U.S.-CANADA TRADE LIBERALIZATION AND MNC PRODUCTION LOCATION

Susan E. Feinberg and Michael P. Keane*

Abstract—Using confidential firm-level panel data from the Bureau of Economic Analysis, we examine how the bilateral trade flows of U.S. multinational corporations (MNCs) and their Canadian affiliates responded to U.S.-Canadian tariff reductions from 1983 to 1992. We find that Canadian affiliate sales to the United States are negatively correlated with Canadian tariffs, but U.S. parent sales to Canadian affiliates have little association with Canadian tariffs. These results contradict the notion that Canadian tariff reductions would lead to a "hollowing out" of Canadian manufacturing. We also find substantial heterogeneity in MNC responses to tariff changes within narrowly defined manufacturing industries. Overall, bilateral trade liberalization is trade-creating, as U.S. MNCs integrated their North American production such that Canadian affiliates increased sales to the United States and reduced domestic sales.

I. Introduction

THE decisions of multinational corporations (MNCs) about where to locate production have been at the forefront of the debate regarding free trade. As the recent debates in the United States over the NAFTA and in Canada over the Canada-U.S. Free Trade Agreement (FTA) showed, there is great concern over the question of whether MNCs systematically prefer to locate production in countries with lower wages and factor costs, as predicted by factor-proportions theories of trade, or whether MNCs prefer to concentrate production in a single country to exploit economies of scale, as predicted by IO-based theories. The concern is that these tendencies will be magnified if trade liberalization enables firms to easily ship goods that are produced abroad back to the United States (or Canada).

At the heart of this debate is the notion that, when trade is liberalized, all firms within the same industries or sectors of particular countries respond similarly based upon either the factor intensities of the products they produce or upon other characteristics of their production technologies (such as economies of scale). In this paper, we examine the extent to which trade liberalization induced U.S. MNCs with affiliates in Canada to adjust their production locations, and the extent to which these adjustments are predicted by characteristics of the industries in which the MNCs operate. Production-location decisions are inferred from trade flows: bilateral flows to and from Canadian affiliates and their U.S. parents, Canadian affiliates' sales to unaffiliated buyers in the United States, and affiliates' domestic sales in Canada. Changes in these trade flows indicate, for example, whether bilateral trade liberalization affected the amount of product that U.S. MNCs produced in the United States to sell to Canada or produced in Canada to sell locally or to the United States.

There is considerable evidence that U.S. MNCs and their Canadian affiliates organize their production in quite different ways, even when they are in the same industry, at the same point in time, and facing the same prices. To illustrate, charts A through D in figure 1 show the four trade flows for affiliates in a single industry in 1983, 1988, and 1992 using confidential firm-level data from the BEA. To preserve confidentiality, we cannot reveal the industry or the number of firms in the industry. However, the within-industry heterogeneity in trade flow patterns and the changes in trade flows between 1983 and 1992 are broadly representative of the industries in our sample. In charts A through D, the horizontal axis represents the ratio of each trade flow to total affiliate sales. (Ratios are specified in ranges of one-tenth.) The vertical axis gives the percentage of affiliates in the industry with flows in the specified ranges. So, for example, in chart D, we can see that, in 1992, approximately 44% of affiliates have Canadian sales in the range of 0.9 to 1. These affiliates are therefore selling nearly all their production locally. Note, however, that the 44% figure has dropped from nearly 80% in 1983—indicating that, between 1983 and 1992, Canadian affiliates in this industry reorganized to sell less of their production locally. From the charts, considerable organizational differences between MNCs are evident, particularly with regard to the two intrafirm trade flows (charts A and B). We argue that such differences in the organization of production reflect unobserved differences in market power or technology, which cause firms in the same industry facing the same prices to choose different configurations of production. The extent to which such choices differ systematically, both within and across industries, is the principal focus of this research. In particular, we expect that MNCs within the same industry that organize production quite differently prior to a tariff change may also respond quite differently to a tariff reduction.

The causal link between firms' production technologies and their responses to trade liberalization can only be ascertained by estimating a structural model of MNC behavior that includes parameters of individual firms' technology and product demand. No such estimated models exist in the empirical trade literature. Grubert and Mutti (1991) and Brainard (1995) discuss difficulties in previous empirical work in which researchers have essentially regressed indicators of direct investment activity, such as MNC sales, on endogenous variables, such as MNC exports. Grubert and Mutti avoid this problem by examining the effect of exog-

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* University of Maryland and New York University, respectively.

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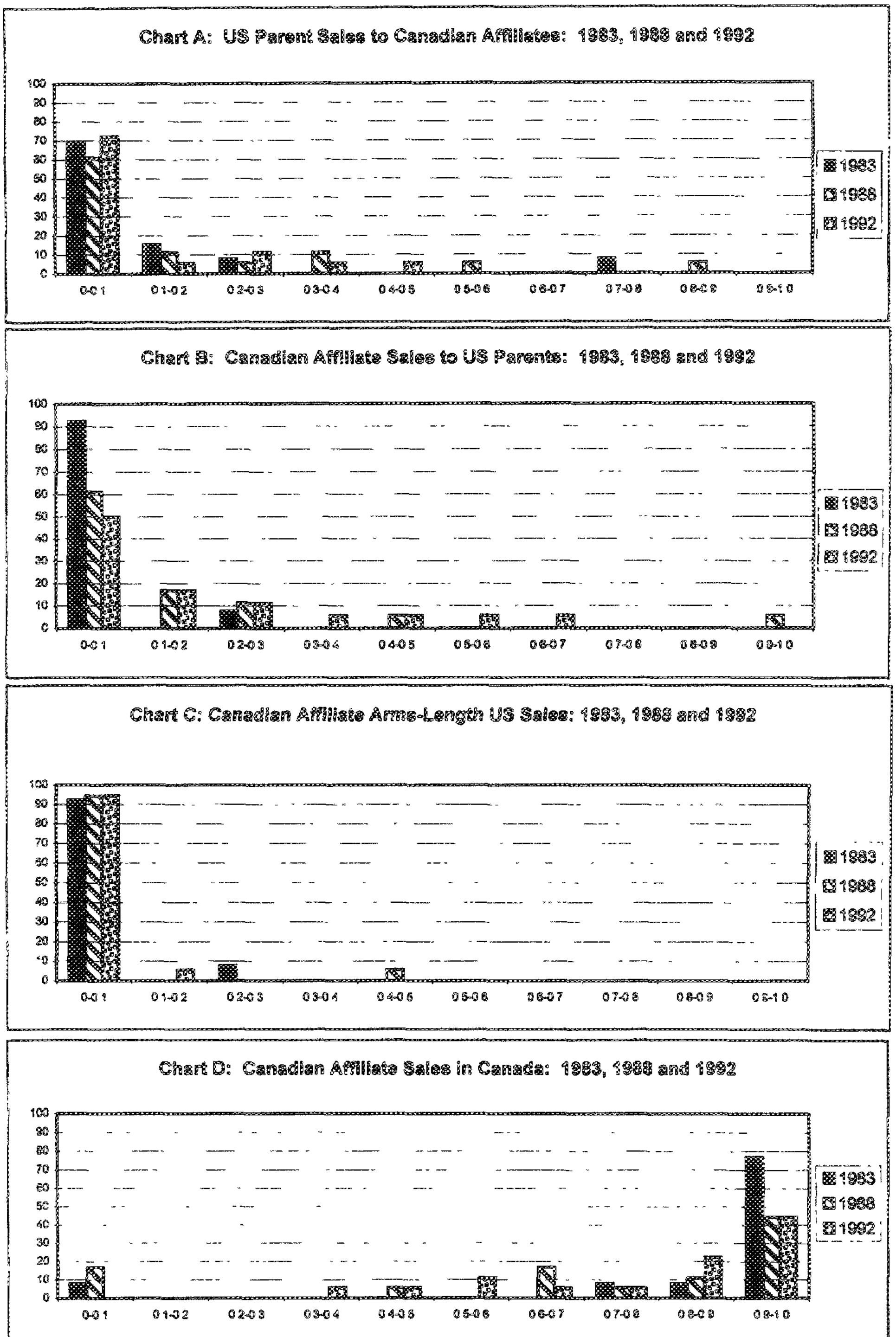
FIGURE 1.—CHARTS A–D: TRADE FLOWS-TO-TOTAL AFFILIATE SALES IN 1983, 1988, AND 1992



Notes: The horizontal axis represents the ratio of the trade flow to total affiliate sales. The vertical axis represents the percentage of affiliates in the industry with flows in the specified ranges. The identity of the industry and number of firms is confidential.

enous variables such as taxes, tariffs, and per capita GDP on MNCs' capital stock. Similarly, Brainard (1993) examines the effect of variables such as trade and investment policies and transportation costs on the share of total MNC sales accounted for by affiliate sales and exports. Using disaggregated panel data, we take a similar approach to Grubert and Mutti (1991) and Brainard (1993), but our aim is to examine the effect of changing tariff levels on MNC production-location decisions over time.

