

Economic Integration and Regional Industrial Specialization Patterns: New Evidence from the Canadian–U.S. FTA Experience

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Abstract

In this paper, we investigate the impact of trade integration between Canada and the United States on the degree of industrial specialization of the Canadian regions. Using provincial trade flows with the United States, we compute for each Canadian province the measures of absolute and relative industrial specialization relative to 290 industrial sectors. We captured trade integration through the decrease of trade-weighted tariffs that was boosted by the FTA implementation. We found strong evidence supporting 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. Consistent with this result, we found no evidence that the free trade agreement had an impact on the dissimilarity in industrial specialization among Canadian provinces.

Keywords: Regional specialization, trade integration, FTA, core-periphery structure, panel data

JEL classification: F14, F15, R12, R15

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1. Introduction

The effect of trade integration on economies of countries and regions has always been one of the most important issues studied in the area of international economics. For a couple of decades, the traditional theories supporting the view of trade integration as a welfare-improving process have been questioned by new theories resting on more realistic hypotheses. Trade integration is also of renewed interest to policy makers because important trade agreements have been implemented in many areas of the world over the last decades. Noted examples are the European integration process initiated in 1993; the Canadian–U.S. Free Trade Agreement (FTA) in 1989, followed by the North American Free Trade Agreement (NAFTA) in 1993; the recent Free Trade Area of the Americas project; and more recent free trade agreements among major East Asian countries. Not surprisingly, these developments have raised many questions in official and academic circles.

Trade integration obviously generates economic benefits. It is reported to induce significant efficiency gains, to promote growth, and to increase 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 of these suspected impacts is that trade integration may increase industrial specialization and therefore enhance or create 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 major aim of this paper is to assess the relationship between trade integration and industrial specialization patterns, using the Canadian–U.S. FTA experience. Using 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 FTA case is interesting in several respects. First, under the FTA, tariff (and non-tariff) trade barriers between Canada and the United States were gradually eliminated between January 1, 1989, and January 1, 1998. The FTA 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 Canada international trade share to GDP was relatively constant. However, between 1990 and 2000, this share increased steadily from 0.51 to 0.86 of GDP. 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). Second, despite this achievement, this project has raised many concerns in official Canadian circles, one key issue being the important 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 of such a development.

Basing our investigation on the FTA also 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 trade integration is captured. As Trefler (2002) emphasized, the FTA is a “clean” policy experiment that allows an exogenous measure of trade integration to be constructed. Such a feature has been called for extensively in the literature (Rodriguez and Rodrik 1999, among others) to permit meaningful estimates of the effects. Second, unlike previous contributions (Sawchuk and Trefler 2002; Trefler 2002), we 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-section 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 a pure cross-section analysis. A third contribution of our study is also related to the cross-section and time-series nature of the data. It is well known that it is very difficult to get information regarding the dynamics of change only from pure cross-sectional evidence. Likewise, obtaining precise estimates of the nature of the dynamic adjustment process from single time-series data is very cumbersome, given the strong multicollinearity among the explanatory variables (for example, see Hsiao 1986, Section 1.1). In a panel set-up, the cross-section dimension reduces the collinearity problem considerably while the time-series dimension allows the explicit disentanglement of the short- and long-run effects of shocks like the FTA. As advocated by some authors (Trefler 2002), the distinction between short-run adjustment costs and long-run efficiency gains is of primary importance for

assessing the impact of trade liberalization. Regarding the impact of the FTA in terms of industrial specialization, such a distinction turns out to be of overwhelming importance.

A final contribution of our study lies in the detailed attention devoted to specialization concepts. Here, we analyze the impact in terms of indicators of both absolute and relative specialization. Each analysis sheds a different light on the issue of trade integration. The first analysis (on indicators of absolute specialization) allows an investigation whether Canadian provinces are becoming more specialized than before, increasing their exposure to specific macroeconomic shocks. The second (on indicators of relative specialization) allows an investigation whether Canadian provinces are becoming more dissimilar from each other. This question is of primary importance given the decentralized nature of the Canadian federation and the extensive reliance on an interprovincial redistribution scheme operated through a federal tax transfer system that aims at smoothing out the consequences of specific regional disturbances.

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. Section 3 presents the data and the computation of indicators relevant to this study. 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, rest 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 overcomes these limitations and provides new insights on the impact of trade integration.

New trade theories have devoted 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 strong decrease in these costs can result in geographical dispersion of economic activities when mobility across sectors exhibits a finite cost. 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 strong 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.

Along with trade integration, other forms of economic integration can induce regions to increase or decrease the extent of their industrial specialization. Several theoretical arguments state that greater risk sharing among regions related to greater capital market integration can lead regions to specialize more in production (Helpman and Razin 1978; Kalemli-Ozcan et al. 2004). The argument is based on the notion that specialization generates obvious benefits but

also increases risk exposure to specific shocks. With the increase in the degree of risk sharing, producers can specialize and reap higher rewards from their activities. Kalemli-Ozcan et al. (2004) provide empirical support for this argument, showing for instance that the regions within a country that usually benefit from higher risk sharing are more specialized than the country as a whole.

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. Consistent with this finding, De Grauwe and Vanhaverbeke (1993) show that divergence in output growth rates and employment are higher for regions within countries than across countries.¹ These preliminary analyses thus provide evidence of a positive link between integration and specialization. While interesting, they do not, however, capture the integration process in an accurate way and remain with the low level of industrial classification. Subsequent studies aimed mainly at 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 experiences 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.

The first set of studies relies on production or employment data to characterize relative 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

¹ A related empirical literature focuses on the relationship between the asymmetry of macroeconomic fluctuations and trade integration. Frankel and Rose (1998) find that more trade makes these fluctuations more symmetrical. Kalemli-Ozcan et al. (2001) show that industrial specialization is negatively correlated with the symmetry of GDP fluctuations. On the whole, this emphasizes potential positive effects between integration and industrial diversification.

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, except for France.

Evidence on the evolution of relative specialization is also mixed. Based on the evolution of a K-spec index and using production data for 36 sectors and 14 European countries, Midelfart-Knarvik et al. (2000) found that relative specialization decreased between 1970 and 1980 but increased until 1997. However, they also report that the pattern diverged after 1990 when using trade data since the K-spec indices are now roughly constant. Furthermore, using 17 production sectors over 119 regions, Hallet (2000) finds a steady and significant decrease of relative specialization for 85 regions over the 1980–1995 period. To sum up, the European evidence of increased specialization due to trade integration appears the most mixed.

A set of recent studies gathered in Harris (2001) investigated the impact of the changes in North American linkages (FTA and NAFTA) on the structure of the Canadian economy. Acharya et al. (2001) isolate moderate (but statistically insignificant) evidence of an increase in the intra-industry trade between the United States and Canada, in line with previous 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, in contrast to an increasing specialization pattern observed for Mexico. This result is broadly consistent with the evidence provided by Head and Ries (2001). 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 point out trade integration's weak effect 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 in several directions. 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 assume either that integration follows a linear process and therefore investigate the evolution of specialization over time (as done by numerous European analysis), 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 progressively lowered since 1980, suggesting that trade integration had taken place well before the FTA. Second, the previous analysis ignores the high degree of heterogeneity across Canadian provinces in general and its core-periphery structure in particular. Canadian provinces, because of their difference in product specialization—especially when the primary sector is taken into account—face different dynamics in their trade-weighted tariffs. This in turn might have contrasted 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 the 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 and import data at the industry SIC 4 level that are 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.² The data bank used in this paper was constructed from the raw data with the valuable assistance of François Rimbaud from Industry Canada. We used all SIC 4 codes (from SIC number 0111 to 3999) for which at least one entry for one province was non-null. The final data used were 290 series for export data and 291 series for import data. Data were computed for exports to and imports from the United States (including U.S. territories), and exports to and imports from the rest of the world.

² 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>.

Before exposing the computation of the specialisation indexes, three important comments are in order. First, the export data used in this paper are measured as shipments from the province for which the goods were grown, extracted, or manufactured. In this sense, they are based on the province of origin and not on total shipments which would induce strong bias in favour of shipment locations. Second, export data net out re-exports of goods to the United States which have been previously imported and are leaving the country without having been processed in Canada. The third point concerns the use of exports data rather than output and/or employment data to compute specialisation indexes. Basically, the two approaches are used in the empirical literature. The choice is often dictated by data availability. In the Canadian case, the use of export data yields two advantages. First, trade flows at the industry level is measured more accurately than employment and output. It also allows using a higher level of desegregation, making possible to characterize in a more precise way the exact shape of the production structure. Second, the use of trade data allows to abstract from the non-traded goods sector and to exacerbate the degree of heterogeneity between provinces. Our specialisation indexes reflect more directly the part of the provincial industrial structure that is likely to be affected by a change in trade costs.

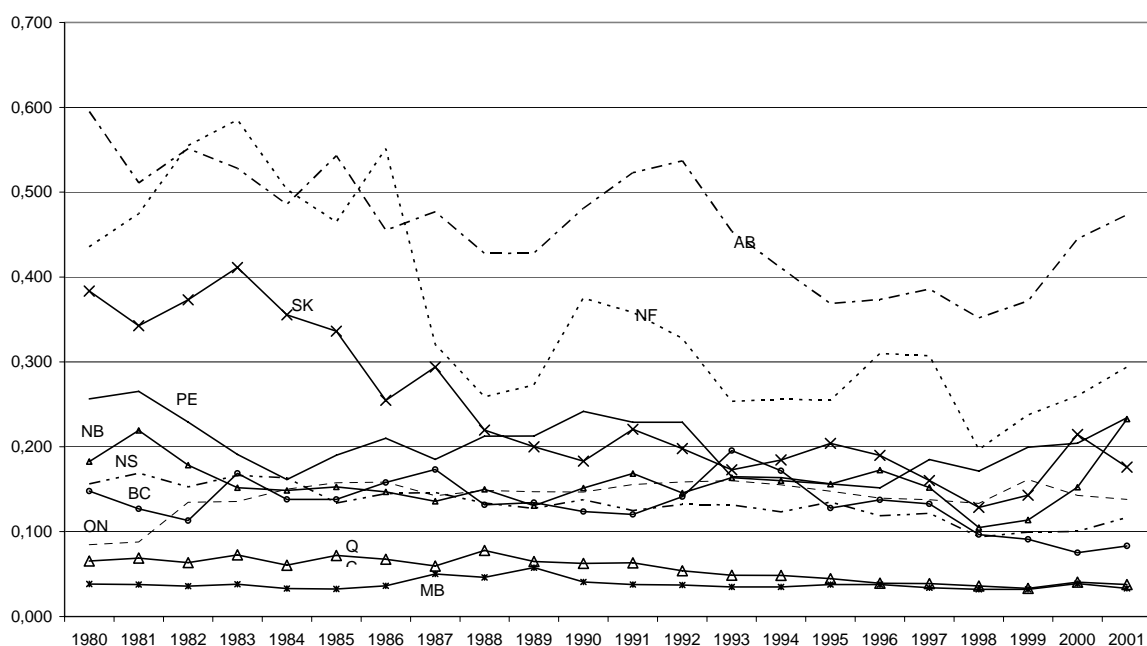
The Herfindahl (SH), the K-spec (SK), and the Grubel-Lloyd (SG) are the three indexes used in this study to capture complementary characteristics of the provincial industrial structure. The Herfindahl index captures the degree of absolute specialization and might be used for the purpose of international 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 was computed as:

$$SH_{i,t} = \sum_{k=1}^{290} (s_{i,t}^k)^2,$$

$$\text{where } s_{i,t}^k = \frac{x_{i,t}^k}{\sum_{k=1}^{290} x_{i,t}^k}$$

Here, $s_{i,t}^k$ 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 squared of the shares over all industries. By definition, $SH_{i,t}$ will therefore be between $1/290$ and 1 . The smaller (bigger) the number, the more diversified (specialized) is the industrial structure of the involved province.

Figure 1: Herfindahl indexes for Canadian provinces



The series for the Herfindahl index 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 *SH* 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 imputed to the oil and natural gas that account for 67% of Alberta’s total exports to the United States during the period. For Newfoundland, the high 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 (see Coulombe 1999). Between 1980 and 1986, fish products accounted on average for 68% of Newfoundland’s total exports. Since 1987, this industry’s share has fallen to an average of only 23%. Note that the upsurge of Newfoundland Herfindahl between 1998 and 2001 is 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, which has accounted for about 38% of Ontario’s total export to the United States since 1990.

