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# Economic integration and the diversification of regional exports: evidence from the Canadian–U.S. Free Trade Agreement

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## Abstract

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

**Keywords:** Regional specialization, Free Trade Agreement, core-periphery structure, panel data
**JEL classifications:** F14, F15, R12

## 1. Introduction

The 1989 Canadian–U.S. Free Trade Agreement (CUSFTA) is considered one of the most spectacular examples of a trade integration process. It is well known that Canada's trade soared after CUSFTA was implemented. Indeed, Coulombe (2004) reports that the Canadian international trade share to GDP was relatively constant between 1980 and 1990, but between 1990 and 2000, the international trade share to GDP increased steadily from 0.51 to 0.86. Despite this, CUSFTA has raised many concerns in official Canadian circles. One key issue is the significant short-run adjustment costs in terms of job losses and worker displacement (Trefler, 2004). Another concern was that Canada could become a large peripheral region of the United States, specializing mainly in the production of primary products.

This last concern is related to Krugman's hypothesis of a positive link between trade integration and economic specialization. The major aim of this article is to use the CUSFTA experience to provide an empirical assessment of this hypothesis concerning the export structure. We show that while such a hypothesis had grounds in theory, there is no evidence that Canadian regions have moved to greater specialization regarding their export structure. Export data are preferred here over output or employment data for statistical purposes. While the dynamics of export might be different from the evolution of output, Canada is a small open economy with a high share of

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tradable sectors. This study should be seen as one of the first econometric analysis testing for such a relationship, in both the short and long run.

At a theoretical level, the nature of the relationship between trade integration and economic specialization is not obvious for, basically, two reasons. First, new trade theories that paid extensive attention to the agglomeration process of economic activities remain undecided regarding consequences of decreased trade costs. Early contributions (Krugman, 1991a; Krugman and Venables, 1995) emphasize lower transport costs as a strong agglomerating force. But recent contributions (Puga, 1999, for example) show that a significant decrease in these costs can result in further geographical dispersion of economic activities when labor mobility across sectors exhibits a finite cost [rather than an infinite cost as assumed in Krugman (1991b) for instance]. As underlined by Fujita and Thisse (1997), a non-linear relationship emerges between trade costs and agglomeration. Second, the concentration of activities and specialization are two different concepts.[1]

While most theoretical studies focus on the relationship between trade costs and production specialization, we focus here on export specialization, as done by a significant subset of empirical studies. A major reason of this choice is that export and import data are of much better quality than production data and display much less measurement errors. Conceptually, production and exports at the sector level can be expected to display different dynamics as a result of major shocks like the CUSFTA. Nevertheless, since Canada is a small open economy with 55% of manufacturing shipments exported in 2000, Canadian production and exports should be expected to go more hand in hand. Furthermore, focusing on export heterogeneity allows assessing the degree of exposition to external shocks faced by Canada as a whole and by its constituent regions. Once more, given the high degree of openness, this assessment is of overwhelming importance.[2]

A set of empirical studies focused on the nature of the relationship between trade integration and economic specialization, using the European experience or focusing on North American trade linkages.[3] The empirical approaches display a high degree of heterogeneity. They greatly vary according to the type of specialization indices (relative versus absolute specialization), the data sources (trade versus production or employment), and the level of sectoral desegregation. Unsurprisingly, the results are very mixed with papers documenting some increase in specialization (Amiti, 1998) and others finding relative stability over time (Sapir, 1996; Head and Ries, 2001) or

1 The nature of the link between agglomeration and specialization is non-linear and might vary over time. For example, a decrease in trade costs can initially induce a particular industry to locate in a particular region or to concentrate its activities there. Second, backward/forward linkages may trigger a traditional agglomeration process. As Krugman (1991a) explains, as both consumption and production prefer a variety of goods, there is a circular causality in the formation of clusters of firms and workers. New supply-side variety 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.

2 Importantly, the foreign demand shocks are typical purely exogenous disturbances in the sense that the Canadian economic policy has no mean to respond to them or even to forecast them. A good example is provided by the security measures undergone by the US government after the 11 September 2001 events.

3 Examples of European studies include Amiti (1998, 1999) and Sapir (1996). Studies on North American integration process include Harris (2001), Head and Ries (2001), and Sawchuk and Sydor (2001).

moderate evidence in favor of diversification (Sawchuk and Sydor, 2001). However, one main common feature of these empirical studies is their implicit assumptions about the process of integration. They either assume that integration follows a linear process and therefore investigate the evolution of specialization over time or they follow a one-shot process by capturing trade integration using time dummy variables [Acharya et al. (2001), for instance]. In contrast to the existing literature, we build an explicit measure of trade integration, relying on the dynamics of trade-weighted tariffs and estimate the short- and the long-run impact of this measure on the dynamics of export specialization in Canada. These trade tariffs have been lowered progressively since 1980, suggesting that trade integration had taken place well before the CUSFTA. The evolution of these tariffs also suggests that integration is far from being a linear process that applies equally to all regions.

Besides explicitly measuring the trade integration process, our investigation contributes in other ways to the empirical literature on integration and specialization. First, as Trefler (2004) emphasized, the CUSFTA is a 'clean' policy experiment that allows the construction of an exogenous measure of trade integration. This feature has been called for extensively in the literature (e.g., Rodriguez and Rodrik, 1999) to permit meaningful estimates of the effects. Second, unlike previous authors (Head and Ries, 2001; Trefler, 2004), we can conduct our empirical analysis at the regional level because of the high degree of heterogeneity across Canadian provinces. For the purpose of our investigation, because of discrepancies in initial provincial export structures and the heterogeneous pace of tariff decrease across products, trade liberalization with the United States led to a differentiated integration process for Canadian regions. Therefore, pooling cross-sectional and time-series (TSCS) data multiplies the information about the nature of the integration process. This allows us to estimate econometric relationships that are more consistent than those obtained in pure time-series or pure cross-sectional analyses. Also, the combination of TSCS data produces better estimates of the dynamic adjustment process. The difficulty in getting information on the dynamics of change solely from pure cross-sectional evidence is well known. Likewise, given the strong multicollinearity among the explanatory variables, it is cumbersome to estimate dynamic models precisely from just one time series. But in a TSCS set-up, the cross-sectional dimension reduces the collinearity problem considerably and the time-series dimension allows explicit disentanglement of short- and long-run effects of shocks such as the CUSFTA. As shown by Trefler (2004), the distinction between short-run adjustment costs and long-run efficiency gains is key when assessing the impact of trade liberalization.

