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by
Michel Beine and Serge Coulombe

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uOttawa

Faculté des sciences sociales
Faculty of Social Sciences

CP 450 SUCC. A
OTTAWA (ONTARIO)
CANADA K1N 6N5

P.O. BOX 450 STN. A
OTTAWA, ONTARIO
CANADA K1N 6N5

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by
Michel Beine
Dulbea, University of Brussels, Belgium

And

Serge Coulombe
University of Ottawa, Canada

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Abstract
We investigate the impact of Canada–U.S. trade integration on the degree of industrial specialization of the Canadian regions. Trade integration is captured through the decrease of trade-weighted tariffs that were boosted by the implementation of the Canadian–U.S. Free Trade Agreement. We found strong evidence to support integration's long-run impact on the patterns of absolute industrial diversification. Significantly, this new finding remains robust to the exclusion of the primary sectors and to the potential presence of unit root in the data. Our results lead us to support a positive long-run relationship between trade integration and industrial diversification.

Keywords: Regional specialization, trade integration, FTA, core-periphery structure, panel data
JEL classification: F14, F15, R12, R15

Résumé
Nous analysons l'incidence de l'intégration économique canado-américaine sur le degré de spécialisation industrielle des régions canadiennes. Le phénomène d'intégration commerciale entre le Canada et les États-Unis est envisagé sous l'angle de la décroissance des tarifs pondérés par le commerce à laquelle a donné lieu l'Accord de libre-échange Canada-États-Unis. Les résultats étayent largement le point de vue selon lequel la libéralisation des échanges aurait favorisé, à long terme, la diversification industrielle des provinces. Ces résultats apparaissent significativement robustes lorsqu'on inclut une racine unitaire dans les données et qu'on exclut le secteur primaire. Nos résultats nous amènent à prévoir une relation positive à long terme entre l'intégration commerciale et la diversification industrielle.

Mots-clés : Spécialisation régionale, intégration commerciale, ALE, structure cœur-périphérie, données en panel
Classification JEL : F14, F15, R12, R15

1. Introduction

The effect of trade integration on the economies of countries and regions has always been one of the most important issues studied in the area of international economics. Since the late 1980s, the traditional theories that support the view of trade integration as a welfare-improving process have been questioned by new theories resting on more realistic hypotheses. Policy makers have a renewed interest in trade integration because of the proliferation of trade agreements during the last decade. More than 150 regional trade agreements are currently in force and more than two-thirds of these take the form of free trade areas and customs unions. Significant examples are the European Union, created in 1957; the Canadian–U.S. Free Trade Agreement (CUSFTA), dating from 1989; followed by the North American Free Trade Agreement (NAFTA) in 1993 and the proposed Free Trade Area of the Americas (FTAA).

Trade integration obviously generates economic benefits. It reportedly induces significant efficiency gains, promotes growth, and increases productivity in the long run. However, trade integration is also suspected of exerting a detrimental effect on at least a subset of the newly integrated areas. One suspected impact is the possible increase of industrial specialization and therefore the enhancement or creation of a core-periphery structure inside the integrated area (Krugman 1991a). Such a development in turn increases the exposure of some regions to specific economic shocks, creates discrepancies in regional economic growth, and last but not least, exacerbates political tensions among regions.

The main aim of this paper is to assess the relationship between trade integration and industrial specialization patterns, using the Canadian–U.S. FTA experience. Employing extensive data on provincial Canadian exports by industry over a period that both precedes (from 1980–89) and covers the formal trade liberalization process, we provide new, striking evidence on the nature of the relationship between economic integration and regional specialization in both the short and long run.

The CUSFTA case is interesting in several respects. First, under the CUSFTA, tariff (and non-tariff) trade barriers between Canada and the United States were gradually eliminated between January 1, 1989, and January 1, 1998. The CUSFTA might be considered one of the most spectacular examples of a trade integration process. For example, Coulombe (2004) reports that between 1980 and 1990, the Canadian international trade share to GDP was relatively constant; between 1990 and 2000, the international trade share to GDP increased steadily from 0.51 to 0.86. The effect, however, was not spread evenly across the provinces. The increase in the trade share was definitely more significant for the central provinces of Ontario and Quebec (Beine and Coulombe 2003). Also, despite this achievement, the project has raised many concerns in official Canadian circles, one key issue being the significant short-run adjustment costs in terms of job losses and worker displacement (Trefler 2002). Another concern was that Canada could become a large peripheral region of the United States, specializing mainly in the production of primary products. This paper will show that, while such a concern perhaps had grounds in theory, there is no evidence whatsoever for such a development.

Basing our investigation on the CUSFTA allows us to contribute to the empirical literature that focuses on the relationship between specialization and integration. The first contribution of this paper lies in the way the trade integration is captured. As Trefler (2002) emphasized, the CUSFTA is a "clean" policy experiment that allows the construction of an exogenous measure of trade integration. Such a feature has been called for extensively in the literature (Rodriguez and Rodrik [1999], among others) to permit meaningful estimates of the effects. The second contribution is that we, unlike previous authors (Head and Ries 2001a; Sawchuk and Trefler 2002; Trefler 2002), conduct our empirical analysis at the regional level. Such an analysis stems from the high degree of heterogeneity across Canadian provinces. For the purpose of our investigation, because of discrepancies in tariffs' initial production structure, the reduction in tariffs with the United States led to a differentiated integration process. Consequently, the pooling of cross-sectional and time-series data multiplies the information regarding the nature of the integration process. This in turn allows us to estimate econometric relationships that are more consistent than the ones obtained in a pure time-series or pure cross-sectional analysis. A third contribution of our study is also related to the cross-sectional and time-series nature of the data. It is well known that it is very difficult to get information regarding the dynamics of change solely from pure cross-sectional evidence. Likewise, obtaining precise estimates of the nature of the dynamic adjustment process from data in one time series is very cumbersome, given the strong multicollinearity among the explanatory variables. In a panel set-up, the cross-sectional dimension reduces the collinearity problem considerably while the time-series dimension allows the explicit disentanglement of the short- and long-run effects of shocks such as the CUSFTA. As advocated by some authors (Trefler 2002), the distinction between short-run adjustment costs and long-run efficiency gains is of primary importance when assessing the impact of trade liberalization. Regarding the impact of the CUSFTA on industrial specialization, such a distinction turns out to be of overwhelming importance.

The paper is organized as follows. Section 2 provides a brief overview of the theoretical literature on the impact of trade integration as well as a summary of the empirical studies devoted to Europe and North America. Section 3 presents the data and discusses the computation of the specialization indices as well as the measures of trade integration. Section 4 describes our econometric methodology and reports the most important findings that are valid under the assumption of stationary processes for integration and specialization patterns. Section 5 analyzes the robustness of non-stationary properties of these indicators. Section 6 concludes.

2. Background

2.1 Theoretical background

Traditional international trade theories, based either on comparative advantages (Ricardo) or factor endowments, emphasize a clear, positive relationship between trade integration and specialization. The theories, however, are based on strong assumptions of constant returns to scale and perfect competition. They are, furthermore, at odds with the fact that trade around the world involves similar rather than different products (intra-industry as opposed to inter-industry trade). The new economic geography (Krugman 1991a) obviously avoids these assumptions and provides new insights into the impact of trade integration.

New trade theories have paid extensive attention to the agglomeration process of economic activities. They do not, however, provide a clear-cut prediction about the relationship between trade integration and industrial specialization. There are two main reasons for this. First, while early contributions (Krugman 1991a; Krugman and Venables 1995) emphasize the decrease in transport costs as a strong agglomerating force, recent contributions (Puga [1999], for example) show that a significant decrease in these costs can result in a geographical dispersion of economic activities when labor mobility across sectors exhibits a finite cost (rather than an infinite cost as assumed in Krugman [1991b] for instance). As underlined by Fujita and Thisse (1997), a non-linear relationship emerges between trade costs and agglomeration. An incomplete integration of markets would create a strong agglomeration pattern while complete integration (in the form of very low trade costs) would result in a new dispersion of economic activities.

