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Specialization and Diverging Manufacturing Structures: The Aftermath of Trade Policy Reforms in Developing Countries

*Antoni Estevadeordal**
*Christian Volpe Martincus**

*Inter-American Development Bank

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Antoni Estevadeordal
IDB

Christian Volpe Martincus[♦]
IDB

Abstract

Trade barriers have been declining around the world over the last five decades. Countries reduced their tariffs unilaterally as well as concertedly in the framework of regional integration agreements. As a consequence, trade flows among economies have substantially intensified. According to economic theory, this should have had a significant impact on the countries' specialization patterns. However, to our knowledge, there is no direct robust econometric evidence on the effect of trade policy on the overall degree of countries' specialization. This paper aims at filling this gap in the literature. We focus on ten Latin American countries members of the LAIA (Latin American Integration Association) over the period 1985-1998. These countries are natural case studies because in the last two decades they implemented broad and comprehensive trade liberalization programs, both generally and preferentially, starting from relatively high tariff protection levels. Our econometric results suggest that reducing own MFN tariffs is associated with increasing manufacturing production specialization. Furthermore, we find that preferential trade liberalization and differences in the degree of unilateral openness have resulted in increased dissimilarities in manufacturing production structures across countries. These results are robust to the specialization measure being used, the correction for groupwise heteroscedasticity, cross-sectional correlation, serial correlation and endogeneity biases, and the inclusion of indicators to account for the real exchange misalignment prevailing in the region during the period under examination.

Keywords: Specialization, Trade Policy, Latin America
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[♦] Correspondence Address: Inter-American Development Bank (Stop W0612, Room NW622),, 1300 New York Avenue, NW, Washington, DC 20577, United States of America. E-mail: christianv@iadb.org. Tel: +1 202 623 3199. Fax: +1 202 623 2169.

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The Aftermath of Trade Policy Reforms in Developing Countries

1 Introduction

Trade barriers have been declining around the world over the last five decades. Countries reduced their tariffs unilaterally as well as concertedly in the framework of regional integration agreements. As a result, trade flows among economies have substantially grown. They have increased by a factor of 89 between 1953 and 2003. According to the economic theory, either due to comparative advantage or agglomeration economies, this should have had a significant impact on the countries' specialization patterns. Has the existing empirical evidence confirmed this theoretical prediction?

Several studies present descriptive evidence on the evolution of specialization indicators over periods of declining trade barriers.[1] This evidence mostly concerns developed countries. However, to our knowledge, there is no direct robust econometric evidence on the effect of trade policy on the overall degree of developing countries' specialization.. This paper aims at filling this gap in the literature. Specialization is worth being studied because it affects the level of welfare, the speed of economic growth, and the degree of macroeconomic convergence across economies.[2]

We focus on ten Latin American countries members of the LAIA (Latin American Integration Association) over the period 1985-1998. These countries are natural case studies because in the last two decades they implemented broad and comprehensive trade liberalization programs starting from relatively high tariff protection levels. More specifically, these countries pursued unilateral plans and also engaged in regional integration initiatives. Thus, this set of nations provides a constellation of trade reforms, which is rich enough to assess their repercussions. Second, some of these economies exhibit substantial changes in their specialization degrees over the aforementioned period. On average, production specialization seems to be increasing and manufacturing structures seems to be becoming increasingly different. We can therefore analyze to what extent general and preferential trade liberalization have contributed to shape the evolving specialization patterns in the region.

[1] See, e.g., Midelfart-Knarvik et al. (2000), Brülhart (2001), and Combes and Overman (2003).
[2] See, e.g., Lucas (1988), Quah and Rauch (1990), Eichengreen (1993), Krugman (1993), Frankel and Rose (1998), and Redding (1999).

We estimate measures of overall specialization from sectoral value added to describe the countries' specialization level, both absolute and relative, and we also compute average MFN (Most Favored Nation) and preferential tariffs, which allow us to explicitly characterize the countries' trade policies. Our econometric results suggest that reducing own MFN tariffs is associated with increasing production specialization. Furthermore, we find that preferential trade liberalization and differences in the degree of unilateral openness have resulted in increased dissimilarities in manufacturing production structures across countries. These results are robust to the specialization measure being used, the correction for groupwise heteroscedasticity, cross-sectional correlation, serial correlation and endogeneity biases, and the inclusion of indicators to account for the real exchange misalignment prevailing in the region during the period under examination.

The remaining of the paper is organized as follows. Section 2 describes the dataset. Section 3 derives the estimation equation and addresses relevant econometric issues. Section 4 outlines some basic stylized facts about the trade policy reforms introduced in Latin American countries since the mid-1980s. Section 5 reports some descriptive evidence on the patterns of manufacturing production specialization and their evolution over the sample period and reports our main findings. Section 6 concludes.

2 Data

Our sample includes ten members of the LAIA: Argentina, Bolivia, Brazil, Chile, Colombia, Ecuador, Mexico, Peru, Uruguay, and Venezuela.[3]

We use annual sectoral value added at the 3-digit level of the ISIC, Revision 2, to characterize overall manufacturing production specialization in these countries over the period 1985-1998. These data come from the database PADI prepared by the United Nation's Economic Commission for Latin America and the Caribbean (ECLAC) and International Industrial Statistics made available by the United Nations Industrial Development Organization (UNIDO). Table A1 in the Appendix identifies the specific data sources and time coverage, whereas Table A2 lists the sectors.

Average MFN tariffs have been calculated for each country in the sample over the period 1985-2001. In addition, bilateral preferential tariffs have been computed for each economy over the same lapse. We have therefore data on the average tariff barriers set and

[3] Unfortunately, we do not have sectoral value added for Paraguay.

faced by countries, both on a MFN basis and under bilateral and multilateral regional trade arrangements. In particular, we can distinguish between general trade liberalization and average bilateral preferential trade liberalization in the context of specific blocs such as the Andean Community and MERCOSUR. Table A3 reports country and period coverage of tariff data.

We use GDP per capita as a proxy for both relative endowments and level of development. Data on this variable are expressed at market prices in constant 1995 U.S. dollars and come from the on line socioeconomic database BADEINSO prepared by the UN's ECLAC. We also incorporate the real effective exchange rate for imports, which is an index (1995=100) calculated by the UN's ECLAC and available from its on line macroeconomic database. Finally, we include a measure of real exchange rate misalignment taken from Terra and Valladares (2003). They estimate misalignments as departures from the long run equilibrium exchange rate as obtained following the methodology proposed by Goldfajn and Valdes (1999), i.e., estimating a long run relationship between the real exchange rate and economic fundamentals using cointegration techniques. Table A4 in the Appendix presents additional detailed information on these variables.

3 Empirical Specification and Econometric Strategy

3.1 Empirical Specification

To define the estimation equation and thus the appropriate functional forms as well as the relevant variables to be included, we will follow Harrigan (1997) and Redding (2002). The idea is to derive theory-consistent summary measures of specialization from the standard international trade theory.

Assume a set of small countries, each of them endowed with a fixed amount of factor of productions. These factors are used to produce final goods under constant returns to scale and perfect competition conditions such that the value of output is maximized. This value is given by:

$$X_{ct} = r(p_{ct}, v_{ct}) \tag{1}$$

where $r()$ is the revenue function, p is the vector of final goods prices, v is the vector of production factors, $c=\{1,\ldots,C\}$ indexes countries and t time. As long as the revenue function is twice continuously differentiable, the vector of the economy's profit-maximizing net output is given by:[4]

$$x_c(p_{ct}, v_{ct}) = \partial r(p_{ct}, v_{ct}) / \partial p_{ct} \tag{2}$$

We will further assume Hicks-neutral technology differences across countries, industries, and time, so that the production function takes the following form:

$$x_{cjt} = \theta_{cjt} f_j(v_{cjt}) \tag{3}$$

where θ_{cjt} parametrizes technology in industry *j* of country *c* at time *t*. As shown in Dixit and Norman (1980), in this case, the revenue function is given by:

$$r(p_{ct}, v_{ct}) = r(\theta_{ct} p_{ct}, v_{ct}) \tag{4}$$

where θ_{ct} is an *nxn* diagonal matrix of the technology parameters θ_{cjt}. This formulation implies that industry-specific neutral technological changes have the same effect on revenue as industry-specific price changes.