Although earlier studies on the impact of tariffs on MNC decision-making (Horst, 1972; Grubert & Mutti, 1991) used aggregate and/or single-period data, our study differs in two important dimensions. First, we use confidential firm-level panel data from the Bureau of Economic Analysis (BEA) to examine MNC responses to trade liberalization. Second, we examine the extent to which MNC responses to trade liberalization vary both across industries, and across firms within the same industry.

From a theoretical point of view as well as from a policy perspective, it is important to know whether trade liberalization affects all firms within particular industries in a similar fashion. For example, when governments administer trade policies (such as the Semiconductor Trade Agreement or the Multifiber Arrangement), they typically focus on particular industries that are considered to be vulnerable or to have been harmed by trade. However, if most of the variance in adjustments to trade liberalization is intra- rather than interindustry, policies designed to protect vulnerable industries may be missing the mark. From a theoretical perspective, theories of the MNC (Dunning, 1979; Rugman, 1981) suggest that idiosyncratic characteristics of firms such as the skills they possess or their reputation are an important determinant of their behavior in domestic and foreign markets. The extent to which such firm-specific characteristics may be important in explaining their response to trade liberalization has never been tested. Here we examine the relative magnitudes of the variance in MNCs' responses to trade liberalization explained by characteristics of firms and characteristics of industries.

By using *firm-level* panel data, we can examine the effect of tariff reductions on MNCs' production location decisions within and across industries. Previous work has typically used single-period cross-sectional designs with industry or country-level data to examine the impact of tariffs on MNCs' production location decisions.[1] For example, Culem (1988) examines the determinants of production location among industrialized countries between 1969 and 1982 but uses country aggregate data. Grubert and Mutti (1991) and Horst (1972) examine the impact of tariffs on production locations using only a single year of aggregate data. Specifically, Horst (1972) examines shares of U.S. exports and Canadian affiliate production for 18 2-digit manufacturing industries, and Grubert and Mutti (1991) use country aggregate data on 33 countries. Caves (1990) uses panel data from 1970–1979 to examine adjustments to Canada-U.S. trade liberalization, but his examination focuses on 4-digit industries rather than firms. More recently, Brainard (1993, 1997) uses disaggregated data from the BEA Benchmark Survey to examine factor proportions and market proximity explanations of MNC location decisions, but she too uses a single year of data. Because we examine changes in MNC production-location decisions using firm-level data over a ten-year period of time that includes both the Tokyo Round of the GATT and the Canada-U.S. Free Trade Agreement, we can better identify the influence of tariffs on MNC production-location decisions.

Specifically, we examine the impact of U.S. and Canadian tariff reductions on the production-location decisions of U.S.-based MNCs and their Canadian affiliates from 1983 to 1992. The Canadian context is an interesting and relevant setting for examining the impact of tariff reductions on MNC production decisions. Because of the large share of Canadian manufacturing capacity owned by U.S. MNCs[2] and the ease of cross-border shipping, there was great concern in Canada that U.S. MNC affiliates would leave Canada and serve the Canadian market through exports originating in the United States (Crookell, 1990). Such divestment could take several possible forms. MNCs could exit the Canadian market altogether and simply replace local production and sales with exports from U.S. parents. Alternatively, MNCs could remain in the Canadian market but source most of the components in the products they make from the United States. This latter type of adjustment constitutes the phenomenon of "hollowing out" (Cohen & Zysman, 1987). In manufacturing industries, it was widely predicted that U.S. MNCs would rationalize their Canadian operations by reducing the number of product varieties produced in each plant and increase plant-level economies of scale (Caves, 1990; Cox & Harris, 1985, 1986). Firms within the same industries were predicted to respond similarly based upon industry production technologies.

The data set used in this paper is from the Benchmark and Annual Surveys of U.S. Direct Investment Abroad (USDIA) from the Bureau of Economic Analysis. In this research, we use firm-level data that includes detailed information on the entire population of U.S.-based MNCs and their foreign affiliates from 1983 to 1992. Due to the confidentiality of the firm-level data, access is restricted by the BEA. However, the size of the sample and the relatively long period over which MNCs are observed enable us to rigorously examine a variety of unanswered questions in the empirical trade literature. In this study, we use a panel of 701 majority-owned, U.S.-based MNC parents and their Canadian affiliates in fifty manufacturing industries.[3] The detailed microdata enable us to examine the relationship between tariff reductions and changes in MNC production-location decisions and the extent to which the latter are systematically explained by firm or industry characteristics.

[1] Thompson's (1993) study differs from other research on adjustments to tariff reductions in that she uses stock price data on Canadian firms to measure the sectoral and intraindustry abnormal returns that correspond to important events in the passage of the Canada-U.S. Free Trade Agreement. She does not examine the impact of tariff reductions on production location decisions per se, but she does find evidence of significant intraindustry variance in returns among firms in two-digit industries.

[2] Foreign ownership of Canadian manufacturing capacity peaked during the Trudeau years, reaching a high of 61% in 1970 (Bothwell, 1992). Today, foreign ownership of manufacturing capacity is approximately 40%, although it varies considerably by industry (Rugman, 1989). At the end of 1977, U.S. investors held 77% of the total foreign investment in Canada, down from 81% in 1967.

[3] The 701 unique parent-affiliate pairs are referred to as *firms* in the remainder of this paper unless specified as *parents* or *affiliates*.

In section II, we describe the econometric techniques used in the paper. Section III describes the data set used in the estimation, and section IV contains the empirical results for each of the four trade flows. Section V summarizes our main findings and concludes.

II. Econometric Framework

The basic regression model used in our analysis is as follows:

$$Y_{it} = \beta_0 + \beta_{i1}CT_{it} + \beta_{i2}UT_{it} + \beta_{i3}\text{Trend}_t + \Gamma Z_{it} + \nu_{it} \quad (1)$$

$$i = 1, 2, \ldots, N; \quad t = 1, 2, \ldots T$$

where

$$UT_{it} = \sum_{j=1}^{J} I(i \in j)UT_{jt}$$

and

$$CT_{it} = \sum_{j=1}^{J} I(i \in j)CT_{jt}$$

and Y_{it} is one of the following four trade flows for U.S. parent-Canadian affiliate pair i:

1. sales from Canadian affiliate i to its U.S. parent
2. sales from Canadian affiliate i to unaffiliated buyers in the United States (arms-length sales)
3. sales from U.S. Parent i to its Canadian affiliate
4. domestic ("Canadian") sales of Canadian affiliate i

CT_{it} and UT_{it} are Canadian and U.S. tariffs in industry j at time t. The $I(i \in j)$ are indicator dummies for whether firm i belongs to industry j ($j = 1, \ldots, J$), so that the summation term picks up the relevant tariffs. The vector Z_{it} contains exogenous variables representing transportation costs, demand, relative factor costs, and other miscellaneous time effects. Specifically, Z_{it} includes real U.S. and Canadian GDP and manufacturing wages, a measure of transportation costs in industry j at time t, a time trend, and—to capture time effects on the cost of capital—we include real U.S. interest rates and Canadian and U.S. price-earnings ratios. Variable selection and measurement is discussed in greater detail in section III.

Because interest lies in examining the extent to which variation in MNC responses to trade liberalization are explained by firm and industry characteristics, the model must allow the slope coefficients on the U.S. and Canadian tariff coefficients to differ across firms and industries. Hence, β_{i1} and β_{i2}, the firm-specific coefficients on the Canadian and U.S. tariffs in equation (1), are specified as

$$\beta_{i1} = \beta_1 + \mu_{i1} \quad \mu_{i1} \sim N(0, \sigma_1^2)$$

$$\beta_{i2} = \beta_2 + \mu_{i2} \quad \mu_{i2} \sim N(0, \sigma_2^2)$$

where β_1 and β_2 are the mean tariff coefficients in the population of firms, and the μ_i's capture across-firm heterogeneity in tariff responses.

