade diversion, whereas in the USA it is due to expansion of the subsidized agriculture sector.

We turn next to the terms of trade effects (the second pair of columns in Table 5). We expect that these effects might be largest in those countries where exports surge the most. Thus it is no surprise that the terms of trade deteriorate for Colombia and Peru, as these are countries with very high average tariffs that must export more to offset the large increase in import volume. The terms of trade also deteriorate for Canada, Mexico, Argentina and Other South America. The TOT deterioration for Mexico and Canada is explained by the fact that these countries currently enjoy tariff free access to the largest market in the region for many of their products. When the FTAA is introduced, other countries obtain the same benefit and they displace

Table 5
Welfare (Equivalent Variation) effects of FTAA outcome ($US million)

Region	GTAP V5 EV Result	Welfare impacts with estimated elasticities							
		Efficiency		Terms of trade		EV total		Utility (% Change)	
		Mean	CV	Mean	CV	Mean	CV	Mean	CV
Canada	1215	835.71	0.13	−620.78	−0.12	167.18[#]	0.26	0.03	0.26
USA	6035	−24.61	−2.87	6244.82	0.04	6845.20[#]	0.04	0.09	0.04
Mexico	1568	1256.02	0.19	−1241.76	−0.12	*78.02*[#]	*1.29*	*0.02*	*1.29*
Central America	3773	520.70	0.05	855.32	0.04	1385.57[#]	0.04	1.70	0.04
Colombia	1029	626.70	0.09	−514.99	−0.07	*41.95*[#]	*0.60*	*0.05*	*0.60*
Peru	304	153.37	0.03	−73.39	−0.06	62.72[#]	0.11	0.11	0.11
Venezuela	133	−6.23	−0.18	82.92	0.16	110.38	0.11	0.15	0.11
Other Andean Pact	45	46.45	0.08	*22.10*	*0.57*	63.94	0.27	0.27	0.27
Argentina	−191	*2.73*	*1.56*	−30.69	−0.51	−74.50[#]	−0.28	−0.02	−0.28
Brazil	3790	1358.59	0.03	548.33	0.10	1921.85[#]	0.06	0.28	0.06
Chile	230	21.64	0.04	189.57	0.03	200.41[#]	0.03	0.29	0.03
Uruguay	11	*1.52*	*1.80*	22.32	0.22	16.71	0.44	0.09	0.44
Other South America	−78	2.75	0.30	−42.78	−0.06	−82.85	−0.04	−0.89	−0.04
Asia–Oceania	−3927	−603.43	−0.04	−2625.12	−0.01	−3577.75[#]	−0.01	−0.05	−0.01
European Union	−3796	−694.09	−0.05	−2383.01	−0.01	−3171.74[#]	−0.02	−0.05	−0.02
Other Europe	−385	−102.12	−0.04	−246.90	−0.05	−390.04	−0.03	−0.03	−0.03
Mid-East and Africa	−339	−88.55	−0.07	−232.27	−0.18	−340.07	−0.14	−0.04	−0.14

Notes: Results in italics indicate that the result can not be signed at the 95% confidence level, i.e. the confidence interval includes zero.
[#]Result is significantly different from the GTAP V5 Equivalent Variation result based on the confidence interval.