Second, one should be aware that concentration of activities and specialization are two different concepts. A decrease in trade costs can initially induce a particular industry to locate in a particular region or to concentrate its activities there. At this point, the result is some specialization of activities in the involved region and to a lesser extent, in the other regions as well. In a second step, however, backward/forward linkages are at work and may trigger a traditional agglomeration process. Indeed, as explained by Krugman (1991a), due to the preference by both consumption and production for a variety of goods, there is a circular causality in the formation of clusters of firms and workers. Some new variety on the supply side increases the real income of workers and induces more consumers to come and more firms to set up. As a result, such a move may induce other industries to locate in the involved region, leading to a diversification of economic activities. Such a story shows not only that concentration and diversification do not necessarily go hand in hand but also that there is a marked need to disentangle the short-run effects of trade integration from the long-run effects. It also suggests that the desegregation level of economic activities used in empirical studies is of primary importance.

2.2 The empirical literature

Krugman (1993) provided the first evidence of the link between specialization and integration. Using a low level of industrial classification (SIC 2 level), he emphasizes that U.S. regions are more specialized than European ones and documents some increase of specialization in particular areas such as the state of Massachusetts. Subsequent studies were aimed mainly at collecting better statistical information, improving statistical procedures, and/or analyzing recent integration episodes. The European integration process (the Single Market) and the North American Free trade Agreement provide interesting experiments that were investigated extensively by researchers.

Empirical studies on the whole emphasize the weak tendency of European countries to specialize. This result, however, is far from being general. It varies greatly depending on the type of specialization indices (relative vs. absolute specialization), the data sources (trade vs. production), and the level of desegregation. A set of studies relies on production or employment data to characterize industrial specialization. Using 27 sectors that were observed for 10 countries from 1968 to 1990, Amiti (1998) found an increase in specialization after 1980. Extending the level of desegregation with 65 sectors for 5 countries, Amiti (1999) also documents an increase in specialization over the 1976–1989 period. However, using export data for 4 European countries and 100 sectors over the 1979–1992 period, Sapir (1996) shows that absolute specialization, captured by Herfindahl indices, appears quite constant over time, France being the exception.

A set of recent studies gathered in Harris (2001) investigated the impact of the changes in North American linkages (CUSFTA and NAFTA) on the structure of the Canadian economy. Acharya, Sharma, and Rao (2001) isolate moderate (but statistically insignificant) evidence of an increase in the intra-industry trade between the United States and Canada, in line with the findings of Trefler (2002). Using SIC 2 level data (20 industries), Sawchuk and Sydor (2001) document moderate evidence of diversification of Canadian manufacturing exports between 1990 and 1998. This is in contrast to an increasing specialization pattern observed for Mexico. This result is broadly consistent with the evidence provided by Head and Ries (2001b). Using SIC 4 level employment data over the 1983–2000 period (213 industries), they document some stability over time of the specialization index (Herfindahl). On the whole, these findings illustrate the weak effect of trade integration on the specialization patterns of Canadian production and on the nature of the bilateral trade flows with the United States.

Our paper aims at extending the previous empirical attempts on several fronts. First and most significantly, we intend to capture trade integration in a more accurate way. Most studies in the existing literature make implicit assumptions about the process of integration. They either assume that integration follows a linear process and therefore investigate the evolution of specialization over time (as done by numerous European analyses) or they follow a one-shot process by capturing trade integration by time dummy variables (Acharya et al. 2001, for instance). In contrast, our measure of trade integration rests on the dynamics of trade-weighted tariffs. These trade tariffs have been lowered progressively since 1980, suggesting that trade integration had taken place well before the CUSFTA. Second, the previous analysis ignores the high degree of heterogeneity across Canadian provinces in general and its core-periphery structure in particular. Canadian provinces face different dynamics in their trade-weighted tariffs because of their different product specialization—especially when the primary sector is taken into account. This in turn might have differing effects on their specialization patterns and hence on the core-periphery structure. To overcome this problem, we conduct the analysis at a regional level, linking provincial specialization indexes and provincial trade-weighted tariffs. On the methodological side, such an approach permits a more robust econometric analysis by pooling the cross-sectional and time-series information. Distinguishing the short- and long-run effects of trade integration is the final significant contribution of our analysis. The importance of this “disentangling” comes from the theoretical works that stress the dynamic effects of agglomeration (see below).

3. Data methodology

3.1 Capturing specialization patterns

The indexes of industrial structures used in this study were computed from export data at the industry SIC 4 level, available annually for the 10 Canadian provinces over the 1980–2001 period. The 1980–1989 raw data were obtained from Statistics Canada by Industry Canada. The 1990–2001 data are the raw data used to compute various trade indicators on the Strategis website of Industry Canada.[1] We used all SIC 4 codes (from SIC number

[1] Much information on definitions and methodology regarding trade data can be found on the Strategis website at http://strategis.ic.gc.ca/sc_mrkti/tdst/tdo/tdoDefinitions_30.php#industry_selection_search. The authors constructed the data bank used in this paper from the raw data with the valuable assistance of François Rimbaud from Industry Canada.

0111 to 3999) for which at least one entry for one province was non-null. The final data used involve 290 series per province. Data were computed for exports to United States (including U.S. territories) and for exports to the Rest of the World.

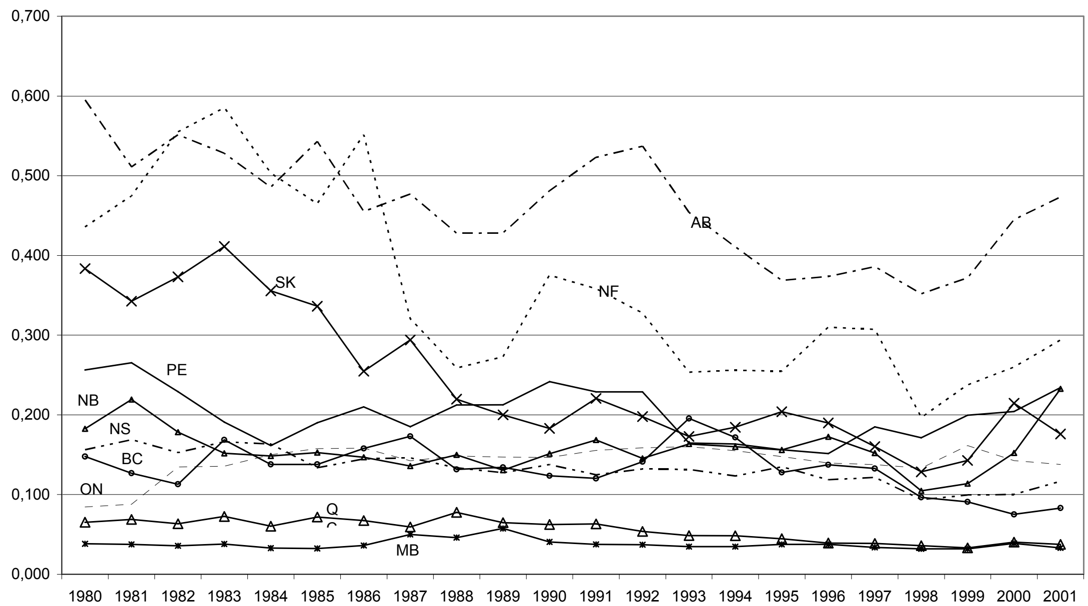
We use the Herfindahl index that captures the degree of absolute specialization and can be used for international comparisons. Its evolution might reveal to what extent a given province is becoming more specialized or diversified, regardless of how the economic structure for Canada as a whole is evolving. For each province *i* and each year *t*, the Herfindahl index (denoted by $S_{i,t}$) was computed as

$$S_{i,t} = \sum_{k=1}^{J} \left(s_{i,t}^{k} \right)^{2},$$

where $s_{i,t}^{k} = x_{i,t}^{k} \Big/ \sum_{k=1}^{J} x_{i,t}^{k}$. Here, J is the number of investigated industries (sectors), $s_{i,t}^{k}$ stands for the share of export $x_{i,t}^{k}$ of industry *k* in the total exports of province *i*. The Herfindahl is the sum of the squares of the shares over all industries. By definition, $S_{i,t}$ will therefore be between ($1/J$) and 1. The smaller (bigger) the number, the more diversified (specialized) is the industrial structure of the province involved. In our subsequent analysis, we make a clear distinction between the all-sectors case in which all industries including primary products are considered ($J = 290$) and the manufacturing case ($J = 213$).