In this type of model, we can estimate the population mean for each β and the variances of the β_i's and test the hypothesis that the variances of the β_i's are equal to zero (Hsaio, 1986). For the purposes of our study, this is not an interesting hypothesis. Trade theory suggests that all the variance in the β_i's should be across industry, not that it should be zero. We can test this hypothesis by constructing estimates of the individual firm betas, a posteriori, and then by decomposing these betas into across- and within-industry components.[4]

In our estimated model, we also allow for heterogeneity in the time trend. Because the U.S. and Canadian tariffs have strong trends, it is possible that heterogeneity in the tariff coefficients might simply pick up the effect of unobserved time-varying factors (such as changes in technology or demand) on individual firms. (For example, because some firms grow over time due to unobserved forces, a random coefficient on any trending variable like tariffs would tend to pick up this feature of the data.) Allowing the trend coefficient to be random deals with this potential problem.

Finally, we specify the error term v_{it} in equation (1) as consisting of two components: $v_{it} = \phi_i + \varepsilon_{it}$, where $\phi_i \sim N(0, \sigma_\phi^2)$ is a vector of unobserved time-invariant firm-specific characteristics, while ε_{it} is assumed to vary over time and across firms. The variance of the firm-specific error component indicates how much of the across-firm variation in the trade flows is due to unobserved characteristics of firms. Incorporating the error component and three random coefficients into equation (1) yields the random effects model (2):

$$Y_{it} = \beta o + (\beta_1 + \mu_{i1})CT_{it} + (\beta_2 + \mu_{i2})UT_{it} + (\beta_3 + \tau_i)\text{Trend}_t + \beta_4 Z_{it} + \phi_i + \varepsilon_{it} \quad (2)$$

where $\tau_i \sim N(0, \sigma_\tau^2)$ is the random component of the time trend coefficient. Finally, we specify the time-varying error component as an AR(1) process: $\varepsilon_{it} = \rho\varepsilon_{it-1} + \eta_{it}$, where η_{it} is i.i.d. over time and across firms. If the time-varying

[4] The procedure works as follows. Adopt the prior that each firm-specific coefficient is distributed in the population with the mean and variance given by our estimates. Then, apply Bayes rule to form the posterior density of the vector of firm-specific parameters. For each firm, solve for the vector of firm-specific parameters that maximizes this posterior density conditional on the firm's observed behavior. We treat this as the a posteriori estimate of the firm-specific parameter values. Given such estimates for each firm, we can find the average of the firm-specific coefficients within each narrowly defined manufacturing industry; we call these the *industry effects*. We can then decompose the total variance of the firm-specific parameters into across- and within-industry variation.

error component is autocorrelated and we fail to account for it, the importance of the time-invariant error components will tend to be exaggerated, because those components would be the only means by which the model could generate persistence. Thus, given our focus on examining properties of the time-invariant error components, accounting for other sources of persistence (such as adjustment lags) is important.

Because some MNCs organize their U.S. and Canadian production to produce and sell all output locally (for instance, affiliates may produce and sell all their output in Canada and not export to, or import from, the United States), a significant proportion of the Y_{it}'s are equal to zero. Therefore, the use of OLS to estimate equation (2) is inappropriate. Rather, we treat equation (2) as a random-effects Tobit model and estimate it using maximum likelihood. Tobit models have the following form:

$$\begin{aligned} y_{it}^* &= \beta' z_{it} + v_{it} \\ y_{it} &= 0 \qquad \text{if } y_{it}^* \leq 0, \\ y_{it} &= y_{it}^* \qquad \text{if } y_{it}^* > 0. \end{aligned}$$

where z_{it} represents the vector of all covariates. In Tobit models, a change in z_{it} has two effects. It affects the conditional mean of y_{it}^* in the positive (nonzero) part of the distribution, and it affects the probability that the observation falls in that part of the distribution (Judge et al., 1985; Greene, 1993). So, given the censoring, the effect of z_{it} on y_{it} is

$$\frac{\partial E[y_{it}|z_{it}]}{\partial z_{it}} = \beta\Phi\left(\frac{\beta' z_{it}}{\sigma}\right)$$

where Φ is a standard normal distribution function and $\Phi(\cdot)$ is the probability that y_{it} is in the uncensored region.

In a censored regression, OLS yields biased parameter estimates. If the model were estimated using only the observations with nonzero y_{it} values, then only terms for which $v_{it} > -\beta z_{it}$ would be included in the sample. This results in a truncated normal distribution with a nonzero mean for the error term. To assess the extent of the bias, we report OLS estimates for the four dependent variables.

III. Data

The data set used in this paper is from the Benchmark and Annual Surveys of U.S. Direct Investment Abroad administered by the Bureau of Economic Analysis. These surveys are the most comprehensive data available on the activities of U.S.-based MNCs and their foreign affiliates. For this study, we use the BEA data disaggregated at the individual foreign-affiliate level for each MNC from 1983 to 1992; to isolate firm and industry effects, we use only single-industry affiliates (those affiliates that reported sales in only one industry). Because trade theory makes no predictions about production location of nonmanufacturing industries and because many nonmanufacturing industries include non-tradeables (such as retailing, real estate, and hotels), we use only manufacturing affiliates. The BEA data includes fifty manufacturing categories, and appendix A provides descriptions of these industries along with mean tariffs and transportation costs. The sample was also screened to include only majority-owned affiliates.

Several alterations were made to the original sample to construct the panel. First, because the BEA conducts two different surveys (the Benchmark and Annual Surveys) with different reporting requirements in terms of affiliate size, reported data are not available for all the affiliates throughout the ten-year period. In particular, the Benchmark Surveys, conducted in 1977, 1982, and 1989, include the whole population of MNCs and their foreign affiliates, and smaller affiliates are required to report. But, in the Annual Surveys, many of the small affiliates that reported data in the 1982 and 1989 Benchmark Surveys are exempt from filing. In cases in which affiliates report data in a Benchmark Survey but are exempt from the Annual Surveys, the BEA carries them forward by estimating data.[5] As a result of this sampling procedure, most of the observations for smaller affiliates were estimated data for most of the ten-year period. We decided to remove affiliates for which most of the data was estimated rather than reported.

The initial screen was to exclude from the data set those affiliates that were observed multiple times but had only one reported (that is, not estimated) observation.[6] The next step was to remove any estimated data for an affiliate that appeared prior to its first reported observation or subsequent to its last reported observation. The rationale for this is as follows: as we noted above, if an affiliate drops out of the sample because it becomes exempt from reporting, the BEA carries the affiliate forward by estimating data. Because the Annual Survey data contains data on affiliates carried forward from the 1982 and 1989 Benchmark Surveys, for many affiliates, the data observed in the middle of the sample period is largely reported, whereas the data at the beginning and end of the sample period is largely estimated. Thus, we decided to eliminate the estimated data at the beginning and end and keep only the middle observations. After both screens, the total number of firm-year observations was reduced from 5,687 to 3,203—of which only 53 were estimated data points.[7,8]

Data were removed four more times to arrive at the final sample. First, affiliates in the same industry with the same parent were combined. Second, in 1987, SIC codes were

[5] Note that the individual affiliates that are carried forward are small and are thus are not likely to have a significant impact on the BEA's published data at the industry or country level.

[6] Note that an affiliate that is observed only once but with reported (that is, not estimated) data survives this screen.

[7] Recall that a "firm-year" observation is a parent-affiliate pair observed in a given year

[8] Note that the 53 estimated data points that remained in the sample after the initial screens were all bracketed by reported observations.

revised for many industries, which resulted for the most part in codes being merged together.[9] After recoding more than a dozen industries for the entire sample period, affiliates in the same industry with the same parent were merged a second time. These two screens reduced the sample to 2,939 observations. Finally, observations were removed for two industries in which there were no Canadian tariff data, and observations were removed in which affiliates reported zero sales. These final screens produced the sample used in this study with 2,881 firm-year observations on 701 firms in fifty manufacturing industries.