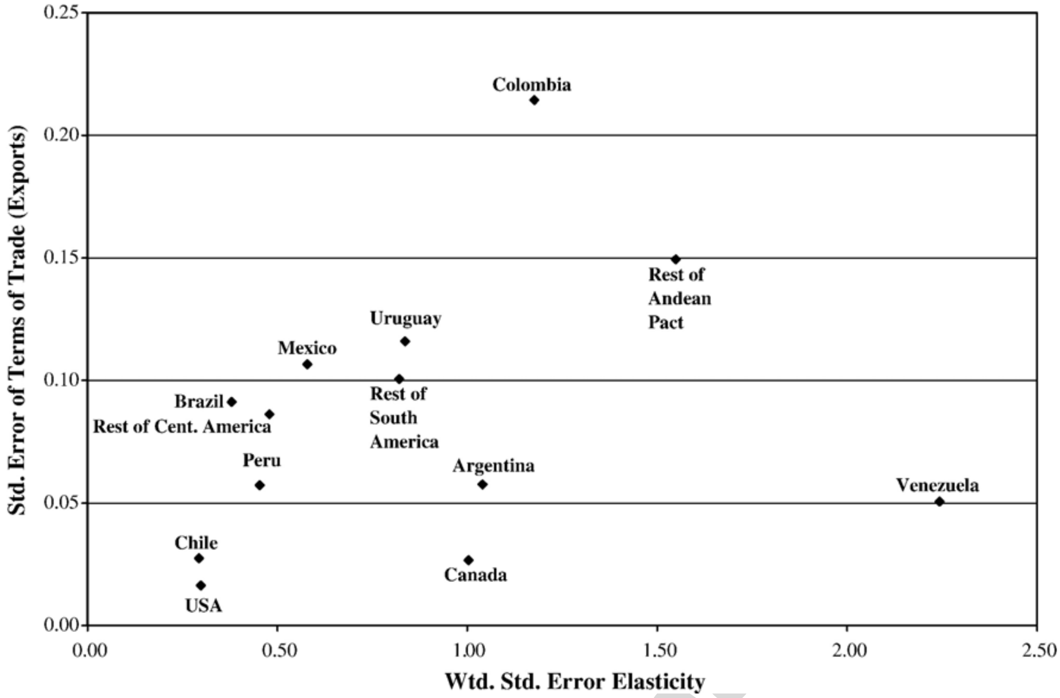

Fig. 2. Uncertainty comparison: trade elasticity and TOT export price effect.

Mexican and Canadian imports. The same phenomenon explains why Argentina's terms of trade decline following her displacement within MERCOSUR markets. Of course, one region's TOT loss is another's gain, and Other Central America is one of the regions showing a strong TOT gain. In this case, it is the strong increase in exports from this region that is driving the import growth reported in Table 2.

Another dimension of this analysis of uncertainty in the terms of trade effect can be observed in Fig. 2. In this figure, we show how uncertainty in each FTAA region's average export demand elasticity translates into uncertainty in the export price component of their terms of trade (see McDougall, 1993 for details on the terms of trade decomposition). Countries that rely heavily on exports of commodities whose substitution elasticities in trade are highly uncertain (e.g., Colombia) are exposed to a great deal of uncertainty in the size of their average export demand elasticity (horizontal axis of Fig. 2) and tend to experience more uncertainty in the export price component of their terms of trade (vertical axis). Venezuela is an exception to this rule. It exhibits a high degree of uncertainty in export demand elasticities due to a heavy reliance on oil and gas exports, which are large and rather uncertain (recall Table 1). On the other hand, the variation in the export price component of its terms of trade is relatively small, due to the relative homogeneity of this product (high elasticity of substitution in Table 1) and the generally low tariffs on oil.

We are now in a position to answer the question: Which countries gain from the FTAA? We see from the final two pairs of columns in Table 5 that nine of the thirteen FTAA countries gain from this free trade agreement based on a 95% confidence interval. On the other hand, Argentina and Other South America are shown to lose from the FTAA. In the case of Argentina, this is driven by a dominant, adverse terms of trade effect. For Other South America, both the terms of trade and allocative effects are negative.

The welfare impact of the FTAA on Colombia and Mexico is uncertain. These are both interesting cases, since the component parts of their aggregate welfare impacts are "certain" in the

Table 6
Regional employment effects due to FTAA (unskilled labor only)

Region	Number of sectors with:		Pct results with 95% confidence (%)
	Employment increase	Employment decrease	
Canada	29	13	83
USA	31	11	88
Mexico	28	14	79
Central America	9	33	93
Colombia	20	22	88
Peru	18	24	88
Venezuela	22	20	88
Other Andean Pact	18	24	76
Argentina	22	20	90
Brazil	23	19	86
Chile	21	21	88
Uruguay	13	29	79
Other South America	19	23	83
Asia–Ocenia	19	23	81
European Union	17	25	88
Other Europe	16	26	86
Mid-East and Africa	17	25	81

sense that the 95% confidence intervals do not include zero. However, the positive allocative efficiency component is offset by an equally large negative terms of trade effect. So while the components of the welfare change are certain, the sign of their summation – the aggregate welfare gain – is *uncertain*. This type of uncertainty is not inherited from uncertainty about the model parameters governing substitution in trade. Rather, it relates to the presence of competing economic forces at work in the determination of the change in national welfare.

There are many other variables in addition to the change in aggregate welfare that we could examine, particularly at the sector level. Here we focus our attention on employment of low skill workers, since the displacement of unskilled workers is often one of the most sensitive topics surrounding any free trade agreement.[22] Table 6 summarizes the directional changes in unskilled employment, by sector, for each region. The first column reports the total number of sectors in which employment of unskilled labor rises and the second column reports the number in which employment falls. Since total employment remains unchanged, by assumption, the relative size of these two numbers is not very meaningful. However, it is interesting to ask how many of these changes are significantly different from zero at the 95% level. This is reported in the next column of Table 6. Here we see that the changes in employment are generally robust to the estimated variation in trade elasticities. We are confident in the sign of the change in sectoral employment for every region in over 75% of the sectors.

Table 7 reports the same employment results as Table 6, with the focus shifted to unskilled employment by sectors across all of the model regions. This allows us to evaluate how the uncertainty in sectoral trade elasticities translates into uncertainty about the employment effects

[22] The impact of the FTAA on unskilled employment could be further exacerbated by biased technological change that might be induced by increased imports. This factor is not considered in our analysis.