Our empirical results support the idea that in the long run, trade integration between Canada and the United States led to more diversification of manufacturing exports of the Canadian regions. Our findings are, to some extent, robust to the inclusion of the primary sectors and to the potential presence of unit root in the data.

The article is organized as follows. Section 2 presents the data and discusses computation of the specialization index and the measures of trade integration. Section 3 describes our econometric methodology and reports the most important findings that are valid under the assumption of stationary processes for integration and export specialization patterns. Section 4 analyzes the robustness of non-stationary properties of these indicators. Section 5 concludes.

## 2. Data methodology

### 2.1. Capturing specialization patterns

The indexes of export structures used in this study were computed from export data at the industry SIC 4 level, available annually for the 10 Canadian provinces for 1980–2001. 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.[4] 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 number of null entries is large in small provinces such as Prince Edward Island and almost zero in the bigger provinces like Ontario and Quebec. The final data used involve 290 series per province in the all-sectors case (all industries including primary products) and 213 in the manufacturing case. Data were computed for exports to United States (including territories) and to the rest of the World.

While the empirical literature relies on export, output, or employment data to compute the index, we compute specialization from export data only. Data availability is the main reason as export data are much less subject to measurement errors at the regional/industry level, compared with employment or production data. Also, as our investigation focuses on the impact of decreasing tariffs, one might expect most of the effect on the tradable sector. Export data might also be viewed as a proxy of the tradable part of the industrial structure.

We measure the degree of absolute specialization with the Herfindahl index, the common measure used in the empirical literature to capture the degree of specialization or concentration. The Herfindahl measures the extent to which distribution of export shares differs from a uniform distribution. For each province $i$ and each year $t$, the Herfindahl index (denoted by $S_{i,t}$) was computed as

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

where $s_{i,t}^{k} = x_{i,t}^{k} \Big/ \sum_{k=1}^{J} x_{i,t}^{k}$. Here, $J$ is the number of investigated industries (sectors), $s_{i,t}^{k}$ stands for the share of export $x_{i,t}^{k}$ of industry $k$ in the total exports of province $i$. By definition, $S_{i,t}$ will thus be between $(1/J)$ and 1. The smaller (bigger) the number, the more diversified (specialized) is the export structure of the province involved.

The series for the Herfindahl index (all-sectors case) computed from export data to the United States is shown in Figure 1 for the 10 provinces for 1980–2001. The economies of Manitoba and Quebec are by far the most diversified with averages $S_{i,t}$ of 0.038 and 0.055, respectively; the economies of Alberta and Newfoundland are the most specialized with average Herfindahl indexes of 0.463 and 0.357 during the period. Not surprisingly, the extreme specialization of Alberta is attributed to oil and natural gas that account for 67% of Alberta's total exports to the United States during the period. For Newfoundland, the high degree of specialization was attributed to the fish products industry. But the dramatic decline of the Atlantic fish stocks

4 http://strategis.ic.gc.ca/sc_mrkti/tdst/tdo/tdoDefinitions_30.php#industry_selection_search. The authors created the data bank used in this paper from the raw data with assistance of François Rimbaud from Industry Canada.

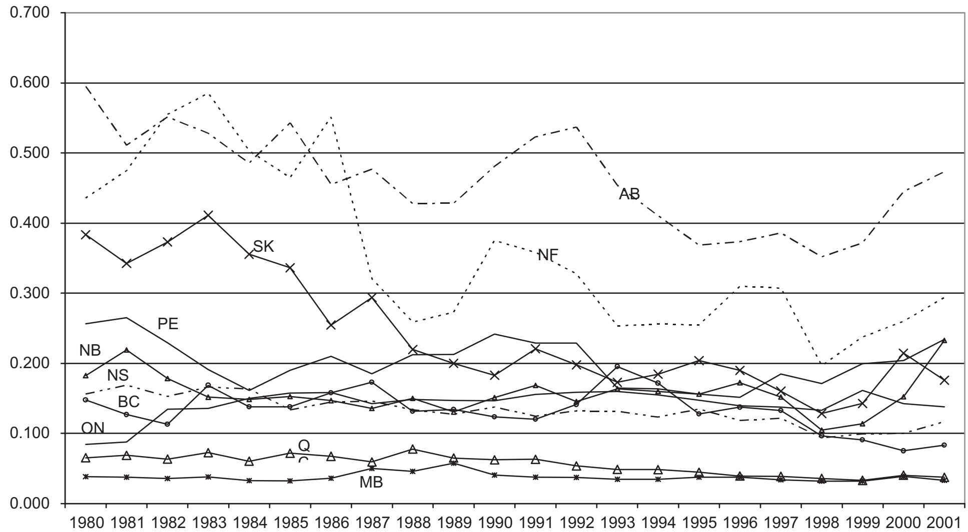


**Figure 1.** Herfindahl indexes for Canadian provinces.

since the end of the 1980s has contributed to the change in the evolution of Newfoundland's specialization index. Between 1980 and 1986, fish products accounted on average for 68% of the province's total exports. Since 1987, this share has fallen to an average of 23%. Note that the upsurge of the Newfoundland Herfindahl index between 1998 and 2001 is attributable significantly to the Hibernia offshore oil field that began producing oil in November 1997. The diversification measure of Ontario, the largest provincial economy, is lowered by the substantial share of the motor vehicle industry that has accounted for about 38% of Ontario's total exports to the United States since 1990.

## 2.2. Capturing trade integration

A straightforward method of capturing (provincial) trade integration could be by using provincial trade flows (exports and imports) between each province and the United States. However, using trade flows is likely to lead to significant endogeneity bias. Trade volumes are likely to be affected by specialization patterns and by the nature of the bilateral trade with the United States (intra versus inter-industrial trade).