Following Woodland (1982), Kohli (1991), and Harrigan (1997), in order to operationalize the model, we assume a translog revenue function:[5]

$$\begin{aligned} lnr(\theta p, v) = \alpha_{00} + \sum_{j=1}^{n} \alpha_{0j} ln\theta_j p_j + \frac{1}{2}\sum_{k=1}^{n}\sum_{j=1}^{n} \alpha_{jk} ln(\theta_j p_j) ln(\theta_k p_k) + \sum_{i=1}^{m} \beta_{ih} lnv_i + \\ + \frac{1}{2}\sum_{i=1}^{m}\sum_{h=1}^{m} \beta_{ih} lnv_i lnv_h + \sum_{j=1}^{n}\sum_{i=1}^{m} \gamma_{ji} ln(\theta_j p_j) ln(v_i) \end{aligned} \tag{5}$$

where *j,k* index goods and *i,h* index factors. Symmetry of cross-effects implies:

$$\alpha_{jk} = \alpha_{kj} \quad and \quad \beta_{ih} = \beta_{hi} \quad \forall\ h, i, j, k \tag{6}$$

Further, linear homogeneity in *v* and *p* requires:

$$\sum_{j=1}^{n} \alpha_{0j} = 1 \quad \sum_{i=1}^{m} \beta_{0i} = 1 \quad \sum_{j=1}^{n} \alpha_{kj} = 0 \quad \sum_{i=1}^{m} \beta_{ih} = 0 \quad \sum_{i=1}^{m} \gamma_{ji} = 0 \tag{7}$$

Differentiating the natural logarithm of the revenue function with respect to each p_j, we obtain the share of industry *j* in country *c*'s GDP at time *t*, s_{cjt}:

$$s_{cjt} = \frac{p_{cjt} x_{cjt}(\theta_{ct} p_{ct}, v_{ct})}{r(\theta_{ct} p_{ct}, v_{ct})} = \alpha_{0j} + \sum_{k=2}^{n} \alpha_{kj} ln \frac{p_{ckt}}{p_{c1t}} + \sum_{k=2}^{n} \alpha_{kj} ln \frac{\theta_{ckt}}{\theta_{c1t}} + \sum_{i=2}^{m} \gamma_{ji} ln \frac{v_{cit}}{v_{c1t}} \tag{8}$$

This equation relates a theory-consistent measure of sectoral specialization to their underlying economic determinants: relative prices, technology, and factor endowments. The

[4] A suficient condition is that there are at least as many factors as goods (see Redding, 2002).

[5] The translog model is frequently interpreted as a second order approximation to an unknown function form (see Greene, 1997).

translog specification implies that the coefficients on the variables are constant across countries and over time.

We are interested in the economies' overall degree of manufacturing specialization. Specialization can be defined as the narrowness of the range of activities developed in a country. Thus, a country is highly specialized if a few industries account for large shares of its overall industrial activity. To measure a country's overall degree of absolute specialization we use the Herfindahl index. Formally:

$$H_{ct} \equiv \sum_{j=1}^{n_1} \left(s_{cjt}^{n_1} \right)^2 \tag{9}$$

where n_1 is the number of manufacturing sectors and $s_{cjt}^{n_1}$ is the share of industry j in country c's total manufacturing value added at time t. This index ranges between $1/n_1$, when all sectors account for the same share of total economic activity, and 1, when the whole activity is accounted by only one sector. In particular, the more unequal the sectoral shares, the more specialized an economy.

Noting that $s_{cjt} = s_{ct}^{n_1} s_{cjt}^{n_1}$, where the first term measures the share of manufacturing in country c's total GDP, and substituting for the country-sectoral shares in Equation (9), we can write:

$$H_{ct} \equiv \sum_{j=1}^{n_1} \left(s_{cjt}^{n_1} \right)^2 = \frac{1}{\left(s_{ct}^{n_1} \right)^2} \sum_{j=1}^{n_1} \left[\alpha_{0j} + \sum_{k=2}^{n} \alpha_{kj} ln \frac{p_{ckt}}{p_{c1t}} + \sum_{k=2}^{n} \alpha_{kj} ln \frac{\theta_{ckt}}{\theta_{c1t}} + \sum_{i=2}^{m} \gamma_{ji} ln \frac{v_{cit}}{v_{c1t}} \right]^2 \tag{10}$$

Taking natural logarithms on both sides of Equation (10) and expanding in a first order Taylor-series around [1…1]', we get:[6]

$$lnH_{ct} = \beta - 2ln\left(s_{ct}^{n_1} \right) + \frac{2}{\sum_{j=1}^{n_1} \left(\alpha_{0j} \right)^2} \sum_{j=1}^{n_1} \left(\alpha_{0j} \right) \left[\sum_{k=2}^{n} \alpha_{kj} ln \frac{p_{ckt}}{p_{c1t}} + \sum_{k=2}^{n} \alpha_{kj} ln \frac{\theta_{ckt}}{\theta_{c1t}} + \sum_{i=2}^{m} \gamma_{ji} ln \frac{v_{cit}}{v_{c1t}} \right] + \varepsilon_{ct} \tag{11}$$

We follow Redding (2002) in assuming non-traded goods prices and technology differences as being drawn from an estimable probability function. Further, we use MFN tariffs to capture cross-country differences in relative prices of traded goods as well as preferential tariffs as additional control (see Anderson and van Wincoop, 2004). Moreover, we proxy cross-country differences in relative endowments using GDP per capita (see Helpman, 1987, and Romalis, 2004). Our estimation equation therefore becomes:

[6] The logarithm of H is interpreted as a function of the logarithm of the underlying variables.

$$lnH_{ct} = \delta_1 ln\left(1+\tau_{ct}^{MFN}\right) + \delta_2\left[ln\left(1+\tau_{ct}^{MFN}\right) - \ln\left(1+\tau_{ct}^{P}\right)\right] + \delta_3 lnGDPPC_{ct} + \eta_c + \rho_t + \varepsilon_{ct} \quad (12)$$

where τ_{ct}^{MFN} is the average MFN tariff *set* by country *c* at time *t;* τ_{ct}^{P} is the average preferential tariff set by country *c* at time *t* (within the Latin American Integration Association LAIA or in the most important sub-regional trade agreement in which the country is member is member), so that the term in brackets is a measure of preferential margin; $GPDPPC_{ct}$ is the Gross Domestic Product per Capita of country *c* at time *t;* η_c are country fixed effects that control for any permanent country-specific barriers to trade (e.g., remoteness), any permanent country-differences in technology (e.g., associated as social infrastructure, see Hall and Jones, 1999) and/or any permanent country-differences in the relative importance of manufacturing; and ρ_t are year fixed effects that capture common changes in relative prices, technologies, factor endowments, and manufacturing shares across countries.

One well known result of the standard international trade theory is that unilateral trade liberalization induces countries to specialize according to their *overall* comparative advantage. Declining external trade barriers are thus associated with a diminished range of commodities domestically produced and an expanded range of goods import from abroad, i.e., increased specialization. This result holds both in the Ricardian as well as in the Heckscher-Ohlin model (see Dornbusch et al., 1977, and Romalis, 2004). We expect then the estimated coefficient on MFN to be significantly negative, i.e., higher tariffs lead to sectoral diversification in production and vice versa. On the other hand, regional trade integration leads economies to specialize according to their *regional* comparative advantage (see Venables, 2003). This would also imply a reduction in the range of goods produced at home. *A priori*, conditional on the degree of external openness, this might further increase specialization, i.e., stronger concentration of manufacturing activity in fewer sectors.

Furthermore, a negative estimated coefficient on GDP per capita is also expected. This can be explained in terms of preference or portfolio arguments. In the presence of non-homothetic preferences, changing consumption patterns towards greater diversity prevails as income growth, which induces matching changes in the structure of production when trade costs are very high (see Imbs and Wacziarg, 2003). The second reason relates to the fact that, because of indivisibilities, sectoral diversification opportunities improve with the aggregate capital stock and the level of development (see Acemoglu and Zilibotti, 1997).

The Herfindahl index is a common measure of sectoral concentration in the international trade and economic geography literatures (see, e.g., Sapir, 1996, Haaland et al., 1999). This is however only one measure among many different ones and, *a priori,* there is no reason to favor one over the other. Therefore, to check the robustness of our results, we follow Imbs and Wacziarg (2003) in using also the Gini coefficient based on sector shares. This coefficient, like the previous indicator, increases with higher inequality in sectoral shares (see Cowell, 2002).

As a result of the trade policy reforms implemented in the region, countries may not only become more or less absolutely specialized but also more or less similar to each other. This corresponds to the notion of relative specialization. In particular, a country is relatively specialized if its economic structure differs from that of some reference benchmark, such as the region as whole or the remaining countries that belong to the relevant economic space, jointly or individually considered. Here we will focus on bilateral relative specialization, i.e., to what extent the sectoral composition of manufacturing value added differs across pairs of countries. To compare countries' economic structures we use the Krugman index (see Midelfart-Knarvik et al., 2000). Formally:

$$K_{cdt} \equiv (1/2)\sum_{j=1}^{n_1}\left|s_{cjt}^{n_1} - s_{djt}^{n_1}\right| \tag{13}$$

This index ranges between *0* if both countries have an identical industrial structure and takes a maximum of *1* if they have no industries in common. Hence, the more unequal the countries' sectoral shares, the greater the relative specialization.