The series for the Herfindahl index (all-sectors case) computed from export data to the United States are shown in figure 1 for the 10 provinces over the 1980–2001 period. The economies of Manitoba and Quebec are by far the most diversified with mean $S_{i,t}$ of .038 and .055 respectively. At the other end of the spectrum, the economies of Alberta and Newfoundland are the two most specialized with average Herfindahl indexes of .463 and .357 during the period. Not surprisingly, the extreme specialization of Alberta is attributed to the oil and natural gas that account for 67 percent of Alberta’s total exports to the United States during the period. For Newfoundland, the high degree of specialization was attributed to the fish products industry. The dramatic decline of the Atlantic fish stocks since the end of the 1980s, however, has contributed to the change in the evolution of Newfoundland’s specialization index. Between 1980 and 1986, fish products accounted on average for 68 percent of Newfoundland’s total exports. Since 1987, this industry’s share has fallen to an average of only 23 percent. Note that the upsurge of the Newfoundland Herfindahl index between 1998 and 2001 is significantly attributable to the Hibernia offshore oil field that began producing oil in November 1997. The increase in the specialization of Newfoundland's exports is likely to continue with the start of the exploitation of the Terra Nova platform in 2002. The diversification measure of Ontario, the largest provincial economy, is lowered by the substantial share of the motor vehicle industry that has accounted for about 38 percent of Ontario's total exports to the United States since 1990.

Figure 1: Herfindahl indexes for Canadian provinces



In order to provide further insight into the dynamics of industrial structures, we also use the well-known K-spec index (Krugman 1991a). In contrast to the Herfindahl index, the K-spec index is a measure of *relative* rather than absolute specialization. The evolution of this index might capture the degree of heterogeneity across Canadian provinces and its evolution along with the trade integration process. In our case, it captures the gap between the industrial structure of province *i* and the average of the industrial structure of the other nine provinces.

3.2 Capturing trade integration

A straightforward method of capturing (provincial) trade integration could be using provincial trade flows (exports and imports) between each province and the United States. As emphasized by Beine and Coulombe (2003), there is much heterogeneity across provinces in international trade patterns throughout the integration process. Nevertheless, using trade flows is likely to lead to significant endogeneity bias. Trade volumes in particular are likely to be affected by specialization patterns and by the nature of the bilateral trade with the United States (intra- vs. inter-industrial trade). Therefore, to determine trade integration, alternative measures that are obviously exogenous to specialization patterns and the share of intra-industrial trade are called for.

Following several authors including Trefler (2002) and Sawchuk and Trefler (2002), our integration measure will be based on the dynamics of tariffs between the United States and Canada. Our starting point is the per industry Canadian tariff against the United States (Trefler 2002; Sawchuk and Trefler 2002). Unlike those studies, however, our trade-weighted tariffs are also computed from all industries, including primary products that play a prominent role in several provinces such as Alberta and Saskatchewan. Another essential difference from the previous studies is that our tariffs are computed at the provincial level using provinces' specific weights based on their industrial structure rather than weights for Canada as a whole. Trade-weighted tariffs are computed for the all-sectors case (J = 290) as well as the manufacturing case (J = 213) and are defined as

$$TW_{i,t} = \sum_{k=1}^{J} w_{i,t}^{k} \tau_{t}^{k}$$

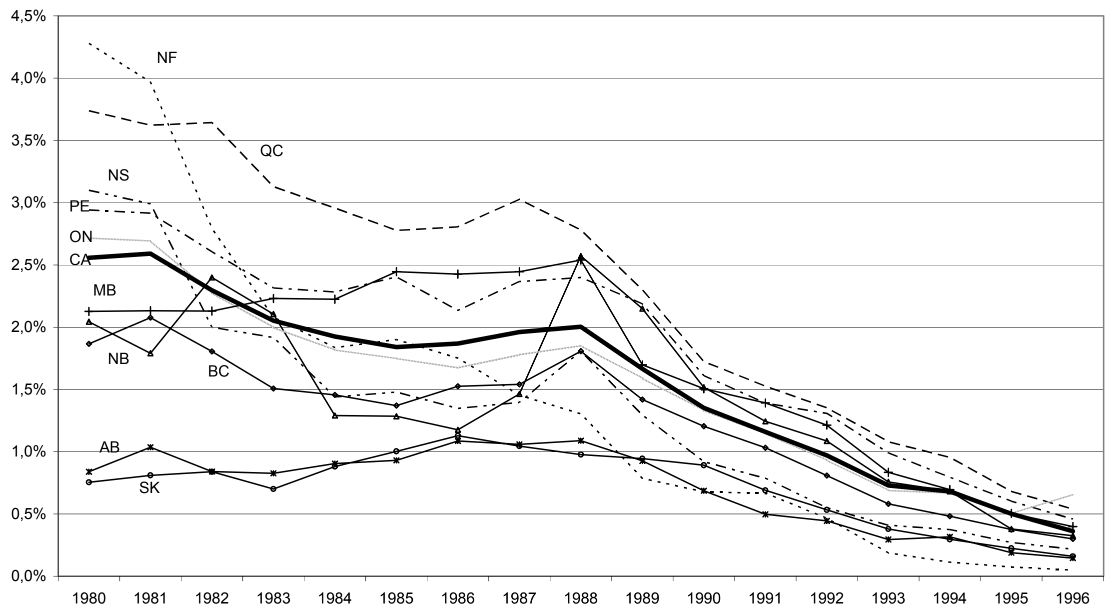
where $TW_{i,t}$ denotes the level of the trade-weighted tariff of province *i* at time *t*, $w_{i,t}^{k}$ is the weight of industry *k* in the total export of province *i* to the United States at time *t,* and τ_{t}^{k} is the tariff relative to industry *k* between Canada and the United States.[2]

Figure 2 reports the evolution over time of the trade-weighted tariffs (computed for all sectors) for Canada as a whole and for each of the provinces. It illustrates the decreasing pattern of the tariffs since 1980. While the CUSFTA induced a quicker rate of decay for the $TW_{i,t}$, figure 2 also shows that trade integration between Canada and the United States had begun well before the implementation of the CUSFTA in 1989. Figure 2 also illustrates an important point that is relevant to our analysis at the regional level: the significant degree of heterogeneity of the integration process across provinces. This heterogeneity comes from the different patterns of decrease in the bilateral tariffs across industries, combined with the strong differences in the (initial) provincial industrial structures. Because of this, differences across provinces are observed for the $TW_{i,t}$ in both

[2] Trefler (2002) uses the tariffs between Canada and the United States for the manufacturing sector over the 1980–1996 period. Those data are the collected duties paid by U.S. exporters to Canada. We complement Trefler's data with the NAFTA tariffs on primary products, provided by S. Rao from Industry Canada. For the 1997–1998 period, we assume that the tariffs on manufacturing goods remained at their 1996 level. As shown by our results relative to the manufacturing sector, this assumption does not affect the main results of this paper.

their levels and their speeds of decay. To illustrate, Newfoundland and Labrador experiences the sharpest decrease in $TW_{i,t}$, with the highest level in 1980 and the lowest one in 1998. In contrast, Alberta and Saskatchewan experienced quite constant rates until 1990 and a modest decrease after the implementation of the CUSFTA. The high degree of heterogeneity of the trade integration process across provinces allows us to combine a purely time-series piece of information with a cross-sectional one. This in turn allows us to overcome the statistical problem of a low number of data points at hand in the pure time-series approach of the CUSFTA.