The K-spec (Krugman 1991a), in contrast to the Herfindahl 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 nine other provinces. It is defined as:

$$SK_{i,t} = \frac{1}{2} \sum_{k=1}^{290} \text{abs}(s_{i,t}^k - \bar{s}_{i,t}^k)$$

where $\bar{s}_{i,t}^k = \sum_{j \neq i} x_{j,t}^k / \sum_{k=1}^{290} \sum_{j \neq i} x_{j,t}^k$

The $SK_{i,t}$ index takes value zero if the industrial structure of province *i* is identical to the average of the other nine provinces. Given the normalization used here, the maximum value of $SK_{i,t}$ is equal to 1 when the export structure of one province differs completely from that of the rest of the country.

Figure 2: K-spec indexes for Canadian provinces

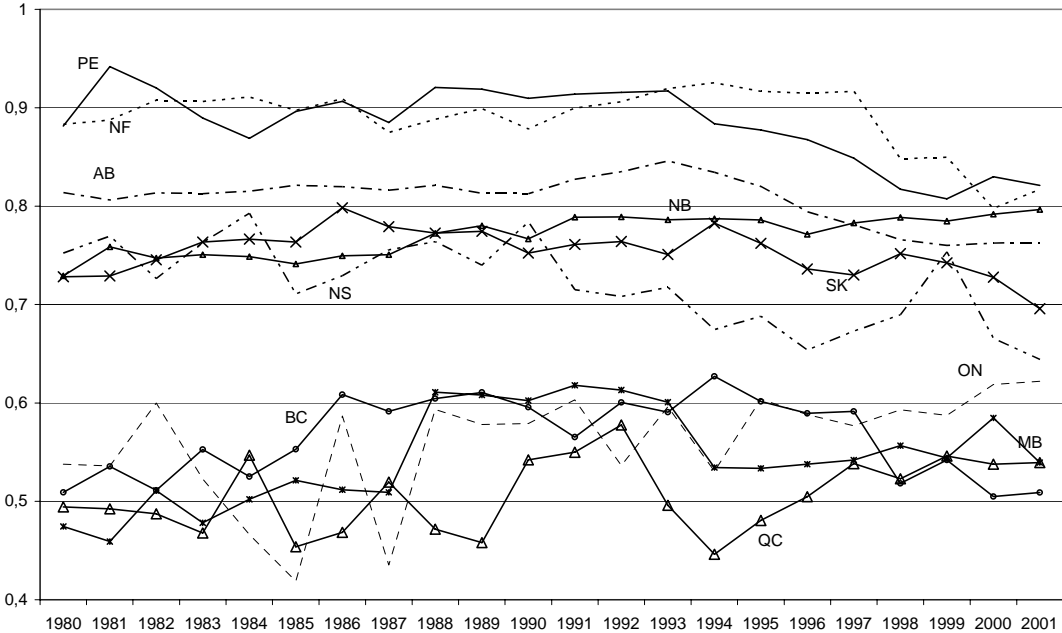


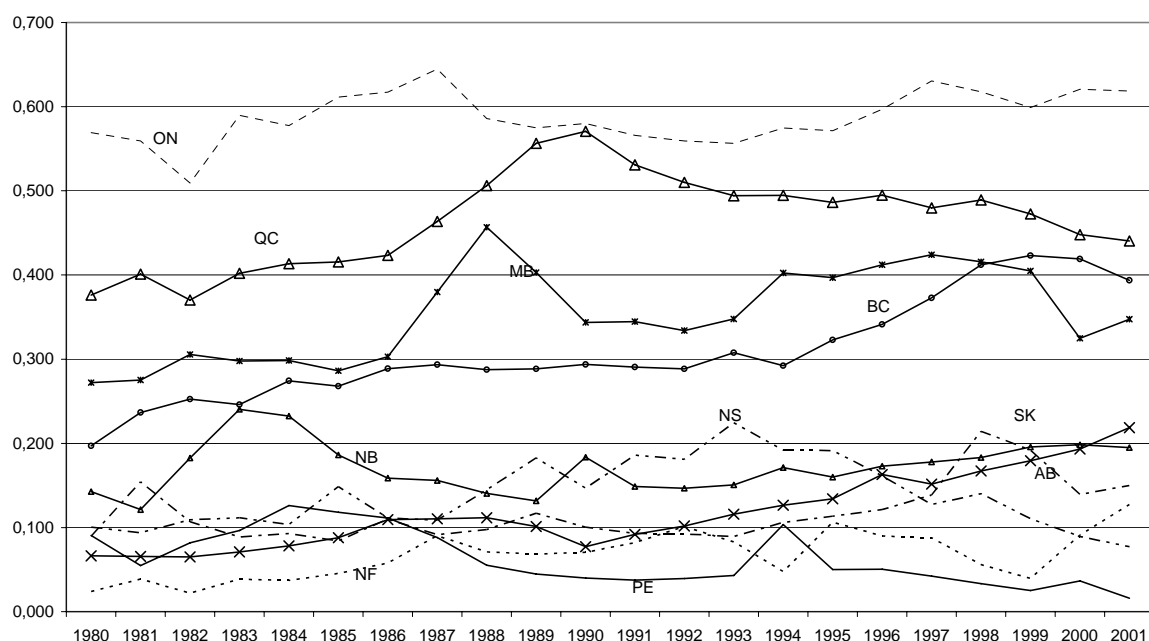
Figure 2 shows the evolution of the K-spec index, calculated from exports to the United States for the 10 provinces. Four provinces, Quebec, followed by Manitoba, Ontario, and British Columbia, share industrial structures that are more similar to the rest of the country than the other six provinces. Interestingly, the K-spec structure appears more stable than the Herfindhal.

Finally, the Grubel-Lloyd index (Grubel and Lloyd 1975) is used as the measure of intra-industry trade. The nature of the trade created by the integration process is of overwhelming importance in assessing the impact of this process on the degree of asymmetry of macroeconomic disturbances. Following the findings of Frankel and Rose (1998), trade integration is thought to increase business cycle symmetry, provided integration raises the share of intra-industry trade rather than inter-industry trade. The Grubel-Lloyd index is defined as:

$$SG_{i,t} = 1 - \frac{\sum_{k=1}^{291} \text{abs}(x_{i,t}^k - m_{i,t}^k)}{\sum_{k=1}^{291} (x_{i,t}^k + m_{i,t}^k)}$$

The Grubel-Lloyd might be viewed as the proportion of intra-industry trade in the total trade of a province. Grubel-Lloyd indexes based on exports to and imports from the United States are shown for the 10 provinces in Figure 3. Not surprisingly, and again due to the automobile industry, Ontario has consistently the highest ratio of intra-industry trade with an average of .588 for the entire period. Quebec is not far behind despite a downward trend in the 1990s. Manitoba and British Columbia also show a substantial degree of intra-industry trade but the other six provinces are well behind.

Figure 3: Grubel-Lloyd indexes for Canadian provinces



3.2 Capturing trade integration

A straightforward way 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, alternative measures that are obviously exogenous from specialization patterns and the share of intra-industrial trade are called for to determine trade integration.

Following several authors including Trefler (2002) and Sawchuk and Trefler (2002), our integration measure will be based on the dynamics of the tariffs between the United States and Canada. Our starting point is the Canadian tariff against the United States per industry of Trefler (2002) and Sawchuk and Trefler (2002). Unlike these studies, however, our trade-weighted tariffs are 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 for weights for Canada as a whole. Trade-weighted tariffs are defined as:

$$TW_{i,t} = \sum_{k=1}^{290} w_{i,t}^k \tau_t^k$$

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.³

Figure 4. Provincial trade weighed tariffs

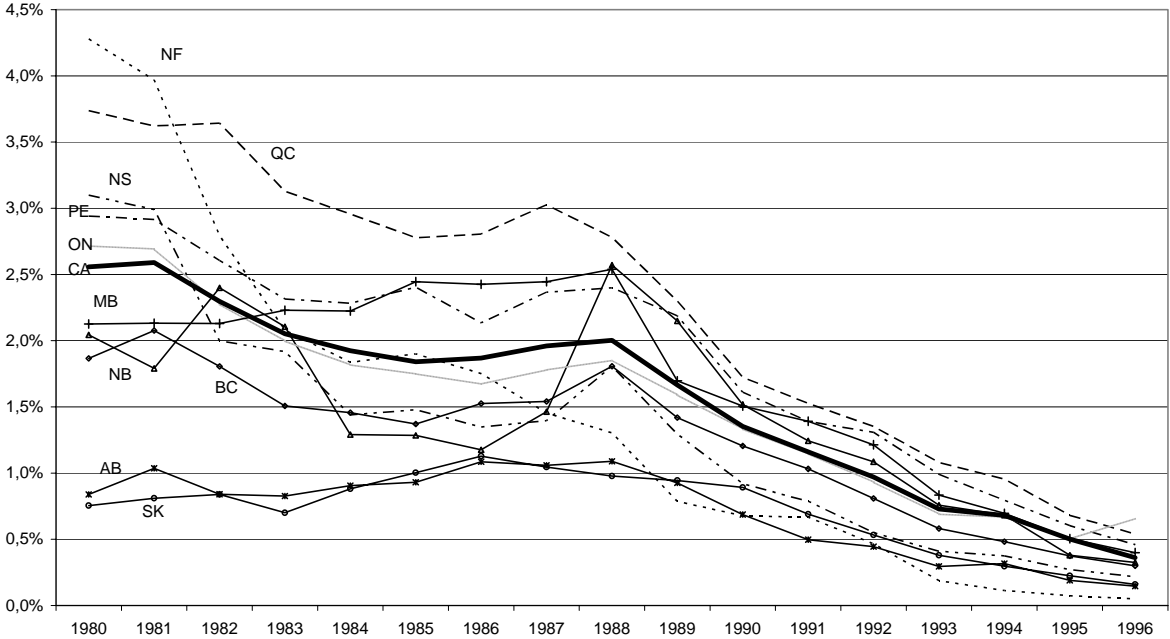


Figure 4 reports the evolution over time of the TW for Canada as a whole and for each of the provinces. It illustrates the decreasing pattern of the tariffs since 1980. While the FTA induced a quicker rate of decay for the TW , Figure 4 also shows that trade integration between Canada and the United States had begun well before the implementation of the FTA in 1989. Furthermore, what remains in effective tariffs in 1996 reflects the fact that some Canadian exporters to the United States prefer to pay the tariff of the most favoured nation instead of

³ Trefler (2002) uses the tariffs between Canada and the United States for the manufacturing sector over the 1980–1996 period. We complement Trefler’s data with the tariffs on primary products that were given by S. Rao from Industry Canada. For the 1997–1998 period, we assume that the tariffs on the 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.

having to incur the administrative and compliance costs associated with the rules of origin requirement that governs all trade under NAFTA.⁴

Figure 4 also reflects an important point that is relevant to our analysis at the regional level, that is, the important degree of heterogeneity of the integration process across provinces. This heterogeneity comes from the combination of 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 for both their levels and their speeds of decay. As an illustration, Newfoundland and Labrador experiences the sharpest decrease in TW , 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 FTA. 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. In turn, it allows us to overcome the statistical problem of a low number of data points at hand in pure time-series approach of the FTA.

Two comments are necessary regarding our measure of trade integration. First, as pointed out by Trefler (2002), the decrease of the $TW_{i,t}$ obviously 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).⁵ In this sense, it captures a broader set of FTA trade-liberalizing policies. The second point concerns the exogenous nature of the $TW_{i,t}$. As emphasized by Trefler (2002), the FTA is a “clean” policy experiment in the sense that it was not considered as part of those policies aimed at smoothing macroeconomic disturbances or counteracting adverse industrial development in Canada or the United States. In this sense, the dynamics of $\tau_{j,t}$ might reasonably be considered as unaffected by the evolution of the indicators of industrial production. However, even if the $\tau_{j,t}$ are exogenous, the $TW_{i,t}$ might be endogenous because the weights $w_{ij,t}$ might depend on the $\tau_{j,t}$. If it is the case, our $TW_{i,t}$ might provide a biased estimate of provincial trade protection. To address this issue, we provide some analysis of robustness in Appendix A, using an alternative computation of

⁴ Kunimoto and Sawchuk (2004) present a broad overview of the issue of NAFTA rules of origin.

⁵ 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.

the $TW_{i,t}$. In the alternative measure, we use constant weights over time, that is,

$$w_{ij,t} = \bar{w}_{ij} = \frac{1}{22} \sum_{t=1980}^{2001} w_{ij,t} \text{ for each } t. \text{ Our results using this alternative measure suggest that}$$

our main findings are unaffected by this problem.

3.3 Other variables

As emphasized by Kalemli-Ozcan et al. (2001), the potential determinants of industrial specialization are primarily those affecting the volume of U.S.–Canada 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.⁶ The annual measure of each cycle is then obtained from the average of the quarterly cycles.