Taking into account that $s_{cjt} - s_{djt} = s_{ct}^{n_1} s_{cjt}^{n_1} - s_{dt}^{n_1} s_{djt}^{n_1}$ and substituting for the country-specific sectoral shares in Equation (13), we obtain:

$$\begin{aligned} K_{cdt} &\equiv (1/2)\sum_{j=1}^{n_1}\left|s_{cjt}^{n_2} - s_{djt}^{n_2}\right| \\ &= (1/2)\sum_{j=1}^{n_1}\left|\lambda_{cd} + \sum_{k=2}^{n}\alpha_{kj}\left(\frac{1}{s_{ct}^{n_1}} ln\frac{p_{ckt}}{p_{c1t}} - \frac{1}{s_{dt}^{n_1}} ln\frac{p_{dkt}}{p_{d1t}}\right) + \sum_{k=2}^{n}\alpha_{kj}\left(\frac{1}{s_{ct}^{n_1}} ln\frac{\theta_{ckt}}{\theta_{c1t}} - \frac{1}{s_{dt}^{n_1}} ln\frac{\theta_{dkt}}{\theta_{d1t}}\right) + \sum_{i=2}^{m}\gamma_{ji}\left(\frac{1}{s_{ct}^{n_1}} ln\frac{v_{cit}}{v_{c1t}} - \frac{1}{s_{dt}^{n_1}} ln\frac{v_{dit}}{v_{d1t}}\right)\right| \end{aligned} \tag{14}$$

where we assume that if *considered* factors are equal across countries, then their production structures differ by a constant summarizing any other influences. Taking natural logarithms on both sides of Equation (14), expanding in a first–order Taylor series around [1…1]', and making similar assumptions as before, we derive the following estimation equation:

$$lnK_{cdt} = \varphi_1 \left| ln\left(1+\tau_{ct}^{MFN}\right) - ln\left(1+\tau_{dt}^{MFN}\right) \right| + \varphi_2 ln\left(1+\tau_{cdt}^{P}\right) + \varphi_3 \left| lnGDPPC_{ct} - lnGDPPC_{dt} \right| + \mu_{cd} + \psi_t + \nu_{ct} \qquad (15)$$

where we use differences in MFN tariffs and the average bilateral (preferential) tariffs to measure differences in relative prices and differences in GDP per capita to account for differences in relative endowments. The country-pair effects μ_{cd} control for any permanent country-pair specific differences in trade barriers (e.g., bilateral distances) and/or any permanent differences in technology (e.g., originated in distinct institutional settings), while the time fixed effects ψ_t capture common changes across countries in relative prices, technology, and factor endowments.

Given that Latin American countries have comparative advantage in a relatively homogenous subset of industries with respect to the rest of the world, our hypothesis is that the sign of the estimated coefficient on the first term (i.e., bilateral differences in MFN tariffs) will be positive, i.e., larger differences in MFN tariffs and thus in the nominal average protection conceded to domestic production will result in increased divergence of industrial structures. On the other hand, the impact of bilateral tariff barriers on the degree of sectoral dissimilarity of manufacturing production will be negative, i.e., for given extra-zone trade impediments, bilateral (preferential) trade liberalization will induce inter-industry specialization so that countries' economic structure will tend to be more dissimilar. Under the relative tariff structure associated with a preferential trade arrangement, some commodities that were originally domestically produced become to be imported from those partners that, even with a comparative disadvantage relative to the rest of the world, have a *regional* comparative advantage, so that sectoral composition of countries' production may be expected to diverge. We are of course aware that opening may favor intra-industry specialization (see Frankel and Rose, 1998). However, as discussed in Venables (2003), we believe that the first case is more likely among developing countries like the ones we are considering here. Finally, we expect the difference in the relative endowments and levels of development (i.e., differences in GDPPC) to be positive.

3.2 Econometric Issues

In estimating Equations (12) and (15), there are several econometric issues that must be addressed. First, our raw dependent variables, the absolute specialization index *H* and the relative specialization index *K*, can only adopt values within [0,1] so that they are truncated. As a consequence, classical estimation will lead to biased estimates. Their natural logarithms range in $(-\infty,0)$ and thus only partially solve the problem. We therefore perform a logistic transformation, similar to Balassa and Noland (1989), to check whether this makes a difference. These variables become then $ln(H/(1-H))$ and $ln(K/(1-K))$ with both ranging in $(-\infty,+\infty)$. Since we do not observe significant differences between those results obtained using the natural logarithms of the specialization indicators and those found using their logistic transformations, we will only report the former ones.

In addition, the standard error component model assumes that the regression disturbances are homoscedastic with the same variance across time and across individuals. This is undoubtedly a very restrictive assumption. Given the panel nature of the data, one can presume that there may be a specific pattern of disturbances associated to the presence of groups of observations. Thus, cross-sectional units may be size-asymmetric and as a result may have different variations (see Baltagi, 1995). Furthermore, the basic model assumes that the error terms are not correlated across individuals. However, economies are not only tied to specific factors, they are also tied to common macroeconomic factors affecting the region as a whole (see Greene, 1997) and likely with differential repercussions across groups of nations. Hence, it seems likely that disturbances could be correlated across countries. Finally, the classical LSDV model assumes that the only correlation over time is due to the presence of the same individual across the panel. In particular, the equicorrelation coefficient is the same no matter how far periods are in time. Clearly, this is also a restrictive assumption for the economic relationships under consideration, as an unobserved shock in the current period might affect the specialization patterns for at least some coming periods (see Baltagi, 1995). Ignoring groupwise heteroscedasticity, cross sectional correlation and/or serial correlation when they are present results in consistent but inefficient estimates of the regression coefficients and biased standard errors. Therefore, we have performed relevant test statistics for identifying such data features.

The modified Wald statistic for groupwise heteroscedasticity in residuals suggests that the null hypothesis of homoscedasticity across panels should be rejected. In addition, the Breusch-Pagan LM test indicates that the null hypothesis of independence of error across panels should be also rejected. Finally, the Baltagi-Li LM test for first order serial correlation in a fixed effects model points out that the null hypothesis of no autocorrelation should be rejected, too. Hence, an estimation strategy that corrects these non-spherical disturbances is required. We remove autocorrelation from the data using the Prais-Winsten transformation and, since the number of cross sectional units is similar to the number of time periods, we then apply LS but replacing LS standard errors with panel-corrected standard errors accounting for heteroscedasticity and contemporaneous correlation across panels as indicated in Beck and Katz (1996).

Furthermore, estimating Equations (12) and (15) without controlling for relevant additional time-varying factors, may result in biased estimates (see Greene, 1997). First, Imbs and Wacziarg (2003) have uncovered a non-monotonic pattern of sectoral concentration of economic activity across the development path. More specifically, higher levels of GDP per capita are associated with lower degrees of absolute specialization up to a certain point and increased specialization thereafter. According to Imbs and Wacziarg (2003), this relationship might emerge as a result of different forces prevailing along the development process: forces pushing for diversification, namely, non-homothetic preferences on the consumption side and portfolio arguments on the investment side, and forces favoring specialization, namely, declining trade barriers as in the Ricardian model and spatial concentration of economic activities with increasing returns to scale as highlighted by new economic geography models (see, e.g., Fujita et al., 1999, and Baldwin et al., 2003). We therefore include a squared GDP per capita term to account for this non-linear relationship in Equation (12) to avoid a likely omitted variable bias. We expect the estimated coefficients on GDPPC and GDPPC squared to be negative and positive, respectively.

Another possible source of omitted variable biases are the episodes of real exchange rate misalignment and, in particular, real appreciation that have been observed in Latin America as countries implemented macroeconomic stabilization programs during the sample period (see Edwards, 1994). In order to control for the influence of those phenomena on the real economy, we add an index of effective exchange rate for imports or a measure of the level of real exchange rate misalignment to Equation (12). A low and particularly an overvalued real exchange rate (i.e., below the equilibrium level) favors imports over domestic production and thus may induce a concentration of economic activity in sectors in which the country has significant comparative advantage, i.e., greater production specialization. We are also aware that this effect may be stronger, the lower the trade barriers. An interaction between the average MFN tariffs and the real exchange rate variables will be therefore also considered. On the other hand, a low/overvalued real exchange rate may facilitate the imports of required inputs (capital goods) by firms in a broader set of sectors that might become internationally competitive producers and thus may foster sectoral diversification.[7] The impact of real exchange rate on absolute production specialization is then an empirical question. We also control for the effect of the exchange rate behavior when examining relative specialization. In particular, we introduce the absolute difference of the aforementioned measures of real exchange rate in Equation (15). These are expected to have a positive impact on the degree of difference of countries' manufacturing structures.