Following Trefler (2004), our integration measure is based on tariff dynamics. Our starting point is the per-industry Canadian tariff against the United States. However, our trade-weighted tariffs, unlike Trefler's measure, are also computed from all industries, including primary products that are prominent in several peripheral provinces. Another key difference is that our tariffs are computed at the provincial level using provinces' specific weights based on their industrial structure rather than weights for Canada as a whole. Trade-weighted tariffs are defined as

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

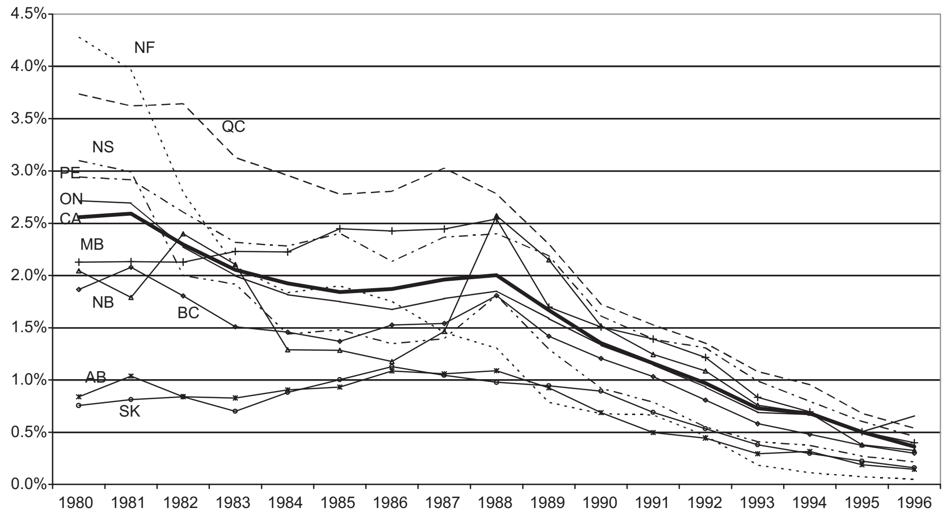


**Figure 2.** Provincial trade weighthed tariffs.

Here $TW_{i,t}$ denotes the level of the trade-weighted tariff of province $i$ at time $t$, $w^k_{i,t}$is the weight of industry $k$ in the total export of province $i$ to the United States at time $t$, and $\tau^k_t$ is the tariff relative to industry $k$ between Canada and the United States.[5]

Figure 2 reports the evolution over time of the trade-weighted tariffs (computed for all sectors) for Canada as a whole and for the 10 provinces. It shows the decreasing pattern of the tariffs since 1980. While the CUSFTA induced a quicker rate of decay for the $TW_{i,t}$, Figure 2 also shows trade integration between Canada and the United States had begun well before its implementation in 1989. Figure 2 also illustrates an important point, relevant to our analysis at the regional level: the significant degree of heterogeneity of the integration process across provinces. This comes from the different patterns of decrease in the bilateral tariffs across industries, combined with strong differences in the (initial) provincial export structures. Because of this, differences across provinces are observed for the $TW_{i,t}$ in both their levels and their speeds of decay. To illustrate, Newfoundland experiences the sharpest decrease in $TW_{i,t}$, with the highest level in 1980 and the lowest in 1998. In contrast, Alberta and Saskatchewan experienced quite constant rates until 1990 and a modest decrease after CUSFTA was implemented. The high degree of heterogeneity of the trade integration process across provinces increases the information in the TSCS empirical framework considerably.

Finally, it is important to bear in mind that, as pointed out by Trefler (2004), the decrease of the $TW_{i,t}$ captures much more than a pure tariff reduction. Basically,

5 Trefler (2004) uses the tariffs between Canada and the United States for the manufacturing sector for 1980–1996. Those data are the collected duties paid by U.S. exporters to Canada. We complement these data with the NAFTA tariffs on primary products, provided by S. Rao from Industry Canada. For 1997–1998, we assume that tariffs on manufacturing goods remained at their 1996 level. As shown by our results relative to the manufacturing sector, this assumption does not affect the main results of this artilce.

it emerges that $\tau_{j,t}$ are well correlated across industries with other non-tariff barriers (NTBs). In this sense, it captures a broader set of CUSFTA trade-liberalizing policies.

### 2.3. Other variables

The potential determinants of export specialization are mainly those affecting Canada–U.S. bilateral trade. In this article, we control for two different sets of variables. The first set 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 the bilateral exchange rate conditions between U.S. and Canadian dollars. We introduce two determinants that have been found to exert detrimental effects on trade volumes. The first is exchange rate volatility that we compute using the measure of integrated volatility proposed by Andersen et al. (2001). The idea of such a measure is to make use of the information available at higher frequencies. Here we build measures of exchange rate volatility at an annual frequency from squared daily returns.[7]

The second determinant is the degree of exchange rate misalignment between the Canadian and U.S. dollars. It has been argued that monetary policy in general plus the overvaluation of the Canadian dollar in the early 1990s had detrimental effects on the macroeconomic dynamics of the Canadian economy (Fortin, 1996). We compute misalignment as the annual absolute deviation of the bilateral exchange rate from its equilibrium value. The equilibrium is assumed to be the purchasing power parity level computed with the ratio of producer prices between Canada and the United States.[8]

## 3. Econometric analysis

### 3.1. Empirical methodology of the TSCS model

The general structure of the TSCS regressions performed in this article is

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

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

$$\varepsilon_{i,t} = c_i + v_{i,t}, \quad \text{and} \quad \varepsilon_{i,t} = c_i + q_t + v_{i,t},$$

6 The data were taken from Beine and Coulombe's (2003) measures of business cycles.
7 The exchange rate data of the CAD–USD at the daily frequency were kindly provided by the Bank of Canada.
8 The producer price index is extracted from the Main Economic Indicators database of the OECD.

where $\nu_{i,t}$ are the idiosyncratic errors, $c_i$ the unobserved individual components, and $q_t$ time dummies.