Variables were included in the model to represent relative factor prices, relative demand, tariffs, and transportation costs. First, we include a ratio of real Canadian-to-U.S. manufacturing wages (*C/U WAGE*). Then, rather than attempting to measure the cost of capital in the two countries, we include both U.S. and Canadian price-earnings ratios and the U.S. real interest rate in the model, while remaining agnostic about how these map into the cost of capital for the MNC and its affiliate.[10] Only the U.S. interest rate was used due to the almost perfect collinearity of the two interest rates over the sample period. All nominal variables (including the dependent variables) were normalized to 1990 U.S. CPI dollars. Canadian dollars were converted to U.S. dollars using annual nominal exchange rates from the *IMF International Financial Statistics Yearbook*.[11]

Another natural control variable to capture changes in relative factor prices would be a real exchange rate (that is, the nominal exchange rate times the ratio of the producer price indices (PPI) for the two countries). However, during our sample period, the PPI ratio is highly correlated with the nominal wage ratio. This induces a very high correlation (0.90) between the real exchange rate and the real-wage ratio, such that inclusion of both variables caused severe collinearity problems. Note that the movements in our real-wage ratio variable are largely driven by changes in the nominal exchange rate. This moves substantially over the sample period, with the Canadian dollar falling from 81 U.S. cents in 1983 to 72 cents in 1986, rising to 87 cents in 1991, and falling to 83 cents in 1992.

U.S. and Canadian real GDP were used as demand shifters. Again, to capture relative changes in demand, we use a ratio of Canadian to U.S. GDP, *C/U-GDP*. To better account for differences in demand cyclicality across different industries, *C/U-GDP* was interacted with the broad industry group of the affiliate: industrial intermediate goods (*II*), industrial machinery (*IM*), consumer durables (*CD*) and consumer nondurables (*CN*). The industry groupings are shown in appendix B. All variables in dollar values were normalized to 1990 CPI dollars, and Canadian dollars were converted to U.S. dollars.

U.S. and Canadian tariffs were measured as annual ratios of the value of duties paid in the United States (Canada) on imports of Canada (U.S.) goods in industry j at time t divided by the total value of imports to the United States (Canada) from Canadian (U.S.) importers in industry j at time t.[12] Similar measures at different levels of aggregation have been widely used in empirical work (Caves, 1990; Grubert & Mutti, 1991). Although the tariff measures used here do not reflect nontariff barriers and are still at a more aggregated level than that at which tariffs are actually imposed, they are more disaggregated than are those measures which are typically used in empirical work (Grubert & Mutti, 1991) and are longitudinal. During the ten-year period in this study, U.S. and Canadian tariffs dropped by approximately 62.5%, the latter dropping from an average of nearly 8% to 3% and the former dropping from 4% to less than 1.5%. There is also considerable cross-sectional variation in tariffs, as can be seen in appendix A. Mean U.S. tariffs for the ten-year period are 3.1%, ranging from a high of 14.7% for tobacco products to a low of 0.11% in motor vehicles and equipment. Similarly, Canadian tariffs average 6.12% over all industries for the sample period and range from a high of 30.27% in the beverages industry to 0.05% in agricultural chemicals. The striking cross-sectional and longitudinal characteristics of the U.S. and Canadian tariff structure indicate how much can be gained by using disaggregated, longitudinal measures.[13]

[9] If two codes existed for two similar industries prior to 1987 and the codes were merged into a single code after 1987, we used the post-1987 code for the entire ten-year period. Similarly, if one code was broken into two after 1987, we used the pre-1987 code for the entire sample period.

[10] If the cost of capital in the United States and Canada differ, say, due to segmented financial markets, then inclusion of separate cost-of-capital variables for each country would be important, as the cost of capital in each country would affect the capital intensity of production in that country. But we view such a scenario as implausible, at least for MNCs. Nevertheless, even if there were no cost-of-capital differences between the two countries (that is, because capital flows are unrestricted)—or if any such differences are irrelevant to MNCs (because they can raise capital in either market)—it remains true that the absolute level of the cost of capital may affect the production location decisions of MNCs. (For example, if world cost of capital rises, an MNC may shift to more labor-intensive production and simultaneously shift production to lower-wage countries). Such an effect would be picked up jointly by the three cost-of-capital variables we have included.

[11] Real U S. interest rates were obtained by subtracting annual inflation rates from average yields on AAA corporate bonds Bond rates were obtained from Moody's Industrial Manual and inflation rates were obtained from the Survey of Current Business. U S. manufacturing wage rates were also obtained from the Survey of Current Business. Canadian manufacturing wage rates were obtained from Statistics Canada's Canadian Economic Observer. U.S. P-E ratios were taken from Citibase, and annual data are twelve-month averages of quarterly data. Canadian P-E ratios were obtained from the Bank of Canada Review.

[12] Tariff and transportation-cost SIC codes matched the codes in which the trade flows were reported. Recall that only single-industry affiliates were included in the sample. However, because most of the U.S. parents were large and diversified companies, it was not possible to limit the sample to single-industry parents as well. It was therefore assumed that parent trade flows were in the same industry as the affiliate. In more than half the cases, the parent's major SIC code was the same as the industry of the affiliate.

[13] U S. tariff data were obtained from the United States Census Bureau, and Canadian tariff data were obtained from Statistics Canada. Canadian tariffs were reported in three-digit Canadian SIC codes and had to be converted to U.S. SIC codes. SIC codes were then converted to ISI codes

TABLE 1.—CONDITIONAL AND NORMALIZED MEANS AND VARIANCES

	CA Sales to U.S. Parent	CA Arms-Length U.S. Sales	U.S. Parent Sales to CA	CA Sales in Canada
Mean	19092	5526	13311	46108
Variance	3.9×10^{10}	6.2×10^{8}	2.1×10^{10}	1.8×10^{10}
Conditional mean[a]	30457	16463	17704	47783
Conditional variance	6.3×10^{10}	1.7×10^{9}	2.7×10^{10}	1.3×10^{8}
Cond. normalized[b] mean	0.1412	0.2323	0.1495	0.7540
Cond. normalized variance	0.0561	0.0858	0.0310	0.1271

[a] Means and variances are conditional (calculated using only nonzero observations)
[b] Normalized means and variances are also conditional and are calculated by dividing the trade flow by each affiliate's total sales averaged over the sample period

An annual measure of transportation costs for each three-digit industry is based on data from the U.S. Census Bureau on freight and insurance charges as reported by exporters to the United States. The data can be further disaggregated by country of exporter. Thus, the measure used here includes the costs for Canadian exporters in each industry j into the United States. Because similar data is not available on the cost of exporting goods to Canada from the United States and no systematic differences in transportation costs were assumed to exist, the same measure is used for sales from U.S. MNC parents into Canada. The measure is expressed as a ratio of the value of shipment costs and insurance on imports in industry j at time t to the total value of shipments. Transportation costs average 1.03% across industries over the sample period, and range from 8.44% in agricultural chemicals to close to 0 in petroleum products. A priori, it was expected that higher transportation costs would reduce cross-border trade and increase the domestic sales of Canadian affiliates.

Finally, because affiliates in the sample varied significantly in terms of size, the four dependent variables were normalized to mitigate problems of heteroskedasticity. For each affiliate, the four trade flows were divided by the mean total sales of the affiliate averaged over the sample period. However, results were not sensitive to this normalization. Table 1 gives the means and variances and conditional means and variances (using only the nonzero observations) and normalized means and variances.

IV. Empirical Results

A. *OLS and Tobit Estimates*

We describe our empirical results in three sections. First, we contrast the OLS and Tobit estimates. Then, we discuss the economic meaning of the estimated tariff effects, and, finally, we discuss the variance decompositions. Table 2 shows the OLS and Tobit results for all four trade flows. The last two rows in table 2 give the R-squared values for the OLS regressions and the number of nonzero observations

used by the BEA. *ISI* codes correspond roughly to the two- or three-digit SIC code level. Correspondence tables are available from the authors

for each trade flow. As can be seen in the last row, the proportion of nonzero observations differs considerably between the four flows. For Canadian affiliate arms-length sales to the United States, only 33.6% of the observations are nonzero, but for affiliate sales to U.S. parents, 62.7% of the observations are positive. This finding is consistent with Rugman's (1990) examination of U.S.-Canada trade patterns in which he notes that approximately 70% of bilateral trade in manufactured products is accounted for by intrafirm sales of MNCs.

The Tobit and OLS parameter estimates are very different. From the discussion in the previous sections, two likely sources of bias are affecting the OLS results. First, the truncation of the dependent variable at zero results in an error distribution with a nonzero mean which depends on β, σ, and x, and which is different for every observation. The second potential source of bias arises from the constraint of equal coefficients in the OLS estimates. In the Tobit panels in table 2, nine of the twelve random coefficients have standard deviations which are significantly greater than zero. This implies that constraining these coefficients to be equal across firms may result in biased estimates.