Finally, some endogeneity problems may be involved. Tariffs, both MFN and preferential, may endogenous. Thus, protectionism could be expected to be fiercer in larger, more diversified economies, since trade liberalization would affect many sectors and larger shares of population. On the other hand, smaller, less diversified economies have traditionally received special treatment in Latin America. These countries have conceded smaller tariff preferences to larger neighbors at least for certain periods. In addition, GDPPC as well as the real exchange rate variables may be endogenous. In particular, we can think of a simultaneity bias. Endogenous growth models highlight that an economy's pattern of international specialization and its rate of economic growth are jointly and endogenously determined (see, e.g., Redding, 1999). The same is also true for the real exchange rate and countries' specialization patterns (see, e.g., Obstfeld and Rogoff, 1996). Moreover, highly specialized countries are more prone to suffer from idiosyncratic business cycles and thus from higher expected exchange rate variability and larger average misalignments.

In order to check the robustness of our results we have therefore carried out GMM estimations and performed the Sargan and Hansen tests for overidentifying restrictions. In particular, two main dynamic panel estimators can be considered: those proposed by Arellano and Bond (1991) ("Differenced GMM") and Blundell and Bond (1998) ("System GMM"). It is well known that for short panels with a large number of cross sectional units highly persistent series lead to severe finite sample bias in the first case, because the lagged levels are weak instruments of the differences. This is not the case for our econometric analysis of absolute manufacturing specialization. The number of time periods (14 years) is larger than the number of panels (10 countries) and, even though there is evidence of persistence, this not strong enough to be a cause of concern.[8] We will therefore only report

[7] Rebelo and Vegh (1995) maintain that exchange rate-based stabilizations have tended to be associated with real exchange rate appreciations and sharp deterioration of external accounts reflecting a large increase of durables and capital goods imports.

[8] The estimated *rho* parameter is around 0.400. The estimated coefficient on the lagged dependent variable according to the bias-corrected LSDV estimator developed by Kiviet (1995) is around 0.600. Further, we find that

estimates based on the method proposed by Arellano and Bond (1991). In contrast, the data used to perform estimations on relative specialization requires a more careful investigation, as the number of cross-sectional units, i.e., country pairs (45), is significantly larger than the number of years (14). Further, the number of observations is relatively large (around 500), which allows using a richer set of instruments. Hence, we will also present results obtained according to the method developed by Blundell and Bond (1998).

4 Trade Policy Reforms in Latin America

During more than 40 years most Latin American countries maintained high tariff barriers with the rest of the world to support a strategy of import-substituting industrialization. Since the mid-1980s these countries started to implement trade policy reforms consisting of sharp cuts of nominal tariffs, reduction of tariff dispersion, and elimination of non-tariff barriers. The pace of these reforms was particularly rapid during the second half of the 1980s and the early 1990s. Thus, simple average MFN tariff in our sample countries fell almost 30 percentage points from 41.57% in 1985 to 13.09% in 2001 with most of this drop taking place between 1985 and 1992. Figure 1 presents the evolution of average MFN tariffs per country over the sample period. Note that in most countries tariffs reached a peak before they begun to be reduced as they were increased in anticipation of the future diminutions to delay effective liberalization and thus smooth the consequent adjustment. At the initial year we can identify three groups of countries with tariffs higher than 50% (Brazil, Colombia, Ecuador, and Peru), countries with intermediate tariffs (Argentina, Mexico, Uruguay, and Venezuela), and countries with average tariffs around and below 20% (Bolivia, Chile, and Paraguay). By 1992, after drastic slashes with varying intensity across countries, the dispersion of MFN tariff in the region had fallen from 17.25 in 1985 to 3.48.

Countries under examination have also opened their economies on a preferential basis. LAIA is an area of economic preferences created in 1980 trough the Montevideo Treaty. In virtue of this arrangement, countries conceded tariff preferences with respect to the rest of the world either to all remaining member nations or to a certain subset of them. Thus, tariffs within the region have been lower than MFN tariffs and have been reduced even more dramatically. The average preferential tariff faced by countries being analyzed

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this coefficient is almost identical when estimating pure autoregressive models using both the Arellano and Bond (1991) and Blundel and Bond (1998) estimators (0.373 and 0.400, respectively).

fell from 39.91% in 1985 to 5.98% in 2001. Figure 2 highlights the evolution of average preferential tariffs set by each county in our sample. Notice that these tariffs experienced a substantial drop during the early 1990s.

The asymmetric path of MFN and preferential tariffs implied an important increase in the average preferential margin, which reached 111.61% in 2001 after beginning with 4.31% in 1985. This regional dimension of trade liberalization was additionally deepened by preferential integration agreements formed by subsets of the countries. The most important initiatives for our sample countries are MERCOSUR, which was established in 1991 by Argentina, Brazil, Paraguay, and Uruguay, and the Andean Community, a trading arrangement formed by Bolivia, Colombia, Ecuador, Peru, and Venezuela. In 2001 average intra-zone tariffs within these blocs ranged between 2% and 3%.

5 The Impact of Trade Policy Reforms on Manufacturing Specialization Patterns

5.1 Trade Policy Reforms and Absolute Manufacturing Production Specialization

Figure 3 plots the trend of the Herfindahl specialization index based on sectoral value added shares for each country in the sample over the period 1985-1998. In general, as expected, the larger countries, Argentina, Brazil, and Mexico, exhibit lower levels of absolute manufacturing specialization. Most Latin American countries are specialized in exploiting their natural resources endowments. The share of food products ranges between 0.10 and 0.25 with the highest relative importance in the industrial structures of Southern Cone countries. In addition, petroleum refineries account for large shares of total manufacturing activity in Argentina, Bolivia, Ecuador, Uruguay, and Venezuela.

Specialization seems to follow an upward trend in most countries. In fact, regressing the Herfindahl Index on a time trend, we find that six out of ten countries experienced significant increases in their overall levels of production specialization: Brazil, Colombia (from 1987 onwards), Ecuador, Peru, Uruguay, and Venezuela.[9] This essentially reflects developments in the countries' larger manufacturing sectors. For example, in Ecuador the combined share of the two largest industries, food products and petroleum refineries, grew

[9] We have tested for non-stationarity to determine whether there are concrete reasons to be concerned about this issue. In particular, we have performed the Levin-Lin-Chu test (Levin et al., 2002) for panel unit roots on the specialization variable for alternative balanced panels. In doing this, we have introduced one lag of this variable to allow for serial correlation in the errors. The null hypothesis of unit root is strongly rejected for all considered configurations, regardless whether we include a time trend or not.

from 0.28 in 1985 to 0.53 in 1996 mainly as a result of the rapid relative expansion of the latter sector. In Uruguay this combined share increased from 0.27 in 1987 to 0.41 in 1998 and in Venezuela from 0.29 to 0.42 over the same lapse. In summary, many Latin American countries displayed increasing absolute manufacturing production specialization during a period characterized by declining trade impediments. To what extent can these changes in specialization patterns be related to the trade policy reforms implemented in this region? The remaining of this subsection aims at answering this question through a formal econometric analysis.

Table 1 reports Prais-Winsten estimations with panel corrected standard errors of different specifications of Equation (12).[10] There is a robust and systematic negative relationship between own MFN tariffs and a country's degree of absolute specialization in industrial production. Hence, unilaterally declining own tariff barriers with respect to the rest of the world is associated with increased sectoral concentration of manufacturing production. On the other hand, average preferential margins conceded by countries in the framework of LAIA (AVPM) or sub-regional integration agreements such as the MERCOSUR and the Andean Community (RIAPM) do not have a significant impact on the level of overall industrial specialization.[11] Regional integration might have triggered changes in the sectoral distribution of manufacturing activity which are different from those promoted by opening up to the rest of the world. In particular, regional trade agreements might have induced countries to become more specialized in sectors in which they have comparative advantage within the region, but not in the external markets. In Latin America, economies have relatively similar comparative advantage patterns vis-à-vis the rest of the world, but have comparative advantage in different industries at the regional level. Further, those sectors which are (not) internationally competitive were initially relatively large (small). Hence, if preferential trade liberalization has fostered specialization in industries with regional comparative advantage more than in those with global comparative advantage and the former were initially smaller, then changes in the distribution of economic activity over sectors will not necessarily lead to a significant increase in the specialization level.

[10] Estimation results with tariff variables lagged one period to account for the possibility that their impact on specialization patterns follow with a lag are essentially the same.These results are available from the authors upon request.

[11] Figure 3 shows that levels of specialization and their evolution over time differ across economies. To exclude the possibility that these findings are driven by specific country experiences, we drop one country at a time from the sample and re-estimate Equation (12). Regression results basically coincide with those shown in Table 1 thus confirming the main messages. We also explore whether estimates are sensitive to changes in the sample period, by successively considering periods beginning in 1986, 1987, 1988, 1989, and 1990. Again, results are essentially the same. These results are not reported, but are available from the authors upon request.