Many alternative estimation techniques are available for pooling TSCS information in a dynamic model. The first point to tackle is the issue of the unobserved components. The use of TSCS data here makes it 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.

Time dummies were included in a subset of regressions. Using time dummies leads to measuring trade barriers using a relative trade integration index, i.e., trade barriers faced by exporters in one province compared with the average trade barriers in all provinces. The inclusion of $q_t$ results in the elimination of the common decreasing trend in tariffs observed are shown in Figure 2. With time dummies, however, we cannot estimate the effect of variables that only have a temporal dimension such as exchange rate volatility and the misalignment variables.

A second important issue is related to the econometric techniques used to tackle various (cross-sectional and time-series) heteroscedasticity problems underlying this type of TSCS analysis. For the first set of results, using pooled least squares (PLS), we rely on white heteroscedasticity consistent standard errors that allow for asymptotically valid inferences in the presence of general heteroscedasticity. The second set of results comes from generalized least-squares (GLS) estimations using cross-sectional weighted regressions.[9] We assess the robustness of our results by using various techniques peculiar to the TSCS data.

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

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

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. In results with time dummies, the $\phi' Z'$ factors are not entered in the regression.[10]

The interesting feature of the dynamic model (1) is that it separates the short- and the long-run effects of economic integration. Assuming stationarity of $S_{i,t}$ and $TW_{i,t}$, model (1) can be estimated directly by PLS or GLS. The long-run effect of economic

9 With seemingly unrelated regression (SUR) estimations, point estimates have the same sign and order of magnitude as with GLS. However, SUR generated extremely high $t$-statistics and the results are not reported. Refer to Beck and Katz (1995) for the tendency of Parks estimator to lead to extreme overconfidence.

10 We also introduced the provincial cycle variable $PC_{i,t}$ that was always found to be insignificant. The results with this variable are very similar to the ones reported in the subsequent tables.

integration that is associated with $TW_{i,t}$ is then given by $(-\hat{\phi}_2/\hat{\phi}_1)$ while the short-run effect is captured by $\hat{\phi}_3$.

As shown in Pesaran and Smith (1995), ignoring parameter heterogeneity as done using model (1) might produce a substantial bias since the regressors of homogenous pooled estimations will be serially correlated. To account for this potential problem, the findings of this first approach are complemented by those obtained by using a pool mean group (PMG) estimator based on heterogeneous regressions. In the spirit of Pesaran et al. (1999), the speed of adjustment $\phi_{1,i}$ and the short-run effect of tariff $\phi_{3,i}$ are allowed to differ across provinces, whereas the long-run effect of tariffs $\phi_2$ is constrained to be identical:

$$\Delta S_{i,t} = \phi_{1,i} S_{i,t-1} + \phi_2 TW_{i,t-1} + \phi_{3,i} \Delta TW_{i,t} + \varepsilon_{i,t} \tag{2}$$

The $\phi' Z'$ factors were not entered in this specification as they were found generally insignificant in the benchmark regressions.

The empirical methodology used is also subject to another potential problem. It is well known (Nickell, 1981) that fixed-effect estimators of the $\phi_1$ and the other $\phi$s are biased in dynamic models when the number of time periods $T$ is small and the number of cross-section $N$ is large, as in traditional microeconometric studies. Kiviet (1995) shows that fixed-effect estimators are consistent only when $T \to \infty$. The usual modern solution to this is to use Generalized Method of Moments (GMM) estimators. However, as shown in Ahn and Schmidt (1995), these are not very efficient and fixed-effect estimators have a much smaller variance (Kiviet, 1995). In our analysis, we rely on fixed-effect estimators for two reasons. First, the number of periods $T$ (18) is relatively large. Kiviet (1995), for example, concentrates on the cases when $T = 3$ and 6. Second, as shown in Kiviet (1995), the bias is reduced when the speed of adjustment is faster, as in our analysis.

## 3.2. Results: the dynamics of export diversification

### *3.2.1 Main results*

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 1–3. Table 1 reports the results of benchmark regressions for the manufacturing sectors [as done in Trefler (2004)]. In Table 2, the analysis is applied to all industries. Table 3 reports results from the PMG estimates.

For the manufacturing case (Table 1), the parameter estimates of the $TW_{i,t-1}$ variable ($\phi_2$) are always positive and significantly different from 0 at the 1% level. Significantly, this positive effect is robust to including time dummies, suggesting that the export structure of the province $i$ reacts to changes in trade-weighted tariffs specific to province $i$. Even more significantly, the PMG estimates of $\phi_2$ parameter (Table 3) are consistent with the benchmark regressions, suggesting this effect is robust to the degree of heterogeneity in the parameter values. The various estimates of the adjustment speed to equilibrium $\phi_1$ are always negative and significant at the 1% level. Therefore, the measures of the long-run elasticity given by $\left(-\hat{\phi}_2/\hat{\phi}_1\right)$ of tariff changes on trade specialization are always positive and significant with $p$-values below the 1% level. The various measures of the long-run elasticity computed from the point estimates are quite

**Table 1.** Dynamics of industrial diversification—Herfindahl (manufacturing sector)

| | Dependent variable: $\Delta\log(S_{i,t})$ | | | |
|---|---|---|---|---|
| | (1) PLS | (2) GLS | (3) PLS | (4) GLS |
| Time dummies | No | No | Yes | Yes |
| $\log(S_{i,t-1})$ | $-0.401^{***}$ (0.069) | $-0.376^{***}$ (0.083) | $-0.387^{***}$ (0.080) | $-0.380^{***}$ (0.078) |
| $\log(1+TW_{i,t-1})$ | $5.323^{***}$ (1.352) | $4.677^{***}$ (1.403) | $6.475^{**}$ (2.785) | $7.883^{***}$ (2.464) |
| $\Delta\log(1+TW_{i,t})$ | 2.600 (4.157) | 3.680 (3.885) | 2.205 (5.063) | 3.010 (4.678) |
| $UC_t$ | −1.220 (1.564) | −0.949 (1.548) | – | – |
| $PC_{i,t}$ | −0.001 (1.886) | −1.190 (1.903) | – | – |
| $VL_t$ | 3.441 (14.283) | 5.526 (12.745) | – | – |
| $MA_t$ | 0.138 (0.232) | 0.146 (0.216) | – | – |
| $R^2$ | 0.251 | 0.280 | 0.240 | 0.266 |
| Adj. $R^2$ | 0.167 | 0.198 | 0.178 | 0.206 |
| S.E. of $R$ | 0.134 | 0.135 | 0.132 | 0.132 |
| Observations | 160 | 160 | 160 | 160 |