The second set of variables is related to 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 one is exchange rate volatility. For instance, Fontagne and Freudenberg (1998) found that exchange rate variability tended to lower the share of intra-industry trade in Europe during the EMS experience. We compute exchange rate volatility 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 return:

$$\sigma_t^2 = \sum_{i=2}^{N_t} \left[\ln(s_{t,i}) - \ln(s_{t,i-1}) \right]^2$$

⁶ 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.

where σ_t^2 denote realized volatility at time (year) t , $s_{t,i}$ is the CAD–USD exchange rate at day i of year t , and N_t is the number of trading days during year t .⁷ As shown by Andersen et al. (2001), this measure provides an unbiased and more efficient estimate than usual measures such as historical volatility. It is an unbiased measure of volatility as soon as exchange rate returns are uncorrelated over time, which is obviously the case at the daily frequency.

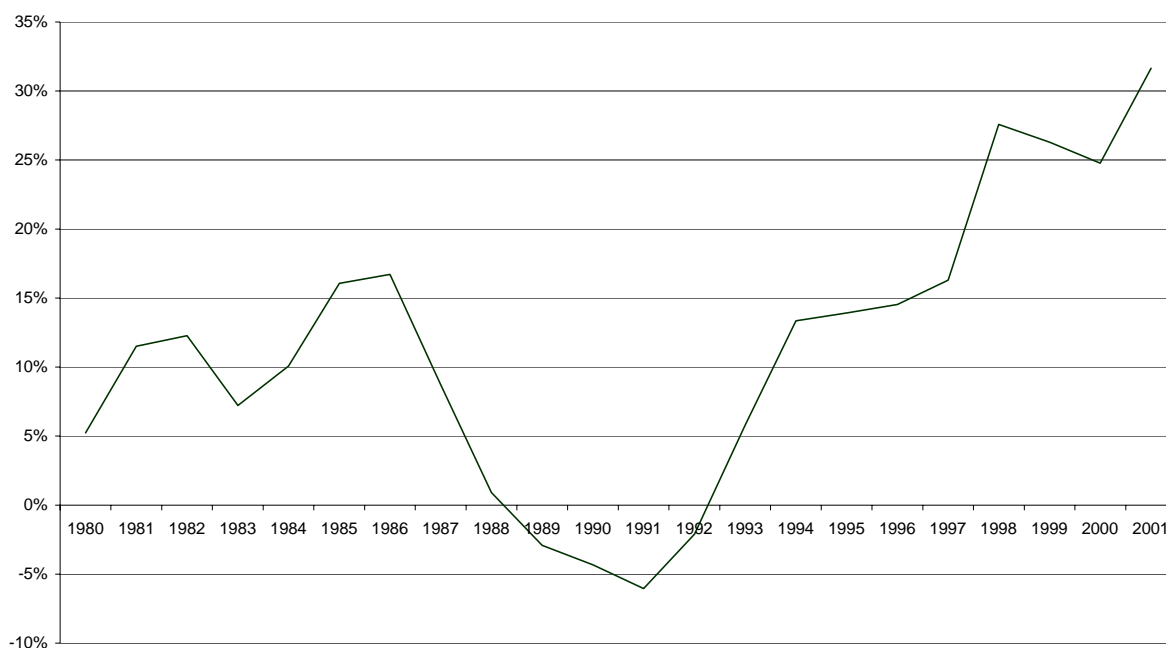
Finally, we introduce the degree of exchange rate misalignment between the Canadian and U.S. dollars. It has been argued that monetary policy in general and the overvaluation of the Canadian dollar in the early 1990s had detrimental effects on the macroeconomic dynamics of the Canadian economy (Fortin 1996). Likewise, Harris (2001) advances that the subsequent undervaluation of the Canadian dollar was partly responsible for the productivity gap between the Canadian and the U.S. economies. We compute misalignment as the annual absolute deviations of the bilateral exchange rate from its equilibrium value.⁸ The equilibrium is assumed to be the purchasing power parity level computed with the ratio of producer prices between Canada and the United States.⁹ Figure 5 plots the degree of misalignment in percentage of the CAD against the USD over the 1980–2001 period. It confirms that over the recent period, the CAD depreciated strongly against the USD, leading to a significant undervaluation of the CAD.

⁷ The exchange rate data of the CAD–USD at the daily frequency were kindly provided by the Bank of Canada.

⁸ In the econometric investigation, we consider both the nominal and the absolute deviations. We present results using only the absolute deviation since the effect of the misalignment variable proved to be better captured empirically with the absolute measure.

⁹ The producer price index is extracted from the Main Economic Indicators database of the OECD.

Figure 5: Degree of misalignment of the CAD-USD exchange rate



4. Econometric analysis under the stationarity assumption

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 pointed out before, the panel data approach takes into consideration the fact that some Canadian provinces, due to their respective industrial structure, were more affected by the change in tariffs. Taking this fact into account multiplies the information compared to a pure time-series investigation based on overall Canadian data. Furthermore, because of their relatively stable cross-sectional distribution, panel data are known to be better able to study the dynamics of adjustment and to isolate the long-run effects of exogenous variables than times-series data (Baltagi 1999, Section 1.2).

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

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

Here, $i=1, \dots, 10$, stands for the Canadian provinces, and $t = 1980$ to 1998 . In $SX_{i,t}$, X refers to one of our three measures (H, K, G) of industrial structure S of province i at time t . $TW_{i,t}$ is the trade-weighted tariff of province i at time t applicable to Canadian imports to the United States (see equation [2]). The $Z_{i,t-p}$ are the other variables that might account for the changes in the industrial structure S during the period under study. The lag p equals either 0 or 1. Most of the Z are aggregate variables and are constant across the i . In some interaction set-ups, alternative X measures of industrial structure are introduced in the $Z_{i,t-p}$ and $\Delta Z_{i,t}$. The $\varepsilon_{i,t}$ are the idiosyncratic errors and the c_i , the unobserved individual components; these are discussed in the following section.

Estimation techniques

Of course, 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 of 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 FGLS estimation of the benchmark equation (1), all provinces' fixed effects are significant well below the 5% level. Conversely, in the same set-up, no time dummies were significant at the 5% level and consequently, time dummies were not used 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 pooled least-squares (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 computed HCCME to allow for asymptotically valid inferences in the presence of the remaining time-series 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.

Finally, we also perform seemingly unrelated regression (SUR) estimations to all regression set-ups. In the panel data model, SUR is also known as the Parks estimation. SUR is the least restricted estimation technique here since it is the FGLS estimator when the residuals are both cross-sectional, heteroscedastic, and contemporaneously correlated. In many cases, however, with SUR, the interactive updating process did not converge. In these cases, SUR generated extremely high t-statistics and, fortunately, points estimates that have the same sign and order of magnitude as with FGLS. Parks estimations are known to be able to produce standard errors that lead to extreme confidence when the number of time series is not sufficiently larger than the number of cross-sections, as it is the case in our study.¹⁰ For this reason, we often do not report the results of SUR estimations.

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 (Hendry 1995). In order to disentangle the long- and short-run dynamics, our benchmark specification relies on what is called a general equilibrium correction specification in the typology adopted by Hendry (1995):

$$\Delta SX_{i,t} = \phi_1 SX_{i,t-1} + \phi_2 TW_{i,t-1} + \phi_3 \Delta TW_{i,t} + \phi' Z' + c_i + \varepsilon_{i,t} \quad (1)$$

where $\phi' Z' = \phi_4 UC_t + \phi_5 PC_{i,t} + \phi_6 VL_t + \phi_7 MA_t$

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. Finally, the c_i are provinces' fixed effects.

¹⁰ Note that this problem does not apply to our FGLS procedure with cross-section weights. Beck and Katz (1995) provide a thorough analysis of problems associated with Parks estimation in time-series cross-section data.

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 $SX_{i,t}$ and $TW_{i,t}$, model (1) can be estimated directly by PLS, FGLS, or SUR. The long-run effect of economic integration that is associated with $TW_{i,t}$ is then given by $(\hat{\phi}_2 / -\hat{\phi}_1)$; the short-run effect is captured by $\hat{\phi}_3$.

The three different indexes X that were used to capture the various characteristics of the industrial structure S are the Herfindahl (H), the K-spec (K), and the Grubel-Lloyd (G). The most significant results of the empirical analysis refer to the effect of the change in tariffs on the degree of specialization of Canadian provinces as captured by the Herfindahl index. Therefore, the analysis that was based on the dynamic evolution of $SH_{i,t}$ was extended to export data of just manufacturing industries, and to export data to the rest of the world.

Interaction specification: In the second regression set-up, the interaction specification, another measure Y of industrial structure is entered in the right-hand side of equation (1) for $X \neq Y$:

$$\Delta SX_{i,t} = \phi_1 SX_{i,t-1} + \phi_2 TW_{i,t-1} + \phi_3 \Delta TW_{i,t} + \theta_1 SY_{i,t-1} + \theta_2 \Delta SY_{i,t} + \phi'Z' + c_i + \varepsilon_{i,t} \quad (2).$$

The purpose of these interaction specifications is to investigate potential causality or spillover effects between the industrial production indexes.

Heterogeneous specification: The third regression set-up, which exhibits three variants, refers to the heterogeneous specification. In the first variant (3), long- and short-run effects of tariff changes ϕ_2 and ϕ_3 were allowed to vary across provinces:

$$\Delta SX_{i,t} = \phi_1 SX_{i,t-1} + \phi_{2,i} TW_{i,t-1} + \phi_{3,i} \Delta TW_{i,t} + c_i + \varepsilon_{i,t} \quad (3)$$

The $\phi'Z'$ factors were not entered in this least restricted set-up given the limited number of degrees of freedom.¹¹ In the second variant (3'), in the spirit of Pesaran et al. (1999), pooled mean group estimation of dynamic heterogeneous panel, the speed of adjustment $\phi_{1,i}$ and the

¹¹ In another set of regressions, we also allow the speed adjustment ϕ_1 to vary across provinces, along with the long run effect ϕ_2 . The regression results are not reported here in order to save space but are available upon request.

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:

$$\Delta SX_{i,t} = \phi_{1,i} SX_{i,t-1} + \phi_2 TW_{i,t-1} + \phi_{3,i} \Delta TW_{i,t} + c_i + \varepsilon_{i,t} \quad (3')$$

The $\phi'Z'$ factors were again not entered in the regression. Finally, in the third heterogeneous set-up (3''), the Z' variables were allowed to differ across provinces, one at a time:

$$\Delta SX_{i,t} = \phi_1 SX_{i,t-1} + \phi_2 TW_{i,t-1} + \phi_3 \Delta TW_{i,t} + \phi'Z' + \phi'_i Z' + c_i + \varepsilon_{i,t} \quad (3'')$$

Here, the $\phi'_i Z'$ term on the right-hand side refers to one of the four Z' variables with a specific provincial effect while the $\phi'Z'$ refers to the other three control variables.

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 LSDV estimators of the ϕ_1 and the other ϕ s are biased when the number of time periods T is small and the number of cross-section N is large as in traditional microeconomic studies. Kiviet (1995) shows that LSDV estimators are consistent only when $T \rightarrow \infty$. The usual solution to this problem was to use instrumental variable (IV) methods (for example, Hsiao 1986), and more recently, Generalized Method of Moments (GMM) estimators (Arellano and Bover 1995). 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 rather 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 with the use of the heterogeneous specifications (3'') 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 these tables, the dependent variable is the measure of industrial specialization based on Canada–U.S. trade. In Tables 1 and 3, the primary sector ($\Delta SH_{i,t}$) is included while in Table 2, the analysis is restricted to industries of the manufacturing sector ($\Delta SHM_{i,t}$), as was done in Trefler (2002). In all but one case, the various parameter estimates of the $TW_{i,t-1}$ variable (ϕ_2) are positive and significantly different from 0 at the 1% level. The one slight exception is the interaction set-up with the intra-industry trade measure in the third column of Table 3 for which the p -value of the ϕ_2 parameter amounts to 1.37%. 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% level. Therefore, the measures of the long-run effect given by $\left(-\frac{\hat{\phi}_2}{\hat{\phi}_1} \right)$ of tariff changes on trade specialization are always highly negative and significant with p -values below 1.4%.

As illustrated by Figure 3, the FTA has accelerated the negative trend in tariffs since 1989. As a consequence, our results imply that the mean (across provinces) effect of trade integration induced by the FTA was a significant increase of the long-run industrial diversification of Canadian provinces. This result is extremely robust and remains invariant across estimation techniques (PLS, FGLS, and SUR) in all benchmark regressions (Tables 1 and 2) as well as in interaction specifications with the Grubel-Lloyd index introduced as a control variable (Table 3).