Figure 2: Provincial trade-weighted tariffs, all sectors



Finally, it is important to bear in mind that, as pointed out by Trefler (2002), the decrease of the $TW_{i,t}$ captures much more than a pure tariff reduction. Basically, it emerges that $\tau_{j,t}$ are well correlated across industries with other non-tariff barriers (NTBs).[3] In this sense, it captures a broader set of CUSFTA trade-liberalizing policies.

[3] Not surprisingly, the data on these NTBs are for the most part unavailable (see Anderson and van Wincoop [2004] for an extensive survey). Even if some data were available, our analysis would entail tariff-equivalent estimates of these NTBs.

3.3 Other variables

As emphasized by Kalemli-Ozcan, Sorensen, and Yosha (2001), the potential determinants of industrial specialization are primarily those affecting the volume of Canada–U.S. bilateral trade. In this paper, we control for two different sets of variables.

The first set of control variables is related to macroeconomic fluctuations. We introduce a measure of the U.S. cycle, as well as a measure of the cycle of each province. In each case, the cyclical part of the GDP is obtained from a traditional HP filtering on quarterly data.[4] The annual measure of each cycle is then obtained from the average of the quarterly cycles.

The second set of variables is related to the bilateral exchange rate conditions between the U.S. and Canadian dollars. We introduce two determinants that have been found to exert detrimental effects on the volume of trade. The first determinant is exchange rate volatility that we compute using the measure of integrated volatility proposed by Andersen et al. (2001). The idea of such a measure is to make use of the information available at higher frequencies. Here we build measures of exchange rate volatility at an annual frequency from squared daily returns.[5]

The second determinant is the degree of exchange rate misalignment between the Canadian and U.S. dollars. It has been argued that monetary policy in general plus the overvaluation of the Canadian dollar in the early 1990s had detrimental effects on the macroeconomic dynamics of the Canadian economy (Fortin 1996). We compute misalignment as the annual absolute deviation of the bilateral exchange rate from its equilibrium value. The equilibrium is assumed to be the purchasing power parity level computed with the ratio of producer prices between Canada and the United States.[6]

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[4] The data were taken from Beine and Coulombe's (2003) measures of business cycles. The interested reader should refer to this paper for details on the cycle extraction procedures and the GDP data sources.

[5] The exchange rate data of the CAD–USD at the daily frequency were kindly provided by the Bank of Canada. As shown by Andersen et al. (2001), this realized volatility provides an unbiased and more efficient estimate than the usual measures such as historical volatility. It is an unbiased measure of volatility when exchange rate returns are uncorrelated over time, which is obviously the case with daily frequency.

[6] The producer price index is extracted from the Main Economic Indicators database of the OECD.

4. Econometric analysis

4.1 Empirical methodology of the panel data model

Structure of panel data regressions

One of the basic reasons for using a dynamic panel data model is to make the best use of the available information by combining cross-sectional (across the 10 Canadian provinces) and time-series data (over the 1980–1998 period). As was pointed out earlier, the panel data approach takes into consideration the fact that some Canadian provinces, due to their respective industrial structures, were more affected by the change in tariffs. Taking this into account multiplies the information compared with a pure time-series investigation based on overall Canadian data. Furthermore it is known that panel data, because of their relatively stable cross-sectional distribution, are better able than time-series data to study the dynamics of adjustment and to isolate the long-run effects of exogenous variables.

The general structure of the panel data regressions performed in this paper is

$$\Delta S_{i,t} = F(S_{i,t-1}, TW_{i,t-1}, \Delta TW_{i,t}, Z_{i,t}, \varepsilon_{i,t}).$$

Here, $i = 1, \ldots, 10$, stands for the Canadian provinces, and $t = 1980$ to 1998 (to 1996 only for the manufacturing sector). In this set-up, $S_{i,t}$ refers to our measure of industrial structure of province i at time t. $TW_{i,t}$ is the trade-weighted tariff of province i at time t. The $Z_{i,t}$ are the other variables that might account for the changes in the industrial structure S during the period under study. Most of the Z are aggregate variables and are constant across the i. The disturbance $\varepsilon_{i,t}$ is modelled as

$$\varepsilon_{i,t} = c_i + v_{i,t},$$

where $v_{i,t}$ are the idiosyncratic errors and the c_i, the unobserved individual components. These are discussed in the following section.

Estimation techniques

Many alternative estimation techniques are available for pooling time-series and cross-sectional information in a dynamic panel data model. The first point to tackle is the issue of the unobserved components. One reason for using panel data here instead of aggregate data for the overall Canadian economy is that it is possible to account for the extensive heterogeneity of Canadian provinces' industrial structure using time-invariant fixed effects. Provinces' fixed effects proved to be very significant and were introduced in all regressions. Therefore, our results come from least-squares dummy-variable (LSDV) estimators. For example, in the pooled least-squares

(PLS) estimation of the benchmark equation (1), all provinces' fixed effects are significant, with associated *p*-values well below the 1 percent level. Conversely, in the same set-up, no time dummies were significant at the 5 percent level. The null hypothesis is not rejected in tests for joint significance of time dummies with *p*-values of .22 and .83 in the all-sectors and the manufacturing sector respectively. Consequently, time dummies were not included in the empirical analysis.

The second important issue is related to the econometric techniques used to tackle various (cross-sectional and time-series) heteroscedasticity problems underlying this type of panel analysis. For the first set of results, using PLS, we rely on white heteroscedasticity consistent standard errors (HCCME) that allow for asymptotically valid inferences in the presence of general heteroscedasticity. The second set of results comes from feasible generalized least-squares (FGLS) estimations using cross-sectional weighted regressions to account for cross-sectional heteroscedasticity. With this technique, we also report consistent standard errors (HCCME) that are robust to heteroscedasticity. With FGLS, we used iterative techniques for updating coefficients and the weighting matrix until convergence. In all cases, convergence was achieved with a number of iterations around 10.[7]

Specific regression set-ups

Benchmark specification: The starting point for our specification procedure is based on the family of dynamic econometric models used in the empirical macroeconomic modelling. In order to disentangle the long- and short-run dynamics, our benchmark specification relies on a general equilibrium correction specification:

$$\begin{aligned} &\Delta \log(S_{i,t}) = \phi_1 \log(S_{i,t-1}) + \phi_2 \log(1+TW_{i,t-1}) + \phi_3 \Delta \log(1+TW_{i,t}) + \phi' Z' + \varepsilon_{i,t} \\ &\text{where } \phi' Z' = \phi_4 UC_t + \phi_5 PC_{i,t} + \phi_6 VL_t + \phi_7 MA_t \end{aligned} \tag{1}$$

[7] Refer to Beine and Coulombe (2004) for results from seemingly unrelated regression (SUR) estimations. SUR, also known as the Parks estimation, is the least restricted estimation technique here since it corresponds to the FGLS estimator for which the residuals are both cross-sectional heteroscedastic and contemporaneously correlated. In many cases, however, SUR generated extremely high t-statistics and fortunately, point estimates that have the same sign and order of magnitude as with FGLS. It is known that Parks estimations are able to produce standard errors that lead to extreme confidence when the number of time series is not that much larger than the number of cross-sections (Beck and Katz 1995).

