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ABSTRACT

Heterogeneous Firm-Level Responses to Trade Liberalization: A Test Using Stock Price Reactions*

This paper presents novel empirical evidence on key predictions of heterogeneous firm models by examining stock market reactions to the Canada-United States Free Trade Agreement of 1989 (CUSFTA). Using the uncertainty surrounding the agreement's ratification, I show that the pattern of abnormal returns of Canadian manufacturing firms was broadly consistent with the predictions of a class of models based on Melitz (2003). Increases in the likelihood of ratification led to stock market gains of exporting firms relative to non-exporters. Moreover, gains were higher in sectors with larger cuts in U.S. import tariffs. Decreases in the likelihood of ratification led to opposite stock market reactions. Results for the impact of Canadian tariff reductions are less conclusive but most specifications suggest that exporters also gained relative to non-exporters in response to such reductions. Translating stock market gains into implied profit changes, I find that CUSFTA increased expected per-period profits of exporters by around 6-7% relative to non-exporters.

JEL Classification: F12, F14 and G14
Keywords: Canada-U.S. free trade agreement, heterogeneous firm models and stock market event studies

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1 Introduction

The last decade has seen a revolution in the theoretical analysis of trade liberalization episodes. Since the seminal contribution by Melitz (2003), models with heterogeneous firms have all but replaced traditional modelling approaches with homogeneous firms. The key innovation of Melitz and subsequent extensions was to show how trade liberalization leads to aggregate productivity gains through intra-industry reallocation.[1] The mechanism underlying this reallocation is the differential impact of trade liberalization on exporting and non-exporting firms. While exporters benefit from increased access to foreign markets, non-exporters suffer lower profits due to increased product and factor market competition. Together with the assumption that exporters are more productive than non-exporters, the ensuing reallocation of market shares towards exporting firms raises aggregate productivity.

Many features of heterogeneous firm models are consistent with stylized facts which have emerged from a large empirical literature over the years. For example, Bernard and Jensen (1999) provide evidence that more productive firms self-select into export markets. Tybout (2003) summarizes several studies which show that market share reallocations were an important part of trade liberalization episodes. A smaller literature also provides more direct evidence on the impact of lower trade costs on the reallocation of market shares between exporters and non-exporters (*e.g.*, Bernard, Jensen and Schott, 2006; Trefler, 2004).

A common feature of all empirical studies to date is their ex-post character. That is, they track the firm- or sector-level variables of interest for a number of years and try to isolate the impact of trade policy changes from a large number of confounding factors. Depending on the specific setting of the liberalization episode, this can pose considerable econometric challenges (see, for example, Trefler, 2004).

In this paper, I take a different approach to providing evidence for the differential impact of trade liberalization across firms. I do so by using stock market reactions surrounding the implementation process of the Canada-United States Free Trade Agreement of 1989 (henceforth, CUSFTA). Under the assumption that unanticipated changes in the likelihood of CUSFTA's implementation are sufficiently rapidly reflected in stock prices, price reactions contain information about changes in future profits and can be used to implement a test of heterogeneous firm models.[2]

One advantage of such an event study approach over traditional ex-post evaluations is that the number of confounding factors is much more limited. Only factors about which expectations change during my one- to two-day event windows will have the potential to contaminate the estimates. Secondly, an event study approach arguably presents a more direct test of heterogenous firm models. These models essentially make predictions about changes in future per-period profits brought about

[1] See, for example, Melitz and Ottaviano (2008) and Chaney (2008). An alternative approach developed by Eaton and Kortum (2002) and Bernard et al. (2003) yields very similar predictions about the effects of trade liberalization.

[2] As I discuss in detail below, the main results of this paper do not require stock markets to be efficient in the sense of immediately reflecting all available information, but only that new events are priced in within a period of one or two days to a statistically detectable extent.

by trade liberalizations. To the extent that expectations about these changes will be reflected in stock prices, analyzing price reactions will be conceptually closer to the models' theoretical predictions than looking at realized firm-level variables ex-post.

CUSFTA is particularly well suited for providing event study evidence on heterogeneous firm models. In particular, the agreement was the main election issue in the Canadian general election of November 1988. Both the election itself as well as a number of events in its run-up provide unanticipated changes in the likelihood of CUSFTA's implementation which can be usefully exploited in an event study. Secondly, CUSFTA was a reciprocal agreement and is as such suitable for analyzing the differential impact of domestic and foreign tariffs. This distinction is a key element of many of the more recent heterogeneous firm models such as Melitz and Ottaviano (2008) or Chaney (2008). Finally, the large variation of tariff cuts across sectors allows the implementation of a differences-in-differences estimation strategy within the event study framework.

My findings are broadly supportive of the predictions of heterogeneous firm models. The election victory of the ruling Progressive Conservatives (a strong supporter of CUSFTA) led to significant stock market gains of exporting firms relative to non-exporting firms. In contrast, opinion polls in the run-up to the election showing a substantial lead for the oppositional Liberal Party (who were opposed to CUSFTA) resulted in negative abnormal returns of exporters compared to non-exporters.

In order to address the possibility that a Conservative election victory may have affected these two groups of firms differently through channels other than CUSFTA, I compare return differences between exporters and non-exporters across industries with different extents of tariff cuts. Consistent with theoretical predictions, I find that the relative gains and losses of exporters were indeed significantly higher in sectors with larger U.S. tariff cuts. These results are robust to including a number of control variables such as changes in intermediate input tariffs and firms' multinational status.

As a further check on my results, I also examine stock market reactions to two earlier events which were directly related to CUSFTA but not the election itself: the reaching of an agreement on CUSFTA after difficult negotiations between the U.S. and Canada in October 1987; and the refusal of the Canadian Senate to ratify the agreement in July 1988. I again find that stock prices of exporters increased relative to those of non-exporters in reaction to the first event, and decreased in response to the second event. As before, reactions were stronger in sectors with higher future U.S. tariff cuts. Finally, I also perform placebo checks by looking at stock market reactions on dates on which no new information about CUSFTA was revealed. Consistent with theoretical predictions, I do not find significant effects in these additional regressions.

My results are less conclusive with respect to the effects of Canadian tariff cuts. The majority of results suggests that exporting firms also gained relative to non-exporting firms in response to such tariff reductions. However, the corresponding coefficient estimates are generally small and have the wrong sign for some specifications and events. Interestingly, as I discuss below, these weaker results correspond to less clear-cut theoretical predictions of heterogeneous firm models with respect to import tariff liberalization (as opposed to export tariff reductions), in the sense that the predictions of existing models seem to partially depend on specific assumptions about demand and cost structures.

To evaluate the quantitative importance and plausibility of the estimated return differences, I also calculate the CUSFTA-induced change in the expected future profits of active firms implied by my estimates. Based on assumptions about the change in the likelihood of CUSFTA's implementation brought about by the Conservative election victory, I estimate that CUSFTA increased exporters' per-period profits by around 6%-7% relative to non-exporters in the most plausible scenarios, and up to 14% under more extreme assumptions.

While stock market event studies are frequently employed in the corporate finance literature, they have rarely been used to test theories of international trade. Exceptions include Grossman and Levinsohn (1989), who use stock market returns to provide evidence in favor of the specific-factor model of trade, and a small number of papers which analyze stock market reactions to trade policy announcements concerning specific industries, such as the imposition of antidumping duties (e.g., Hartigan et al., 1986 and 1989; Hughes et al., 1997). To the best of my knowledge, the present paper is the first to analyze stock market reactions to a broad-based trade liberalization episode and link the results to recent theories of international trade. While my focus is on testing models of heterogeneous firms, some of my robustness checks also provide complementary evidence to existing results from ex-post approaches for the effect of reductions in intermediate input tariffs and the differential impact of trade liberalization on multinational and domestic firms. The use of cross-sectional variation in tariff cuts to implement a difference-in-differences approach within the event study framework is also novel and substantially increases the potential for convincing econometric identification. Finally, the present paper seems to be the first to attempt a quantification of the differential impact of trade liberalizations on the profits of firms within an industry, which is the driving force behind subsequent market share reallocations.

The rest of this paper is structured as follows. Section 2 discusses how stock price reactions can be used to test heterogeneous firm models, and uses a simple model of this class to derive testable predictions for the remaining sections. Section 3 describes CUSFTA and the specific events I study in more detail. Section 4 discusses the event study methodology and describes the data sources used. Section 5 presents the empirical results and Section 6 concludes.