Short-run effect of tariff changes

The other significant highlight of Tables 1 to 3 is that the various point estimates of the short-run effect (the ϕ_3 parameter in specifications [1] and [2]) of tariff changes $\Delta TW_{i,t}$ on industrial diversification $\Delta SH_{i,t}$ and $\Delta SHM_{i,t}$ are always negative. The (reverse) short-run effect, however, is somewhat less robust than the long-run positive effect; it is not significant at the 5% level in a subset of pooled regressions. The effect is usually not significant at the 10% level in PLS regressions and is significant at the 10% level in only one FGLS set-up for the manufacturing sector in Table 2.

Overall, there is some obvious evidence that the short-term effect of tariff reductions differs from the long-run effect. In the short run, economic integration appears to 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 story proposed by some of the models of economy 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.

Dynamic adjustment

As pointed out earlier, 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 tariffs reductions induced by the FTA.

The point estimates of the ϕ_1 parameter relative to the dynamic evolution of $\Delta SH_{i,t}$ and $\Delta SHM_{i,t}$ (Tables 1 through 3) are quite stable across the various specifications: they vary between -0.27 and -0.38. In the case of the benchmark model for the all-sectors Herfindahl index (Table 1), the estimated value of ϕ_1 amounts on average to -0.30. 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 long-run level is adjusted after only 2.3 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 six point estimates amounts to only 2 years. This 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.

The speed of adjustment to the new equilibrium of the industrial structure implies that the contemporaneous negative effect of tariff cuts on regional industrial diversification is quickly overcome. For example, the negative effect of a once and for all decrease in tariffs occurring during year t would be totally offset by the end of year $t+1$ or $t+2$ (depending on the estimates). After that, the economy converges rapidly to a new equilibrium level that is characterized by a higher degree of industrial diversification.

Quantitative effect

The various quantitative effects of tariff cuts on the long-run level of the Herfindahl index

$\left(-\frac{\hat{\phi}_2}{\hat{\phi}_1} \right)$ computed from the point estimates and depicted in Tables 1 and 2 vary significantly

across estimation techniques. They are systematically higher (lower) for PLS (FGLS) estimations for the same regression set-up while SUR effects fall in between. They vary between 1.13 and 3.91 for the all sectors (Table 1) and between 1.51 and 4.04 for the manufacturing sector (Table 2). The patterns of the quantitative effects that are computed from the interaction specification (2) in Table 3 compare closely with those emerging from the analysis of Tables 1 and 2. Overall, even though the long-run effect of tariff reduction on regional diversification is clearly positive, the actual quantitative effect appears awkward to measure with a high degree of precision.

Heterogeneous specification

Table 4 displays the results of the effect of tariff changes on industrial diversification that are captured by heterogeneous specification (3). The dynamic system is estimated using one common coefficient ϕ_1 for the lagged dependent variable $SH_{i,t-1}$; 10 cross-section specific coefficients $\phi_{2,i}$ for the lagged trade-weighted tariff variable $TW - PR_{t-1}$ where PR refer to the province-specific index; and 10 cross-section specific coefficients $\phi_{3,i}$ for the contemporaneous change in tariffs $\Delta TW - PR_t$ (along with 10 cross-section specific fixed effects).

Before interpreting these results, it is important to bear in mind that the effect of tariff changes is here measured much less precisely than with specifications (1) and (2) since it relies only on the time-series information of tariff changes and industrial diversification. Accounting for cross-sectional variance through panel regression analysis is likely to improve

the precision of the estimation, especially when dealing with heterogeneous economies such as the Canadian provinces.

Nevertheless, with the notable exception of the largest and most industrialized province (Ontario), all point estimates of the long-run effect of tariff changes relative to specification (3) are positive, in line with the results yielded by specifications (1) and (2). The negative estimated effect for Ontario using FGLS and SUR exhibits very high p -values. The magnitude of the long-run effect varies widely across provinces. They are higher for the less-diversified economies such as Alberta and Newfoundland. Not surprisingly, the level effect is smaller for the well-diversified provinces of Quebec and Manitoba but is always significant at the 1% and the 5% critical level respectively. Overall, five long-run effects are significant at the 10% level with PLS, six at this significance level with FGLS, and seven at the 5% level with SUR.

Fewer than one-quarter of the short-run tariff changes, province-specific effects are positive and have the reverse sign relative to the corresponding average effect. One of them is significant only at the 5% level—Manitoba with SUR. Many negative short-run specific effects are significant at the 5% level. Interestingly, the negative effect for Ontario is substantial and very significant. Consequently, for the largest and most industrialized Canadian provinces, the effect of tariff changes tends to induce specialization in the short run while the production structure returns around the initial level of diversification in the long run. This combination of short- and long-run dynamics is consistent with a story of initial polarization and subsequent agglomeration of different industries that was presented in the theoretical background section.¹²

Furthermore, in line with the pooled mean group estimation technique of Pesaran et al. (1999), cross-section means and medians for the $\phi_{2,i}$ and $\phi_{3,i}$ point estimates (as well as the asymptotic standard error of the mean) are reported in Table 4. Again, our main result regarding the long-run positive effect of economic integration on diversification appears much more robust than the opposite negative short-run effect.

¹² A formal test of such a story would of course imply an empirical analysis of agglomeration effects of the FTA, which is obviously beyond the scope of this paper.

The point estimates of ϕ_1 are on average 33% higher in absolute value than the ones estimated from the aggregate specifications, thus implying a faster dynamic adjustment. In a dynamic system, when the effect of a shock is allowed to vary across the economies, the estimated speed of adjustment is likely to be higher. Despite the limited abilities of the heterogeneous specification to estimate precisely the effect of tariff changes on industrial diversification, the results generally illustrate the robustness of the findings emerging from Tables 1 through 3.

Comparing LSDV with PMG

As pointed out in the last paragraph of Section 4.1, Pesaran and Smith (1995) argue that LSDV might produce biased estimators when parameter heterogeneity is ignored in a dynamic model. The solution proposed by Pesaran et al. (1999) is the pooled mean group estimation (PMG). Here, short-run effects and the speeds of adjustment are allowed to vary across economic units (the provinces in our case) while estimating a common parameter for the long-run effect. In Table 5, we report results where LSDV estimates of the benchmark model (without the Z 'variables that correspond to column 5 in Table 1) are compared with PMG estimation where both FGLS is used as the common estimation technique.

Both quantitatively and qualitatively, the results of the PMG estimation align very well with the general direction of our main results, with just one exception. The point estimate (mean) of the short-run effect of tariff changes (ϕ_3) is much higher than the one estimated in the benchmark specification. However, the magnitude of the mean of the $\phi_{3,i}$ is driven up by the presence of a couple of large values and the median of the $\phi_{3,i}$ is much closer to its LSDV point estimate counterpart. More importantly, however, 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).

Exports to the rest of the world

Up to now, the effect of tariff changes on industrial diversification has been 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 FTA. Indeed, as emphasized by Trefler (2002), the FTA 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 is directly tested using specifications (1) and (2) for which the dependent variable ($\Delta SHR_{i,t}$) is the Herfindahl index computed from provinces’ exports to the rest of the world. Results of the panel data regressions are displayed in Table 6.

The specialization substitution hypothesis is clearly rejected by our results. No point estimates of the $TW_{i,t-1}$ variable that capture the long-run effects of tariff reductions with the United States are significant and the p -values are extremely high, especially with FGLS. There is, however, some evidence that the specialization substitution hypothesis might hold in the short run: the point estimates of the contemporaneous change in tariff ($\Delta TW_{i,t}$) are positive and significant at the 10% level in three out four cases, with the fourth case displaying a p -value of .107. These results suggest some limited evidence that the contemporaneous increased specialization generated by tariff changes coming from exports to the United States is partly offset by a contemporaneous increased diversification of exports to the rest of the world. The overall short-run effect of tariff changes on industrial specialization is then likely to be smaller than the effect estimated in Tables 1 to 4 (and it was already not that robust).

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 all specifications. Those variables are measures of the U.S. and Canadian provincial business cycles, of the bilateral exchange rate volatility against the U.S. dollar, and of the degree of misalignment between the Canadian and the U.S. dollars.

As shown by the analysis of Tables 1 to 3, with the exception of the exchange rate volatility variable, their effect in specifications (1) and (2) is somewhat limited. It is important to note that most of the significant point estimates come from the SUR estimations. As indicated earlier, the point estimates with SUR should be interpreted with caution since the iterative process did not converge most of the time. When there are a large number of common coefficients to estimate, SUR shows a tendency to generate very low standard errors. We should, therefore, not place too much confidence in results that are significant only when the SUR approach is used.

Having said that, there is still some evidence to suggest that increases in exchange rate variability have a positive effect on industrial diversification. The effect, however, appears to be limited to regressions including the primary sectors. In contrast, there does not seem to be a significant role for exchange rate volatility 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).

In numerous (not reported) variants of regression (3'), we tried to isolate cross-sectional specific effect of the Z' variables. In these cases, SUR estimates are in line with those obtained with FGLS and PLS: the null hypothesis of zero impact was not rejected in almost all cases, with the exception of the exchange rate variability variable.

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 3 illustrate the robustness of our main results regarding both the long- and the short-run effects of tariff changes. The results suggest that introducing other independent variables is unlikely to alter the general direction of the main results. Some of the least significant Z' variables (provincial and U.S. cycles) were deleted from Table 2 and the following tables to save space. The effect of the misalignment variable, however, will prove very interesting in the analysis of the dynamics of intra-industry trade and the K-spec examined in the following sections.

4.3 Results: the dynamics of intra-industry trade

Results regarding the dynamic evolution of intra-industry trade are displayed in Tables 7 to 9. At first glance—as in the Torstensson (1996) study of the determinants of intra-industry trade in 1983 and 1989 in Sweden—we could expect a rather negative relationship (the ϕ_2 parameter) between the Grubel-Lloyd variable and the lagged level of tariffs in specifications (1) and (2). We found the expected negative sign for the ϕ_2 parameter, but with the exception of SUR (non-reported results), the evidence is not robust. In the case of all sectors (Table 7) with FGLS and PLS, the effect is never significant at the 10% level. Furthermore, the short-run effect is found to be insignificant with all estimation techniques. However, the existence

of a positive effect of trade integration on the share of intra-industry trade is much more obvious for the manufacturing sector (Table 8).¹³ On the whole, these results are quite consistent with those of Torstensson (1996) that were obtained for 1989.

One of the most interesting results related to the dynamics of the all-sector Grubel-Lloyd indicators is the positive and significant estimate at the 5% level for the misalignment variable obtained with FGLS and SUR estimations. In the heterogeneous specification (3') (Table 9) where the effect of misalignment is allowed to vary across provinces, PLS results concur with FGLS and SUR. The key point with this specification is that the overall misalignment result is clearly driven by the three western provinces of British Columbia, Alberta, and Saskatchewan. For these provinces, the estimated positive parameters are always significant at least at the 5% level with all estimation techniques. For the other provinces, the sign is often reversed but the misalignment variable is never significant at the 5% level with the three estimation techniques. For Quebec and Nova Scotia, however, the misalignment variable is close to being significant at the 10% level with PLS and FGLS; it is significant at the 5% and 10 % level with SUR for Quebec and Nova Scotia respectively. For these two provinces, the impact goes in the opposite direction compared to the one obtained for the three western provinces. The overall direction of the results emerging from the heterogeneous specification clearly illustrates that the evolution of the bilateral exchange rate in Canada is likely to affect Canadian provinces differently. This finding once more concurs with Beine and Coulombe's (2003) overall analysis, which highlighted the high degree of heterogeneity of the Canadian provinces with respect to the desirability of keeping the prevailing flexible exchange rate regime.

Results for specification (1), restricted to the manufacturing sector, are displayed in Table 8. This leads to an insignificant effect for the misalignment variable on the Grubel-Lloyd. This result suggests that the specialization effect of the misalignment variable on intra-industry trade applies strictly to primary sector industries. In this sector, one might argue that products are less differentiated than in the manufacturing sector. Consequently, it is possible that a higher degree of misalignment between the Canadian and U.S. dollar generates some sort of arbitrage opportunities between firms of the same industry in the two countries. Such

¹³ The analysis of Torstensson (1996) applies only to the manufacturing industries.

opportunities may indeed lead some industries to specialize in certain particular industries or products.

4.4 Results: the core-periphery structures

As well documented in Coulombe (1999) and 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 key interests in analyzing the incidence of tariff changes at the regional level in Canada is to then investigate whether the FTA has induced a change in the core-periphery structure of the Canadian economy. Such an investigation is also of overwhelming importance for Canadian economic policy. If trade integration tends to strengthen the core-periphery structure, this would increase the asymmetry of macroeconomic fluctuations (Kalemli-Ozcan et al. 2001). And this in turn would put more pressure on the federal fiscal mechanisms aimed at smoothing the interregional differences of economic developments.