Then, depending on the initial industrial structure, the impact on overall specialization might be weaker in this case.

Previous estimations control for country specific factors that remain constants over time as well as common changes across countries. However, according to economic theory, there are important additional time-varying country-specific factors that may affect the degree of industrial specialization and thus the relation under examination. One of these factors is relative endowment as proxied by the GDP per capita. Results including these variables are presented in Columns 4 and 5 of Table 1. As expected, the estimated coefficient on GDPPC is negative and significant, while that on the squared GDPPC is positive and significant across the different specifications. Hence, there is non-linear (U-shaped) relationship between sectoral concentration of manufacturing value added and the level of per capital income, i.e., first sectoral diversification occurs, but there is a level of per capital income beyond which countries start to specialize again. This coincides with findings reported in Imbs and Wacziarg (2003).

Another important time-varying factor whose omission may lead to biases estimates is the real exchange rate. Estimates obtained when incorporating this variable are shown in Table 2. Without considering potential nonlinearities, the real effective exchange rate for imports seems to be positively related to the overall level of specialization (Column 1). However, when its interplay with trade policy is accounted for, the estimated coefficient on this exchange rate variable is negative and significant, whereas that on the interaction term is positive and significant (Column 2). Thus, we observe that the higher MFN tariffs and real exchange rate, the greater the absolute manufacturing specialization. In the case of real exchange rate misalignment, the same sign pattern prevail, but estimated coefficients are not significantly different from zero (Column 4). Henceforth, there is some evidence suggesting that a high exchange rate induces manufacturing production diversification when trade barriers are low, but promote industrial specialization when coupled with high tariff barriers.

We check the robustness of our results to the specialization measure being used and the econometric strategy. We first replicate previous regressions using the logarithm of the Gini coefficient calculated on sectoral manufacturing value added instead of the Herfindahl index as dependent variable. Estimations are reported in Tables 3 and 4 and they confirm most of our previous findings. Table 5 presents results obtained performing GMM estimations using the procedure proposed by Arellano and Bond (1991). These estimations aim at addressing possible endogeneity problems in the regressions discusses above. Our main conclusion still holds. Unilateral trade liberalization, i.e., reducing own MFN tariffs with respect to the rest of the world is associated with more concentrated manufacturing production structures. Estimated coefficients on remaining variables are similar to those reported in Table 2, except the one on the interaction between real exchange rate misalignment and tariffs, which now becomes significantly positive thus providing additional evidence of non-monotonicities in the impact of exchange rate on overall absolute specialization.[12]

This sub-section has revealed that trade policy reforms in Latin America have had a substantial impact on countries' international specialization patterns. In particular, unilateral opening has favored increasing manufacturing production specialization. Trade liberalization initiatives may also influence relative specialization across country pairs. More specifically, the magnitude of the differences in the extent to which nations liberalize their trade flows with the rest of the world as well as the level of bilateral (preferential) trade barriers may affect how similar is the sectoral composition of manufacturing production across country pairs. The following sub-section assesses this possibility.

5.2 Trade Policy Reforms and Manufacturing Production Structures: Convergence or Divergence?

We measure relative manufacturing specialization with the Krugman index. This index quantifies the degree of bilateral sectoral disparity of industrial structures. Figure 4 presents the Krugman index for each country pair in our sample over the period 1985-1998. According to a simple regression of this index on a time trend, 22 out of 45 country pairs became more dissimilar in terms of their manufacturing structures and 7 out of 45 do not exhibit significant changes.[13] Interestingly, half of the 16 cases of reductions involve Bolivia,

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[12] There is an additional, more complicated issue in this analysis, namely, if specialization, tariffs, and the level of development are jointly determined, there will be both direct and indirect effects that should be disentangled. For example, real exchange rate misaligments may have direct effects on specialization as well as indirect effects through the influence that they likely exert on choosing the level of tariff protection. We have therefore estimated alternative specifications of systems of equations by 3SLS and GMM where these variables are simultaneously treated as endogenous. This multiple equation strategy generates similar results to those from the single equation one. In addition, we observe that less diversified economies indeed set lower tariffs and tend to have lower levels of GDP per capita. These results are not reported but are available from the authors upon request.

[13] We have tested for non-stationarity of the specialization measure also in thie case. The Levin-Lin-Chu test (Levin et al., 2002) for panel unit roots suggests that series are stattionary. This holds regardless whether we include a time trend or not and the specific balanced panel being considered.

which moved towards convergence in terms of sectoral distribution of economic activity after starting as the most dissimilar nation for each of its Latin American partners.[14]

Table 6 presents Prais-Winsten estimations with panel corrected standard errors of Equation (14). Two main results outstand. As expected, larger differences in the degree of unilateral openness with respect to the rest of the world and deeper bilateral (preferential) trade liberalization foster a higher degree of relative specialization, i.e., more dissimilar industrial structures. This is consistently true from the basic specification also when the level of development and variable reflecting the level of real exchange rate are introduced. In this sense, we should mention that larger differences in the real effective exchange rate for imports are associated with larger disparities in the sectoral distribution of manufacturing production.

Table 7 shows that the core findings are confirmed after implementing GMM procedures to correct endogeneity biases: the Arellano and Bond (1991) estimator and the Blundell and Bond (1998) estimator. Two considerations deserve being made. First, while differences in the real exchange rate for imports seem to be an important factor influencing cross-country differences in manufacturing structures according to the former estimator, differences in the degree of real exchange rate misalignment is the relevant one when the latter estimator is used. Second, although both sets of estimates are consistent as suggested by the respective specification tests, these distinct results across methods recommend to perform a carefully comparison. As mentioned before, highly persistent series may generate weak instrument problems and thus serious finite sample biases when applying the Arellano and Bond estimator. This may be detected by comparing the estimated coefficient on the lagged dependent variable to those from OLS, which is upward biased, and LSDV (Within), which is downward biased (see Bond, 2002). In our case, these coefficients are 0.854 and 0.497, respectively, for the model specification shown in Columns 2 and 5. Estimates reported in Table 7 indicate that there is indeed evidence that the Arellano and Bond estimator is affected by finite sample bias. In contrast, the Blundel and Bond estimator produces an estimate which is well below the OLS one and well above the LSDV one thus appearing as our preferred estimation strategy.

Hence, regional trade integration seems to have spurred inter-industry specialization across countries. This has profound macroeconomic implications. If countries

[14] Bolivia is one the poorest country in the region and suffered from a severe hyperinflation episode during the first part of the 1980s.

become more dissimilar in terms of their production structures and thus more sensitive to specific industry shocks, more idiosyncratic business cycles would prevail (see Kenen, 1969, Eichengreen, 1992, and Krugman, 1993) and, if exchange rates are used as an adjustment mechanism to dampen cyclical fluctuations, higher bilateral exchange rate variability should be expected. This, in turn, might act as channel of agglomeration of economic activities in the larger countries in the region (see Ricci, 1998) and might promote reversions in the integration process in the form of reinsertion of protectionist measures (see Eichengreen, 1993, and Fernández-Arias et al., 2002).

6 Concluding Remarks

This paper has aimed at answering one main question: Did Latin American countries become more and differently specialized as a consequence of trade policy reforms? Our econometric analysis shows that the answer is *yes*.

Unilateral trade liberalization has resulted in increased absolute manufacturing production specialization. Weinhold and Rauch (1999) show that, at least for developing countries, this might have a positive impact, since specialization appears to be positively and significantly correlated with manufacturing productivity growth. One possible explanation for this result comes from models with endogenous growth through learning-by-doing. In this framework, increased openness to international trade can lead to increased specialization, which in turn accelerates productivity growth by more fully realizing dynamic economies of scale. Of course, not only the degree, but also the nature of specialization is important (see Redding, 1999, and Bensidoun et al., 2001). In Latin America, industrial activity has on average become increasingly concentrated in sectors using intensively natural resources endowments. What does this imply for this region? According to Perri et al. (2001), the experience of other countries such as Australia, Canada, Sweden and Finland show that these rich endowments, when properly combined with policies stimulating the adoption of new technologies, are a proven growth recipe.

Preferential trade liberalization has favored a broadening of the disparities between countries' economic structures. This has important implications for the regional macroeconomics and the sustainability of ongoing integration processes. Inter-industry specialization and thus higher sensitivity to industry specific shocks are associated with more idiosyncratic business cycles and higher expected exchange rate variability, which in turn affect locational incentives and trade flows generating pressures for reintroducing protectionist measures.

The previous findings seem to be quite robust. They hold regardless the specialization measure being used, the inclusion of control variables such the real effective exchange rate for imports and the real exchange rate misalignment, and remain valid after using GMM procedures to correct biases originated in serial correlation and endogeneity.