*Notes*: Sample: 1980–1996. All regressions include provincial fixed effects. Estimation techniques: pooled least-squares (PLS), feasible generalized least-squares (GLS). White heteroscedasticity standard errors are shown in parentheses below the estimated coefficients. For columns (3) and (4), reported $R^2$ and adjusted $R^2$ exclude time dummies. $^{*}$significance at 10% level; $^{**}$significance at 5% level; $^{***}$significance at 1% level.

**Table 2.** Dynamics of industrial diversification—Herfindahl (all sectors)

| | Dependent variable: $\Delta\log(S_{i,t})$ | | | |
|---|---|---|---|---|
| | (1) PLS | (2) GLS | (3) PLS | (4) GLS |
| Time dummies | No | No | Yes | Yes |
| $\log(S_{i,t-1})$ | $-0.342^{***}$ (0.070) | $-0.340^{***}$ (0.075) | $-0.344^{***}$ (0.065) | $-0.340^{***}$ (0.067) |
| $\log(1+TW_{i,t-1})$ | $5.384^{***}$ (1.443) | $5.065^{***}$ (1.579) | 2.944 (2.576) | 3.114 (2.275) |
| $\Delta\log(1+TW_{i,t})$ | −5.268 (4.317) | −6.116 (4.345) | −10.844 (7.272) | −11.518 (7.202) |
| $UC_t$ | −1.732 (1.495) | −1.988 (1.341) | – | – |
| $PC_{i,t}$ | 1.325 (1.599) | 1.280 (1.479) | – | – |
| $VL_t$ | $-26.690^{**}$ (11.588) | −15.700 (9.663) | – | – |
| $MA_t$ | −0.261 (0.171) | $-0.262^{*}$ (0.141) | – | – |
| $R^2$ | 0.279 | 0.314 | 0.244 | 0.272 |
| Adj. $R^2$ | 0.208 | 0.246 | 0.189 | 0.220 |
| S.E. of $R$ | 0.129 | 0.130 | 0.124 | 0.124 |
| Observations | 180 | 180 | 180 | 180 |

*Notes*: Sample: 1980–1998. See also notes for Table 1.

comparable across estimation techniques and specifications. They vary between 11 and 20 for the manufacturing sectors.

As illustrated in Figure 2, the CUSFTA accelerated the negative trend in tariffs since 1989. As a consequence, our main results imply that the mean (across provinces) effect of trade integration induced by the CUSFTA was a significant, substantial increase in the long-run industrial diversification of Canadian provinces. This result is

**Table 3.** Dynamics of industrial diversification—Herfindahl, heterogeneous model: PMG estimates

| | Benchmark model | | With time dummies | |
|---|---|---|---|---|
| | All sectors | Manufacturing | All sectors | Manufacturing |
| $\text{Log}(S_{i,t-1})$ | −0.409*** (0.073) | −0.423*** (0.053) | −0.453*** (0.075) | −0.419*** (0.032) |
| Median $\log(S_{i,t-1})$ | −0.461 | −0.424 | −0.443 | −0.421 |
| $\log(1+TW_{i,t-1})$ | 7.727*** (1.521) | 4.703*** (1.392) | 7.761*** (2.759) | 6.327*** (2.615) |
| $\Delta\log(1+TW_{i,t})$ | −20.469** (6.981) | −0.041 (4.693) | −25.151** (10.211) | 2.828 (8.446) |
| Median $\Delta\log(1+TW_{i,t-1})$ | −19.894 | −1.766 | −14.774 | 0.930 |
| Long-run effect of $\log(1+TW_{i,t-1})$ | 18.892*** | 11.118*** | 16.029*** | 15.100*** |

*Notes*: Sample: 1980–1998, all sectors case. Sample: 1980–1996, manufacturing case. Columns 2 and 3 report the results without time dummies while columns 4 and 5 provide the findings of estimates with time dummies included in the model. Asymptotic standard error reported for mean (PMG estimator).

extremely robust and remains invariant across estimation techniques (PLS, FGLS, and PMG) and specifications in the manufacturing sector.

Estimates for all the sectors including the primary sectors lead broadly to the same conclusion. For most estimations (Tables 2 and 3), the long-run parameters are also positive and significant. Some exception to this robustness is found, however, in the benchmark regressions with time dummies (Table 2, last two columns) for which, while still positive, estimates of $\phi_2$ are not significantly different from 0 at conventional significance levels. Using the PMG estimates with time dummies (Table 3), however, we find positive and highly significant estimates of $\phi_2$, suggesting that parameter heterogeneity might lower the quality of PLS and GLS estimates. In all, this result might nevertheless suggest that, for the primary sectors, the change in export structure might be driven more by the general decrease of tariffs than by specific provincial components. Another implication is that for primary sectors, changes in export structure might also be more driven by non-tariff variables. As described in Section 2, changes in the Herfindahl indexes are also clearly related to supply-side factors. For instance, increase in the export intensity of the oil sector at the end of the 1980s due to deregulation of natural gas and oil prices in 1988 led to a significant increase in the degree of specialization for Alberta. Another illustration: the dramatic decline of the Atlantic fish stocks since the end of the 1980s has also contributed to the change in the evolution of Newfoundland's specialization index. The very specific nature of the primary sectors makes the impact of these supply-side factors much more important than for the manufacturing sectors.