2 Theoretical Predictions

This section demonstrates the link between the predictions of heterogeneous firm models and stock market reactions to expected changes in trade costs, and derives testable predictions for the subsequent analysis. In a first step, I discuss how stock market prices are linked to firm-level profits and what assumptions are needed for my approach (Section 2.1). I then discuss the predictions of heterogeneous firm models with respect to how profits change in response to reductions in domestic and foreign tariffs. In Section 2.2, I focus on a heterogeneous firm model based on Chaney (2008) which is simple enough to clearly demonstrate the key mechanisms at work, yet sufficiently flexible to accomodate asymmetric countries and tariff barriers, two key features CUSFTA. Section 2.3 discusses to what extent the insights from this model carry over to more general settings, and Appendix A presents analytical results for two popular extensions of Melitz (2003), the original Chaney (2008) model and the model by Melitz and Ottaviano (2008), both of which allow for asymmetric countries and tariff barriers, while still delivering closed-form solutions for relative profit changes.

2.1 Linking stock prices to expected profits

The standard approach to linking stock prices and expected profits is the dividend discount model (see Brealey and Myers, 2000). The dividend discount model states that the price of firm i's shares at time t equals the net present value of its future stream of dividends per share:

$$p_{it} = \sum_{s=1}^{\infty} \frac{E(DIV_i|I_t)}{(1+e_i)^s} = \frac{E(DIV_i|I_t)}{e_i}$$

where $E(DIV_i|I_t)$ is the expected value of future per-period dividends per share of firm i, given information available on date t (I_t), and e_i is the expected return on securities in the same risk class as firm i. Assuming that firms disburse all profits as dividends, or that profits are reinvested at an internal rate of return equal to e_i, share prices are simply the net present value of expected future profits per share:[3]

$$p_{it} = \sum_{s=1}^{\infty} \frac{E(\pi_i|I_t)}{(1+e_i)^s} = \frac{E(\pi_i|I_t)}{e_i} \tag{1}$$

Now consider any two firms, e.g., an exporter (X) and a non-exporter (NX). The difference in stock market returns between these two firms upon the arrival of new information at time $t+\varepsilon$ will be (assuming e_i stays constant for both groups):

$$r_X - r_{NX} = \frac{E(\pi_X|I_{t+\varepsilon})}{E(\pi_X|I_t)} - \frac{E(\pi_{NX}|I_{t+\varepsilon})}{E(\pi_{NX}|I_t)} \tag{2}$$

What matters for the difference in stock market returns is thus the change in expected future profits of exporters relative to non-exporters upon the arrival of new information (regarding the likelihood of CUSFTA's implementation in the present case). Since models of heterogeneous firms make explicit predictions about these profit changes, stock market returns can be used to implement empirical tests of this class of models.

Note that for testing the *qualitative* predictions of heterogeneous firm models, the assumptions underlying the above derivations can be substantially relaxed. For example, one could easily allow for more complex connections between dividends and profits, as long as the positive correlation between changes in both variables is preserved. Likewise, it is not required that stock prices fully

[3]It is straightforward to allow for growth in expected dividends or positive net present value projects (see Brealey and Myers, 2000). Since this would not add any new insights for the purpose of this paper, I abstract from such complications.

and immediately reflect all relevant information. All that is needed is that new information about the likelihood of CUSFTA's implementation is priced in to a statistically detectable extent within a period of one or two days (which will be the standard length of my event windows).[4] Given the importance of CUSFTA in the Canadian election campaign of 1988 and for the Canadian economy more generally, it seems reasonable that at least some market participants reacted quickly to the Progressive Conservatives' election victory and were able to judge CUSFTA's impact on firm profits, at least in terms of the direction of the change if not its exact magnitude.[5]

2.2 Firm-Level Profits and Trade Liberalization

I now turn to a discussion of how profit changes after trade liberalization vary across different types of firms in models with heterogeneous firms. I use a version of Chaney (2008) to illustrate the main points and discuss to what extent these results carry over to alternative heterogenous firm models in the next section.

Consider a setting with N potentially asymmetric countries. A representative consumer in each country derives utility from the consumption of goods from $S+1$ sectors. The first S sectors each produce a continuum of differentiated goods (Q_s) and the remaining sector provides a single homogenous good (A):

$$U_n = \sum\nolimits_{s=1}^{S} \mu_s \ln Q_{ns} + A_n, \qquad Q_{ns} = \left[\int_{\gamma \in \Gamma_{sn}} q_{ns}(\gamma)^{\frac{\sigma_s - 1}{\sigma_s}} d\gamma \right]^{\frac{\sigma_s}{\sigma_s - 1}} \tag{3}$$

where Γ_{sn} presents the set of available varieties of good Q_{ns} in country n, which will be endogenously determined, and $\sigma_s > 1$ is the elasticity of substitution between any two varieties in sector s. Associated with Q_{ns} is a price index $P_{ns}^{1-\sigma_s} = \left[\int_{\gamma \in \Gamma_{ns}} p_{ns}(\gamma)^{1-\sigma_s} d\gamma \right]$,where $p_{ns}(\gamma)$ is the price of variety γ in sector s, country n. Good A is freely traded and I choose its price as the numeraire. With this setup, expenditure per consumer on Q_{ns} is fixed at $E_{ns} = P_{ns} Q_{ns} = \mu_{ns}$ and demand for individual varieties is $q_{ns}(\gamma) = p_{ns}(\gamma)^{-\sigma_s} P_{ns}^{\sigma_s - 1} \mu_{ns}$.

I choose parameter values such that all countries produce positive amounts of the numeraire. Labor is mobile between sectors, and the numeraire sector operates under perfect competition and with linear production function $A_n = l_{A_n} \theta_{A_n}$, where θ_{A_n} is labor productivity and l_{A_n} labor employed in the numeraire sector in country n. Profit maximization implies that wages in country n are equal to labor productivity, $w_n = \theta_{A_n}$.

The differentiated goods are produced using labor as the only factor of production. Firms vary in productivity levels, γ, and have unit labor requirements of $l(\gamma) = q/\gamma$. In order to ship goods from country i to country j, firms further incur variable trade costs τ_{ij}^{s} of the standard iceberg

[4] In this sense, the general critique that stock market event studies always present a joint test of both the theory in question and the efficient market hypothesis (e.g., Campbell et al., 1997) only applies to a lesser extent to the present paper.

[5] Market efficiency, rational expectations of market participants, and the exact link between dividends and profits do become important, however, when I try to quantify the profit impact of tariff changes in Section 5.3. For this, I will need expression (1) to hold exactly.

type. Finally, a firm in country i selling goods to country j in sector s has to pay a fixed cost of f_{ij}^s in terms of the numeraire.

Each firm in the differentiated goods sectors is a monopolist for the variety it produces and sets prices at a constant markup over marginal costs, $p_{ij}^s(\gamma) = \frac{\sigma_s}{\sigma_s - 1}\frac{\tau_{ij}^s w_i}{\gamma}$. Profits at this price from sales in market j are $\pi_{ij}^s(\gamma) = p_{ij}^s(\gamma)\, q_{ij}^s(\gamma)\, \sigma_s^{-1} - f_{ij}^s$. There are a large number, M_{ns}, of potential entrants in each country and sector which have to decide in which of the N countries to sell. Productivity levels γ are known to firms before entry. In equilibrium, only firms which can earn non-negative profits in a given market will become active on that market, leading to market-pair specific productivity cutoffs, $\gamma_{ij,s}^*$. Finally, as in Chaney (2008) and Melitz and Ottaviano (2008), I assume that firm-level productivity γ in country n, sector s, is Pareto distributed with density $v_{ns}(\gamma) = a_s\,(k_{ns})^{a_s}\,\gamma^{-(a_s+1)}$, where $k_{ns} > 0$, $a_s > \sigma_s - 1$ and $\gamma \geq k_{ns}$. For notational ease, I focus on a single sector and drop the relevant subscript s from now on.