Figure 1

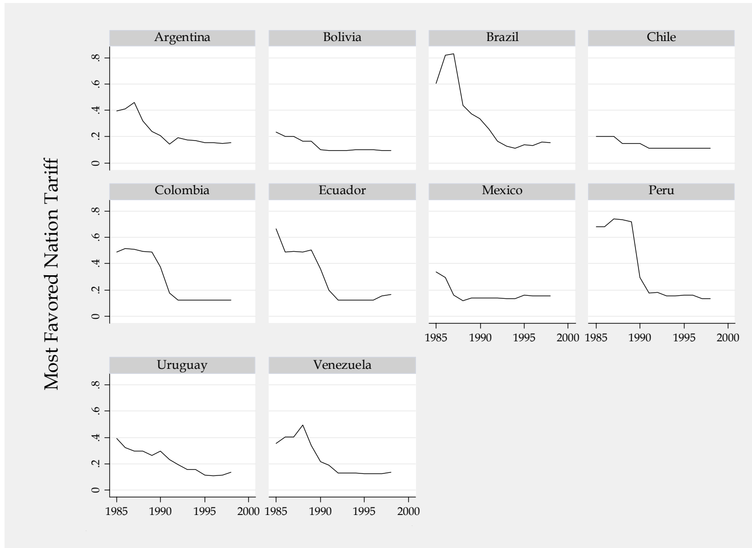


Figure 2

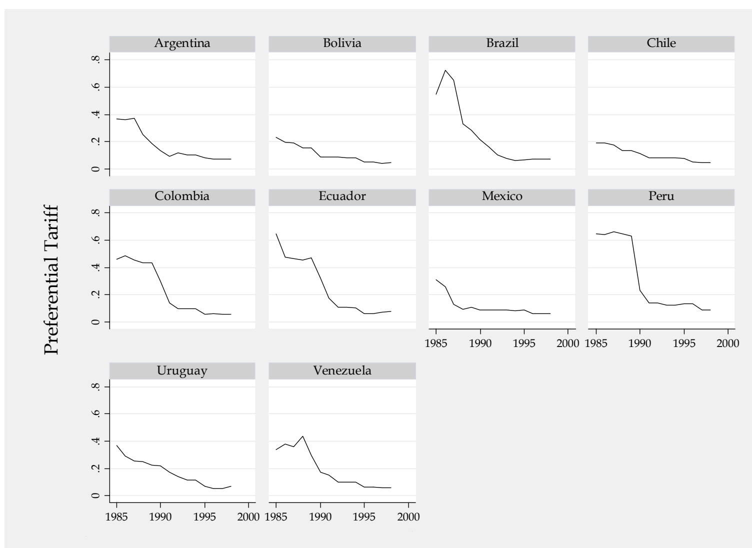
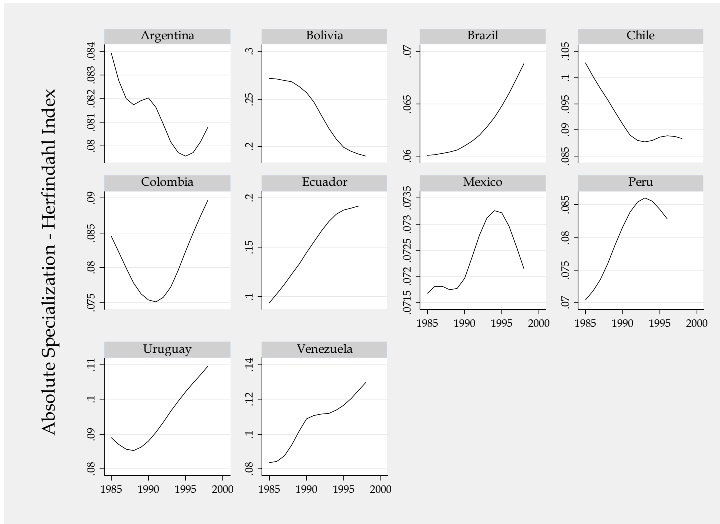
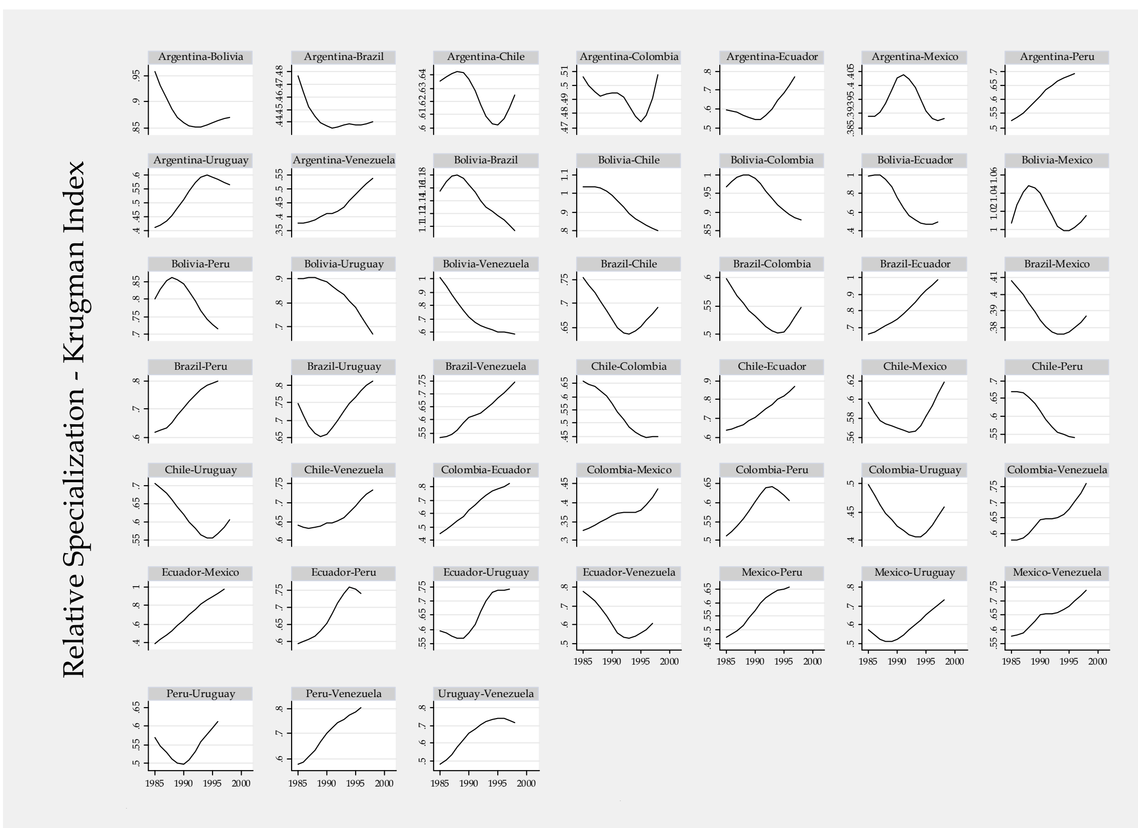


Figure 3



The figure shows the trend of the Herfindahl Index for each of the sample countries as obtained using the filter proposed by Hodrick and Prescott (1997).

Figure 4



The figure shows the trend of the Krugman Index for each of the sample country pairs as obtained using the filter proposed by Hodrick and Prescott (1997).

Table 1: The Impact of Unilateral and Preferential Trade Policy on Absolute Specialization

	(1) H	(2) H	(3) H	(4) H	(5) H
MFN	-0.245***	-0.308***	-0.301***	-0.228***	-0.220***
	(0.061)	(0.080)	(0.068)	(0.077)	(0.075)
AVPM		0.098			
		(0.077)			
RIAPM			0.124	0.096	0.097
			(0.079)	(0.079)	(0.082)
GDPPC				-0.559**	-7.347**
				(0.229)	(3.342)
GDPPC2					6.485**
					(3.238)
Country-FE	Yes	Yes	Yes	Yes	Yes
Year-FE	Yes	Yes	Yes	Yes	Yes
Observations	137	137	137	137	137

The table reports Prais-Winsten estimates with panel corrected standard errors as suggested in Beck and Katz (1996) (correctection for groupwise heteroscedasticity, cross-sectional correlation, and serial correlation). Reported estimated coefficients are standardized, i.e., slopes are multiplied by the standard deviation of the respective explanatory variable and divided by the standard deviation of the dependent variable. Sample size is defined as in Table A1 in the Appendix. The dependent variable, H, is the natural logarithm of the overall level of specialization as measured by the Herfindahl Index. MFN is the natural logarithm of the average Most Favored Nation Tariff set by the country plus one. AVPM is the average preferential margin conceded by the country to Latin American partners, i.e., the natural logarithm of MFN plus one minus the natural logarithm of AVPT plus one, where AVPT is the average preferential tariff applied on trade flows with members of the LAIA. RIAPM is the average preferential margin conceded by the country within the most important RIA with Latin American partners, i.e., the natural logarithm of MFN plus one minus the natural logarithm of RIAPT plus one, where RIAPT is the average preferential tariff applied on trade flows with members of this agreement. GDPPC is the natural logarithm of Gross Domestic Product per capita (GDPPC2: squared). * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 2: Absolute Specialization, Trade Policy, Level of Development, and Exchange Rate