In this section, we first test directly this potentially important policy issue in Canada by estimating the effect of tariff changes on the relative specialization structure of the Canadian provinces through an analysis of the dynamics of the K-spec indicator. We then investigate whether a measure of the core-periphery structure based on export data can explain some significant part of the dynamics of the specialization patterns.

The dynamics of the K-spec

Results regarding the effect of tariff changes on the dynamics of the K-spec indicator, using specifications (1) and (2), are reported in the first four columns of Table 10. The main result is that the evolution of the K-spec tariff has not been affected by tariff changes. In the first two columns, both the long- and the short-run effects (the parameters associated with the $TW_{i,t-1}$ and the $\Delta TW_{i,t}$ variables respectively) are not significant. This result does not change when the Grubel-Lloyd index is used in the interaction specification (2)—third and fourth columns. The null hypothesis of no effect is not even rejected with SUR estimations in all specifications, confirming the robustness of this finding. The analysis therefore suggests that the reduction in tariffs fostered by the FTA has not affected the relative degree of industrial

diversification of the Canadian provinces. From a Canadian economic policy perspective, one might consider this as relatively good news.

Interestingly, when the dependent variable is the change in the K-spec indicator (first four columns of Table 10), the estimated parameter of the misalignment variable always has a negative sign that is significant at the 1% level. Contrary to the effect of the misalignment variable in the Grubel-Lloyd specifications in Section 4.3 (not reported), the results obtained with some heterogeneous specification indicate that this effect is not driven by a subgroup of provinces such as the peripheral western provinces. Consequently, the analysis suggests that a high degree of misalignment helps alleviate the differences in the industrial structures of Canadian provinces.

Ratio of primary to manufacturing exports

One way to capture the core-periphery structure of the Canadian economy is to consider the ratio of primary to manufacturing exports per province (variable $RMP_{i,t}$ in Table 12) as in Beine and Coulombe (2003). We computed this variable from our SIC 4 industry database. The time averages over the 1980–2001 period for the 10 provinces are shown in Table 11. The numbers in this table illustrate well the huge cross-sectional variance of the $RMP_{i,t}$ variable. Provinces' averages range from peaks of 3.131 and 2.512 for Saskatchewan and Alberta respectively to lows of 0.013 and 0.041 for Ontario and Quebec.

With such heterogeneity across provinces, a key issue is verifying whether the $RMP_{i,t}$ variable can account for the dynamics of industrial diversification across the Canadian regional economies. The variables $RMP_{i,t-1}$ and $\Delta RMP_{i,t}$ were introduced as a Z' variable in equation (1) and the results of the estimations depicted in Table 11. The other Z' variables were eliminated from the exercise as they were not significant.

The results in the first two columns of Table 12 show a striking fit for the all-sector Herfindahl index. The long- and short-run effects (parameters of $RMP_{i,t-1}$ and $\Delta RMP_{i,t}$) are positive and significant at the 1% level with both PLS and FGLS as is the case with all other specifications and with SUR (non-reported results). Interestingly, the introduction of the $RMP_{i,t}$ illustrates the robustness of the variable of our main result regarding the long-run

effect of tariffs on industrial diversification: point estimates of the $TW_{i,t-1}$ variable remain significant at the 1% level (with the expected sign) in all specifications.

However, when the dependent variable is restricted to the Herfindahl index of the manufacturing industries, the long-run effect of the $RMP_{i,t-1}$ variable disappears completely (columns 3 and 4 of Table 11). It is important to point out here that these results should be interpreted with great care since the ratio of primary to manufacturing exports cannot be interpreted simply as an exogenous variable with respect to both the integration process and the level of industrial diversification. For this reason, despite the excellent fit, specifications using the $RMP_{i,t-1}$ variable were not considered as the “benchmark” specification for the all-sector Herfindahl model. Nevertheless, the analysis indicates that the ratio of primary to manufacturing exports is an important determinant of the degree of specialization as measured by the all-sector Herfindahl index. The long-run effect is positive, indicating that a high (low) level of primary exports coincides with industrial specialization (diversification). The effect is negligible, however, on the diversification of the manufacturing sector.

4.5 Results: interaction between indicators of industrial structure

We complement the analysis of the dynamic panel data model by focusing on interactions among the various industrial structures indicators used in the empirical analysis. The results for the interaction effects are in Tables 3, 6, and 10 in the columns referring to specifications (2). These estimations have already been discussed in the previous sections since they help illustrate the general robustness of our most significant findings. In this section, we focus on the point estimates of the θ_1 and θ_2 parameters in the general interaction specification (2), described in Section 4.1 above.

Only one long-run (parameter θ_1) interaction channel emerges as significant in a robust way: the Herfindahl index has a significant (at the 1% level) and positive effect with both PLS and FGLS on the K-spec indicator as shown in columns 3 and 4 of Table 10. The short-run effect is also positive and is significant at the 1% level. Thus the effect of the Herfindahl index on the K-spec appears in a contemporaneous and persistent way. 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 θ_1 parameter is not significant in the reversed interaction set-up (the last two columns of Table 6). These results confirm to some extent that causality runs from specialization structure to the core-periphery structure, and not the other way around.

Not surprisingly, given the share of Canada's total exports that go to the United States, the effect of the Herfindahl, measured for the rest of the world, on the Herfindahl measures for exports to the United States is negligible (see the first two columns of Table 3). Some weak evidence exists, however, that the reverse interaction is positive (last two columns of Table 6). This indicates that an increase in the diversification (or specialization) of export structures to the United States might generate a kind of spillover effect on the exports to the rest of the world. This effect, however, is not significant with PLS.

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 imply any specific test. In contrast, the question of 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 does not permit a straightforward answer. The issue is made even trickier since we are dealing with data that have been observed not only over time but also for various regions. Fortunately, since the beginning of the 1990s, econometric literature dealing with the issue of non-stationarity in panel data has emerged. In this section, we employed some of the most well-known econometric tools of this literature in order 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, i.e., follow non-stationary process in level, then estimation of ϕ_1 and ϕ_2 can suffer from some problems. As explained by Pedroni (2003), even with homogeneous panels for which the slope estimates are found to be consistent for the number of time periods and cross-sections, ignoring unit root properties of the variables will still produce problems in the inference procedure. 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 permit investigation 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.

5.1 Panel unit root tests

Since the seminal work of Levin and Lin (1993), several tests have been developed to test for the presence of a unit root in panel data. In this respect, we use three different test statistics, each based on the residuals $\hat{v}_{i,t}$ obtained from the two following generic regressions:

$$\Delta y_{i,t} = \alpha + \rho y_{i,t-1} + \sum_{j=1}^{p_i} \mu_{ij} \Delta y_{i,t-j} + v_{i,t} \quad (4)$$

$$\Delta y_{i,t} = \rho_i y_{i,t-1} + \sum_{j=1}^{p_i} \mu_{ij} \Delta y_{i,t-j} + v_{i,t} \quad (4')$$

where the $y_{i,t}$ is one of our interest variables. In contrast to model (4), which assumes some homogeneous dynamics across provinces, model (4') allows for some heterogeneous regression coefficients. Under the null hypothesis of a unit root, $H_0 : \rho = 0$ (while the alternative hypothesis of a stationary process) is written as: $H_0 : \rho < 0$. The three test statistics used to test for H_0 are the so-called Levin-Lin rho statistics ($z_{\hat{\rho}}$), the Levin-Lin t-rho statistics ($z_{t\hat{\rho}}$), and the Levin-Lin ADF statistics (t_{adf}).¹⁴ Given an appropriate adjustment in the computation of the statistics and the value of the variance of the limiting distribution, they are all asymptotically distributed as a standard normal distribution $N(0,1)$.

Given their definition, it is important to pay attention to the way our variables are expressed. Indeed, all the specialization measures and trade-weighted variables are between 0 and 1, which leads to conceptual problems for integrated processes. One way to overcome this

problem is to conduct the unit root tests on the logistic transformation of the $y_{i,t}$ variables:

$z_{i,t} = \ln\left[\frac{y_{i,t}}{1-y_{i,t}}\right]$. We conduct the unit root tests for three different maximum numbers

of lags: $p=0, 1$, and 2 . For each province, the number of lags is selected according to the significance of the μ_{ij} coefficients. In the end, our investigation procedure leads to eight different cases for each of the test statistics. These are reported in Tables 13 and 14 for the homogeneous and heterogeneous cases respectively.

On the whole, the tests do not allow us to draw a clear-cut picture of the statistical properties of the variables. In general, the homogeneous case leads to conclusions more in favour of the $I(1)$ process than the heterogeneous case. Tests conducted on the logistic transformation are slightly more prone to favour the presence of a unit root process, compared with the analysis using the original variables. In general, the two Herfindahl indexes are found to be stationary but there is still some evidence of unit roots in the homogeneous case. In contrast, the trade-weighted tariffs, for either all goods or for the manufacturing sector, appear to exhibit unit root properties. It is significant that in no case can we favour stationary processes for any variable with a high degree of certainty. This in turn calls for some cointegration analysis.

5.2 Cointegration tests

Panel cointegration tests permit proper testing for the significance of the coefficient linking variables $SH_{i,t}$ and $TW_{i,t}$ when they are best described by an $I(1)$ process. In this paper, we use the testing procedures proposed by Pedroni (1999; 2003). Like for the unit root test procedures, the tests allow for both homogeneous and heterogeneous cointegration relationships across provinces. The two static cointegrating relationships are based on the following panel data models:

$$SH_{i,t} = \gamma_i + \beta TW_{i,t} + \eta_{i,t} \quad (5)$$

$$SH_{i,t} = \gamma_i + \beta_i TW_{i,t} + \eta_{i,t} \quad (5')$$

¹⁴ The interested reader may look for more details in Levin and Lin (1993) and Pedroni (2003).

Model (5) captures some homogeneous dynamics while model (5') considers a full heterogeneous relationship across provinces. Pedroni (1999) develops seven different test statistics computed from auxiliary regressions involving the residuals $\hat{\eta}_{i,t}$. Basically, these procedures test for the presence of a unit root in the $\hat{\eta}_{i,t}$. As for the unit root tests described above, we consider three different maximum numbers of lags, i.e. 0, 1, and 2. We also allow for the definitions of the $SH_{i,t}$ and $TW_{i,t}$ both in levels and in logistic form. Given the number of test statistics, we report the conclusion only in terms of positive or negative evidence for a cointegrating relationship between the two variables. Positive (negative) evidence is considered when a majority of the seven test statistics proposed by Pedroni (1999) leads to a rejection (acceptation) of the null hypothesis of I(1) residuals $\hat{\eta}_{i,t}$ at a 5% significance level.¹⁵

Results displayed in Table 15 provide some strong evidence in favour of a cointegrating relationship between industrial diversification and trade-weighted tariffs. Such evidence tends to hold in both the homogeneous and the heterogeneous case. Therefore, these estimations provide robustness in favour of a long-run relationship between industrial diversification and economic integration.

5.3 Error correction modelling

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):

$$\Delta SH_{i,t} = \phi_3 \Delta TW_{i,t} + \phi_4 UC_t + \phi_5 PC_{i,t} + \phi_6 VL_t + \phi_7 MA_t + \delta_i + \lambda (ECM_{i,t-1}) + \varepsilon_{i,t} \quad (6)$$

in which $-\lambda$ will capture the adjustment speed toward the long-run equilibrium between the industrial structure SH and tariffs TW . In the homogeneous case that corresponds to our benchmark model (1), $ECM_{i,t} = SH_{i,t} - \alpha_i - \beta TW_{i,t}$ captures deviations from this long-run equilibrium. The β coefficient may be directly estimated by PLS, allowing for the presence of individual fixed effects α_i to rule out any heterogeneity bias. Interestingly, model (6) turns out to be a re-parameterization of the benchmark model (1) under the assumption of I(1)

¹⁵ The panel unit root tests and cointegration tests were carried out using RATS Codes written by Peter Pedroni. These are available on www.estima.com.

variables: indeed, it can be seen that $\lambda = \phi_1$ and $\phi_2 = -\lambda\beta$ and $c_i = \delta_i - \lambda\alpha_i$. Similarly, under the assumption of heterogeneity across provinces previously captured by model (3), the ECM specification is given by $ECM_{i,t} = SH_{i,t} - \alpha_i - \beta_i TW_{i,t}$.