In table 2, panel 1, columns 1 and 2 report the Tobit and OLS estimates for Canadian affiliate sales to U.S. parents. As expected, the U.S. tariff coefficients are negative and significant, and the standard deviation of the random U.S. tariff coefficient is significant in the Tobit estimate. An interesting difference in the OLS and Tobit estimates for Canadian affiliate sales to U.S. parents is the marginal significance in the OLS model of all the GDP and relative wage coefficients. None of these estimated parameters are significant in the Tobit model.

Columns 1 and 2 of panel 2 report the Tobit and OLS results for Canadian affiliate arms-length sales to the United States. As expected, the U.S. tariff coefficient is significant and negative in both models, but the standard deviation of the random U.S. tariff coefficient is not significantly greater than zero. A comparison of the Tobit and OLS results shows striking differences in the estimates for the other parameters. In particular, the Canadian tariff coefficient, which is positive and insignificant in the OLS results, becomes significant and negative in the Tobit results. Five other coefficients flip signs as well.

Columns 1 and 2 of panel 3 give the Tobit and OLS results for U.S. parent sales to Canadian affiliates. The Canadian tariff coefficient is negative but achieves only modest significance levels in both models, and the standard deviation of the random Canadian tariff coefficient is also not significantly greater than zero. As we discuss in the next section, the magnitude of these estimates is quantitatively quite small, and this contradicts with the view that trade liberalization would "hollow out" Canadian manufacturing. The U.S. tariff coefficient is insignificant in both the OLS and Tobit models.

TABLE 2.—TOBIT AND OLS RESULTS FOR TRADE FLOWS

	CA Sales to U.S. Parents		CA Arms-Length Sales to U.S.		U.S. Parent Sales to CA		CA Sales in Canada	
	TOBIT	OLS	TOBIT	OLS	TOBIT	OLS	TOBIT	OLS
CAN-TARIFF-β_1	0.0016	0.0032[c]	−0.0132[c]	0.0013	−0.0024	−0.0016[b]	0.0045	−0.0015
	(0.0022)	(0.0009)	(0.0049)	(0.0009)	(0.0018)	(0.0008)	(0.0035)	(0.0017)
US-TARIFF-β_2	−0.0194[c]	−0.0175[c]	−0.0137[a]	−0.0125[c]	0.0028	0.0019	0.0199[c]	0.0322[c]
	(0.0040)	(0.0020)	(0.0077)	(0.0020)	(0.0034)	(0.0017)	(0.0069)	(0.0038)
C/U-WAGE	−0.0017	−0.0092[b]	−0.0027	−0.0023	−0.0003	0.0028	0.0032	0.0068
	(0.0021)	(0.0045)	(0.0077)	(0.0045)	(0.0028)	(0.0037)	(0.0060)	(0.0084)
C/U-GDP IND = CN	0.0018	0.0831[a]	−0.0261	0.0030	0.0008	−0.0332	0.0132	−0.0333
	(0.0239)	(0.0428)	(0.0732)	(0.0433)	(0.0265)	(0.0356)	(0.0577)	(0.0804)
C/U-GDP IND = CD	0.0118	0.0866[b]	0.0124	0.0158	0.0107	−0.0262	−0.0020	−0.0473
	(0.0237)	(0.0428)	(0.0733)	(0.0434)	(0.0268)	(0.0356)	(0.0575)	(0.0805)
C/U-GDP IND = II	0.0144	0.0899[b]	0.0024	0.0101	0.0031	−0.0286	0.0003	−0.0427
	(0.0234)	(0.0427)	(0.0730)	(0.0433)	(0.0266)	(0.0355)	(0.0574)	(0.0803)
C/U-GDP IND = IM	0.0154	0.0891[b]	−0.0036	0.0089	0.0146	−0.0230	0.0068	−0.0413
	(0.0235)	(0.0428)	(0.0737)	(0.0434)	(0.0266)	(0.0356)	(0.0573)	(0.0805)
TRANS COST	0.0121[c]	0.0210[c]	−0.0079	0.0036	−0.0131[c]	−0.0129[c]	−0.0135[b]	−0.0305[c]
	(0.0041)	(0.0035)	(0.0142)	(0.0036)	(0.0037)	(0.0029)	(0.0052)	(0.0066)
TREND	0.0003	0.0186	−0.0122	0.0002	0.0074	0.0026	0.0076	−0.0022
	(0.0063)	(0.0127)	(0.0215)	(0.0129)	(0.0077)	(0.0106)	(0.0165)	(0.0239)
US INT. RATE	−0.0050	0.0118	−0.0265[a]	−0.0099	0.0038	−0.0001	−0.0071	−0.0159
	(0.0058)	(0.0114)	(0.0160)	(0.0116)	(0.0062)	(0.0095)	(0.0129)	(0.0215)
CAN P-E RATIO	−0.0003	−0.0096	−0.0018	−0.0065	0.0006	0.0037	0.0078	0.0056
	(0.0033)	(0.0094)	(0.0149)	(0.0095)	(0.0063)	(0.0078)	(0.0126)	(0.0177)
US P-E RATIO	−0.0011	−0.0049	0.0005	−0.0027	−0.0004	−0.0002	−0.0015	−0.0012
	(0.0018)	(0.0050)	(0.0078)	(0.0050)	(0.0033)	(0.0041)	(0.0068)	(0.0093)
Constant-β_0	0.1432[a]	0.1920	0.4084	0.3889	−0.0047	0.0725	0.2911	0.4669
	(0.0829)	(0.2761)	(0.3779)	(0.2797)	(0.1613)	(0.2296)	(0.3352)	(0.5191)
$\sigma(\mu_{i1})$	0.0015		0.0136[c]		0.0003		0.0070[a]	
	(0.0035)		(0.0039)		(0.0026)		(0.0043)	
$\sigma(\mu_{i2})$	0.0136[c]		0.0056		0.0080[b]		0.0292[c]	
	(0.0040)		(0.0162)		(0.0035)		(0.0058)	
$\sigma(\phi_i)$	0.1522[c]		0.2989[c]		0.1301[c]		0.1881[c]	
	(0.0087)		(0.0250)		(0.0069)		(0.0218)	
$\sigma(\tau_i)$	0.0096[c]		0.0160[b]		0.0133[c]		0.0271[c]	
	(0.0031)		(0.0073)		(0.0018)		(0.0036)	
AR(1)ρ	0.6505[c]		0.5093[c]		0.3316[c]		0.5385[c]	
	(0.0309)		(0.0425)		(0.0223)		(0.0226)	
Model error ($\sigma\varepsilon_i$)	0.1162[c]	0.1941	0.1929[c]	0.1966	0.1048[c]	0.1614	0.2434[c]	0.3649
	(0.0013)		(0.0033)		(0.0010)		(0.0024)	
R-Squared		0.0584		0.0568		0.0552		0.06531
Number of nonzero observations		1806		967		2166		2780

Sample size is 2,881 for all estimations
Numbers in parentheses are standard errors
[a] Significant at the 10% level
[b] Significant at the 5% level
[c] Significant at the 1% level

Finally, columns 1 and 2 of panel 4 give the results for Canadian affiliate sales in Canada. In the Tobit results, both sets of tariffs are positive (U.S. tariffs are significant), and all the random coefficients have significant standard deviations. The main contrast with the OLS results is that there the Canadian tariff coefficient is of the opposite sign and insignificant. From the last row of table 2, we find that the number of zero observations is small for this trade flow. Thus, it seems likely that most of the bias in the OLS results for this trade flow arises out of the equal-coefficients restriction rather than from the truncated distribution.