The variable UC_t captures the U.S. business cycle at time *t* and the $PC_{i,t}$ are measures of Canadian provinces' *i* business cycles. VL_t is the bilateral exchange rate volatility against the U.S. dollar and MA_t is the degree of misalignment between the Canadian and the U.S. dollars. We also report results for which the $\phi' Z'$ factors were not entered in the regression.

The interesting feature of the dynamic model (1) is that it separates the short- and the long-run effects of economic integration. Assuming stationarity of $S_{i,t}$ and $TW_{i,t}$, model (1) can be estimated directly by PLS or FGLS. The long-run effect of economic integration that is associated with $TW_{i,t}$ is then given by $(-\hat{\phi}_2/\hat{\phi}_1)$ while the short-run effect is captured by $\hat{\phi}_3$.

Pooled mean group specification: In the spirit of Pesaran, Shin, and Smith (1999), the second regression set-up refers to pooled mean group estimation (PMG) of dynamic heterogeneous panel. The speed of adjustment $\phi_{1,i}$ and the short-run effect of tariff changes $\phi_{3,i}$ are allowed to differ across provinces whereas the long-run effect of tariffs ϕ_2 is constrained to be identical:

$$\Delta \log(S_{i,t}) = \phi_{1,i} \log(S_{i,t-1}) + \phi_2 \log(1+TW_{i,t-1}) + \phi_{3,i} \Delta \log(1+TW_{i,t}) + \varepsilon_{i,t} \qquad (2)$$

The $\phi' Z'$ factors were not entered in this specification.

Potential problems of the empirical methodology

The empirical methodology used in this paper is subject to two potential problems. First, it is well known (Nickell 1981) that least-squares dummy-variable (LSDV) estimators of the ϕ_1 and the other ϕs are biased when the number of time periods *T* is small and the number of cross-section *N* is large, as in traditional microeconometric studies. Kiviet (1995) shows that LSDV estimators are consistent only when $T \to \infty$. The usual solution to this problem was to use instrumental variable (IV) methods and more recently, Generalized Method of Moments (GMM) estimators. However, as shown in Ahn and Schmidt (1995), GMM estimations are not very efficient and LSDV estimators have a much smaller variance (Kiviet 1995). In our analysis, we rely on LSDV estimators for two reasons. First, the number of periods *T* (18) is relatively large. Kiviet (1995), for example, concentrates on the bias of LSDV in the cases when *T* = 3 and 6. Second, as shown in Kiviet (1995), the bias of LSDV is reduced when the speed of adjustment is faster, as it is the case in our analysis.

Second, even when the number of time periods is relatively large, LSDV estimations of dynamic panel reveal another danger. As shown in Pesaran and Smith (1995), ignoring parameter heterogeneity might produce a substantial bias since the regressors of homogenous pooled estimations will be serially correlated. In this paper, this potential problem is tackled by the use of the heterogeneous specifications (2) whose structure was suggested by Pesaran et al. (1999). These results are reported in the following section. From both the qualitative and the quantitative points of view, our main results regarding long-run effects are robust to alternative heterogeneous specifications. This suggests that the problem raised by Pesaran and Smith (1995) does not alter the substance of the main findings of our analysis.

4.2 Results: the dynamics of industrial diversification

Long-run effect of tariff changes

The general direction of our results regarding the long-run effect of economic integration on provincial industrial diversification emerges from specifications (1) and (2) in tables 1 to 3. In table 1, the primary sector is included while in table 2, the analysis is restricted to industries of the manufacturing sector, as was done in Trefler (2002). The results from LSDV (benchmark) and PMG for the all-sectors case are compared in table 3.

In all cases, the various parameter estimates of the $TW_{i,t-1}$ variable (ϕ_2) are positive and significantly different from 0 at the 1 percent level. Furthermore, in all panel specifications reported in tables 1 to 3, the various estimates of the adjustment speed to equilibrium ϕ_1 are always negative and significant at the 1 percent level. Therefore, the measures of the long-run elasticity given by $\left(-\hat{\phi}_2/\hat{\phi}_1\right)$ of tariff changes on trade specialization are always positive and significant with p-values below the 1 percent level. The various measures of the long-run elasticity computed from the point estimates are quite comparable across estimation techniques and specifications. They vary between 15 and 19 for the all-sectors (table 1) and between 12 and 13 for the manufacturing sectors (table 2). They are also systematically slightly higher (lower) for PLS (FGLS) estimations for the same regression set-up.

Table 1. Dynamics of industrial diversification – Herfindahl (all sectors)

Dependent variable: $\Delta\log(S_{i,t})$				
	PLS	FGLS	PLS	FGLS
$\log(S_{i,t-1})$	-0.342***	-0.340***	-0.328***	-0.324***
	(0.070)	(0.075)	(0.071)	(0.080)
$\log(1+TW_{i,t-1})$	5.384***	5.065***	6.320***	5.865***
	(1.443)	(1.579)	(1.449)	(1.593)
$\Delta\log(1+TW_{i,t})$	-5.268	-6.116	-6.975	-8.350*
	(4.317)	(4.345)	(4.331)	(4.522)
UC_t	-1.732	-1.988	–	–
	(1.495)	(1.341)		
$PC_{i,t}$	1.325	1.280	–	–
	(1.599)	(1.479)		
VL_t	-26.690**	-15.700	–	–
	(11.588)	(9.663)		
MA_t	-0.261	-0.262*	–	–
	(0.171)	(0.141)		
R^2	.279	.314	.233	.289
Adj. R^2	.208	.246	.178	.238
S.E. of R.	.129	.130	.132	.132
observations	180	180	180	180

Notes: Sample: 1980–1998. All regressions include provincial fixed effects. Estimation techniques: pooled least-squares (PLS), iterated feasible generalized least-squares (FGLS). White heteroscedasticity standard errors are shown in parentheses below the estimated coefficients. * = significant at 10% level; ** = significant at 5% level; *** = significant at 1% level.

Table 2. Dynamics of industrial diversification – Herfindahl (manufacturing sector)

Dependent variable: $\Delta\log(S_{i,t})$				
	PLS	FGLS	PLS	FGLS
$\log(S_{i,t-1})$	-0.401***	-0.376***	-0.399***	-0.383***
	(0.069)	(0.083)	(0.069)	(0.085)
$\log(1+TW_{i,t-1})$	5.323***	4.677***	5.305***	4.872***
	(1.352)	(1.403)	(1.344)	(1.409)
$\Delta\log(1+TW_{i,t})$	2.600	3.680	2.077	3.340
	(4.157)	(3.885)	(3.489)	(3.090)
UC_t	-1.220	-0.949	-	-
	(1.564)	(1.548)		
$PC_{i,t}$	-0.001	-1.190	–	–
	(1.886)	(1.903)		
VL_t	3.441	5.526	–	–
	(14.283)	(12.745)		
MA_t	0.138	0.146	–	–
	(0.232)	(0.216)		
R^2	.251	.280	.245	.264
Adj. R^2	.167	.198	.183	.204
S.E. of R.	.134	.135	.133	.133
observations	160	160	160	160

Notes: Sample: 1980–1996. See also notes for table 1.

As illustrated in figure 2, the CUSFTA has accelerated the negative trend in tariffs since 1989. As a consequence, our main results imply that the mean (across provinces) effect of trade integration induced by the CUSFTA was a significant and substantial increase of the long-run industrial diversification of Canadian provinces. This result is extremely robust and remains invariant across estimation techniques (PLS, FGLS, and PMG) and specifications.