Table (16) reports the estimates of the long-run relationship under both the homogeneity (β estimate) and heterogeneity (β_i estimates) assumptions, for all sectors and the manufacturing sectors.¹⁶ Estimates of β for the all-sector and the manufacturing sector cases in the homogeneous case can be compared with the estimates of the long-run effects in the benchmark model (1) captured by $(\hat{\phi}_2 / -\hat{\phi}_1)$. In this respect, the long-run estimates are strikingly in line with the ones implied by the benchmark model: depending on the specification, the estimates of $(\hat{\phi}_2 / -\hat{\phi}_1)$ range between 1.15 and 3.85 and between 1.49 and 4.14 for the all sectors and manufacturing cases respectively.

Table (17) reports the estimates of the ECM models (6). Overall, the estimations provide striking support for an ECM type of mechanism. For the all-sector case, the estimates of λ suggest that production structures take between about 2.5 years (heterogeneous specification) and about 4 years (homogeneous specification) to adjust to deviations to long-run equilibrium. Again, as in Section 4.2, the speed of adjustment is found to be higher in the case of manufacturing, ranging between about 2 years (heterogeneous specification) and about 3 years (homogeneous specification).

6. Conclusion

In this paper, we have investigated some of the consequences of trade integration between Canada and the United States for the 10 Canadian provinces. In particular, we have focused on the impact of Canada–U.S. economic integration on the dynamics of diversification of the provincial industrial structures. To do this, we have built various indexes of production diversification and indexes of the share of intra-industry trade using import and 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 permits the econometric analysis to combine

pure time-series and the cross-sectional information and hence to increase the estimation of the various effects of trade integration.

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

Our main result concerns the differing impact over time of the trade integration process on the evolution of absolute diversification of the provincial economic structures. Typically, we found that on average, trade integration slightly increases absolute industrial specialization in the short run but leads to more diversification in the long run. Such a dynamic process is consistent with a story of some initial relocation process for a given industry followed by the agglomeration of other industries in order to take benefit of the backward-forward linkages.

While interesting, such an interpretation should nevertheless be further tested by some specific analysis of the agglomeration processes. We see extensions in two different directions. First, one could replicate the econometric analysis relative to the Herfindahl indexes at a lower desegregation level (say at the SIC-3 level) in order to identify which sectors are directly involved in the process of long-run agglomeration. Second, it might be interesting to extend the analysis to the location of trade at the industry level. The point here would be to capture the short and the long-run effect of tariff changes the degree of dispersion

¹⁶ Of course, under the unit root hypothesis, the related t-statistics of the β and β_i do not follow the standard

of industries across the Canadian provinces. Both analyses could lead to a better understanding of the mechanisms at work in the agglomeration process. We leave this for future research.

Interestingly, trade integration was not found to increase the differences of production structure across the provinces. These findings are consistent with the predictions of some recent analysis belonging to the new “economic geography” approach of international trade popularized by Krugman (1991b). In particular, these are consistent with simulations of the Krugman and Venables (1995) model showing that the occurrence of very low trade costs (such as the ones implied by the FTA) lead in the long run to some dispersion of economic activities.

In Canada, more than in the United States, a substantial part of the risk-sharing between the regional economies is operated explicitly by the insurance mechanism (equalization and unemployment insurance programs) and implicitly through taxes and transfers by a sophisticated fiscal federalism system. From an economic policy perspective, one could have expected that the pressures on the interprovincial redistribution scheme would be higher to the extent that economic integration is thought to induce specialization at the regional level. Specialization of economic activities at the regional level, which might be necessary to capture the gains from trade, nevertheless increases the exposure of Canadian provinces to specific shocks and generates higher variability of regional output and employment. As our main result indicates, this story appears to be clearly rejected by the facts. The impact effect of economic integration on regional specialization is reversed in the long run. It suggests that through trade integration, Canadian provincial economies can capture the gains from economic integration without becoming more vulnerable.

Finally, as the evolution of the trade-weighted tariffs in Figure 4 clearly illustrates, most of the reductions in tariffs have 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 FTA and NAFTA was completed by the eve of the new millennium. We think however that the analysis is also of interest from a prospective

student-t distribution. We therefore report only the point estimates.

economic policy point of view since the process of Canada-U.S. economic integration is not completed. Canada as a whole still face important trade costs with the United States which can be reduced by policy actions. To start with, a certain number of sectors are still subject to non negligible tariffs. We found that, 89, 52 and 15 of the SIC 4 industries faced tariffs rates higher than 1, 2 and 5% respectively in 1996. This fact is an indication that many Canadian exporters prefer to pay the relevant duties rather than having to meet NAFTA rules of origin requirements. It also indicates that trade under NAFTA is costly because firms have to meet the requirements. Furthermore, the currency barrier generates substantial trade costs with the United States. The barrier entails for instance uncertainty costs due to exchange rate volatility and hedging costs for exporters and importers. Anderson and van Wincoop (2004) extensively discuss the computation of tariff equivalents of the reduction in trade barriers from adopting a common currency. While the levels might be highly variable for instance across pairs of countries and assumptions made in the computation, they argue that currency barriers amount on average to a 14% tariff equivalent. We think that our analysis could be used to assess the short and long-run effects on the industrial structure of policy actions aiming at decreasing trade costs with our most important trading partner. It is also of interest for those who want to assess the effect of not eliminating what remains of trade barriers between Canada and the United states.

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

Dependent variable: $\Delta SH_{i,t}$						
	(1) PLS	(1) FGLS	(1) SUR	(1) PLS	(1) FGLS	(1) SUR
$SH_{i,t-1}$	-0.271*** (0.092)	-0.342*** (0.072)	-0.280*** (0.040)	-0.266*** (0.095)	-0.348*** (0.073)	-0.275*** (0.037)
$TW_{i,t-1}$	0.933*** (0.331)	0.386*** (0.107)	0.591*** (0.141)	1.041*** (0.297)	0.404*** (0.101)	0.521*** (0.1516)
$\Delta TW_{i,t}$	-1.746 (1.105)	-1.107*** (0.352)	-2.283*** (0.294)	-1.938* (1.093)	-1.042*** (0.302)	-2.50*** (0.300)
UC_t	-0.623* (0.366)	-0.143 (0.160)	-0.514*** (0.128)	–	–	–
$PC_{i,t}$	0.503 (0.428)	0.045 (0.186)	0.360*** (0.116)	–	–	–
VL_t	-4.971* (2.583)	-0.486 (0.871)	-4.197*** (1.010)	–	–	–
MA_t	-0.012 (0.043)	-0.000 (0.014)	-0.009 (0.015)	–	–	–
R^2	.228	.300	.220	.195	.295	.181
Adj. R^2	.153	.23	.143	.137	.244	.122
S.E. of R.	.032	.034	.032	.033	.034	.033
observations	180	180	180	180	180	180

Notes: Sample: 1980–1998. The regressions include province fixed effects. Estimation techniques: pooled least-squares (PLS), iterated feasible generalized least-squares (FGLS), seemingly unrelated regression (SUR). Estimated equation is shown in parentheses before the estimation technique. 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: $\Delta SHM_{i,t}$						
	(1) PLS	(1) FGLS	(1) SUR	(1) PLS	(1) FGLS	(1) SUR
$SHM_{i,t-1}$	-0.335*** (0.079)	-0.369*** (0.074)	-0.358*** (0.040)	-0.321*** (0.080)	-0.385*** (0.076)	-0.322*** (0.032)
$TWM_{i,t-1}$	1.356*** (0.390)	0.580*** (0.146)	1.093*** (0.141)	1.272*** (0.388)	0.580*** (0.142)	1.006*** (0.137)
$\Delta TWM_{i,t}$	-1.455 (1.115)	-1.093** (0.518)	-1.961*** (0.256)	-1.234 (1.056)	-0.897* (0.462)	-1.742*** (0.214)
UC_t	0.179 (0.345)	-0.153 (0.174)	-0.128 (0.107)	–	–	–
VL_t	0.063 (3.122)	0.705 (1.240)	0.301 (1.030)	–	–	–
MA_t	0.098* (0.05)	0.020 (0.022)	0.074*** (0.017)	–	–	–
R ²	.254	.331	.241	.232	.315	.228
Adj. R ²	.170	.256	.157	.159	.259	.165
S.E. of R.	.030	.031	.031	.030	.031	.030
observations	160	160	160	160	160	160

Notes: Sample: 1980–1996. The regressions include province fixed effects. Estimation techniques: pooled least-squares (PLS), iterated feasible generalized least-squares (FGLS), seemingly unrelated regression (SUR). Estimated equation is shown in parentheses before the estimation technique. 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 3. Dynamics of industrial diversification – Herfindahl (all sectors): Interaction with Herfindahl for exports to the rest of the world and Grubel-Lloyd

Dependent variable: $\Delta SH_{i,t}$				
	(2) PLS	(2) FGLS	(2) PLS	(2) FGLS
$SH_{i,t-1}$	-0.267*** (0.093)	-0.348*** (0.072)	-0.268*** (0.089)	-0.346*** (0.074)
$TW_{i,t-1}$	0.912*** (0.321)	0.393*** (0.105)	0.862** (0.372)	0.436*** (0.131)
$\Delta TW_{i,t}$	-1.829 (1.140)	-0.991*** (0.361)	-1.702 (1.083)	-1.010*** (0.366)
$SHR_{i,t-1}$	-0.061 (0.064)	0.006 (0.016)		
$\Delta SHR_{i,t}$	0.008 (0.063)	-0.016 (0.015)		
$SG_{i,t-1}$			-0.066 (0.058)	0.013 (0.019)
$\Delta SG_{i,t}$			-0.256*** (0.086)	0.000 (0.038)
VL_t	-4.839* (2.586)	-0.609 (0.888)	-5.055** (2.529)	-0.612 (0.891)
MA_t	-0.021 (0.048)	0.000 (0.015)	-0.004 (0.044)	0.002 (0.014)
R ²	.239	.311	.260	.305
Adj. R ²	.154	.234	.177	.227
S.E. of R.	.032	.034	.032	0.035
observations	180	180	180	180

Notes: Sample: 1980–1998. The regressions include province fixed effects. Estimation techniques: pooled least-squares (PLS), iterated feasible generalized least-squares (FGLS), seemingly unrelated regression (SUR). Estimated equation is shown in parentheses before the estimation technique. 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 4. Dynamics of industrial diversification – Herfindahl (all sectors), heterogeneous model

Dependent variable: $\Delta SH_{i,t}$						
	(3)PLS		(3)FGLS		(3)SUR	
$SH_{i,t-1}$	-0.359***	(0.147)	-0.442***	(0.084)	-0.402***	(0.052)
$TW - AB_{t-1}$	5.814**	(2.838)	6.593**	(2.613)	5.153**	(2.286)
$TW - BC_{t-1}$	0.278	(1.028)	0.235	(1.018)	0.273	(0.940)
$TW - MB_{t-1}$	0.298**	(0.134)	0.315**	(0.130)	0.301**	(0.021)
$TW - NB_{t-1}$	0.459	(0.923)	0.490	(0.920)	0.513	(0.664)
$TW - NF_{t-1}$	3.084**	(1.300)	3.688***	(1.331)	2.828**	(1.430)
$TW - NS_{t-1}$	0.831*	(0.433)	0.956**	(0.374)	0.748**	(0.336)
$TW - ON_{t-1}$	0.001	(0.515)	-0.170	(0.439)	-0.031	(0.285)
$TW - PE_{t-1}$	0.856	(0.728)	1.058	(0.668)	1.441**	(0.612)
$TW - QC_{t-1}$	0.453***	(0.167)	0.521***	(0.134)	0.521***	(0.136)
$TW - SK_{t-1}$	4.633	(3.016)	5.564*	(2.894)	4.589**	(2.305)
Mean $TW - PR_{t-1}$	1.671**	(0.657)	1.925**	(0.773)	1.633**	(0.060)
Median $TW - PR_{t-1}$	0.645		0.739		0.635	
$\Delta TW - AB_t$	-15.170**	(7.197)	-13.402**	(6.715)	-13.871***	(4.626)
$\Delta TW - BC_t$	-5.001	(3.372)	-4.581	(3.247)	-5.201**	(2.309)
$\Delta TW - MB_t$	-1.055**	(0.469)	-1.130	(0.437)	1.109**	(0.355)
$\Delta TW - NB_t$	-0.391	(0,858)	-0.307	(0.696)	-0.693	(0.996)
$\Delta TW - NF_t$	-2.132	(4.431)	-1.575	(4.757)	-1.665	(5.162)
$\Delta TW - NS_t$	0.347	(1.174)	0.387	(1.134)	0.516	(0.855)
$\Delta TW - ON_t$	-3.972**	(1.708)	-3.991**	(1.740)	-4.034***	(1.049)
$\Delta TW - PE_t$	0.994	(4.080)	0.583	(3.858)	0.775	(1.228)
$\Delta TW - QC_t$	-0.723	(0.613)	-0.763	(0.553)	-1.516***	(0.560)
$\Delta TW - SK_t$	-1.730	(12.749)	2.339	(11.448)	-1.530	(4.519)
Mean $\Delta TW - PR_t$	-2.884*	(1.484)	-2.244	(1.399)	-2.833*	(1.357)
Median $\Delta TW - PR_t$	-1.392		-0.947		-1.522	
R ²	.308		.395		.298	
Adj. R ²	.169		.273		.157	
S.E. of R.	.032		.032		.032	
observations	180		180		180	

Notes: Sample: 1980–1998. The regressions include province fixed effects. Estimation techniques: pooled least-squares (PLS), iterated feasible generalized least-squares (FGLS), seemingly unrelated regression (SUR). Estimated equation is shown in parentheses before estimation techniques. White heteroscedasticity standard errors are shown in parentheses on the right of the estimated coefficients. Asymptotic standard error reported for Mean (PMG estimator) * = significant at 10% level; ** = significant at 5% level; *** = significant at 1% level. Provinces are Alberta (AB), British Columbia (BC), Manitoba (MB), New Brunswick (NB), Newfoundland and Labrador (NF), Nova Scotia (NS), Ontario (ON), Prince Edward Island (PE), Quebec (QC), and Saskatchewan (SK).