Under the above assumptions I obtain a solution for the entry cutoffs γ_{ij}^* in each sector as:

$$\gamma_{ij}^* = \left(\frac{f_{ij}}{\mu_j}\frac{a\sigma}{a-\sigma+1}\right)^{1/a}\left(\sum_n \left(\frac{\tau_{nj}w_n}{\tau_{ij}w_i}\right)^{-a} M_n k_n^a \left(\frac{f_{nj}}{f_{ij}}\right)^{\frac{\sigma-a-1}{\sigma-1}}\right)^{1/a} \tag{4}$$

If a firm from i is active in market j, its profits there can be expressed as a function of the relevant entry cutoff:

$$\pi_{ij}(\gamma) = \max\left(\gamma^{\sigma-1}\left[\left(\gamma_{ij}^*\right)^{1-\sigma} - \gamma^{1-\sigma}\right] f_{ij}, 0\right) \tag{5}$$

Total profits of a Canadian firm with productivity γ are:

$$\pi_i(\gamma) = \sum_n \pi_{in}(\gamma) = \gamma^{\sigma-1}\sum_n \max\left(\left[\left(\gamma_{in}^*\right)^{1-\sigma} - \gamma^{1-\sigma}\right] f_{in}, 0\right)$$

I look at the impact of tariff reductions between Canada (i) and the United States (j) on Canadian firms' profits. This corresponds to a reduction of τ_{ij} and τ_{ji} in the model, where τ_{ij} denotes the U.S. import tariff Canadian firms face and τ_{ji} is the Canadian import tariff U.S. firms have to pay when exporting to Canada. Note that because of quasi-linear preferences and the assumption of a fixed number of incumbents, third market profits of Canadian firms will not be affected by changes in U.S. or Canadian import tariffs (see expression (4)). Thus, it is sufficient to analyse changes in domestic profits (π_{ii}) and in profits from exports to the U.S. (π_{ij}).[6] For firms which export both before and after liberalization, we have:

$$\frac{\Delta\pi_X\left(\gamma, \tau_{ij}, \tau_{ij}'\right)}{\pi_X(\gamma)} = \frac{\gamma^{\sigma-1} f_{ij}\left(\gamma_{ij}^{*\prime 1-\sigma} - \gamma_{ij}^{*1-\sigma}\right)}{\pi_X(\gamma)} > 0$$

where τ_{ij} denotes the initial tariff and τ_{ij}' the new (lower) tariff. The relative profit change for

[6] In the following, I assume parameter values such that $\gamma_{in}^* > \gamma_{ii}^*$ for all n. Thus, all active firms serve the domestic market whereas only the more productive firms export to the U.S. and other markets (which is the empirically relevant case).

existing exporters is positive because the domestic cutoff is not affected and the U.S. export cutoff falls, $\gamma_{ij}^{*'} < \gamma_{ij}^{*}$. For firms which export neither before nor after the tariff reduction, U.S. profits (π_{ij}) are zero and the percentage change in profits after a lowering of U.S. tariffs is also zero because the domestic cutoff is not affected (see 4):

$$\frac{\Delta\pi_{DOM}\left(\gamma,\tau_{ij},\tau'_{ij}\right)}{\pi_{DOM}\left(\gamma\right)} = \frac{\gamma^{\sigma-1}f_{ii}\left(\gamma_{ii}^{*'1-\sigma} - \gamma_{ii}^{*1-\sigma}\right)}{\pi_{DOM}\left(\gamma\right)} = 0$$

Finally, for firms which start exporting only after U.S. tariffs have been reduced, we have:

$$\frac{\Delta\pi_{S}\left(\gamma,\tau_{ij},\tau'_{ij}\right)}{\pi_{S}\left(\gamma\right)} = \frac{\left(\gamma_{ij}^{*'}\right)^{1-\sigma}\gamma^{\sigma-1}f_{ij} - f_{ij}}{\pi_{S}\left(\gamma\right)} > 0$$

Thus, existing and new exporters observe stronger relative profit increases than purely domestic firms. From (2), we should thus observe a positive difference in stock market returns between new and existing exporters and non-exporters upon the arrival of new information making an implementation of CUSFTA more likely.[7]

Next, consider a reduction in Canadian tariffs from τ_{ji} to τ'_{ji}. From (4), the export cutoff γ_{ij}^{*} will not be affected whereas the domestic entry cutoff γ_{ii}^{*} will rise ($\gamma_{ii}^{*'} > \gamma_{ii}^{*}$). Thus, only domestic profits will be affected. The implied change in total profits of exporting firms will be:

$$\frac{\Delta\pi_{X}\left(\gamma,\tau_{ji},\tau'_{ji}\right)}{\pi_{X}\left(\gamma\right)} = \frac{\gamma^{\sigma-1}f_{ii}\left(\gamma_{ii}^{*'1-\sigma} - \gamma_{ii}^{*1-\sigma}\right)}{\pi_{X}\left(\gamma\right)} < 0$$

For non-exporters which continue to serve the Canadian market we have:

$$\frac{\Delta\pi_{DOM}\left(\gamma,\tau_{ji},\tau'_{ji}\right)}{\pi_{DOM}\left(\gamma\right)} = \frac{\gamma^{\sigma-1}f_{ii}\left(\gamma_{ii}^{*'1-\sigma} - \gamma_{ii}^{*1-\sigma}\right)}{\pi_{DOM}\left(\gamma\right)} < \frac{\Delta\pi_{X}\left(\gamma,\tau_{ji},\tau'_{ji}\right)}{\pi_{X}\left(\gamma\right)}$$

So both exporters and non-exporters lose but losses are more severe for non-exporters. Intuitively, the part of exporters' total profit derived from the U.S. market is not affected by Canadian tariff cuts, so that the relative decline in total profits is smaller. Secondly, exporters are more productive and spread the market-specific fixed costs over a larger amount of sales. The percentage decline in domestic profits alone will thus also be smaller.

Finally, the least productive Canadian firms will exit the domestic market after the reduction in τ_{ji}:

[7] Note that it is not possible to unambiguously rank the relative profit changes of existing exporters and new exporters. While the most productive new exporter will have a higher percentage profit change than all existing exporters, the least productive new entrant will have a relative change lower that that of all firms already exporting. In contrast, *absolute* profit increases (i.e., $\Delta\pi$ rather than $\Delta\pi/\pi$) are smallest for the least productive new exporter and increase monotonically with productivity, yielding an unambiguous ranking. I will return to this point in my robustness checks in Section 5.2.

$$\frac{\Delta\pi_{EXIT}\left(\gamma,\tau_{ji},\tau'_{ji}\right)}{\pi_{EXIT}\left(\gamma\right)}=\frac{0-\left(\gamma_{ii}^{*1-\sigma}\gamma^{\sigma-1}f_{ii}-f_{ii}\right)}{\pi_{EXIT}\left(\gamma\right)}=-1<\frac{\Delta\pi_{DOM}\left(\gamma,\tau_{ji},\tau'_{ji}\right)}{\pi_{DOM}\left(\gamma\right)}$$

So, to summarize, Canadian tariff reductions will reduce profits of all Canadian firms but exporters will be less affected than both continuing and exiting domestic firms. We should thus observe a positive difference in stock market returns between exporters and non-exporters upon the arrival of new information making an implementation of CUSFTA more likely. In contrast, *absolute* profit increases ($\Delta\pi$, rather than $\Delta\pi/\pi$) are smallest for the least productive new exporter and monotonically rise with productivity, yielding an unambiguous ranking. I will provide evidence on this additional prediction in the robustness checks in Section 5.2.

2.3 Discussion

To what extent do these results carry over to alternative modeling frameworks? Chaney (2008) introduces income effects in an otherwise identical model by letting his utility function take a Cobb-Douglas rather than a quasi-linear form. This changes the magnitude of the profit responses but leaves the qualitative predictions of the previous section intact, as I demonstrate in Appendix A.

Another simplifying assumption of the above model is that wages are exogenously fixed and that there are therefore no factor market interactions. In contrast, such interactions are crucial for the results in Melitz (2003). While tariffs (or more generally, variable trade costs) are assumed to be symmetric in his model, the general intuition is clear. Lower foreign tariffs lead exporters to expand which puts upward pressure on domestic wages. Non-exporters thus face higher input costs but do not benefit from increased access to foreign markets. In the present context, U.S. tariff cuts would thus increase the profits of existing and new exporters relative to continuing and exiting non-exporters, similar to the predictions from the last section.

A third simplification which is more critical for the previous results, especially with respect to domestic tariff reductions, is the assumption of a fixed number of potential entrants. For example, Melitz and Ottaviano (2008) present a version of their model with long-run entry in which expected profits are reduced to zero. In this case, lower U.S. tariffs again increase access to the U.S. market and raise the profits of exporting firms. This effect is now reinforced through the exit of U.S. firms which leave their market in the long-run because of reduced domestic profit opportunities. However, lower U.S. tariffs now also trigger entry into the Canadian market by new domestic firms. This increases competition there and lowers the domestic profits of both exporters and non-exporters. But because demand is assumed to be linear, the percentage profit reduction is again smaller for the more productive and thus larger exporters (see Appendix A).

In contrast, domestic (Canadian) tariff reductions lead to a reduction in long-run entry which increases profits for the remaining firms. At the same time, better access to the Canadian market leads to increased entry of U.S. firms which also serve their domestic market. This makes it more difficult for Canadian exporters to sell there, lowering profits from exporting. Linear demand again implies that the less productive non-exporters will see a stronger increase in their domestic profits than exporters. In addition, they do not suffer a reduction in their export profits. Thus, in the free-entry version of Melitz and Ottaviano, Canadian tariff reductions favor those non-exporters in Canada which do not exit the market entirely (see Appendix A).