Up to now, specification (1) has seemingly ignored spillover effects across provinces. Spillover effects might be due to the fact that a specific variation of tariffs in a given province might induce a change in the degree of export diversification in other provinces because of factor mobility among Canadian provinces. Labor mobility is known to be quite important within Canadian provinces compared with other parts of the world such as Europe. A natural way to account for these spillovers is to augment specification (1) by including some average measure of other provinces' trade-weighted tariffs. Using population weight to compute this measure, we re-estimated model (1) and found positive estimates of $\phi_2$. The level of significance was nevertheless much lower.[11] This drop in significance is clearly due to high correlation between the trade-weighted tariffs of province $i$ and the average of the nine other provinces, leading to a classical problem of multicollinearity.[12] The inclusion of time dummies, however, is an alternative, efficient way to check for robustness to the spillover effects; including $q_t$ is equivalent to including any average of the trade-weighted tariffs for the 10 Canadian provinces. Therefore, the fact that our results are robust to including time dummies provides further evidence that our findings are robust to the issue of spillovers across provinces.

Our estimation results are very robust to the inclusion of the various control variables ($Z$ variables). Those variables are supposed to explain part of the short-run variations in the export diversification. Most of them were insignificant. There is limited evidence to suggest that increases in exchange rate variability have a positive effect on export diversification in the all-sectors case. This result, however, is not robust to the exclusion of primary sectors. All in all, the results suggest first, that exchange rate fluctuations and monetary arrangements can exert asymmetric effects on the industrial production structures of provincial economies (Beine and Coulombe, 2003) and second, that the inclusion of control variables is unlikely to alter the general direction of the main results.

We find limited evidence for short-run effects of tariff changes ($\phi_3$). This impact is found to be negative for the all-sectors case, suggesting trade liberalization might lead to more export specialization in the short run. However, this effect is not robust to the exclusion of exports of primary products: the short-run effect of tariff changes is insignificant for the manufacturing sector. Again, the results obtained using benchmark regressions are fully consistent with those based on the PMG estimates. That the short-run impact of trade liberalization differs from the long-run effect emphasizes the need to disentangle the short- and long-run components of the tradable sector response. For example, the negative effect of a one-time decrease in tariffs occurring during year $t$ in the all-sectors case would be totally offset by the end of year $t+1$ or $t+2$. Static model estimates might thus reach misleading conclusions about trade liberalization's impact. The results also suggest export structures need time to adjust to tariff changes, although the adjustment is quite fast.

### 3.2.2. Interpretation: what drives the results?

An important question is why trade integration tends to increase export diversification in the long run. One possible channel is that trade liberalization implies a change in production structures and the creation of new exporting sectors. This could be driven by the agglomeration process and the backward-forward linkages emphasized by the new economic geography (Fujita et al., 1999). A second channel could be that the decrease in tariffs results in a more direct increase in exports of existing sectors. This increase might be triggered by a general increase in production or/and by

11 To save space, the results are not reported here. Depending on the inclusion of the primary sectors, $p$-values of the $\phi_2$ estimates lie between 0.05 and 0.085.

12 Regressing $TW_{i,t}$ on this average measure results with and without fixed individual effects in a regression coefficient of 0.89 (resp. 0.90) with $t$-statistics above 27 (resp. 40), suggesting that the degree of correlation is very high.

substituting production for the local market with production for the United States. Our estimations do not provide direct answers to this question, but indirect evidence tends to support the second channel as the operative one here.

First, our estimation suggests that the speed of adjustment of export structures to a shock such as the decrease in tariffs is quite fast. The point estimates of the $\phi_1$ parameter in the all-sectors and manufacturing cases (Tables 1 and 2) are quite stable across the various specifications and estimation techniques; they vary between $-0.32$ and $-0.40$. In the case of the benchmark model for the manufacturing index (Table 1), the estimated value of $\phi_1$ is $-0.38$ on average. This implies that, following a shock on the long-run equilibrium of the degree of export specialization, half the gap between initial situation and new long-run level is adjusted in a bit less than two years approximately. This quick adjustment suggests that the change in the degree of export diversification involves existing sectors rather than new ones whose implementation would require much more time. Second, existing studies (Coulombe, 2004) suggest that international and interprovincial trade complement rather than substitute. In other terms, the strong increase in trade volumes between Canadian provinces and the United States observed over our investigation period was not at the expense of production for national consumption. This tends to shed some doubt on the substitution hypothesis and to favor the view that the increase of exports was driven by a general increase in production of existing tradable sectors. To complement this assessment, we have also run a cross-section regression involving the long-run change in Canadian exports at the sector level over the 1980–1996 period.[13] The change in exports is regressed on the initial value of export in 1980 (to capture potential catch-up effects) and the change in production over the same period of time. Interestingly, we obtain negative estimates for the first variable, suggesting some diversification process of Canadian exports and moderate positive estimates for the change in production, suggesting that the increase of exports occur in sectors where the change in production was relatively higher.[14] Once more, this tends to reject the so-called substitution hypothesis. Obviously, more evidence is called for to improve this evaluation but this is clearly beyond the scope of this article. We therefore leave that for future work.

The effect of tariff changes on industrial diversification is captured from Herfindahl indexes computed from the provinces' exports to the United States. One could argue that in the long run, the increased diversification might be the result of some sort of export substitution at the industry level—from the rest of the World to the United States after the arrival of the CUSFTA. Indeed, as emphasized by Trefler (2004), the CUSFTA mandated tariff concessions to the United States that were preferential compared with the rest of the World. We refer to this as the 'specialization substitution hypothesis.' This hypothesis was directly tested using the benchmark specification for which the dependent variable is based on the Herfindahl index computed from provinces' exports to the rest of the World. Non-reported results of the panel data regressions indicate that the specialization substitution hypothesis is clearly rejected. No point estimates of the long-run effect were significant and the *p*-values were

13 The regression includes 204 sectors only since some production data are missing for some sectors either in 1980 or in 1996.

14 The results can be obtained from the corresponding author upon request.

extremely high, especially with FGLS. There is, however, some evidence (non-robust) that the specialization substitution hypothesis might hold in the short run.