Table 5. Dynamics of industrial diversification – Herfindahl (all sectors), heterogeneous model: LSDV (1) versus PMG (3') estimators

	Dependent variable: $\Delta SH_{i,t}$	
	(1) FGLS LSDV	(3')FGLS PMG
$SH_{i,t-1}$	-0.348*** (0.073)	-0.401*** (0.082)
Median $SH_{i,t-1}$	–	-0.432
$TW_{i,t-1}$	0.404*** (0.101)	0.560*** (0.121)
$\Delta TW_{i,t}$	-1.042*** (0.302)	-6.091** (2.671)
Median $\Delta TW_{i,t}$	–	-2.490
Long-run effect of $TW_{i,t-1}$	1.161***	1.397***

Notes: Sample: 1980-1998. The regressions include province fixed effects. Estimation technique, iterated feasible generalized least-squares (FGLS). Estimated equation is shown in parentheses before estimation techniques. White heteroscedasticity standard errors are shown in parentheses on the right of the estimated coefficients. Asymptotic standard error reported for Mean (PMG estimator) * = significant at 10% level; ** = significant at 5% level; *** = significant at 1% level.

Table 6. Dynamics of industrial diversification – Herfindahl (all sectors, rest of the world)

Dependent variable $\Delta SHR_{i,t}$				
	(1) PLS	(1) FGLS	(2) PLS	(2) FGLS
$SHR_{i,t-1}$	-0.503*** (0.121)	-0.281*** (0.051)	-0.505*** (0.124)	-0.304*** (0.057)
$TW_{i,t-1}$	0.635 (0.458)	0.072 (0.218)	0.456 (0.457)	-0.023 (0.222)
$\Delta TW_{i,t}$	2.640* (1.359)	1.564* (0.919)	2.562* (1.386)	1.501 (0.927)
$SH_{i,t-1}$	–	–	0.069 (0.132)	0.101** (0.043)
$\Delta SH_{i,t-1}$	–	–	0.019 (0.150)	0.102* (0.060)
UC_t	-0.414 (0.544)	-0.176 (0.223)	-0.403 (0.756)	-0.111 (0.371)
VL_t	3.690 (3.735)	3.249* (1.939)	4.179 (4.045)	3.582* (1.864)
MA_t	-0.087* (0.049)	-0.063* (0.030)	-0.088* (0.048)	-0.052* (0.030)
R^2	.293	.263	.296	.277
Adj. R^2	.230	.196	.217	.197
S.E. of R.	.049	.051	.050	.051
observations	180	180	180	180

Notes: Sample: 1980-1998. The regressions include province fixed effects. Estimation techniques: pooled least-squares (PLS), iterated feasible generalized least-squares (FGLS). Estimated equation is shown in parentheses before estimation techniques. 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 7. Dynamics of intra-industry trade – Grubel-Lloyd (all sectors)

	Dependent Variable $\Delta SG_{i,t}$			
	(1) PLS	(1) FGLS	(2) PLS	(2) FGLS
$SG_{i,t-1}$	-0.260*** (0.067)	-0.205*** (0.061)	-0.258*** (.067)	-0.231*** (0.062)
$TW_{i,t-1}$	-0.244 (0.318)	-0.240 (0.291)	-0.121 (0.329)	-0.108 (0.289)
$\Delta TW_{i,t}$	0.190 (0.972)	-0.119 (0.837)	-0.106 (0.923)	-0.436 (0.795)
$SH_{i,t-1}$	–	–	-0.031 (0.042)	-0.053* (0.030)
$\Delta SH_{i,t}$	–	–	-0.161*** (0.053)	-0.145*** (0.042)
UC_t	0.070 (0.297)	0.096 (0.227)	-0.033 (0.281)	-0.069 (0.206)
VL_t	-0.401 (2.30)	-1.334 (1.952)	-1.113 (2.25)	-2.471 (1.921)
MA_t	0.029 (0.034)	0.066** (0.028)	0.027 (.034)	0.079** (0.028)
R^2	.145	.156	.181	.204
Adj. R^2	.061	.073	.089	.115
S.E. of R.	.026	.026	.025	.025
observations	180	180	180	180

Notes: Sample: 1980-1998. The regressions include province fixed effects. Estimation techniques: pooled least-squares (PLS), iterated feasible generalized least-squares (FGLS). Estimated equation is shown in parentheses before estimation techniques. 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 8. Dynamics of intra-industry trade – Grubel-Lloyd (manufacturing sector)

	Dependent Variable $\Delta SG_{i,t}$		
	(1) PLS	(1) FGLS	(1) SUR
$SG_{i,t-1}$	-0.352*** (0.066)	-0.358*** (0.061)	-0.388*** (0.054)
$TW_{i,t-1}$	-0.587 (0.397)	-0.772** (0.349)	-0.716*** (0.227)
$\Delta TW_{i,t}$	0.349 1.134	-0.185 (0.895)	0.085 (0.571)
UC_t	0.163 (0.360)	0.197 (0.300)	0.284 (0.161)
VL_t	-1.169 (3.143)	-2.444 (2.666)	-1.502 (1.422)
MA_t	-0.000 (0.049)	0.017 (0.040)	0.006 (0.024)
R^2	.195	.223	.193
Adj. R^2	.105	.136	.102
S.E. of R.	.032	.032	.032
observations	160	160	160

Notes: Sample: 1980-1998. The regressions include province fixed effects. Estimation techniques: pooled least-squares (PLS), iterated feasible generalized least-squares (FGLS), seemingly unrelated regression (SUR). Estimated equation is shown in parentheses before estimation techniques. 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 9. Dynamics of intra-industry trade – Grubel-Lloyd (all sectors), heterogeneous model

Dependent variable: $\Delta SG_{i,t}$						
	(3)PLS		(3)FGLS		(3)SUR	
$SG_{i,t-1}$	-0.315***	(0.064)	-0.275***	(0.055)	-0.345***	(0.046)
$TW_{i,t-1}$	-0.431	(0.313)	-0.480	(0.291)	-0.459***	(0.130)
$\Delta TW_{i,t}$	0.123	(0.964)	-0.127	(0.756)	0.015	(0.404)
$MA - AB_t$	0.238**	(0.103)	0.234**	(0.102)	0.241***	(0.080)
$MA - BC_t$	0.195***	(0.060)	0.187***	(0.053)	0.201***	(0.053)
$MA - MB_t$	-0.002	(0.133)	-0.002	(0.134)	-0.002	(0.120)
$MA - NB_t$	-0.018	(0.076)	-0.027	(0.076)	-0.015	(0.098)
$MA - NF_t$	-0.081	(0.065)	-0.078	(0.068)	-0.081	(0.069)
$MA - NS_t$	-0.144	(0.096)	-0.143	(0.098)	-0.145*	(0.096)
$MA - ON_t$	0.106	(0.084)	0.103	(0.088)	0.109	(0.095)
$MA - PE_t$	0.011	(0.074)	0.008	(0.071)	0.013	(0.088)
$MA - QC_t$	-0.174	(0.108)	-0.163	(0.103)	-0.181**	(0.083)
$MA - SK_t$	0.112***	(0.039)	0.105***	(0.037)	0.117***	(0.045)
R^2	.242		.266		.240	
Adj. R^2	.135		.163		.134	
S.E. of R.	.025		.025		.025	
observations	180		180		180	

Notes: Sample: 1980-1998. The regressions include province fixed effects. Estimation techniques: pooled least-squares (PLS), iterated feasible generalized least-squares (FGLS), seemingly unrelated regression (SUR). Estimated equation is shown in parentheses before estimation techniques. 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 10. Dynamic of industrial diversification, K-spec (all sectors)

Dependent variable →	$\Delta SK_{i,t}$				$\Delta SH_{i,t}$	
	(1) PLS	(1) FGLS	(2) PLS	(2) FGLS	(2) PLS	(2) FGLS
$SK_{i,t-1}$	-0.630*** (0.115)	-0.401*** (0.066)	-0.631*** (0.117)	-0.389*** (0.062)	0.005 (0.070)	-0.020 (0.025)
$\Delta SK_{i,t}$	—	—	—	—	0.141** (0.069)	0.002 (0.022)
$SH_{i,t-1}$	—	—	0.100*** (0.035)	0.100*** (0.024)	-0.273*** (0.093)	-0.331*** (0.073)
$\Delta SH_{i,t-1}$	—	—	0.134*** (0.050)	0.092*** (0.035)	—	—
$TW_{i,t-1}$	-0.177 (0.307)	0.110 (0.219)	-0.465 (0.351)	-0.369 (0.234)	0.874** (0.351)	0.337*** (0.124)
$\Delta TW_{i,t}$	-0.659 (0.912)	0.060 (0.521)	-0.536 (0.910)	-0.133 (0.505)	-1.900* (1.127)	-1.283*** (0.393)
UC_t	-0.027 (0.303)	-0.065 (0.168)	0.048 (0.307)	.024 (0.151)	-0.612 (0.357)	-0.139 (0.179)
VL_t	2.957 (2.740)	0.722 (1.733)	4.002 (2.756)	1.936 (1.518)	-5.268** (2.590)	-0.389 (0.943)
MA_t	-0.124*** (0.041)	-0.090*** (0.028)	-0.125*** (0.040)	-0.089*** (0.024)	-0.004 (0.047)	-0.004 (0.015)
R ²	.329	.253	.349	.323	.248	.314
Adj. R ²	.263	.180	.276	.247	.165	.237
S.E. of R.	.032	.033	.031	.033	.032	.034
observations	180	180	180	180	180	180

Notes: Sample: 1980-1998. The regressions include province fixed effects. Estimation techniques: pooled least-squares (PLS), iterated feasible generalized least-squares (FGLS). Estimated equation is shown in parentheses before estimation techniques. 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 11. Ratio of primary to manufacturing exports
Mean 1980–2001**

Alberta	2.512
British Columbia	0.136
Manitoba	0.258
New Brunswick	0.099
Newfoundland and Labrador	0.240
Nova Scotia	0.291
Ontario	0.013
Prince Edward Island	0.457
Quebec	0.041
Saskatchewan	3.131

Primary exports refer to SIC 4 industries with a 0 as the first number. Manufacturing refers to SIC 4 industries with either 1, 2, or 3 as the first number.

TABLE 12: Dynamics of Industrial Diversification – Herfindahl Interaction with Ratio of Primary to Manufacturing Exports $RMP_{i,t}$

Dependant Variable →	$\Delta SH_{i,t}$		$\Delta SHM_{i,t}$	
	(1) PLS	(1) FGLS	(1) PLS	(1) FGLS
$SH_{i,t-1}$	-0.324** (0.131)	-0.377*** (0.078)	–	–
$SHM_{i,t-1}$	–	–	-0.318*** (0.081)	-0.376*** (0.077)
$TW_{i,t-1}$	1.158*** (0.336)	0.476*** (0.113)	1.249*** (0.390)	0.573*** (0.143)
$\Delta TW_{i,t}$	-1.705 (1.010)	-0.968*** (0.365)	-1.497 (1.078)	-1.087** (0.446)
$RMP_{i,t-1}$	0.023*** (0.008)	0.029*** (0.006)	-0.001 (0.004)	0.000 (0.003)
$\Delta RMP_{i,t}$	0.040*** (0.008)	0.048*** (0.007)	-0.012** (0.005)	-0.009** (0.005)
R ²	.300	.474	.243	.339
Adj. R ²	.240	.429	.169	.275
S.E. of R.	.031	.032	.030	.031
observations	180	180	160	160

Notes: Sample: 1980-1998 (first two columns), 1980-1996 (last two columns). The regressions include province fixed effects. Estimation techniques: Pooled least squares (PLS), Iterated Feasible Generalized Least Squares (FGLS). Estimated equation is shown in parenthesis before estimation techniques. White heteroscedasticity standard errors are shown in parentheses below the estimated coefficients. *: significant at 10% level; ** at 5% level; *** at 1% level.