Before concluding this discussion, we note three common features of the full set of results. First, in all the Tobit models, the standard deviation of ϕ_i, the firm-specific error component, is significant at the 1% level, indicating that unobserved characteristics of MNCs account for a significant portion of the variation in the levels of the trade flows across firms that we documented in figure 1. Second, as expected, the standard deviation of the random time trend, τ_i, was also significant at the 1% level in all of the Tobit estimates. This further implies that unobserved characteristics of MNCs explain a significant part of the firm-specific variation in the trade flows over time. However, our estimates of the means and variances of the tariff coefficients were remarkably insensitive to whether the random time trend was included in the model. Third, the AR(1) coefficients range from 0.33 to 0.65 and are all highly significant. These values are in a sense rather small, implying that

TABLE 3.—EXPECTED SIGN PATTERNS AND COEFFICIENT ESTIMATES FOR THE TRADE FLOWS

		CA Sales to U.S. Parent	CA Arms-Length U.S. Sales	U.S. Parent Sales to CA	CA Sales in Canada
Expected Signs	Canadian Tariff[a]	?	?	—	?
	U.S. Tariff[b]	—	—	?	?
Estimates	Canadian Tariff	+0.0016	−0.0132[c]	−0.0024	+0.0045
	U.S. Tariff	−0.0194[e]	−0.0137[c]	+0.0028	+0.0199[e]

[a] The sample mean of the Canadian tariff is 5.852, and the sample standard deviation is 5.538.
[b] The sample mean of the U.S. tariff is 2.967, and the sample standard deviation is 2.459.
[c] Significant at the 10% level.
[d] Significant at the 5% level.
[e] Significant at the 1% level.

transitory deviations in the dependent variable largely die out in a couple of years. Finally, most of the control variables were insignificant. The coefficients on transportation costs—the only control variable for which strong effects were expected a priori—were significant for two of the trade flows. However, the signs of the transportation-cost coefficients were in the opposite direction as expected for Canadian affiliate sales to U.S. parents and for Canadian affiliate sales in Canada. (This result is explained in greater detail in the next section.)

B. *Tariffs and Production Location*

Table 3 gives the expected sign patterns for U.S. and Canadian tariffs and the Tobit coefficient estimates for the four trade flows. For the three two-way trade flows (Canadian affiliate sales to U.S. parents, Canadian affiliate arms-length sales to the United States, and U.S. parent sales to Canadian affiliates), there is a strong prediction for the sign of the tariff of the receiving country. We would expect lower Canadian tariffs to lead U.S. MNCs to increase their sales from the United States to Canada. Similarly, lower U.S. tariffs should increase Canadian affiliate sales to the United States.

These simple two-way predictions were borne out for the three two-way trade flows. Using the sample means of the trade flows and the estimated coefficients, we can calculate the effect of tariff changes. For example, a one-percentage point drop in the U.S. tariff increases Canadian affiliate sales to U.S. parents by 13.8%. Interestingly, the effect of a drop in U.S. tariff levels is smaller for affiliate arms-length sales to the United States. For arms-length sales, a one-percentage point drop in the U.S. tariff raises sales to the United States by 5.9%.

Although the U.S. tariffs were significant and negative for both trade flows into the United States, several other results for the intrafirm and arms-length trade flows differ. First, transportation costs were unexpectedly significant and positive for Canadian affiliate sales to U.S. parents (see panel 1 in table 2) and were insignificant and negative for arms-length sales. Examination of the cross-sectional characteristics of affiliate exports reveals that many are concentrated in industries with high transportation-costs such as: pulp, paper, and board mills (ISI 262); lumber and wood products (ISI 240); and stone, clay and concrete (ISI 329). It appears that Canada has a sufficiently large advantage in resource-based sectors to compensate for the relatively high cost of transporting the products.

A second difference in the results for intrafirm and arms-length sales is the positive (but marginally insignificant) Canadian tariff coefficient in the former and significant negative coefficient in the latter. A priori, we would not expect to see differences in the impact of Canadian tariffs on affiliates' decisions to export goods back to their U.S. parents or to unaffiliated buyers in the United States. It must be that underlying differences in MNC technologies lead to the systematic differences in the export patterns that we observe here. Indeed, an examination of the cross-sectional characteristics of the intrafirm and arms-length trade flows reveals a greater concentration of manufactured and finished goods (such as farm machinery, construction machinery, and several types of refined chemical products) being exchanged intrafirm as contrasted with a higher concentration of resource-based goods (such as primary metals and paper) being sold to unaffiliated U.S. buyers.[14]

A more interesting and unexpected result is the negligible effect of Canadian tariff reductions on U.S. parent sales to Canadian affiliates. Indeed, the coefficient on the Canadian tariff variable was unexpectedly small (−0.0024), which implies a one-percentage point reduction in Canadian tariffs increases U.S. parent sales to Canadian affiliates by only 1.6% on average. Furthermore, this coefficient attains significance only at the 20% level. This result is surprising because of commonly expressed concerns in Canada that industry would be "hollowed out" if trade with the U.S. were liberalized (Crookell, 1990). Hollowing-out implies that, rather than exiting a market, firms maintain some local presence but import most of their value-added parts from abroad. In the case of U.S.-Canada trade, if U.S. MNCs had indeed hollowed out their Canadian operations, we would expect to see considerable increases in sales from U.S. parents to Canadian affiliates (along with probable reductions in sales from Canada to the United States) as a result of trade liberalization. This was not borne out in our findings.

[14] More detail on the composition of the trade flows is available from the authors.

Finally, table 3 gives the estimates for the Canadian and U.S. tariff coefficients for Canadian affiliate sales in Canada. The U.S. tariff coefficient is positive and significant. A priori, there were no strong predictions for the effect of either tariff on Canadian affiliates' domestic sales. Our results indicate that a one-percentage point reduction in U.S. tariffs reduced affiliates' Canadian sales by 3.7%. This, combined with our previously noted findings that U.S. tariff reductions were associated with increases in affiliates' sales to U.S. parents of 13.9% and to unaffiliated U.S. buyers of 5.9%, suggests that with U.S. tariff reductions the affiliates became more export-oriented. This is intuitively plausible if MNCs restructured to integrate U.S. and Canadian production as a result of tariff reductions.

The positive Canadian tariff coefficient would seem to support the tariff wall theory that, if tariffs are set high enough, firms substitute domestic production for imports. But this coefficient is insignificant and small in magnitude. It is true that some high-tariff Canadian industries would suffer from severe import penetration if those tariffs were lowered (good examples being the two highest tariff industries: alcohol and tobacco). But other high-tariff Canadian industries in which average affiliate sales are high include those industries that use resources in which Canada is abundantly endowed. In general, industries with high affiliate sales in Canada seem to fall into three categories: high-tariff, regulated industries such as tobacco and alcohol; resource-based industries such as food products, metals, furniture, and paper, in which either tariffs or transportation costs are high; and industries related to the production of automobiles (such as industrial chemicals, metal stamping, glass, and engines and turbines) or agriculture (such as farm and garden machinery). Many of the industries in the second and third categories correspond with those identified by Rugman (1989, 1990) as having country-specific and firm-specific advantages in Canada. For political reasons, Canadian tariffs may have been set high in industries that use abundant domestic resources to prevent firms from shipping out raw materials and providing value-added activities abroad. This would help to explain our earlier finding that lower Canadian tariffs are associated with increased arms-length sales to the United States—such sales being more likely in resource-based industries.

Our findings in this section showed that reductions in U.S. tariffs led to increased Canadian production for export to the United States but lower Canadian production for domestic sales. These findings are consistent with the pattern (for one industry) shown in charts B and D of figure 1, wherein the proportion of affiliates' total sales destined for the local market shrinks considerably between 1983 and 1992, but the proportion of affiliates' sales to U.S. parents increases by nearly as much.[15] Similarly, lower Canadian tariffs led to a small increase in U.S. sales into Canada, a larger increase in affiliate arms-length sales to the United States, and little reduction in Canadian production for domestic sales. Overall, we find that trade liberalization appears to have been trade-creating, as lower tariffs increased bilateral trade flows, but we find no evidence that reduced Canadian tariffs led to a substantial increase in U.S. parent sales to Canadian affiliates, as would be expected in a hollowing-out scenario.

A possible challenge to our findings arises because our random-effects Tobit model identifies tariff coefficients from both cross-sectional and over-time variation in tariffs and trade flows. It is possible that a negative cross-sectional correlation between tariffs and trade flows existed prior to trade liberalization, but that trade liberalization led to no change in trade flows. In that case, our estimated tariff coefficients in the trade flow equations are still negative, solely because of the negative cross-sectional correlation between tariffs and trade flows induced by the initial conditions.[16]

The typical way to deal with this type of problem is to estimate a fixed-effects model. The fixed effects would pick up time-invariant influences on the trade flows, and the tariff coefficients would pick up only the association between tariff changes and trade flow changes. Unfortunately, fixed-effects Tobit models are inconsistent with small T, which is the situation here, and their estimation is prohibitively computationally burdensome. However, in the present case, a sensible alternative is to include the initial period (1983) preliberalization tariff levels as additional covariates in the Tobit model. Suppose that, prior to trade liberalization, tariffs were set low in industries with large trade flows and vice versa. Suppose further that, as tariffs were lowered with trade liberalization, this basic trade flow pattern re-

[15] As mentioned earlier, the reorganization depicted in figure 1 is broadly representative of the industries in our sample. Indeed, Canadian affiliate sales to U.S. parents as a percentage of total affiliate sales increased from 8.36% in 1983 to 10.47% in 1992. Similarly, the ratio of Canadian affiliate arms-length U.S. sales as a percentage of total affiliate sales increased from 7.07% in 1983 to 9.35% in 1992. As in figure 1, Canadian affiliates' Canadian sales as a percentage of total sales fell from 76.75% in 1983 to 70.46% in 1992. Across all industries, the pattern is clearly one of increasing Canadian affiliate export-orientation. Bilateral trade also increased with trade liberalization: U.S. parent sales to Canadian affiliates as a percentage of total Canadian affiliate sales increased from 9.5% in 1983 to 12.87% in 1992.