	(1) H	(2) H	(3) H	(4) H
MFN	-0.271**	-0.681***	-0.213**	-0.230***
	(0.086)	(0.162)	(0.083)	(0.080)
RIAPM	0.169	0.121	0.032	0.028
	(0.106)	(0.104)	(0.081)	(0.085)
GDPPC	-9.725**	-10.632***	-8.034**	-8.767**
	(4.120)	(3.810)	(3.626)	(3.500)
GDPPC2	8.825**	9.485***	7.190**	7.835**
	(3.866)	(3.552)	(3.422)	(3.294)
REERM	0.062*	-0.148*		
	(0.038)	(0.077)		
RERMIS			0.055*	-0.054
			(0.033)	(0.074)
MFN x REERM		0.005***		
		(0.002)		
MFN x RERMIS				0.111
				(0.070)
Country-FE	Yes	Yes	Yes	Yes
Year-FE	Yes	Yes	Yes	Yes
Observations	117	117	134	134

The table reports Prais-Winsten estimates with panel corrected standard errors as suggested in Beck and Katz (1996) (correctection for groupwise heteroscedasticity, cross-sectional correlation, and serial correlation). Reported estimated coefficients are standardized, i.e., slopes are multiplied by the standard deviation of the respective explanatory variable and divided by the standard deviation of the dependent variable. Sample size is defined as in Table A1 in the Appendix. The dependent variable, H, is the natural logarithm of the overall level of specialization as measured by the Herfindahl Index. MFN is the natural logarithm of the average Most Favored Nation Tariff set by the country plus one. RIAPM is the average preferential margin conceded by the country within the most important RIA with Latin American partners, i.e., the natural logarithm of MFN plus one minus the natural logarithm of RIAPT plus one, where RIAPT is the average preferential tariff applied on trade flows with members of this agreement. GDPPC is the natural logarithm of Gross Domestic Product per capita (GDPPC2: squared). REERM is the real effective exchange rate for imports calculated by the ECLAC for the period 1987-2001. REERMIS is the real exchange rate misalignment estimated by Terra and Valladares (2003) for the period 1985-1998. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 3: The Impact of Unilateral and Preferential Trade Policy on Absolute Specialization (Alternative Specialization Measure)

	(1) G	(2) G	(3) G	(4) G	(5) G
MFN	-0.207***	-0.283***	-0.261***	-0.144**	-0.138**
	(0.056)	(0.073)	(0.064)	(0.065)	(0.063)
AVPM		0.120*			
		(0.070)			
RIAPM			0.120*	0.076	0.074
			(0.073)	(0.070)	(0.072)
GDPPC				-0.954***	-6.513**
				(0.220)	(2.717)
GDPPC2					5.314**
					(2.524)
Country-FE	Yes	Yes	Yes	Yes	Yes
Year-FE	Yes	Yes	Yes	Yes	Yes
Observations	137	137	137	137	137

The table reports Prais-Winsten estimates with panel corrected standard errors as suggested in Beck and Katz (1996) (correctection for groupwise heteroscedasticity, cross-sectional correlation, and serial correlation). Reported estimated coefficients are standardized, i.e., slopes are multiplied by the standard deviation of the respective explanatory variable and divided by the standard deviation of the dependent variable. Sample size is defined as in Table A1 in the Appendix. The dependent variable, G, is the natural logarithm of the overall level of specialization as measured by the Gini Coefficient. MFN is the natural logarithm of the average Most Favored Nation Tariff set by the country plus one. RIAPM is the average preferential margin conceded by the country within the most important RIA with Latin American partners, i.e., the natural logarithm of MFN plus one minus the natural logarithm of RIAPT plus one, where RIAPT is the average preferential tariff applied on trade flows with members of this agreement. GDPPC is the natural logarithm of Gross Domestic Product per capita (GDPPC2: squared). * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 4: Absolute Specialization, Trade Policy, Level of Development, and Exchange Rate (Alternative Specialization Measure)

	(1) G	(2) G	(3) G	(4) G
MFN	-0.188**	-0.554***	-0.147**	-0.165**
	(0.073)	(0.132)	(0.073)	(0.070)
RIAPM	0.146	0.109	0.036	0.026
	(0.096)	(0.090)	(0.063)	(0.066)
GDPPC	-8.960***	-9.875***	-7.038**	-7.633***
	(2.926)	(2.728)	(2.800)	(2.666)
GDPPC2	7.796***	8.500***	5.893**	6.403***
	(2.690)	(2.486)	(2.605)	(2.473)
REERM	0.059*	-0.121*		
	(0.032)	(0.065)		
RERMIS			0.053*	-0.064
			(0.027)	(0.062)
MFN x REERM		0.004***		
		(0.001)		
MFN x RERMIS				0.117*
				(0.062)
Country-FE	Yes	Yes	Yes	Yes
Year-FE	Yes	Yes	Yes	Yes
Observations	117	117	134	134

The table reports Prais-Winsten estimates with panel corrected standard errors as suggested in Beck and Katz (1996) (correctection for groupwise heteroscedasticity, cross-sectional correlation, and serial correlation). Reported estimated coefficients are standardized, i.e., slopes are multiplied by the standard deviation of the respective explanatory variable and divided by the standard deviation of the dependent variable. Sample size is defined as in Table A1 in the Appendix. The dependent variable, G, is the natural logarithm of the overall level of specialization as measured by the Gini Coefficient. MFN is the natural logarithm of the average Most Favored Nation Tariff set by the country plus one. RIAPM is the average preferential margin conceded by the country within the most important RIA with Latin American partners, i.e., the natural logarithm of MFN plus one minus the natural logarithm of RIAPT plus one, where RIAPT is the average preferential tariff applied on trade flows with members of this agreement. GDPPC is the natural logarithm of Gross Domestic Product per capita (GDPPC2: squared). REERM is the real effective exchange rate for imports calculated by the ECLAC for the period 1987-2001. REERMIS is the real exchange rate misalignment estimated by Terra and Valladares (2003) for the period 1985-1998. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 5: Absolute Specialization – GMM Estimations

	(1) H	(2) H	(3) H
MFN	-0.168*	-0.611**	-0.220*
	(0.094)	(0.250)	(0.111)
RIAPM		0.082	0.071
		(0.077)	(0.041)
GDPPC		-9.468**	-7.289*
		(4.133)	(3.864)
GDPPC2		8.464**	6.603*
		(3.834)	(3.620)
REERM		-0.142**	
		(0.061)	
RERMIS			-0.136
			(0.092)
MFN x REERM		0.005**	
		(0.002)	
MFN x REERMIS			0.265*
			(0.156)
Year-FE	Yes	Yes	Yes
Sargan Test	86.940	79.250	93.660
[p-value]	[1.000]	[1.000]	[1.000]
Test for 2nd O. A.	-0.160	-0.070	0.190
[p-value]	[0.872]	[0.945]	[0.846]
Observations	117	107	114

The table reports one step GMM estimations according to the procedure proposed by Arellano and Bond (1991) with all explanatory variables treated as endogenous. A lagged dependent variable is included (not reported). Reported estimated coefficients are standardized, i.e., slopes are multiplied by the standard deviation of the respective explanatory variable and divided by the standard deviation of the dependent variable. Sample size is defined as in Table A1 in the Appendix. The dependent variable, H, is the natural logarithm of the overall level of specialization as measured by the Herfindahl Index. MFN is the natural logarithm of the average Most Favored Nation Tariff set by the country plus one. RIAPM is the average preferential margin conceded by the country within the most important RIA with Latin American partners, i.e., the natural logarithm of MFN plus one minus the natural logarithm of RIAPT plus one, where RIAPT is the average preferential tariff applied on trade flows with members of this agreement. GDPPC is the natural logarithm of Gross Domestic Product per capita (GDPPC2: squared). REERM is the real effective exchange rate for imports calculated by the ECLAC for the period 1987-2001. REERMIS is the real exchange rate misalignment estimated by Terra and Valladares (2003) for the period 1985-1998. Robust standard errors are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 6: Relative Specialization, Trade Policy, Level of Development, and Exchange Rate

	(1) K	(2) K	(3) K	(4) K	(5) K
DIFF MFN	0.055*	0.104***	0.112***	0.111***	0.119***
	(0.032)	(0.036)	(0.036)	(0.033)	(0.037)
AVPT		-0.277***	-0.276***	-0.228***	-0.332***
		(0.087)	(0.085)	(0.084)	(0.088)
DIFF GDPPC			0.261***	0.303***	0.236***
			(0.090)	(0.087)	(0.091)
DIFF REERM				0.081***	
				(0.031)	
DIFF RERMIS					-0.017
					(0.020)
Country Pair-FE	Yes	Yes	Yes	Yes	Yes
Year-FE	Yes	Yes	Yes	Yes	Yes
Observations	604	604	604	514	578