## 3.3. Robustness

### *3.3.1. Constant weight tariffs*

It is possible to argue that our baseline measure of trade-weighted tariffs is partly endogenous since the evolution of provinces' weights is affected by the (endogenous) evolution of the industrial structure. To address this particular point, we build an alternative tariff variable using constant weights based on the mean structure of exports over the 1980–2001 period. The results (Table 4, columns 1 and 4) clearly indicate that our main result regarding the long-run effect of economic integration is not driven by the change in industrial structure on provinces' weights. Estimates of the $\phi_2$ parameter in the all-sectors and the manufacturing cases with the alternate measure are very comparable with the ones found with flexible weights and are highly significant. However, the measures of the long-run elasticity of tariff changes are slightly lower in the constant-weight case.

### *3.3.2. Serial correlation*

The possibility that residuals might be serially correlated could lead to serious problems of estimator inconsistency in dynamic TSCS with fixed effects. We have re-estimated the benchmark specification under the assumption that the disturbance follows either an AR(1) or an AR(2) process using non-linear least squares. Results from PLS and FGLS the all-sectors and the manufacturing sector are reported Table 4, columns 2, 3, 5, and 6, for the AR(1) specification. Non-reported results for the AR(2) specification go in the same direction.

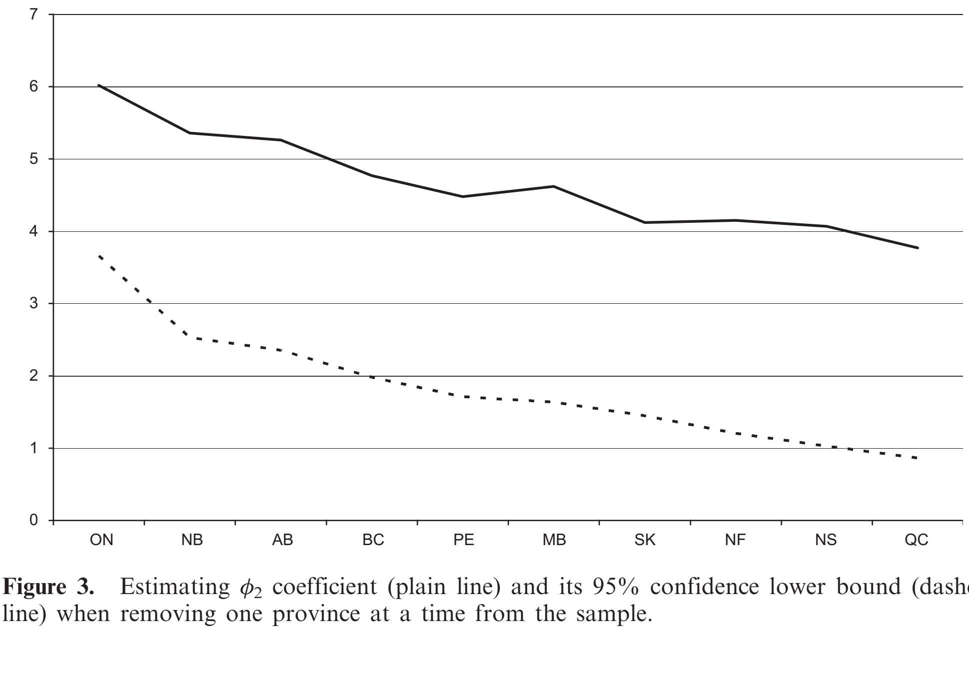


**Figure 3.** Estimating $\phi_2$ coefficient (plain line) and its 95% confidence lower bound (dashed line) when removing one province at a time from the sample.

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

| | Dependent variable: $\Delta\log(S_{i,t})$ | | | | | |
|---|---|---|---|---|---|---|
| | All sectors | | | Manufacturing sectors | | |
| | Constant weights | AR(1) PLS | AR(1) FGLS | Constant weights | AR(1) PLS | AR(1) FGLS |
| $\log(S_{i,t-1})$ | $-0.346^{***}$ (0.074) | $-0.366^{***}$ (0.094) | $-0.359^{***}$ (0.073) | $-0.337^{***}$ (0.078) | $-0.451^{***}$ (0.126) | $-0.408^{***}$ (0.071) |
| $\log(1+TW_{i,t-1})$ | $4.131^{***}$ (1.279) | $5.396^{***}$ (1.739) | $5.089^{***}$ (1.244) | $2.785^{***}$ (1.048) | $5.484^{***}$ (1.775) | $4.099^{***}$ (0.979) |
| $\Delta\log(1+TW_{i,t})$ | 6.715 (7.297) | −4.021 (4.406) | −4.830 (3.803) | 4.548 (4.832) | 2.539 (4.594) | −4.083 (3.933) |
| $\mathrm{UC_t}$ | −2.209 (1.347) | −2.273 (1.615) | −1.979 (1.273) | −0.772 (1.476) | −1.818 (1.605) | −0.498 (1.346) |
| $\mathrm{PC_{i,t}}$ | 1.261 (1.519) | 2.144 (1.665) | 1.491 (1.498) | −1.234 (1.897) | 1.012 (2.018) | −1.399 (1.526) |
| $\mathrm{VL_t}$ | −14.260 (10.291) | $-27.812^{**}$ (11.670) | −12.884 (8.961) | 5.794 (12.349) | 1.673 (14.347) | 6.347 (9.912) |
| $\mathrm{MA_t}$ | $-0.415^{**}$ (0.160) | −0.282 (0.175) | $-0.213^{*}$ (0.121) | 0.076 (0.207) | 0.105 (0.233) | 0.239 (0.141) |
| AR1 | – | −0.039 (0.130) | $-0.187^{*}$ (0.111) | – | −0.009 (0.151) | $-0.188^{*}$ (0.111) |
| $R^2$ | 0.285 | 0.304 | 0.427 | 0.269 | 0.279 | 0.249 |
| Adj. $R^2$ | 0.215 | 0.227 | 0.363 | 0.188 | 0.186 | 0.153 |
| S.E. of $R$ | 0.133 | 0.129 | 0.132 | 0.136 | 0.134 | 0.137 |
| Observations | 180 | 170 | 170 | 180 | 150 | 150 |

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

Accounting for serial correlation does not change the general direction of the results: both magnitude and significance level of key parameters remain quite stable. This suggests that our benchmark model succeeds in capturing a significant part of the dynamics of the disturbances.