[16] A scenario that would generate such a pattern is as follows. Suppose in each industry that there exists a characteristic organization of MNC production that is fixed over time. That is, some industries require large two-way trade flows (both intrafirm and arms-length) whereas others do not, and that the magnitude of these flows is stable over time (except for idiosyncratic, year-to-year fluctuations). Further suppose that, through the political process, preliberalization tariff levels were set low in industries that require large trade flows, and vice versa. This scenario has a negative cross-sectional correlation between tariffs and trade flows, but trade liberalization leads to no change in the trade flows. Another, perhaps more plausible, scenario arises if adjustment costs in production are substantial, so that trade flows respond very slowly to tariffs. Then, if preliberalization tariff levels were set low in industries with large trade flows, our tariff coefficients may be largely picking up the negative cross-sectional correlation induced by the initial conditions, rather than any change in trade flows over time. Note, however, that neither of these scenarios seems consistent with the data patterns exhibited in figure 1, or the aggregate statistics discussed in footnote 15.

TABLE 4.—U.S. AND CANADIAN TARIFF COEFFICIENT ESTIMATES CONDITIONAL ON 1983 TARIFFS

	CA Sales to U.S. Parents		CA Arms-Length U.S. Sales		U.S. Parent Sales to CA		CA Sales in Canada	
Canadian tariff	0.0016	0.0035	−0.0132[c]	−0.0115	−0.0024	−0.0026	0.0045	−0.0003
	(0.0022)	(0.0029)	(0.0049)	(0.0089)	(0.0018)	(0.0025)	(0.0035)	(0.0048)
U.S. tariff	−0.0194[c]	−0.0166[c]	−0.0137[a]	−0.0105	0.0028	0.0001	0.0199[c]	0.0169[b]
	(0.0040)	(0.0052)	(0.0077)	(0.0104)	(0.0034)	(0.0042)	(0.0069)	(0.0080)
Canadian tariff (1983)	—	−0.0015	—	−0.0001	—	−0.0013	—	0.0025
		(0.0027)		(0.0074)		(0.0025)		(0.0049)
U.S. tariff (1983)	—	−0.0068	—	−0.0052	—	0.0068	—	0.0105
		(0.0063)		(0.0109)		(0.0042)		(0.0085)

[a] Significant at the 10% level
[b] Significant at the 5% level
[c] Significant at the 1% level

TABLE 5.—PERCENTAGE OF VARIANCE IN RANDOM EFFECTS EXPLAINED BY WITHIN- AND ACROSS-INDUSTRY EFFECTS

	CA Sales to USP		CA A-L U.S. Sales		USP Sales to CA		CA Canadian Sales	
	Within	Across	Within	Across	Within	Across	Within	Across
Canadian tariff $\sigma\mu_{i1}$	82.6%	17.4%	83.4%[a]	16.6%[a]	78.1%	21.9%	80.5%[a]	19.5%[a]
U.S. tariff $\sigma\mu_{i2}$	84.3%[a]	15.7%[a]	83.4%	16.6%	83.2%[a]	16.8%[a]	85.3%[a]	14.7%[a]
Time trend $\sigma\tau_i$	82.3%	17.7%	78.5%	21.5%	86.3%	13.7%	73.2%	26.8%
Firm effect $\sigma\phi_i$	85.8%	14.2%	78.8%	21.2%	77.6%	22.4%	75.6%	24.4%

CA sales to USP are Canadian affiliates' sales to U.S. parents. CA A-L U.S. sales are arms-length sales, USP sales to CA are U.S. parents' sales to Canadian affiliates, and CA Canadian sales are the affiliates' sales in Canada.
[a] Denotes significant tariff-coefficient variances.

mained unchanged. In that case, the 1983 tariffs should be correlated with trade flows. But, conditional on the 1983 tariffs, the post-1983 contemporaneous tariff levels should be uncorrelated with the trade flows.

Table 4 reports the estimates of the tariff coefficients for models that include the 1983 tariff levels as additional control variables. Note that the 1983 tariff levels are in all cases insignificant. Furthermore, the coefficients on the contemporaneous tariffs are little affected by their inclusion. The key coefficients on U.S. tariffs in the equations for affiliate sales to the U.S. parent and the affiliate domestic sales equations remain highly significant. The current tariff coefficients in the equation for affiliate arms-length sales to the U.S. do lose their significance, but this is due primarily to an increase in their standard errors.

Based on these results (and the aggregate statistics reported in footnote 15), we conclude that there is strong evidence that trade flows are actually increasing as tariffs fall. We are not just picking up a cross-sectional correlation that arises because industries with larger trade flows had lower initial tariff levels prior to trade liberalization. In the next section, we examine the extent to which heterogeneity in MNC trade flow adjustments is explained by characteristics of the industries in which the MNCs operate or by idiosyncratic firm characteristics.

C. *Variance Decompositions*

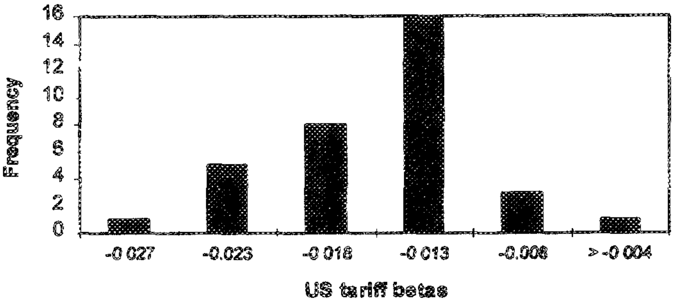
In section II, we indicated that, from estimates of the Tobit models, we could construct estimates of the individual firm and industry betas a posteriori. We can examine the relative magnitude of firm and industry effects in MNC responses to tariff changes by decomposing the variance of the random coefficients into within- and across-industry variance. Schmalensee (1985) used a similar approach to evaluate the relative contribution of firm and industry effects to the total variance in firm profitability. We depart here from the standard variance components models in that we decompose the variance in the random coefficients in addition to the unexplained error. This allows us to evaluate the relative importance of firm and industry effects in MNC responses to changing tariff levels.[17]

In table 5, we report the percentage of variance in the three random coefficients and the firm effect explained by within-industry and across-industry variation. As can be seen in panels 1 through 4, within-industry (firm) effects explain much more of the variance in the random tariff coefficients, time trend, and firm effect than do across-industry effects.[18] With regard to the tariff coefficients, across-industry effects explain only approximately 15% to

[17] It is important to note that our a posteriori estimates of the firm-specific betas are based on a fairly small number of observations per firm. Recall that there are 2,881 firm-year observations on 701 affiliates, so the average number of observations per affiliate is a little over four. The main reason for this is that small affiliates typically do not report data for every year in the sample period, as they are not always required to report. In using a Bayesian updating rule to estimate the firm-specific betas, we start from a prior mean on the betas that is the same for all firms. Because in many cases a fairly small number of observations are used to update those means, the posteriors are compressed towards the means. This reduces both the within- and across-industry variances of the firm-specific betas. However, we see no a priori reason to expect that either component of the variance would be relatively more compressed. Thus, we hope that our estimates of the fraction of variance due to each source are not misleading. Still, this issue should be revisited when a longer panel is available.