Finally, even without long-run entry the result that exporters see their domestic profits fall by relatively less than non-exporters in response to import tariff reductions seems to be at least in part due to specific assumptions about demand and cost structures. In the model from the previous section, it is the presence of fixed costs which causes the relatively smaller fall of domestic profits for more productive firms, and in Melitz Ottaviano (2008) it is the assumption of linear demand. While the existing literature has not yet explored this issue, one could easily imagine a demand curve with more curvature than CES. This would imply a stronger percentage reaction in domestic profits for the more productive exporters and might reverse some of the above results. (With CES and in the absence of fixed cost, relative domestic profit changes are identical for firms with different levels of productivity.)

In summary, the results of the previous section with respect to export tariff reductions seem robust across a range of heterogeneous firm models. Intuitively, the direct effect of lower export tariffs is to reduce the effective cost of supplying goods for the subset of firms which already export or start exporting while leaving domestic firms (initially) unaffected. The result that the former group of firms increase their profits relative to the latter in the new equilibrium should thus be quite general. In contrast, the relative effect of import tariff reductions on exporters and non-exporters appears to be less robust, and might well be different in more general frameworks than the one presented here.

3 Description of Events

A key element of any event study is the identification of suitable events. In the present context, I am looking for points in time at which the likelihood of CUSFTA's implementation changed substantially. This is a potentially difficult challenge, given that the negotiation and ratification process covered a period of over two years, from the start of negotiations in May 1986 until the eventual ratification by the Canadian parliament in December 1988. Given that the idea of liberalizing trade between Canada and the United States had also been around for some time before CUSFTA, the successful conclusion of negotiations and the subsequent signing and ratification of the agreement might have been anticipated to a large degree.

Fortunately, the Canadian general election on 21 November 1988 provides a more sharply defined event which can be usefully exploited for event study evidence.[8] The first reason for this is that the ratification of CUSFTA was extremely contentious among the main Canadian political parties, with

[8] Brander (1991) and Thompson (1993) also evaluate whether there were significant stock market reactions to CUSFTA but are not primarily interested in testing theories of international trade. A lack of tariff data also prevents them from differentiating CUSFTA's influence more clearly from other contemporaneous factors such as the Conservative election victory.

the governing Progressive Conservatives (who had negotiated the agreement) in favor, and broad sections of the main opposition parties (the Liberals and the New Democratic Party) opposed. Indeed, the Liberal Party's leader, John Turner, publicly vowed as late as October 1988 that he would dismantle CUSFTA in case of victory in the elections. The fate of CUSFTA thus directly depended on the election outcome on November 21.

Secondly, CUSFTA received an unprecedented amount of attention in the election campaign and was indeed the single-most important issue in voters' minds. In opinion polls taken in the month before the election, over 80% of the electorate cite CUSFTA as the most important election issue. Traditional areas of concerns such as inflation, unemployment, the budget deficit, welfare spending or national unity all were each mentioned by at most 2% of voters (Frizzell et al., 1989). One would thus expect that market reactions to a Conservative or Liberal victory in the elections would be predominantly determined by the consequences for the implementation of CUSFTA.

Finally, the outcome of the election was highly uncertain. Given the particularities of the Canadian electoral system, the Conservatives needed a vote share of slightly more than 40% to obtain a parliamentary majority (see Johnston et al., 1992). As late as the week before the vote on November 21, however, opinion polls showed Liberals and Conservatives head-to-head at 35% of the vote each.[9] Such an outcome would most likely have given Liberals and New Democrats a majority of parliamentary seats and would thus have meant that CUSFTA would not be ratified. The turning point came only with the publication of three nationwide polls on November 19, the Saturday before the election. All three polls put the Conservatives at over 40% and clearly ahead of the Liberals. These predictions proved indeed to be almost exactly correct, and on November 21 the Conservatives won the election with 43% of the popular vote, compared to 32% for the Liberal Party and 20% for the New Democrats.

Besides the election itself, I will look at three earlier events which also changed the likelihood of CUFTA's implementation. The second event is the reaching of an agreement on CUSFTA between Canada and the U.S. on 3 October 1987.[10] Negotiations had been difficult and were only brought to a successful conclusion hours before the deadline on October 3. Thus, the reaching of an agreement was to some extent unexpected. At the same time, many of the key elements of CUSFTA (including the extent of the tariff reductions) had been agreed already so that market participants were probably aware of most of its consequences.

The third event is again related to CUSFTA's ratification. On the morning of 20 July 1988, John Turner, the Liberal Party's leader, announced at a press conference that he had instructed the Liberal majority in the Senate to block the ratification of CUSFTA until a general election, which was expected to be called within the next months. This was seen by many as a move to revive the electoral prospects of his party which was trailing in the opinion polls (Johnston et al., 1992). By delaying the ratification, John Turner effectively turned the general election into a referendum on

[9] All opinion polls quoted in this section are taken from Frizzell et al. (1989).

[10] The information in this paragraph is based on the extensive coverage of the negotations in the Canadian newspaper *The Globe and Mail* from 5 October 1987. Also see Thompson (1993).

CUSFTA. This move destroyed any hopes for a quick ratification and even raised the possibility that CUSFTA might not be implemented at all, given the hostility of Liberals and New Democrats to the agreement.

Finally, I also use a particularly dramatic change in opinion polls in the run-up to the election. After it had become clear that the Senate would not ratify CUSFTA, prime minister Brian Mulroney called a general election on October 1. In the initial phase of the election campaign, the Conservatives had a clear lead in the opinion polls with a predicted vote share of over 40%. As discussed above, this was enough to guarantee a parliamentary majority sufficient for CUSFTA's ratification. An important turning point came with the only two televised debates between the main parties' leaders on October 24 and 25. Against expectations, John Turner emerged as the clear winner from these debates and electoral fortunes started to change. The most dramatic and unexpected event in this phase of the campaign was the publication of a Gallup poll on the morning of November 7, putting the Liberals at 43% of the vote, compared to only 31% for the Conservatives and 22% for the New Democrats. While opinion polls had been gradually shifting since the debates, this presented a massive increase in support for John Turner's party and for the first time made a Liberal victory look likely.[11] In response, the Conservatives undertook a radical overhaul of their campaign strategy, enabling them to catch up in the opinion polls again (Frizzell et al., 1989). However, it was only with the above-mentioned publication of three nationwide opinion polls on November 19 that it became clear that the Conservatives would win.

Table 1 summarizes these events. My principal event is the election day itself (November 21) and the first trading day after the election (November 22). While markets could only react to the election results on November 22, the publication of the opinion polls on November 19 had already made a Conservative victory very likely.

The remaining three events are less important shifts in the likelihood of CUSFTA's implementation but are very useful as robustness checks. In particular, events three and four imply a decrease in the likelihood of ratification and should lead to opposite stock market reactions from the election event. Finally, events two and three present changes in the probability of CUSFTA's implementation which are unrelated to the election outcome. They will provide additional evidence that market reactions were indeed due to CUSFTA rather than a Conservative election victory.[12]

[11] See Brander (1991) and Frizzell et al. (1989). The surprise at the extent of the Liberals' lead is also evident in the press coverage of November 8. For example, *The Globe and Mail* titled "Confusion, disbelief greet poll showing strong Liberal surge" on November 8 and highlighted market particants' concerns that a Liberal government would tear up the free trade agreement.

[12] As discussed in more detail below, my identification strategy will also control for additional effects of a Conservative victory by relying on variation in tariff cuts across sectors. Moreover, the overwhelming importance of CUSFTA during the election campaign makes it likely that market reactions on November 21 and 22 were mainly due to the implications of a Conservative victory for CUSFTA, rather than for other policies.