Dynamic adjustment

As was just pointed out, the point estimate of the ϕ_1 parameter in the dynamic panel data model is important as it is a determinant of the long-run effect. This parameter captures the speed of adjustment toward long-run equilibrium when this equilibrium relationship has been affected by shocks such as further tariff reductions induced by the CUSFTA.

The point estimates of the ϕ_1 parameter in the all-sectors and manufacturing cases (tables 1 and 2) are quite stable across the various specifications and estimation techniques; they vary between –0.32 and –0.40. In the case of the benchmark model for the all-sectors index (table 1), the estimated value of ϕ_1 amounts on average to -0.33. This implies that following a shock on the long-run equilibrium of the degree of industrial specialization, half of the gap between the initial situation and the new long-run level is adjusted after just 2.1 years, approximately. The speed of adjustment is even faster in the benchmark specification that is related solely to the manufacturing industries (table 2) since the associated mean half-life of adjustment for the four point estimates amounts to only 1.8 years. The explanation for this is that the adjustment process in the primary sector is slowed because primary production is dependent on the localization of natural resources.

Effect of aggregate variables (Z')

Four other variables, thought to provide some explanation for short-run variations in industrial diversification, were introduced as independent variables in the benchmark specification. Those variables are measures of the U.S. and Canadian provincial business cycles, the bilateral exchange rate volatility against the U.S. dollar, and the degree of misalignment between the Canadian and the U.S. dollars.

As shown by the analysis of tables 1 and 2, their effect is somewhat limited and not robust. Having said that, there is still some evidence to suggest that increases in exchange rate variability or in misalignment have a positive effect on industrial diversification. The effects, however, appear to be limited to regressions including the primary sectors. In contrast, there does not seem to be a significant role for exchange rate variables in diversification patterns of the manufacturing industries. Such a finding is consistent with the view that exchange rate fluctuations and monetary arrangements can exert asymmetric effects on the industrial production structures of the provincial economies (Beine and Coulombe 2003).

It might be concluded that most of the dynamic evolution of industrial diversification has been driven by the dynamics of trade-weighted tariffs combined with provinces' specific fixed effects. The results for the Z' variables in tables 1 to 2 also illustrate the robustness of our main results regarding the long-run effect of tariff changes. The results suggest that introducing other independent variables is unlikely to alter the general direction of the main results.

Pooled mean group specification

In the PMG specification, short-run effects and the speeds of adjustment are allowed to vary across provinces. In table 3, we report results where LSDV estimates of the benchmark model (without the Z' variables that correspond to column 4 in table 1) are compared with PMG estimation in the all-sectors case where FGLS is used as the common estimation technique. In the PMG case, we report cross-sectional means and medians for the $\phi_{2,i}$ and $\phi_{3,i}$ point estimates as well as the asymptotic standard error of the mean.

Both quantitatively and qualitatively, the results of the PMG estimation align very well with the general direction of our main results. The long-run elasticity of tariff changes estimated from both the PMG and LSDV specifications is almost identical and the mean long-run effect from PMG is highly significant. This analysis clearly indicates that the main results of this paper regarding the long-run effect of tariff changes on the industrial structure appear very robust to the potential problem raised by Pesaran and Smith (1995).

Table 3. Dynamics of industrial diversification – Herfindahl (all-sectors), heterogeneous model: LSDV versus PMG estimators

Dependent variable: $\Delta\log(S_{i,t})$		
	LSDV	PMG
$\Delta\log(S_{i,t-1})$	-0.324*** (0.080)	-0.409*** (0.073)†
Median $\log(S_{i,t-1})$	–	-0.461
$\log(1+TW_{i,t-1})$	5.865*** (1.593)	7.727*** (1.521)
$\Delta\log(1+TW_{i,t})$	-8.350* (4.522)	-20.469** (6.981)†
Median $\log(1+TW_{i,t-1})$	–	-19.894
Long-run effect of $\log(1+TW_{i,t-1})$	18.101***	18.892***

Notes: Sample: 1980–1998, FGLS estimations. See also notes for table 1.
†Asymptotic standard error reported for Mean (PMG estimator).

Short-run effect of tariff changes

Interestingly, the point estimate (mean) of the short-run effect of tariff changes (ϕ_3) is significant at the 5 percent level with PMG estimations. Furthermore, the point estimate of the mean is not biased by some outliers since its magnitude is very close to the median. This result concurs with the various point estimates of the short-run effect of tariff changes on industrial diversification in the all-sectors case of table 1, which are always negative. This (reverse) short-run effect, however, is less robust than the long-run positive effect since it is significant at the 10 percent level only in the restricted FGLS set-up from specification (1), and in the manufacturing sector, the short-run effect is always positive but never significant.

Results from PMG bring some evidence, however, in favour of a negative short-run effect of tariff changes on industrial diversification in the all-sectors case. In the short run, economic integration might stimulate industrial specialization at the Canadian regional level but the effect is clearly reversed in the long run. These findings are consistent with the theoretical scenario proposed by some of the models of economic geography that were presented in Section 2.1. This once more emphasizes the need to separate the short- from the long-run effects of trade interaction.

The fast estimated speed of adjustment to the new equilibrium implied by the estimates of ϕ_1 suggests however, that the potential contemporaneous negative effect of tariff cuts on regional industrial diversification is quickly overcome. For example, the negative effect of a one-time decrease in tariffs occurring during year *t* would be totally offset by the end of year t+1 or t+2. After that, the economy converges rapidly to a new equilibrium level that is characterized by a higher degree of industrial diversification.

Further analyses

The effect of tariff changes on industrial diversification is captured from Herfindahl indexes computed from the provinces' exports to the United States. One could argue that in the long run, the increased diversification might be the result of some sort of export substitution at the industry level—from the Rest of the World to the United States after the arrival of the CUSFTA. Indeed, as emphasized by Trefler (2002), the CUSFTA mandated tariff concessions to the United States that were preferential compared with the Rest of the World. We refer to this as the "specialization substitution hypothesis." This hypothesis was directly tested using the benchmark specification for which the dependent variable is based on the Herfindahl index computed from provinces' exports to the Rest of the World. Non-reported results of the panel data regressions indicate that the specialization substitution hypothesis is clearly rejected. No point estimates of the long-run effect were significant and the p-values were extremely high, especially with FGLS. There is, however, some (non-robust) evidence that the specialization substitution hypothesis might hold in the short run.

We also perform tests with the benchmark specification using the relative specialization structure of the Canadian provinces from a K-spec indicator as a measure of the industrial structure $S_{i,t}$. As was well documented in Beine and Coulombe (2003), among others, the Canadian economy is best characterized by a core-periphery structure. Strikingly, Krugman (1991a) used the Canadian case as the typical example of this type of structure. Most manufacturing industries are located in the Quebec City–Windsor corridor that dominates the export data for Quebec and Ontario, the two most important regional economies in Canada. One of the interests in analyzing the incidence of tariff changes at the regional level in Canada is to investigate whether the CUSFTA has induced a change in the core-periphery structure of the Canadian economy. Non-reported results clearly indicate that the evolution of the K-spec tariff has not been affected by tariff changes. The analysis therefore suggests that the reduction in tariffs fostered by the CUSFTA has not affected the relative degree of industrial diversification of the Canadian provinces. This result is important in the Canadian context, given the decentralized nature of the Canadian federation and the extensive reliance on an interprovincial redistribution scheme operated through a sophisticated federal tax transfer system aimed at smoothing out the consequences of specific regional disturbances.