[18] We also estimated versions of our models in which the random coefficients were restricted to be homogeneous within industries (as trade theory would suggest). For all four trade flows, this led to very substantial deterioration in the likelihood functions, providing further evidence of

FIGURE 2.—FIRM-SPECIFIC U.S. TARIFF COEFFICIENTS FOR CANADIAN AFFILIATE SALES TO U.S. PARENTS: INDUSTRIAL CHEMICALS INDUSTRY



25% of the variance across firms. These results imply that firms within the same industry respond quite differently to tariff changes, presumably because of idiosyncratic firm characteristics such as differences in technology or organization. Such a finding is counterintuitive in the context of trade theory, in which factor-based or technological characteristics of industries determine adjustment patterns.[19] Of course, if industries were defined more narrowly, the across variance would increase relative to the within variance, and in the limit each firm defines its own industry and all variance is across. But it is also true that, as industries are defined successively more narrowly, both neoclassical and IO-based trade theories become meaningless.

To illustrate the heterogeneity of firm betas in a single industry, figure 2 shows the U.S. tariff coefficients in the industrial chemicals industry for Canadian affiliate sales to U.S. parents. The horizontal axis represents the size of the estimated betas (the range in this industry is from −0.04255 to −0.0037, and the mean industry beta is −0.01879). The vertical axis gives the frequency of firms with betas within the specified range. Recall that the overall mean beta across all firms in all industries is −0.0194.

Although there is much less variance across industries in tariff coefficients, some interesting differences do emerge among the three two-way trade flows. For instance, from a factor-proportions standpoint, one would expect to see Canadian affiliate sales to the United States (both arms-length and to U.S. parents) increase in industries in which Canada is relatively factor abundant. And, indeed, the U.S. tariff coefficients for Canadian affiliate sales to the United States (both to U.S. parents and arms-length buyers) are among the largest in magnitude in: furniture and fixtures (250); pulp, paper, and board mills (262); paper and allied products (265); leather products (310); and bakery products (205). These industries correspond with those identified by Leamer (1984) as resource-abundant industries in Canada. However, other industries with high U.S. tariff coefficients for Canadian affiliate sales to the United States (both arms-length and to U.S. parents) include: preserved fruits and vegetables (203); textile mill products (220); and soap, cleaners, and toilet goods (284). It is not obvious that Canada would have a comparative advantage in these industries, but these industries do have in common the feature of high initial U.S. tariffs.

V. Discussion and Conclusions

The results presented in the previous sections clearly demonstrate the importance of tariff reductions to MNC production-location decisions. Reductions in U.S. tariffs led to greater affiliate production for sales into the United States (both to parents and to unaffiliated buyers) and reductions in affiliates' Canadian sales. Similarly, reductions in Canadian tariffs had a positive but unexpectedly small relation with U.S. parents' sales into Canada. The surprisingly small impact of the Canadian tariff on U.S. MNC sales into Canada contradicts the conventional wisdom in Canada that free trade with the United States would lead to a hollowing-out of Canadian industry.

Additionally, we find that firms within narrowly defined industries responded quite differently to tariff changes. Such a pattern has not previously been demonstrated empirically. From a theoretical standpoint, this result should not simply be interpreted to mean that adjustments to trade liberalization were primarily of the intraindustry type originally modeled by Grubel (1970) and others. Theories of intraindustry trade still predict that industry characteristics such as economies of scale and existence of differentiated products will be the primary determinant of adjustment patterns. We would therefore expect, if these theories held, that there would be greater differences between industries in adjustments than within industries (because technologies differ from industry to industry). Instead, our finding that firm characteristics explained more of the variance in adjustments points to a potentially different explanation for production-location choices: one based upon characteristics of firms such as their international configuration of technology. From a policy standpoint, such a finding implies that government action that is designed to protect vulnerable sectors from trade liberalization might be altering patterns of domestic competition rather than helping entire industries.

substantial within-industry heterogeneity. The results of the industry-level estimations are available upon request from the authors.

[19] To put the results in a broader context, one might note, for instance, that human-capital earnings functions typically can explain only 20% of the variance of log earnings across individuals based on education and experience differences, and yet the human-capital theory is considered quite powerful.

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APPENDIX A

INDUSTRY AVERAGE U.S. AND CANADIAN TARIFFS AND TRANSPORTATION COSTS

IND	Industry Description	U.S. Tariff	Canadian Tariff	TransCost
201	Meat products and packaging	0.82%	1.80%	0.41%
202	Dairy products and processing	6.61%	8.30%	3.89%
203	Preserved fruits and vegetables	7.29%	5.91%	1.25%
204	Grain mill products	2.33%	4.72%	1.34%
205	Bakery products	0.56%	5.32%	0.44%
208	Beverages	3.08%	30.27%	1.01%
209	Other food and kindred products	3.30%	3.98%	1.01%
210	Tobacco products	14.74%	24.83%	0.87%
220	Textile mill products	8.11%	13.27%	0.64%
230	Apparel and other textile products	10.56%	20.82%	0.71%
240	Lumber and wood products	0.43%	2.56%	4.01%
250	Furniture and fixtures	1.72%	9.35%	0.66%
262	Pulp, paper, and board mills	0.16%	3.31%	2.59%
265	Other paper and allied products	2.29%	6.49%	0.84%
270	Newspapers, printing, and publishing	3.15%	1.41%	1.32%
281	Industrial chemicals and synthetics	3.50%	3.41%	2.22%
283	Drugs	3.14%	5.05%	0.37%
284	Soap, cleaners, and toilet goods	4.64%	10.73%	0.62%
287	Agricultural chemicals	0.23%	0.05%	8.44%
289	Chemical products, n.e.c.	2.79%	5.48%	0.71%
299	Petroleum and coal products	4.52%	0.91%	0.00%
305	Rubber products	3.32%	5.95%	0.69%
308	Miscellaneous plastics products	4.34%	8.87%	0.70%
310	Leather and leather products	5.77%	10.83%	0.72%
321	Glass products	2.03%	4.56%	0.59%
329	Stone, clay, and concrete	2.11%	3.79%	2.39%
331	Primary metal industries, ferrous	3.34%	4.38%	1.12%
335	Primary metal industries, nonferrous	2.27%	1.19%	0.42%
341	Metal cans, forgings and stampings	0.65%	6.12%	1.27%
342	Cutlery, hardware, and screw products	2.23%	7.23%	0.58%
343	Heating equipment and plumbing fixtures	3.91%	9.65%	0.79%
349	Metal services products, ordnance, n.e.c	2.38%	5.79%	0.60%
351	Engines and turbines	1.46%	9.32%	0.15%
352	Farm and garden machinery	0.19%	0.11%	0.79%
353	Construction, mining, and machinery	1.96%	3.38%	0.32%
354	Metalworking machinery	3.36%	5.87%	0.34%
355	Special industry machinery	2.78%	3.16%	0.31%
356	General industrial machinery	2.25%	3.20%	0.43%
357	Computer and office equipment	0.51%	1.08%	0.77%
358	Refrigeration and service industry machinery	2.45%	4.28%	0.40%
359	Industrial and commercial machinery, n.e.c.	2.16%	2.67%	0.61%
363	Household appliances	2.63%	8.82%	0.76%
366	Household audio and video and communications	3.71%	4.55%	0.48%
367	Electronic components and accessories	1.92%	6.79%	0.78%
369	Electrical machinery, n.e.c.	2.38%	4.11%	0.44%
371	Motor vehicles and equipment	0.11%	0.37%	0.39%
379	Other transportation and equipment	0.51%	1.71%	0.22%
381	Measuring, scientific, and optical instruments	2.19%	2.16%	0.35%
384	Medical instruments and supplies	4.46%	1.89%	0.49%
390	Miscellaneous manufacturing industries	3.66%	6.40%	0.50%
	AVERAGES	3.10%	6.12%	1.03%

Appendix B
Industry Categories and Observations

IND	Total Observations	Individual Category
201	3	CN
202	5	CN
203	11	CN
204	22	CN
205	10	CN
208	46	CN
209	75	CN
210	22	CN
220	57	II
230	70	CD
240	61	II
250	58	CD
262	66	II
265	70	II
270	80	II
281	169	II
283	131	CN
284	102	CN
287	13	II
289	121	II
299	3	II
305	59	II
308	94	II
310	14	CN
321	38	II
329	57	II
331	74	II
335	54	II
341	19	II
342	46	II
343	83	IM
349	118	IM
351	10	IM
352	3	IM
353	65	IM
354	41	IM
355	47	IM
356	44	IM
357	16	IM
358	59	IM
359	13	IM
363	28	CD
366	34	CD
367	63	CD
369	154	CD
371	196	CD
379	53	CD
381	63	IM
384	43	IM
390	98	CD

CN are consumer non-durables CD are consumer durables II are industrial intermediates IM are industrial machinery