4 Methodology, Data and Descriptive Statistics

Methodology. Testing the theoretical predictions from Section 2 requires a model of "normal" stock returns which adjusts for differences in risk and other characteristics of stocks. A standard approach in the literature is to use the so-called market model which relates the return on security i at time t to a stock-specific constant and the return of the market portfolio, R_{mt} (Campbell et al., 1989; Binder, 1998):

$$r_{it} = \alpha_i + \beta_i R_{mt} + \varepsilon_{it} \tag{6}$$

The error term ε_{it} captures "abnormal" returns which in the present context could be caused by the arrival of unexpected news about the implementation of CUSFTA. A straightforward way to measure abnormal returns related to CUSFTA is to directly model the error term in equation (6) according to the theoretical discussion from Section 2:[13]

$$r_{it} = \alpha_i + \beta_i R_{mt} + \sum\nolimits_{e=1}^{E} d_{et} \left(d_j + \beta_{1e} d_{ix} \right) + \eta_{it} \tag{7}$$

where the d_{et} are a set of dummy variables, each taking a value of one for one particular day during event window E. The d_j are industry fixed effects, and d_{ix} is a dummy variable which equals one if firm i exported to the U.S. in the year the event took place. The coefficient estimate $\hat{\beta}_{1e}$ thus represents the average abnormal return difference between exporters and non-exporters on event day e, after controlling for industry fixed effects. In the case where an event takes place over several days (as is the case for the first event in Table 1), I calculate cumulative average abnormal returns (CAARs) which are defined as:

$$CAAR_E = \sum\nolimits_{e=1}^{E} \hat{\beta}_{1e}$$

As already discussed, one concern with (7) is that my main event (the general election) not only changed the likelihood of CUSFTA's implementation but also expectations about other policies. For example, a conservative victory might have been seen as particularly advantageous for exporting firms. I thus make use of the sectoral variation in tariff cuts implemented under CUSFTA by estimating the following specification:

$$r_{it} = \alpha_i + \beta_i R_{mt} + \sum\nolimits_{e=1}^{E} d_{et} \left(d_j + \beta_{1e} d_{ix} + \beta_{2e} d_{ix} d\tau_{CAN,j} + \beta_{3e} d_{ix} d\tau_{US,j} \right) + \eta_{it} \tag{8}$$

where $d\tau_{CAN,j}$ and $d\tau_{US,j}$ denote Canadian and U.S. tariff reductions in industry j between 1988 and 1996, respectively.[14] Recall from the earlier discussion that exporters should benefit more from higher U.S. tariff cuts than non-exporters (i.e., $\beta_3 < 0$, given that higher reductions imply a more negative $d\tau$). In the model presented in Section 2 this is also true for Canadian tariff cuts, although

[13] See Binder (1998) for the advantages of measuring abnormal returns in a regression framework.

[14] 1996 is the last year for which I have tariff data. Manufacturing tariffs were phased out linearly over a period of up to ten years under CUSFTA and were close to zero in 1996 (see Trefler, 2004).

it was noted that this prediction might not survive in other heterogeneous firm models.

Introducing variation in tariff cuts into the modeling of abnormal returns means that I only require the weaker identifying assumption that the differential impact of a Conservative victory on exporters and non-exporters does not vary systematically with the extent of U.S. or Canadian tariff cuts. I thus use (8) as my main specification.

Data. Estimation of (7) and (8) requires data on daily returns on individual stocks and the market portfolio, the tariff cuts implemented under CUSFTA, as well as information on whether a firm exports to the U.S. For comparability with the existing literature and because of the availability of information on tariff reductions, I focus my analysis on firms in the manufacturing sector. Because of the tradability of its output, this is also the sector most directly affected by CUSFTA and the one that corresponds best to the theoretical model from Section 2.

I use daily stock returns from Datastream for all Canadian manufacturing firms listed on one or several Canadian or U.S. stock exchanges for which I have a least one year of return data prior to the event studied. This is the standard length in the event study literature for the pre-event window used to estimate the market model's parameters (see Binder, 1998). I also follow a large part of the literature by using the value-weighted CRSP portfolio as a proxy for the market portfolio.[15]

Tariff data are from Trefler (2004) who provides U.S. and Canadian ad-valorem tariffs for manufacturing industries at the four-digit level of the Canadian Standard Industrial Classication of 1980. I map these tariffs into the industry classification used by Datastream (the Industry Classification Benchmark, ICB) which sorts manufacturing firms into 20 broad industries.[16]

Finding a suitable proxy for export status is more challenging. One issue is that my data only contain data on firms' exports for a minority of firms. It is also not clear whether actual export status would be a good proxy even with perfect data availability. Recall from Section 2 that firms which start exporting in response to U.S. tariff reductions belong conceptually to the same group of firms as exporters – both observe profit increases relative to firms which never export. In the present case, new exporters accounted for a large fraction of all exporters. For example, Baldwin and Gu (2003) report that the fraction of exporters among manufacturing firms increased by almost 70% during the implementation period of CUSFTA. On the other hand, it is impossible to know whether all of these firms started exporting because of CUSFTA or would have taken up exporting anyway. Thus, focusing on actual export status risks selecting an inappropriate mix of firms for

[15] I obtain data on CRSP portfolio returns from the Wharton Research Data Services (wrds.wharton.upenn.edu). Using the CRSP portfolio should be less susceptible to endogeneity concerns, given that the firms in my sample represent a large share of the overall market capitalization in purely Canadian-based portfolios such as the S&P/TSX Composite Index. Also note that CRSP contains a number of Canadian firms quoted on U.S. stock exchanges (but which only account for a small fraction of overall U.S. market capitalization).

[16] See Table 2 for a list of these industries. I use detailed descriptions of individual industries obtained from Datastream and Statistics Canada to construct a mapping from Trefler's 213 Canadian Standard Industrial Classification (CANSIC) industries to the 20 ICB industries used in this paper. This mapping was unique in 90% of cases, in the sense that each CANSIC industry could be mapped into one ICB industry only. I aggregate the tariff data to the ICB level by taking weighted averages across all CANSIC categories mapping into an ICB industry, using 1988 output shares of CANSIC industries as weights. Output data are also from Trefler (2004).

treatment and control groups.

Instead, I rely on the theoretical model from Section 2 to derive an alternative proxy. This model, and all other heterogeneous firm models discussed so far, display a strict hierarchy of (current or future) export status with respect to productivity and sales. The positive correlation between firm size as measured by sales and export status is also one of the most robust empirical findings in the literature on exporter premia (e.g., Bernard and Jensen, 1999). In my baseline specification, I thus proxy the export dummies in (7) and (8) by the log of the value of a firm's sales. Using a continuous measure avoids taking a stance on the exact cutoff value of sales which separates exporters and non-exporters both before and after CUSFTA. Using log sales also facilitates the inclusion of number of binary control variables in later robustness checks which are often highly correlated with firm size (such as multinational status). In extensive robustness checks in Section 5, I experiment with a large number of alternative export proxies, including measures based on information on actual export status available in my data.[17]

Sales and export data are also available from Datastream. I complement this information with data from Compustat North America whenever Datastream has missing values. This yields a sample of 247 publicly traded Canadian companies with primary activities in manufacturing for which I have information on sales and stock prices, and a smaller sample of 54 firms for which I also observe the value of exports.

Descriptive Statistics and Figures. Table 2 provides summary statistics for the number of firms, firm sales and tariff reductions by industry. I note two main points. First, there is a strong variation in sales within industries, ranging from small start-ups with sales of less than a million Canadian dollars to big corporations with several billion dollars in turnover. Given the strong empirical correlation between sales and export status, these figures suggest that there should be substantial variation in pre- and post-CUSFTA export status within industries, which is a prerequisite for precise identification in the econometric analysis carried out below.

Secondly, tariff cuts also show substantial sectoral variation despite the relatively aggregate industry classification used here (columns 6-7). Canadian tariff cuts range from sectors which basically enjoyed free trade before CUSFTA to over 25% for "Beverages". U.S. tariff cuts are lower on average but still show strong sectoral differences, with tariff cuts between 0% and close to 10%.

Figure 1 takes a closer look at the data by visualizing the difference-in-differences identification strategy embodied in my key specification, equation (8). I focus on my main event, the general election on November 21. However, to fully appreciate the high degree of uncertainty surrounding the election outcome, it is useful to look at a slightly longer window, starting a week before the televised debates between the main parties' leaders on October 24 and 25. For this period, I plot cumulative average return (CAR) differences between large and small firms, defined here simply

[17] In general, deviations of my proxies from actual and future export status can be thought of as classical measurement error which will tend to bias the results against finding significant effects.

as firms with sales above and below the 50th percentile in each industry, respectively.[18] I plot CAR differences for two groups of firms. Those belonging to the 50% of industries with the highest U.S. tariff cuts implemented under CUSFTA, and those with the 50% lowest tariff cuts.[19] CAR differences are normalized to zero for both groups one week before the televised debates on October 24 and 25.

The figure clearly shows a sharp divergence in the CAR differences between high- and low-tariff cut industries in the aftermath of the debates, as the Liberal Party's standing in the polls starts to improve. Note that this divergence is particularly dramatic on the day of the publication of the Gallup poll, November 7. Also visible in the graph is the stabilization in CAR differences between large and small firms, and between high- and low-tariff cut industries, after the Conservatives catch up in the polls again. (The week beginning November 14 brought a couple of opinion polls showing the parties head-to-head again.) Finally, the difference between high- and low-tariff cut industries narrows sharply on election day, November 21, and to a lesser extent on November 22.