We also investigate the interaction channels between absolute (Herfindahl) and relative (K-spec) diversification using interaction specifications where the alternative measures of industrial structure were added as additional regressors in the benchmark specification of the other diversification index. Non-reported results indicate that only one long-run interaction channel emerges as significant in a robust way: the Herfindahl index has a significant (at the 1 percent level) and positive long-run effect with both PLS and FGLS on the K-spec indicator. Interestingly, this result indicates that an increase in industrial specialization at the provincial level increases the asymmetry of the industrial structure of Canadian regional economies. In other words, at the Canadian level at least, absolute specialization goes hand-in-hand with relative specialization. In this sense, trade integration exerts an indirect effect (through absolute diversification effects) on relative industrial production. Interestingly, the parameter of the long-run effect is not significant in the reversed interaction set-up where the Herfindahl is the dependent variable. These results confirm to some extent that causality runs from specialization structure to the core-periphery structure, not the reverse.

4.3 Robustness

Table 4. Dynamics of industrial diversification, robustness analysis – all-sectors and manufacturing sectors

Dependent variable: $\Delta \log(S_{i,t})$						
	All sectors			Manufacturing sectors		
	Constant weights	AR(1) PLS	AR(1) FGLS	Constant weights	AR(1) PLS	AR(1) FGLS
$\log(S_{i,t-1})$	-0.346***	-0.366***	-0.359***	-0.337***	-0.451***	-0.408***
	(0.074)	(0.094)	(0.073)	(0.078)	(0.126)	(0.071)
$\log(1+TW_{i,t-1})$	4.131***	5.396***	5.089***	2.785***	5.484***	4.099***
	(1.279)	(1.739)	(1.244)	(1.048)	(1.775)	(0.979)
$\Delta \log(1+TW_{i,t})$	6.715	-4.021	-4.830	4.548	2.539	-4.083
	(7.297)	(4.406)	(3.803)	(4.832)	(4.594)	(3.933)
$UC_{i,t}$	-2.209	-2.273	-1.979	-0.772	-1.818	-0.498
	(1.347)	(1.615)	(1.273)	(1.476)	(1.605)	(1.346)
$PC_{i,t}$	1.261	2.144	1.491	-1.234	1.012	-1.399
	(1.519)	(1.665)	(1.498)	(1.897)	(2.018)	(1.526)
VL_t	-14.260	-27.812**	-12.884	5.794	1.673	6.347
	(10.291)	(11.670)	(8.961)	(12.349)	(14.347)	(9.912)
MA_t	-0.415**	-0.282	-0.213*	0.076	0.105	0.239
	(0.160)	(0.175)	(0.121)	(0.207)	(0.233)	(0.141)
AR1	-	-0.039	-0.187*	-	-0.009	-0.188*
		(0.130)	(0.111)		(0.151)	(0.111)
R^2	.285	.304	.427	.269	.279	.249
Adj. R^2	.215	.227	.363	.188	.186	.153
S.E. of R.	.133	.129	.132	.136	.134	.137
observations	180	170	170	180	150	150

Notes: Samples: 1980–1998, all sectors; 1980–1996, manufacturing sector. See also notes for table 1. Constant weights means FGLS regression involving tariffs computed with constant trade weights equal to the average over 1980–1998. AR1 means regression with autoregressive term of order one included in the error term.

Constant weight tariffs

It is possible to argue that our baseline measure of trade-weighted tariffs is partly endogenous since the evolution of provinces' weights might be affected by the (endogenous) evolution of the industrial structure. To address this particular point, we build an alternative tariff variable using constant weights based on the mean structure of exports over the 1980–2001 period. The results reported in columns 1 and 4 of table 4 indicate clearly that the main orientation of our results regarding the long-run effect of economic integration is not affected by the indirect effect of the change in the industrial structure on provinces' weights. Estimates of the ϕ_2 parameter in the all-sectors and the manufacturing cases using this new measure are very comparable with the ones found with our flexible weight $TW_{i,t}$ variable and are highly significant. The measures of the long-run elasticity of tariff changes is, however, slightly lower in the constant weight case.

Serial correlation

The possibility that residuals might be serially correlated could lead to serious problems of estimator inconsistency in dynamic panels with fixed effects. Dealing with this issue, we have re-estimated the benchmark specification under the assumption that the disturbance follows either an AR(1) process or an AR(2) process using non-linear least squares. Results from PLS and FGLS estimations in the all-sectors and the manufacturing sector are reported in columns 2, 3, 5, and 6 of table 4 for the AR(1) specification. Results for the AR(2) specification go along the same direction and are not reported.

In the two specifications with PLS, the coefficient of the AR(1) is not significant at the 10 percent level. With FGLS, however, the coefficient is negative and significant at the 10 percent level. Accounting for serial correlation through an extended specification does not change the general direction of the results: both the magnitude and the significance level of the key parameters remain quite stable. This suggests that our benchmark model succeeds in capturing a significant part of the dynamics of the disturbances.

Reduced cross-sectional sample

To further assess the sensitivity of our results, we estimated the benchmark equation, removing a different province each time from the sample. Such an analysis checks whether the results are driven by the dynamics of a particular province. Such an effect might occur due to the relatively low cross-sectional dimension (10) of the panel. Results (non-reported) indicate that the estimates of the speed of adjustment are remarkably stable and are always significant at the 1 percent level. The measures of the long-run effect of trade-weighted tariffs are also stable and significant at the 1 percent level in nine cases. In the other cases, the *p*-value is very close to 1 percent.

Interaction with the Rest of the World

We also test the robustness by using interaction specifications. Here, the K-spec or the Herfindahl measured from exports to the Rest of the World were added to the benchmark specification in the list of regressors to capture part of the dynamics of industrial diversification. Given the share of Canada's total exports going to the United States, it is not surprising that the effect of the Herfindahl, measured for the Rest of the World, on the dynamics of the Herfindahl measures for exports to the United States is negligible.

5. Accounting for non-stationarity

The previous econometric analyses, of either the benchmark model or the alternative frameworks, implicitly assume stationarity properties of the data. For many of the variables involved in our analysis—such as economic cycles, exchange rate volatility, or misalignment—this assumption obviously holds and does not require any specific test. In contrast, there is no straightforward answer to the question whether indexes of industrial specialization (the Herfindahl indexes for all-goods and for the manufacturing sectors) and the trade-weighted tariffs exhibit some stochastic trends. The issue is made even more difficult since we are dealing with data that have been observed not only over time but also for various regions. Fortunately, econometric literature dealing with the issue of non-stationarity in panel data has emerged since the beginning of the 1990s. In this section, we employed some of the best-known econometric tools in this literature to assess whether the previous results are robust to the question of non-stationarity.

Accounting for non-stationarity in this analysis is important for at least two major reasons. The first is related to the well-known problem of spurious regressions. If the data exhibit some unit root, that is, follow a non-stationary process in level, then estimating ϕ_1 and ϕ_2 can have some problems. As explained by Pedroni (2003), ignoring variables' unit root properties will still produce problems in the inference procedure even with homogeneous panels for which the slope estimates are found to be consistent for the number of time periods and cross-sections. The second related reason is that if both series follow I(1) processes, their long-run relationships are best captured in a cointegration framework. Specific cointegration tests allow one to investigate whether such a long-run relationship holds in the case of non-stationary processes. In turn, accounting for the presence of such relationships allows the dynamic benchmark model (1) to be re-written in the form of an error correction model.

When dealing with potentially I(1) variables, the first step is to test for the presence of unit roots. Since the seminal work of Levin and Lin (1993), several tests have been developed to perform such a task in panel data. In this regard, we use three different test statistics, namely the so-called Levin-Lin rho statistics, the Levin-Lin t-rho statistics, and the Levin-Lin ADF statistics.[8] Given the fact that the specialization measures and trade-weighted variables are between 0 and 1, the analysis is also conducted on the logistic transformation of these variables. The results of these tests are not reported here due to space constraints but all details can be found in Beine and Coulombe (2004).