The table reports Prais-Winsten estimates with panel corrected standard errors as suggested in Beck and Katz (1996) (correctection for groupwise heteroscedasticity, cross-sectional correlation, and serial correlation). Sample size is defined as in Table A1 in the Appendix. Reported estimated coefficients are standardized, i.e., slopes are multiplied by the standard deviation of the respective explanatory variable and divided by the standard deviation of the dependent variable. The dependent variable, K, is the natural logarithm of the relative level of specialization as measured by the Krugman Index. DIFF MFN is the absolute difference of the natural logarithms of Most Favored Nation Tariffs set by the countries plus one. AVPT is the natural logarithm of the average bilateral preferential tariff plus one. DIFF GDPPC is the absolute difference of natural logarithms of the Gross Domestic Product per capita. DIFF REERM is the absolute difference of the real effective exchange rate for imports calculated by the ECLAC for the period 1987-2001. DIFF REERMIS is the absolute difference of the real exchange rate misalignments estimated by Terra and Valladares (2003) for the period 1985-1998. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 7: Relative Specialization – GMM Estimations

	(1) K	(2) K	(3) K	(4) K	(5) K	(6) K
K(-1)	0.430***	0.390***	0.503***	0.567***	0.638***	0.550***
	(0.066)	(0.055)	(0.067)	(0.220)	(0.192)	(0.127)
DIFF MFN	0.094**	0.074	0.097**	0.139**	0.075	0.157***
	(0.044)	(0.046)	(0.040)	(0.059)	(0.063)	(0.051)
AVPT	-0.218***	-0.170**	-0.237***	-0.442***	-0.285***	-0.308**
	(0.073)	(0.069)	(0.078)	(0.153)	(0.130)	(0.146)
DIFF GDPPC	0.277	0.290*	0.297	0.241*	0.051	0.174
	(0.173)	(0.175)	(0.189)	(0.146)	(0.107)	(0.153)
DIFF RERM		0.069**			-0.125*	
		(0.034)			(0.070)	
DIFF RERMIS			0.034			0.353***
			(0.032)			(0.111)
Year-FE	Yes	Yes	Yes	Yes	Yes	Yes
Sargan/Hansen Test	30.420	35.440	30.610	16.530	24.670	24.180
[p-value]	[1.000]	[1.000]	[1.000]	[0.168]	[0.214]	[0.189]
Test for 2nd O. A.	-0.600	-1.050	0.570	0.910	-0.800	1.350
[p-value]	[0.546]	[0.296]	[0.571]	[0.361]	[0.424]	[0.177]
Observations	514	469	488	559	514	533

Columns (1)-(3) report one-step GMM estimations according to the procedure proposed by Arellano and Bond (1991) with MFN and AVPT treated as predetermined and remaining variables as endogenous. Sample size defined as in Table A1 in the Appendix. The Sargan Test is based on two-step estimations (Arellano and Bond Estimations). Columns (4)-(6) report one-step GMM estimations according to the procedure proposed by Blundell and Bond (1998). Instrumented used are 1-5 lags of DMFN and AVPT, and 3-6 lags of DGDPPC, DRERM, DRERMIS in the level equation; and 2-6 lags of DMFN and AVPT, and 4-7 lags of DGDPPC, DRERM, DRERMIS in the difference equation. Reported estimated coefficients are standardized, i.e., slopes are multiplied by the standard deviation of the respective explanatory variable and divided by the standard deviation of the dependent variable. The dependent variable, K, is the natural logarithm of the relative level of specialization as measured by the Krugman Index (K(-1): lagged one year). DIFF MFN is the absolute difference of the natural logarithms of Most Favored Nation Tariffs set by the countries plus one. AVPT is the natural logarithm of the average bilateral preferential tariff plus one. DIFF GDPPC is the absolute difference of natural logarithms of the Gross Domestic Product per capita. DIFF REERM is the absolute difference of the real effective exchange rate for imports calculated by the ECLAC for the period 1987-2001. DIFF REERMIS is the absolute difference of the real exchange rate misalignments estimated by Terra and Valladares (2003) for the period 1985-1998. * significant at 10%; ** significant at 5%; *** significant at 1%.

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Appendix

Table A1

Value Added: Countries, Sectors, Time Coverage, and Sources				
Country	**Sectoral Coverage**	**Number of Sectors**	**Time Coverage**	**Source**
Argentina	Manufacturing	28	1985-1998	PADI (ECLAC)
Bolivia	Manufacturing	28	1985-1998	IIS (UNIDO)
Brazil	Manufacturing	28	1985-1998	PADI (ECLAC)
Chile	Manufacturing	28	1985-1998	PADI (ECLAC)
Colombia	Manufacturing	28	1985-1998	PADI (ECLAC)
Ecuador	Manufacturing	28	1985-1997	IIS (UNIDO)
Mexico	Manufacturing	28	1985-1998	PADI (ECLAC)
Peru	Manufacturing	28	1985-1996	PADI (ECLAC)
Uruguay	Manufacturing	28	1985-1998	PADI (ECLAC)
Venezuela	Manufacturing	28	1985-1998	PADI (ECLAC)

Table A2

International Standard Industrial Classification (ISIC), Revision 2, 3 digits	
Code	**Description**
311	Food products
313	Beverages
314	Tobacco
321	Textiles
322	Wearing apparel, except footwear
323	Leather and leather products, except footwear and wearing apparel
324	Footwear, except vulcanized or moulded rubber or plastic footwear
331	Wood and wood and cork products, except furniture
332	Furniture and fixtures, except primarily of metal
341	Paper and paper products
342	Printing, publishing and allied industries
351	Industrial chemicals
352	Other chemicals product
353	Petroleum refineries
354	Miscellaneous products of petroleum and coal
355	Rubber products
356	Plastic products not elsewhere classified
361	Pottery, china, and earthenware
362	Glass and glass products
369	Other non-metallic mineral products
371	Iron and steel
372	Non-ferrous metals
381	Fabricated metal products
382	Machinery, except electrical
383	Electrical machinery apparatus
384	Transport equipment
385	Professional, scientific, measuring, controlling, photographic and optic equipment
390	Other manufacturing industries

Table A3

MFN and Preferential Tariffs: Countries, Time Coverage, and Sources			
Country	**Variable**	**Time Coverage**	**Source**
Argentina	MFN	1985-1998	IDB/INT
	PT	1985-1998	IDB/INT
Bolivia	MFN	1985-1998	IDB/INT
	PT	1985-1998	IDB/INT
Brazil	MFN	1985-1998	IDB/INT
	PT	1985-1998	IDB/INT
Chile	MFN	1985-1998	IDB/INT
	PT	1985-1998	IDB/INT
Colombia	MFN	1985-1998	IDB/INT
	PT	1985-1998	IDB/INT
Ecuador	MFN	1985-1998	IDB/INT
	PT	1985-1998	IDB/INT
Mexico	MFN	1985-1998	IDB/INT
	PT	1985-1998	IDB/INT
Peru	MFN	1985-1998	IDB/INT
	PT	1985-1998	IDB/INT
Uruguay	MFN	1985-1998	IDB/INT
	PT	1985-1998	IDB/INT
Venezuela	MFN	1985-1998	IDB/INT
	PT	1985-1998	IDB/INT

Table A4

Real Exchange Rate Measures: Countries, Time Coverage, and Sources			
Country	**Variable**	**Time Coverage**	**Source**
Argentina	REERM	1987-1998	ECLAC
	RERMIS	1985-1998	Terra and Valladares (2003)
Bolivia	REERM	1987-1998	ECLAC
	RERMIS	1985-1998	Terra and Valladares (2003)
Brazil	REERM	1987-1998	ECLAC
	RERMIS	1985-1998	Terra and Valladares (2003)
Chile	REERM	1987-1998	ECLAC
	RERMIS	1985-1998	Terra and Valladares (2003)
Colombia	REERM	1987-1998	ECLAC
	RERMIS	1985-1998	Terra and Valladares (2003)
Ecuador	REERM	1987-1998	ECLAC
	RERMIS	1985-1994	Terra and Valladares (2003)
Mexico	REERM	1987-1998	ECLAC
	RERMIS	1985-1998	Terra and Valladares (2003)
Peru	REERM	1987-1998	ECLAC
	RERMIS	1985-1998	Terra and Valladares (2003)
Uruguay	REERM	1987-1998	ECLAC
	RERMIS	1985-1998	Terra and Valladares (2003)
Venezuela	REERM	1987-1998	ECLAC
	RERMIS	1985-1998	Terra and Valladares (2003)