This graphic analysis provides some first suggestive evidence that stock prices reacted to news about CUSFTA in a way consistent with the predictions of heterogeneous firm models. To see whether these findings hold up in a more thorough econometric analysis, I now turn to the estimation of the baseline equations (7) and (8).

5 Results

5.1 Baseline Results

Column 1 of Table 3 reports results based on specification (7), using log sales as the proxy for export status. The results indicate that larger firms experienced significantly higher abnormal returns – about 0.3 percentage points per log point of sales. This is consistent with the predictions of heterogeneous firm models which predict such a differential effect across exporters and non-exporters, and smaller and larger firms. As already mentioned, this result could also capture a more positive impact of a Conservative election victory on larger firms.

In column 2, I include the tariff interaction terms as in (8). As predicted, the sign on the U.S. tariff interaction is negative and significant. Thus, larger firms observed stronger positive abnormal returns in sectors with larger U.S. tariff cuts. This is strongly supportive of a Melitz-type story in which exporters benefit from increased export opportunities.

Exporters also benefited from higher Canadian tariff cuts relative to non-exporters. This is

[18] The cumulative average return of a group of stocks G between t_1 and t_2 is defined as $CAR_{t_1 t_2} = \sum_{s=t_1}^{t_2} \frac{1}{N_G} \sum_{i \epsilon G} r_{is}$, where r_{is} is the return of stock i at time s and N_G is the number of stocks in group G. The difference in $CARs$ between exporters and non-exporters in high tariff cut industries, for example, is then simply $CAR_{Xhigh} - CAR_{NXhigh}$. Using abnormal rather than simple returns yields a similar picture.

[19] I focus on U.S. tariff cuts since the theoretical predictions are unambiguous here. Graphs using Canadian tariff cuts yield a broadly similar if less clear-cut picture. This similarity reflects the positive correlation between U.S. and Canadian tariff concessions. As we shall see in the econometric analysis below, only U.S. tariff cuts have a robust impact on abnormal return patterns.

consistent with the model outlined in Section 2 as well as with Chaney (2008) and the "short-run" version of Melitz and Ottaviano (2008). While this effect is also highly statistically significant, it is much smaller in absolute magnitude than the effect of U.S. tariff reductions, even after taking into account that Canadian tariff cuts were on average twice as large as U.S. tariff cuts (see Table 2).

5.2 Robustness Checks

Alternative Export Status Definitions As a first robustness check, I experiment with a number of alternative proxies for export status. Table 4 shows results for several indicators which are also based on firm sales but which now take a binary form, classifying a firm as an exporter if its sales exceed a given industry-specific threshold value. Table 5 uses actual export information which are available for a subsample of 54 firms in my data.

In columns 1 and 2 of Table 4, I classify firms as exporters if their sales are above the 30th percentile of an industry's sales distribution. This threshold was chosen to match the fraction of exporters for the subsample of 54 firms for which I observe exports in 1988. Using this alternative export proxy yields qualitatively identical results to my baseline specification. Exporting firms experienced abnormal returns which were 0.9 percentage points higher than those of non-exporters, with the difference being highly statistically significant (column 1). In column 2, I include the tariff interaction terms which are again negative and significant for both U.S. and Canadian tariffs. The abnormal return difference between exporters and non-exporters increases by 0.9 percentage points for each percentage point in U.S. tariff reductions, and by 0.2 percentage points for each percentage point in Canadian tariff reductions.

The 30th percentile threshold is my preferred binary export proxy but I also present results for cutoffs based on more extreme assumptions, ranging from the 20th to the 80th percentile of industry-specific sales distributions. The 20th percentile threshold rule is again derived from the fraction of exporting firms in the subsample with export information, but this time also classifies firms as exporters if they have positive export sales in either 1988 or in any year of CUSFTA's implementation period (1989-1997). Implicitly, this assumes that all of these new exporters entered the export market because of CUSFTA. Since this is a strong assumption, the 20th percentile threshold should be seen as an upper bound on the true fraction of pre- and post-CUSFTA exporters. At the other end of the range of the thresholds used in Table 4 is the 80th percentile cutoff (columns 9-10), which classifies only 20% of firms as exporters. This figure corresponds to the fraction of exporters among Canadian manufacturing plants in the pre-CUSFTA period reported in Baldwin and Gu (2003). Since most of these units of production are substantially smaller than the publicly traded firms in my sample, and since firm and plant size are strongly correlated with export status, the 80th percent threshold is clearly a lower bound on the number of exporters.[20]

[20] Baldwin and Gu (2003) also show the fraction of exporters among plants surveyed for the Annual Surveys of Manufactures (ASM) in 1984-1996, which are substantially bigger than the average Canadian manufacturing plant and thus correspond more closely to my sample of publicly traded firms. The fraction of exporters among these plants rose from 31% in 1984 to 55% in 1996, also within the range of thresholds reported in Table 4.

Again, the results in columns 3-10 are qualitatively similar to my baseline specification with the exception of the results based on the 20th percentile threshold, where the Canadian tariff interaction is positive (albeit small and only marginally significant). The magnitude of the coefficient estimates is also relatively stable across specifications, with most estimates in the range -0.7 to -1.3 for the U.S. tariff cut interaction and around -0.05 to -0.20 for the Canadian tariff interaction variable. Clearly, the pattern that larger firms gained relative to smaller firms, and more so in sectors with higher U.S. tariff cuts is robust to a wide range of sales-based proxies for export status. The results related to Canadian tariff reductions also mostly confirm my baseline results, although the magnitude of the reported effects is again smaller than for U.S. tariff cuts.

Next, I use information on the fraction of exporters per Canadian industry published in Statistics Canada (2000) to introduce sectoral variation in the percentile threshold. As discussed above, it is likely that exporting is more common among the firms in my sample. Thus, I normalize the average fraction of exporters across industries to equal 30% as in my binary baseline specification, but preserve the sectoral variation present in the Statistics Canada data. This yields export thresholds ranging from the 90th percentile of the sales distribution in Media to the 5th percentile in Technology Hardware and Equipment (i.e., the fraction of exporters varies between 10% and 95%).[21] Again, the corresponding results are similar to my baseline binary export proxy which used the 30th percentile uniformly across industries.

Finally, I make use of the more limited information on export sales available in my data. In columns 1-4 of Table 5, I reestimate equations (7) and (8) for the 54 firms for which I observe actual exports.[22] In columns 1-2, I classify firms as exporters if they report positive export sales in 1988. In columns 3 to 4, I extend this definition to also include firms which report positive exports during at least one year of CUSFTA's implementation period (1989-1997). As described above, these classifications yield exporter shares of 70% and 80%, respectively. The results for these specifications are again qualitatively similar to before, with exporters experiencing higher abnormal returns than non-exporters, with the difference being stronger in sectors with larger U.S. tariff cuts.

Note that the small size of these two subsample precludes the use of industry fixed effects. Together with the change in sample structure, this makes a direct comparison of coefficient magnitudes with Table 4 difficult. I thus reestimate equations (7) and (8) for this smaller sample, excluding industry fixed effects and using the two binary export proxies based on sales thresholds at the 20th and 30th percentile. The results in columns 5-8 are surprisingly similar to columns 1-4 which use actual export status. Note that Canadian tariff cuts are now estimated to have led to lower relative returns of exporters, in contrast to most of the results from Table 3 and 4. However,

[21] I always classify the firm with the lowest sales in an industry as non-exporting, in order to avoid having industries consisting only of exporters which would then be dropped from the estimation. However, results are similar if I allow for exporter ratios of 100% and use variation from the remaining industries only (available upon request).

[22] I only observe the value of total exports, not the value of exports to the United States. However, given that over 80% of Canadian exports between 1988-1997 went to the U.S., any firm that exported during this period is very likely to have served the U.S. market.

this result is obtained both when using actual export status and when using my binary proxy based on sales, again with almost identical coefficient magnitudes. In conclusion, it seems that using sales as a proxy for export status yields estimates which are very close to proxies based on actual export information. Results using log sales for this smaller sample are harder to compare quantitatively to the results for actual export status because of the different functional form used. But as seen in columns 9 and 10, results are again qualitatively similar.

Longer Event Period In the first column of Table 6, I switch back to my baseline export proxy (log sales) but extend the event period to include the week before the elections (November 14-22). This allows me to evaluate to what extent the election results had been anticipated by market participants. As seen, the size of the coefficient estimates for the U.S. tariff interaction increases by around 25%, so the election outcome seems to have been priced in to a certain degree already.[23] This is not entirely surprising, given that the Conservative Party had been catching up in the opinion polls in the week prior to the elections. Note, however, that the increase in the coefficient magnitude is only about 0.025 per additional event day. This is substantially below the comparable coefficient magnitude for November 21 and 22. Also note that the coefficient on the Canadian tariff cut interaction only changes very little with the extension of the event period. I thus focus on the more sharply defined event of the election itself (November 21 and 22) for the results reported in this paper.