In general, referring to the homogeneous case, we found mixed evidence of non-stationarity in the two Herfindahl indexes.[9] In contrast, the trade-weighted tariffs appear to exhibit unit root properties for both the all-goods and the manufacturing sectors. It is significant that in no case can we conclude in favor of stationary processes for any variable with a high degree of certainty. This in turn calls for some cointegration analysis.

Panel cointegration tests permit proper testing for the significance of the coefficient linking variables $S_{i,t}$ and $TW_{i,t}$ when these variables are best described by an I(1) process. In this paper, we use the testing procedures proposed by Pedroni (1999, 2003). Pedroni (1999) develops seven different test statistics computed from auxiliary regressions involving the residuals $\hat{\eta}_{i,t}$ of $\log(S_{i,t}) = \gamma_i + \beta \log(TW_{i,t}) + \eta_{i,t}$ which might be called the static cointegrating regression. Under the null of a cointegrating relationship, the long-run elasticities of tariffs β might be estimated directly using PLS or FGLS techniques, allowing for the presence of individual fixed effects α_i to rule out any heterogeneity bias. The results of Pedroni's (2003) seven cointegrating tests provide some strong evidence in favor of a cointegrating relationship between industrial diversification and trade-weighted tariffs. Depending on the specification regarding the inclusion of the lagged variables (see Beine and Coulombe [2004] for further details), five or six of the seven tests suggest that a cointegrating relationship holds.

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[8] The interested reader may find more details in Levin and Lin (1993) and Pedroni (2003).

[9] In Beine and Coulombe (2004), two different statistical processes are allowed for. The first one involves coefficients (autoregressive parameters and long-run elasticities) that are common across provinces and refers to the homogeneous case. The second process involves coefficients that differ across provinces and refers to the heterogeneous case. For the sake of brevity, we investigate here only the homogeneous model that is the direct counterpart of the benchmark model discussed in Section 4.

Under the hypothesis of a cointegrating relationship between industrial diversification and trade-weighted tariffs, the dynamics are best captured by the so-called error correction model (ECM)

$$\Delta \log(S_{i,t}) = \phi_3 \Delta \log(1+TW_{i,t}) + \phi_4 UC_t + \phi_5 PC_{i,t} + \phi_6 VL_t + \phi_7 MA_t + \delta_i + \lambda(ECM_{i,t-1}) + \varepsilon_{i,t} \quad (3)$$

in which $-\lambda$ will capture the adjustment speed toward the long-run equilibrium between the industrial structure *S* and tariffs *TW*. In the homogeneous case that corresponds to our benchmark model (1), $ECM_{i,t} = \log(S_{i,t}) - \alpha_i - \beta \ \log(1+TW_{i,t})$ captures deviations from this long-run equilibrium. Interestingly, model (6) turns out to be a re-parameterization of the benchmark model (1), under the assumption of I(1) variables. Indeed, it can be seen that $\lambda = \phi_1$ and $\phi_2 = -\lambda\beta$ and $c_i = \delta_i - \lambda\alpha_i$.

The long-run estimates are mostly in line with the ones implied by the benchmark model, although slightly lower. Depending on the specification, the estimates of $(-\hat{\phi}_2 / \hat{\phi}_1)$ range between 14.897 and 19.268 for the all-sectors case and between 12.472 and 13.274 for the manufacturing case.

Table 5. Error correction models – Herfindahl, homogeneous case: all-sectors and manufacturing sectors

Dependent variable : $\Delta \log(S_{i,t})$				
	All sectors		Manufacturing sectors	
$ECM_{i,t-1}$	-0.349***	-0.341***	-0.381***	-0.387***
	(0.076)	(0.083)	(0.084)	(0.086)
$\Delta TW_{i,t}$	-6.542	-9.912**	3.072	2.748
	(4.323)	(4.55)	(3.852)	(3.050)
VL_t	-16.242*	-	5.219	-
	(9.572)		(12.654)	
UC_t	-2.014	-	-0.933	-
	(1.340)		(1.544)	
$PC_{i,t}$	1.289	-	-1.329	-
	(1.493)		(1.891)	
MA_t	-0.295**	-	0.142	-
	(0.131)		(0.212)	
R^2	.313	.284	.278	.260
Adj. R^2	.250	.237	.203	.205
S.E. of R.	0.130	0.132	0.135	0.133
observations	180	180	160	160

Notes: Samples: 1980–1998, all sectors; 1980-1996, manufacturing sector. See also notes for table 1. $ECM_{i,t}$ is the error correction term at time *t* for region *i* obtained with long-run elasticities of trade-weighted tariffs equal to 12.283 and 9.636 for the all-sectors and the manufacturing sectors cases respectively. See text for further details.

Table 5 reports the estimates of the ECM models (3). Overall, the estimations provide striking support for an error correction type of mechanism. For the all-sectors case, the estimates of λ suggest that production structures take about three years to adjust to half of the deviation to long-run equilibrium. Again, as in Section 4, the speed of adjustment is found to be higher in the case of manufacturing with a half-life of about 2.5 years.

6. Conclusion

Does trade integration make regions or countries more dependent on a few industries? This paper provides a negative answer by investigating the impact of the Canada–U.S. economic integration process on the dynamics of the Canadian provincial industrial structures. We compute specialization indices based on export data between Canada and the United States at the SIC 4 level (290 industries). The trade integration process was captured by the evolution over time of trade-weighted tariffs and turns out to be quite different across Canadian provinces. This allows us to combine pure time-series and the cross-sectional information in the econometric analysis and thus to improve the estimation of the various effects of trade integration.

It is obvious from the evolution of trade-weighted tariffs that Canadian provinces were not affected equally by the gradual elimination of trade barriers that was observed over the last 20 years. From an economic policy point of view, this indicates that the increased economic integration should be able to account for some of the heterogeneous evolution of the Canadian regional economies. From a methodological point of view, this illustrates the point that many things that are not necessarily intrinsically related to the field of regional economics could be learned by analyzing provincial/regional data. For example, because of the relative stability of the cross-sectional (cross-provincial) variances compared with the time-series variances, in this paper, we are able to disentangle the short- from the long-run effects of economic integration on diversification in a rather robust way. Given that the number of available time series is limited and that we are dealing with only one aggregate shock (CUSFTA), it would have been rather heroic to try this exercise looking at only one time series.

Our main result concerns the different impact over time of the trade integration process on the evolution of absolute diversification of the provincial economic structures. We found that on average, trade integration clearly leads to more diversification in the long run. We found quite limited evidence of some short-run effect in the other direction, although this impact is not robust across sectors. Such a dynamic process is consistent with the story of an initial relocation process for a given industry, followed by the agglomeration of other industries to take benefit of the backward-forward linkages. Interestingly, it was not found that trade integration increased the differences in production structures across the Canadian provinces, thereby putting more pressure on the federal insurance mechanism operated through taxes and transfers.

Finally, as the evolution of the trade-weighted tariffs in figure 2 clearly illustrates, most of the reductions in tariffs had been implemented by 1996. Given the relatively fast adjustment speed found in the empirical analysis, one could argue that the main contribution of the paper was to show that the adjustment of the Canadian industrial structure toward more diversification at the regional level following CUSFTA and NAFTA was completed by the eve of the new millennium. We think, however, that the analysis is also of interest from a prospective economic policy point of view since the process of Canada–U.S. economic integration is not completed. Canada as a whole still faces significant trade costs in its trade with the United States, costs that can be reduced by policy actions such as those related to the choice of the exchange rate regime. Furthermore, a certain number of sectors are still subject to significant tariffs. We found that 89, 52, and 15 of the SIC 4 industries faced tariff rates higher than 1%, 2%, and 5% respectively in 1996. Given the way tariff data are collected in our study, this fact indicates that many Canadian exporters prefer to pay the relevant MFN duties rather than meeting NAFTA rules-of-origin requirements.

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