Table 13. Panel unit root tests - homogeneous case

Variable		Herf		Herf Man		TW tariffs		TW tariffs Man	
	Stat	$y_{i,t}$	$z_{i,t}$	$y_{i,t}$	$z_{i,t}$	$y_{i,t}$	$z_{i,t}$	$y_{i,t}$	$z_{i,t}$
p=0	$z_{\hat{\rho}}$	-2.38***	-1.87**	-2.34***	-1.87**	1.79	3.31	1.73	4.21
	$z_{t\hat{\rho}}$	-0.54	-0.01	-0.87	-0.01	1.89	5.09	2.18	6.56
	t_{adf}	-0.40	0.10	-0.38	0.10	1.96	5.07	2.25	6.49
p=1	$z_{\hat{\rho}}$	-1.87**	-1.25	-2.79***	-1.25	1.63	3.01	1.57	3.73
	$z_{t\hat{\rho}}$	-0.23	0.37	-1.16	0.37	1.88	4.41	2.14	5.47
	t_{adf}	0.09	0.63	-0.99	0.63	1.98	4.40	2.38	4.57
p=2	$z_{\hat{\rho}}$	-1.82**	-1.17	-2.36***	-1.17	1.63	2.78	1.58	3.73
	$z_{t\hat{\rho}}$	-0.14	0.47	-0.86	0.47	1.88	3.98	2.15	5.47
	t_{adf}	-0.51	1.10	0.01	1.10	1.98	2.50	2.38	4.57

*, **, *** denote rejection of the null hypothesis of a unit root respectively at the 10%, 5%, and 1%

Table 14. Panel unit root tests – heterogeneous case

Variable		Herf		Herf Man		TW tariffs		TW tariffs Man	
	Stat	$y_{i,t}$	$z_{i,t}$	$y_{i,t}$	$z_{i,t}$	$y_{i,t}$	$z_{i,t}$	$y_{i,t}$	$z_{i,t}$
p=0	$z_{\hat{\rho}}$	-5.35***	-5.29***	-4.17***	-2.98***	2.20	4.95	2.04	6.45
	$z_{t\hat{\rho}}$	-2.37***	-2.19**	-2.04**	-1.61*	-0.07	1.10	-0.09	2.40
	t_{adf}	-2.16**	-1.99**	-1.99**	-1.40*	0.06	1.21	-0.06	2.48
p=1	$z_{\hat{\rho}}$	-5.34***	-5.11***	-3.56***	-2.98***	1.24	3.99	-1.65	5.72
	$z_{t\hat{\rho}}$	-2.37***	-2.15**	-1.92**	-1.61*	-0.30	0.82	-0.17	2.09
	t_{adf}	-1.85**	-1.74**	-1.67**	-1.40*	-0.73	0.82	-0.39	2.13
p=2	$z_{\hat{\rho}}$	-6.24***	-5.51***	-3.50***	-3.46***	0.88	3.39	-1.67	5.72
	$z_{t\hat{\rho}}$	-2.53***	-2.21**	-1.91***	-1.70**	-0.37	0.63	-0.15	2.09
	t_{adf}	-2.16**	-1.72**	-1.70**	-1.54*	-0.77	0.39	-0.29	2.13

*, **, *** denote rejection of the null hypothesis of a unit root respectively at the 10%, 5%, and 1%

Table 15. Testing for a cointegrating relationship between $S_{i,t}$ and $TW_{i,t}$

Lags	Homogeneous panel		Heterogeneous panel	
	All sectors	Manufacturing sectors	All sectors	Manufacturing sectors
Level				
p=0	5/7	5/7	5/7	4/7
p=1	6/7	5/7	3/7	4/7
p=2	6/7	5/7	3/7	4/7
Logistic transformation				
p=0	6/7	5/7	5/7	5/7
p=1	6/7	5/7	5/7	5/7
p=2	6/7	5/7	5/7	5/7

Notes: The figures report the proportion of conclusions in favour of a cointegrating relationship drawn from the seven different tests proposed by Pedroni using a 5% significance level.

Table 16. Estimates of long-run relationship between industrial diversification and tariffs under the cointegration hypothesis

Homogeneous case			Heterogeneous case		
	All sectors	Manufacturing sectors		All sectors	Manufacturing sectors
β	2.66	2.80	β_{AB}	11.17	0.22
			β_{BC}	-0.04	-0.79
			β_{MB}	0.24	0.66
			β_{NB}	0.78	0.51
			β_{NF}	6.56	6.50
			β_{NS}	1.64	1.09
			β_{ON}	-1.80	-2.13
			β_{PE}	2.58	8.74
			β_{QC}	0.90	1.03
			β_{SK}	14.53	8.04

Notes: The coefficients are obtained from bivariate regressions using pooled OLS with fixed effects (homogeneous case) and OLS (heterogeneous case).

Table 17. Error correction models – Herfindahl (all sectors)

Dependent variable: $\Delta SH_{i,t}$						
	Homogeneous case			Heterogeneous case		
$ECM_{i,t-1}$	-0.27*** (0.05)	-0.28*** (0.05)	-0.28*** (0.05)	-0.41*** (0.06)	-0.41*** (0.06)	-0.41*** (0.06)
$\Delta TW_{i,t}$	-1.90* (1.06)	-1.94* (1.04)	-2.10** (1.03)	-1.30 (1.02)	-1.35 (0.99)	-1.50 (0.99)
VL_t	-5.11* (2.85)	-5.05* (2.85)	-5.61** (2.82)	-6.09** (2.71)	-6.09** (2.70)	-6.64** (2.68)
UC_t	-0.64* (0.35)	-0.44* (0.35)	-	-0.57* (0.34)	-0.44 (0.31)	-
$PC_{i,t}$	0.52 (0.39)	-	-	0.35 (0.37)	-	-
MA_t	-0.019 (0.038)	-	-	-0.018 (0.036)	-	-
R^2	.23	.22	.21	.30	.30	.29
Adj. R^2	.16	.16	.15	.24	.24	.24
S.E. of R.	0.032	0.032	0.032	0.030	0.031	0.031
observations	180	180	180	180	180	180

Notes: Sample: 1980–1998. The regressions include provincial fixed effects. Estimation techniques: pooled least-squares (PLS). 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 18. Error correction models – Herfindahl (manufacturing sectors)

Dependent variable: $\Delta SH_{i,t}$				
	Homogeneous case		Heterogeneous case	
$ECM_{i,t-1}$	-0.33*** (0.05)	-0.33*** (0.05)	-0.49*** (0.06)	-0.48*** (0.06)
$\Delta TW_{i,t}$	-1.76* (1.01)	-1.80* (1.00)	-1.17 (0.95)	-1.28 (0.91)
MA_t	0.09* (0.05)	0.09* (0.05)	0.09* (0.05)	0.08* (0.04)
VL_t	0.05 (3.03)	-	-1.05 (2.83)	-
UC_t	0.16 (0.34)	-	0.12 (0.31)	-
$PC_{i,t}$	-0.32 (0.38)	-	-0.35 (0.35)	-
R^2	.25	.21	.34	.33
Adj. R^2	.17	.15	.28	.28
S.E. of R.	0.030	0.032	0.028	0.028
observations	160	160	160	160

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

Appendix A

Results using trade-weighted tariffs with constant weights

As discussed in Section 3.2, 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 (see Section 3.2) of exports over the 1980–2001 period. The results presented in Table A1 below clearly indicate 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. The short-run effect exhibits the same negative sign but the estimate is not significant at the 10% level with PLS and SUR.

Table A1. Dynamics of industrial diversification – Herfindahl (all sectors, trade-weighted tariffs with constant weights)

Dependent variable: $\Delta SH_{i,t}$						
	(1) PLS	(1) FGLS	(1) SUR	(1) PLS	(1) FGLS	(1) SUR
$SH_{i,t-1}$	-0.248*** (0.090)	-0.330*** (0.070)	-0.313*** (0.046)	-0.242*** (0.093)	-0.318*** (0.073)	-0.339*** (0.044)
$TW_{i,t-1}$	0.515** (0.238)	0.230*** (0.083)	0.448*** (0.105)	0.647*** (0.213)	0.279*** (0.078)	0.386*** (0.110)
$\Delta TW_{i,t}$	-3.401 (2.792)	-1.214 (0.834)	-2.602*** (0.675)	-3.711* (2.387)	0.281 (0.619)	-3.501*** (0.612)
UC_t	-0.549 (0.379)	-0.125 (0.183)	-0.459*** (0.117)	–	–	–
$PC_{i,t}$	0.548 (0.437)	0.016 (0.206)	0.301** (0.131)	–	–	–
VL_t	-5.469** (2.602)	-0.283 (1.059)	-4.196*** (0.947)	–	–	–
MA_t	-0.014 (0.048)	-0.037* (0.019)	-0.012 (0.014)	–	–	–
R^2	.205	.300	.190	.171	.220	.141
Adj. R^2	.127	.23	.111	.111	.164	.079
S.E. of R.	.033	.034	.033	.033	.034	.034
observations	180	180	180	180	180	180

Notes: Sample: 1980–1998. The regressions include province fixed effects. Estimation techniques: pooled least-squares (PLS), iterated feasible generalized least-squares (FGLS), seemingly unrelated regression (SUR). Estimated equation is shown in parentheses before estimation techniques. White heteroscedasticity standard errors are shown in parentheses below the estimated coefficients. * = significant at 10% level; ** = significant at 5% level; *** = significant at 1% level.

Appendix B

Robustness analysis

To further assess the sensitivity of our results, we estimated the benchmark equation (1) including the full set of independent variables but removing from the sample a different province each time. Such an analysis aims at checking 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 are presented in Table B1. The Z' variables were not significant and results are not reported for them.

The estimates of the speed of adjustment are remarkably stable and are always significant at the 1% level. The measures of the long-run effect of trade-weighted tariffs are also stable and significant at the 1% level in nine cases. The only exception is when Quebec is removed from the estimation. In this case, however, the p -value of the long-run effect amounts to .0144. The short-run effect of tariff changes is also significant at the 1% level in nine cases out of ten. When Manitoba is removed from the estimation, the p -value is 0.134.

This exercise illustrates strikingly that the main result of our empirical analysis dealing with the long-run diversification effect of economic integration is not driven by the particular evolution of one province. For example, as shown and discussed in Section 3.1, the decrease in the Herfindhal index for Newfoundland since 1987 should be partly attributed to the decline of Atlantic fish stocks. This phenomenon of course is not related to the FTA. However, removing Newfoundland from the sample does not change the general direction of our results.

Finally, the exercise to some extent also illustrates the fact that the result dealing with the short-run impact of economic integration on specialization is not as robust as the long-run result. The precision of the estimation substantially decreases when Manitoba, one of the most diversified provinces, is removed from the sample.

Table B1. Dynamics of industrial diversification – Herfindahl (all sectors)

Dependent variable: $\Delta SH_{i,t}$			
Province removed from regression	Speed of adjustment: $-\hat{\phi}_1$	Long-run effect of TW : $\begin{pmatrix} -\hat{\phi}_2 \\ \hat{\phi}_1 \end{pmatrix}$	Short-run effect of TW : $\hat{\phi}_3$
Alberta	0.350*** (0.079)	1.111***	-1.085*** (0.348)
British Columbia	0.327*** (0.073)	1.192***	-1.075*** (0.359)
Manitoba	0.324*** (0.075)	1.231***	-1.075 (0.558)
New Brunswick	0.331*** (0.073)	1.184***	-1.222*** (0.360)
Newfoundland	0.357*** (0.077)	1.059***	-1.061*** (0.350)
Nova Scotia	0.324*** (0.071)	1.219***	-1.338*** (0.335)
Ontario	0.376*** (0.066)	1.098***	-0.939*** (0.358)
PEI	0.327*** (0.074)	1.144***	-1.169*** (0.365)
Quebec	0.315*** (0.071)	1.131**	-1.174*** (0.407)
Saskatchewan	0.405*** (0.085)	1.007***	-1.079*** (0.326)

Notes: Sample: 1980–1998. Iterated feasible generalized least-squares estimation of equation (1) with 162 observations. The regressions include province fixed effects. White heteroscedasticity standard errors are shown in parentheses below of the estimated coefficients. * = significant at 10% level; ** = significant at 5% level; *** = significant at 1% level.

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