Fama-French Portfolios. In the second column of Table 6, I consider a different abnormal returns model. One concern with the standard market model approach is that it does not control for some important systematic return differences across firms. For example, Fama and French (1992) show that firm size (as proxied by market capitalization) and book-to-market equity ratios are important determinants of the cross-sectional variation in average stock returns. The fact that size by itself is a good predictor of stock returns is potentially problematic, given that I use measures based on firm sales as my export status proxy in most specifications. One way to address this issue is to directly control for the role of size in calculating abnormal returns. I do so by using additional portfolios in the abnormal returns regressions, as suggested by Fama and French (1993):

$$r_{it} = \alpha_i + \beta_{i1} R_{mt} + \beta_{i2} SMB_t + \beta_{i3} HML_t + \sum_{e=1}^{E} d_{et} \left(d_j + \beta_{1e} d_{ix} + \beta_{2e} d_{ix} d\tau_{CAN,j} + \beta_{3e} d_{ix} d\tau_{US,j} \right) + \eta_{it}$$

where SMB_t is the difference in the returns on portfolios of small and large stocks, and HML_t is the difference in the returns of portfolios of high and low book-to-market equity stocks.[24]

[23] Here and in the remaining sections of the paper, I focus on my main specification (8) for the sake of brevity. Results for specification (7) are available upon request. The general pattern of the omitted results is consistent with the predictions discussed in section 2. Events that increased the likelihood of CUSFTA's implementation always led to positive abnormal returns for exporters relative to non-exporters, and events that lowered the likelihood of implementation led to opposite results.

[24] As Fama and French, I further subtract the one-month treasury bill rate from individual stock returns and the return to the market portfolio, R_{mt}. Data on all three factors were taken from Kenneth French's web page at

The results in column 2 are very close to my baseline specification. The most likely explanation for this similarity to the results based on the simpler market model is that systematic differences in abnormal stock returns only become clearly apparent over longer event horizons. For the two-day window considered here, different abnormal return definitions yield almost identical results (also see the related discussion in Andrade et al., 2001).

Outliers and Changes in Sample Composition. In the third column of Table 6, I reestimate my baseline specification (8) but use log returns instead of simple returns as the dependent variable. This provides a natural way of reducing the importance of return outliers. Again, results are basically identical to my baseline specifications.

Columns 4 and 5 also deal with potentially influential observations by dropping sectors with large tariff reductions from my sample. Column 4 drops the Personal Goods sector which saw the most substantial reduction in U.S. tariffs and the second most important reduction in Canadian tariffs. Column 5 removes the Beverage industry which is a strong outlier in terms of Canadian tariff cuts (26.6% compared to the next biggest cut of 12.7% in the Personal Goods sector). The most notable change in results resulting from these regressions is a reduction in the magnitude of the U.S. tariff interaction, and a corresponding increase in the Canadian tariff interaction term, when dropping the Beverage industry. Qualitatively, however, the result pattern is again similar to before, with larger U.S. and Canadian tariff reductions leading to higher abnormal return differences between larger and small firms.

Finally, column 6 excludes three sectors which combine manufacturing and non-manufacturing activities as defined by the Canadian Standard Industrial Classification on which my tariff data are based. These are Oil Equipment & Services (which includes production of construction and mining machinery but also services related to oil extraction), Healthcare Equipment & Services (which includes the production of medical equipment and supplies but also services such as operating hospitals and clinics), and Media (which includes printing and publishing but also broadcasting, advertising and public relations). Excluding these sectors only leads to minor changes in the baseline coefficient estimates.

Alternative Tariff Measures. Column 7 of Table 6 experiments with only using the part of bilateral U.S. and Canadian tariff reductions which exceeds changes in the tariffs between these countries and the rest of the world. Market participants might have used expected tariff changes due to multilateral initiatives such as the Uruguay round of the General Agreement on Tariffs and Trade (GATT) as the most likely scenario in the case of a non-ratification of CUSFTA, rather than no tariff changes at all. Thus, I follow Trefler (20004) by using interaction terms based on $d\tau'_{US,j} = d\tau_{US,j} - d\tau_{USROW,j}$ and $d\tau'_{CAN,j} = d\tau_{CAN,j} - d\tau_{CANROW,j}$ as regressors in my baseline equation, where $d\tau_{USROW,j}$ and $d\tau_{CANROW,j}$ denote average tariff reductions between the U.S. and the rest of the world (excluding Canada), and between Canada and the rest of the world (excluding

Dartmouth which also contains additional information on their construction.

the U.S.). These average tariff reductions are again from Trefler (2004). As seen, using the new adjusted tariffs only slightly changes the baseline estimates. This is probably not surprising, given that the correlation between simple and adjusted tariff reductions is around 90%.

Input Tariffs and Multinational Status. In Table 7, I consider two potential alternative explanations for my results.

A first concern is that output tariff reductions under CUSFTA might partially pick up the impact of intermediate input tariff reductions. As Amiti and Konings (2007) showed for Indonesia, lower tariffs on imported intermediate inputs can lead to significant increases in firm-level productivity. In their sample, these gains were particularly pronounced among firms importing intermediates directly. In the present case, Canadian tariff reductions lowered the costs of inputs imported from the U.S. This should have increased profits of Canadian firms and potentially more so for importers. If importers are among the largest firms in each industry (as the empirical literature on firm-level imports does suggest), my interactions of tariff cuts and firm sales could simply be picking up the effect of cheaper imported intermediates. This is particularly true given the generally positive correlation between input and output tariffs.[25]

To control for this possibility, I rerun my baseline specification but include an additional interaction term between reductions in Canadian intermediate input tariffs and log sales. I construct input tariffs by using the Canadian input-output matrix together with the information on Canadian tariff reductions used previously. In analogy to Amiti and Konings, I construct the input tariff for a given industry j as the weighted average of the Canadian output tariffs of all industries k supplying this industry:

$$input\_tariff_j = \sum_k w_{kj} \times tariff_k$$

where w_{kj} is the cost share of industry k in the production of goods in industry j in 1988. I construct input tariffs for 1988 and 1996 and use the difference as my measure of input tariff reductions due to CUSFTA.

A second potential omitted variable is multinational status. Given that multinational enterprises (MNEs) tend to be among the largest firms in all sectors, my sales proxy for export status is likely to be positively correlated with MNE status. Again, my results might thus simply pick up a differential impact of tariff reductions on MNEs and non-MNEs. Fortunately, my data contain information on foreign affiliate sales and assets for about 80% of firms in my baseline sample, so that I can separately identify the impact of export status (log sales) and MNE status.[26]

Column 1 in Table 7 presents results controlling for intermediate input tariffs, column 2 for multinational status, and column 3 includes both control variables together. As expected, stronger

[25] In my sample, the correlation of Canadian input tariffs with Canadian output tariffs is 28%, and the correlation with U.S. output tariffs is 47%. See below for how import tariffs were constructed.

[26] A firm is classified as an MNE if it either reports positive local affiliate sales abroad or owns assets outside of Canada. Using alternative definitions based on either of these two variables yield almost identical results.

reductions in input tariffs further increase the abnormal return difference between exporters and non-exporters as proxied by firm sales. In contrast, MNE status tends to lower abnormal returns in sectors with higher tariff reductions, ceteris paribus, although the effect is only statistically and economically significant for U.S. tariff reductions. This is consistent with, for example, a tariff-jumping motive for foreign direct investment, in which Canadian MNEs establish U.S. production sites to avoid export duties on their sales there. As U.S. tariffs are eliminated, the value of this local presence is diminished.

Finally, note that the results relating to U.S. tariff cuts are robust to the inclusion of the above control variables, and coefficient magnitudes are similar to our baseline specification. In contrast, the Canadian tariff interaction term becomes insignificant or even slighlty positive once we control for MNE status. This reinforces the impression from the previous robustness checks that the findings related to Canadian tariff reductions are less robust to changes in the estimation equation.

Placebo Checks. I now turn to settings for which I would *not* expect to find significant abnormal return differences between exporters and non-exporters, nor a strong variation of these differences across industries with high and low tariff cuts. Specifically, I estimate specification (8) for dates between 1 November 1987 and 30 June 1988, a period during which the likelihood of CUSFTA's implementation did not vary substantially. I repeatedly draw two consecutive dates from this period at random and estimate (8) for these dates. I then calculate cumulative average abnormal returns (CAARs) based on my estimates of $\hat{\beta}_{1e}$, $\hat{\beta}_{2e}$ and $\hat{\beta}_{3e}$ for these random two-day event windows. I repeat this procedure 1,000 times, thus obtaining a set of 1,000 CAARs estimates comparable to the ones presented in Table 3. I report means, standard deviations and percentiles of the resulting distributions in Table 8.