Table 7
Sectoral employment effects due to FTAA (unskilled labor only)

Sector	Number of regions with:		Pct results with 95% confidence (%)
	Employment increase	Employment decrease	
Paddy rice	9	8	65
Wheat	6	11	59
Cereal grains	9	8	65
Vegetables, fruit, nuts	6	11	71
Oil seeds	10	7	76
Sugar	9	8	82
Plant-based fibers	4	13	71
Other Crops	7	10	76
Cattle	12	5	76
Animal products	9	8	94
Raw milk	8	9	94
Wool	7	10	88
Forestry	8	9	82
Fishing	9	8	94
Coal	4	13	59
Oil	8	9	82
Gas	8	9	82
Minerals	12	5	76
Bovine meat products	13	4	82
Other meat products	7	10	88
Veg. oils and fats	11	6	94
Dairy products	6	11	100
Processed rice	8	9	35
Processed sugar	10	7	76
Food products	8	9	100
Bev. and tobacco	10	7	88
Textiles	7	10	94
Wearing apparel	9	8	88
Leather products	11	6	100
Wood products	7	10	82
Paper products	8	9	100
Petroleum, coal products	8	9	71
Chemical, rubber, plastic	6	11	100
Mineral products	9	8	88
Ferrous metals	5	12	100
Metals	9	8	100
Metal products	7	10	94
Motor vehicles and parts	5	12	100
Transport equipment	12	5	100
Electronic equipment	7	10	100
Machinery	6	11	94

for a given sector. Here we see that for all but five of the forty-two sectors, we are confident in the direction of change in employment for a given sector in seventy percent or more of the regions. The exceptions here are four primary commodities: paddy rice, wheat, other grains, and coal, as well as for processed rice. Not surprisingly, these sectors have some of the largest standard deviations relative to the size of the estimated trade elasticity (recall Table 1).

7. Summary and conclusions

Computable General Equilibrium analysis is often criticized for its lack of econometric foundations (McKitrick and Ross, 1998). The goal of this paper is to show that it is indeed possible to provide substantial statistical underpinning to policy analyses conducted using the CGE framework. We focus our attention on analysis of Free Trade Agreements – specifically, the Free Trade Agreement of the Americas – for which the key behavioral parameter is the elasticity of substitution among imports from different countries. This governs the extent to which non-FTAA regions will be displaced by the preferential reduction in tariffs on imports from FTAA countries and we show that the welfare results are far more sensitive to this parameter than all other parameters combined.

Historically, estimation of the import substitution elasticity has been difficult, due to insufficient observed variation in relative prices. In this paper, we capitalize on a unique data set and approach developed by Hummels (1999), in which variation in bilateral transport costs is combined with bilateral tariff variation in order to enhance the observed variability of relative prices for imports from different sources in six FTAA countries and one non-FTAA country. Elasticities are estimated at the GTAP commodity level to facilitate subsequent incorporation into our CGE model. The resulting estimates of the elasticity of substitution among imports are all significant at the 95% level. These estimates, together with their standard errors, are used in the subsequent policy simulations.

The FTAA analysis takes explicit account of the fact that we do not know the true trade elasticities with certainty. Rather, we sample from a distribution of parameter values, constructed based on our econometric results. The outcome of this systematic sensitivity analysis is a distribution of model results, from which we can construct confidence intervals with which to answer the basic question posed in the title of this paper. We find that imports increase in all regions of the world as a result of the FTAA, and this outcome is robust to variation in the trade elasticities. Nine of the thirteen FTAA regions experience a welfare gain in which we are more than 95% confident. Two regions, Argentina and rest of South America experience welfare losses as they are displaced from existing markets in which they currently enjoy preferential access. Finally, the welfare impacts of the FTAA in Mexico and Colombia are uncertain due to offsetting efficiency and terms of trade effects. We also examine the robustness of our employment effects. With the exception of several primary products, where the trade elasticity is relatively uncertain, we can be confident in the sign of the sectoral employment effects in the majority of regions.

Of course all of these findings are conditional on the underlying model structure. Variations in that structure will change both the econometric procedures as well as the CGE model itself. Given the uncertainty surrounding the appropriate structure for international trade modeling, and the diversity of outcomes that such changes in structure can engender, we must view the confidence intervals in this paper as being on the narrow side. Future work should focus on discriminating among these alternative model structures for purposes of establishing a firmer foundation for CGE analysis of trade policies (e.g., Hummels and Klenow, 2005).

In summary, we conclude that there is great potential for combining econometric work with CGE-based policy analysis in order to produce a richer set of results that are likely to prove more satisfying to the sophisticated policy maker. In the end, decision makers and their advisors increasingly ask: How robust are the policy findings? In this paper we have found that some of the FTAA conclusions are robust, while others are not. This is important information for those seeking to make key political decisions based in part on results from quantitative economic models.

Appendix A. Supplementary data

Supplementary data associated with this article can be found, in the online version, at doi:10.1016/j.econmod.2006.12.002.

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