### 3.3.3. Removing one province at a time

To further assess the sensitivity of our results, we estimated the benchmark equation, removing a different province each time from the sample. This analysis checks whether the results are driven by dynamics of a particular province. This effect might occur due to the relatively low cross-sectional dimension (10) of the panel. Figure 3 reports the value of the estimated long-run coefficients $\phi_2$ (plain line) with its 95% confidence lower bound (dashed line) when one particular province is removed from the sample (indicated on the horizontal axis). The results clearly indicate that they are not driven by one particular region, all estimated coefficients being significant at least at the 5% level.

## 4. Accounting for non-stationarity

The previous econometric analyses implicitly assume stationarity properties of the data. For many of the variables in our analysis—such as economic cycles, exchange rate volatility, or misalignment—this assumption obviously holds. In contrast, indexes of export specialization and the trade-weighted tariffs might exhibit stochastic trends. In this section, we employed some recent econometric tools of the recent literature devoted to the stationarity issue in TSCS data to assess the robustness of our previous results.

Accounting for non-stationarity in this analysis might be important for two reasons. The first is related to the problem of spurious regressions. If the data follow a non-stationary process in level, then estimating $\phi_1$ and $\phi_2$ can lead to problems. Second, if both series follow I(1) processes, their long-run relationships are best captured in a cointegration framework. Cointegration tests allow one to investigate whether a long-run relationship holds for non-stationary processes.

The first step is testing for the presence of unit roots. Since Levin and Lin (1992), several tests have been developed to perform such a task in panel or TSCS data. We use three different test statistics: the Levin–Lin rho statistics, the Levin–Lin *t*-rho statistics, and the Levin-Lin ADF statistics.[15] Given that the specialization measures and trade-weighted variables are between 0 and 1, the analysis is also conducted on the logistic transformation of these variables. The results of these tests are not reported here due to space constraints but all details can be found in Beine and Coulombe (2004).

We found mixed evidence of non-stationarity in the two Herfindahl indexes.[16] In contrast, the trade-weighted tariffs appear to exhibit unit root properties for both the all-goods and the manufacturing sectors. In no case can we conclude in favor of

15 The interested reader may find more details in Levin and Lin (1993) and Pedroni (2004).

16 In Beine and Coulombe (2004), two different statistical processes are allowed for. The first involves coefficients that are common across provinces and refers to the homogeneous case. The second involves coefficients that differ across provinces and refers to the heterogeneous case. We investigate here only the homogeneous model that is the direct counterpart of the benchmark model discussed in Section 3.

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

| | Dependent variable: $\Delta\log(S_{i,t})$ | | | |
|---|---|---|---|---|
| | All sectors | | Manufacturing sectors | |
| $\mathrm{ECM}_{\mathrm{i},t-1}$ | −0.349*** (0.076) | −0.341*** (0.083) | −0.381*** (0.084) | −0.387*** (0.086) |
| $\Delta TW_{i,\mathrm{t}}$ | −6.542 (4.323) | −9.912** (4.55) | 3.072 (3.852) | 2.748 (3.050) |
| $\mathrm{VL_t}$ | −16.242* (9.572) | – | 5.219 (12.654) | – |
| $\mathrm{UC_t}$ | −2.014 (1.340) | – | −0.933 (1.544) | – |
| $\mathrm{PC_{i,t}}$ | 1.289 (1.493) | – | −1.329 (1.891) | – |
| $\mathrm{MA_t}$ | −0.295** (0.131) | – | 0.142 (0.212) | – |
| $R^2$ | 0.313 | 0.284 | 0.278 | 0.260 |
| Adj. $R^2$ | 0.250 | 0.237 | 0.203 | 0.205 |
| S.E. of $R$ | 0.130 | 0.132 | 0.135 | 0.133 |
| Observations | 180 | 180 | 160 | 160 |

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

stationarity for any variable with a high degree of certainty. This in turn calls for some cointegration analysis.

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

Under the hypothesis of a cointegrating relationship between export diversification and the $TW_{i,t}$, the dynamics are best captured by the error correction model (ECM)

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

in which $-\lambda$ captures the adjustment speed toward the long-run equilibrium between $S_{i,t}$ and $TW_{i,t}$. The $ECM_{i,t} = \log(S_{i,t}) - \alpha_i - \beta\log(1 + TW_{i,t})$captures deviations from this long-run equilibrium. Model (3) is a re-parameterization of the benchmark model (1), under the assumption of I(1) variables where $\lambda = \phi_1$, $\phi_2 = -\lambda\beta$ , and $c_i = \delta_i - \lambda\gamma_i$.

The long-run estimates are mostly in line with the ones implied by the previous regressions. Depending on the specifications, the estimates of $-\hat{\phi}_2\big/\hat{\phi}_1$ range between 14.897 and 19.268 for the all-sectors case and between 12.472 and 13.274 for the manufacturing case. Table 5 reports the estimates of the ECM models (3).

The estimates of $\lambda$ are strikingly close to the previous estimates of $\phi_1$ in both sectors. To sum up, the results obtained here suggest that the main result of this article is robust to the potential presence of a stochastic trend in the data

## 5. Conclusion

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

From the evolution of trade-weighted tariffs, Canadian provinces obviously were not affected equally by the gradual elimination of trade barriers observed over the last 20 years. Due to the relative stability of cross-sectional (cross-provincial) variances compared with time-series variances, in this article we are able to separate short- and long-run effects of economic integration on export diversification in a rather robust way. Given the limited number of available time series and the fact that we are dealing with only one aggregate shock (CUSFTA), it would have been rather heroic to try this exercise with only one time series.

We found that, on average, trade integration between Canada and the United States clearly leads to more export diversification in the long run. This result holds especially for the manufacturing sector. It turns out to be robust to the presence of spillover effects across provinces, the issue of heterogeneous parameters in regression models, the potential presence of stochastic trends in the trade-weighted tariffs data, the issue of endogeneity of the trade weights used to compute this latter measure, and a set of other statistical concerns such as the occurrence of serial correlation.

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