In the light of the earlier theoretical discussion, one would not expect export status to matter much as a determinant of abnormal returns in this earlier period, both on its own and when interacted with tariff cuts. On the other hand, if my results so far were picking up some general characteristics of firms or sectors correlated with export status and tariff cuts, one would expect parameter estimates of the same magnitude as in my baseline results to show up more frequently than expected from pure sampling variation. For example, if large firms in sectors with high future U.S. tariff cuts systematically experienced above average abnormal returns, my baseline and additional results might be due to some (unknown) omitted factor. Table 8 shows that this is not the case, at least for the U.S. tariff cut interaction. The probability of observing two-day U.S.-tariff-related CAARs on randomly chosen dates which are as large or larger than the magnitudes reported in Table 3 is only about 3%. In contrast, the probability of randomly generating two-day Canadian tariff-related CAARs larger than in Table 3 is somewhat higher at around 30%. In both cases, however, I am unable to reject the hypothesis that that the mean of my generated CAARs is equal to zero (see column 3).[27]

[27] In unreported results, I also used equation (7) to compare abnormal return differences between large and small firms in manufacturing and non-manufacturing industries, where the latter where supposedly less affected by CUS-

Absolute Price Changes Instead of Returns. The model from Section 2 also provides an interesting additional testable prediction related to absolute price changes which I briefly discuss here. Recall that in response to Canadian tariff reductions, domestic Canadian firms were predicted to see a relatively larger fall in profits than exporters *relative* to initial profits. However, the *absolute* decline in profits (i.e., $\Delta\pi$ rather than $\Delta\pi/\pi$) is smallest for the least productive firms and largest for the most productive ones. So absolute price changes (Δp rather than $\Delta p/p$) should be more negative for the more productive exporters than for purely domestic firms.[28] In contrast, the ranking of absolute profit changes of Canadian firms remains unchanged when looking at U.S. tariff reductions. Existing and new exporters see stronger increases than non-exporters, thus implying that the former should see stronger absolute price increases than the latter.[29]

I test this additional prediction by using absolute price changes ($p_t - p_{t-1}$) rather than returns as the dependent variable in a specification based on (8). Using absolute price changes has of course the strong disadvantage that the methodological framework of event studies no longer applies. In particular, the inclusion of stock-specific correlations with the market portfolio no longer has a theoretical basis. Thus, I estimate an adhoc variant of (8) of the form:

$$p_{it} - p_{it-1} = \alpha_i' + \sum\nolimits_{e=1}^{E} d_{et}\left(d_j' + \beta_{1e}' d_{ix} + \beta_{2e}' d_{ix} d\tau_{CAN,j} + \beta_{3e}' d_{ix} d\tau_{US,j}\right) + \eta_{it}' \tag{9}$$

In Table 9, I show results for my log-sales proxy as well as for the preferred binary export proxy from Section 5.2, which uses the 30th percentile of industry sales to separate exporters and non-exporters. Interestingly, the Canadian tariff interaction does indeed change sign although it is only significant for the binary export proxy. Also consistent with the model's predictions, the coefficient on the U.S. tariff interaction remains positive and highly significant. Thus, although the theoretical foundations of these additional results are less robust than that of my baseline specification, they provide additional support for the predictions of heterogeneous firm models.

Additional Events. In conclude my robustness checks by presenting results for the three additional events discussed in Section 3. In Table 10, column 1, I focus on the first trading day after the

FTA. This comparison is not unproblematic since it precludes the use of tariff variation in the identification (tariff data are of course not available for the service sector). CUSFTA also included a number of provisions which might have led to differential abnormal return reactions between large and small firms in non-manufacturing sectors, such as initiatives to liberalize trade in services or make government procurement procedures more accessible to foreign firms (see CUSFTA, 1988, Parts 3-5). Nevertheless, regressing returns on log sales and an interaction term between log sales and a dummy for manufacturing yielded a positive and significant coefficient on the interaction, indicating higher abnormal return differences in manufacturing.

[28] From (1), $p_{it} - p_{it-1} = e_i^{-1}\left(E(\pi_i|I_t) - E(\pi_i|I_{t-1})\right)$. Since $\Delta\pi\left(\gamma, \tau_{ij}, \tau_{ij}'\right) = \gamma^{\sigma-1} f_{ii}\left({\gamma_{ii}^{*'}}^{1-\sigma} - {\gamma_{ii}^{*}}^{1-\sigma}\right)$, and $\gamma_{ii}^{*'} > \gamma_{ii}^{*}$, prices should decline by more for more productive firms. Note, however, that discount rates e_i do not cancel out when looking at absolute price changes. So for exporters to see stronger absolute price declines, I need the additional assumption that differences in e_i are either unrelated to productivity or at least not sufficiently higher for more productive firms.

[29] This again assumes that there are no systematic and sufficiently large differences in discount rates (see the previous footnote). Also note that, in contrast to relative profit changes, the ranking of new and existing exporters is now unambiguous, with the most productive existing exporters experiencing the strongest absolute profit and price increases (compare footnote 7).

successful conclusion of negotiations on October 3, 1987. Similar to the election outcome itself, this event increased the likelihood of an implementation of CUSFTA. Consistent with the theoretical discussion from Section 2, I find stronger abnormal returns of exporters relative to non-exporters in industries with higher U.S. tariff cuts. The same is also true for Canadian tariff reductions, although the size of the corresponding coefficient is again an order of magnitude smaller.

In column 2, I look at the effect of John Turner's announcement that he had instructed the Liberal majority in the Canadian Senate to block CUSFTA until after a general election. In column 3, I focus on the impact of the publication of the Gallup poll on November 7 which predicted a twelve percentage point lead for the Liberal Party. Both events lowered the likelihood of a ratification of CUSFTA. According to the theoretical predictions, one would thus expect to see an effect opposite to the first two events. This is indeed what I find. The positive coefficient estimates on all the U.S. tariff interactions indicate indeed that exporters experienced more negative abnormal returns than non-exporters in sectors in which CUSFTA foresaw higher tariff cuts. The coefficients for the Canadian tariff cut interaction are also positive and statistically significant, albeit only at the 5% level on July 20.[30]

Interestingly, the magnitude of the coefficient estimates for all three additional events is smaller than that of the estimates relating to my baseline event, the Conservative election victory on November 21-22 (see Table 3, column 2). This is consistent with the idea that the latter event presented the most significant change in CUSFTA's implementation probability, given that its ratification by the Canadian parliament was far from assured just before the election but almost certain right after the Conservative victory (see below for a more detailed discussion of changes in implementation probabilities).

5.3 Quantification of Results

I now analyze the quantitative importance and plausibility of the estimated abnormal return differences more closely. I present two sets of figures. First, predicted abnormal returns are easily computed using a simple transformation of my baseline equation (8):

$$E\left(ar_{iE}\right) = \sum\nolimits_{e=1}^{E} d_{et}\left(d_j + \beta_{1e} d_{ix} + \beta_{2e} d_{ix} d\tau_{CAN,j} + \beta_{3e} d_{ix} d\tau_{US,j}\right) \tag{10}$$

where $E\left(ar_{it}\right)$ denotes the expected value of the abnormal returns of stock i during event window E (here, the election victory of the Progressive Conservatives on November 21 and 22).

Secondly, under further assumptions about the probability of CUSFTA's implementation prior to and after the Conservative election victory, I can also compute implied profit changes. To see this, I use the link between returns and profits implicit in equation (1), and solve for profit changes as a function of abnormal returns and ex-ante and ex-post implementation probabilities:

[30] In unreported results, I show that export status as proxied by log sales is also correlated with abnormal returns in the expected way when not relying on variation in tariff cuts (as in Table 3, column 1, for my baseline event). That is, estimating equation (7) for these additional events yields a positive and significant coefficient on log sales for event 2, and a negative and significant one for events 3 and 4.

$$
\begin{aligned}
1+ar_{it+\varepsilon} &= \frac{E\left(\pi_i|I_{t+\varepsilon}\right)}{E\left(\pi_i|I_t\right)} = \frac{prob_{Ct+\varepsilon}\pi_{iC}+\left(1-prob_{Ct+\varepsilon}\right)\pi_{iNC}}{prob_{Ct}\pi_{iC}+\left(1-prob_{Ct}\right)\pi_{iNC}} \\
\Leftrightarrow & \quad \frac{\left(\pi_{iC}-\pi_{iNC}\right)}{\pi_{iNC}} = \frac{ar_{t+\varepsilon}}{prob_{Ct+\varepsilon}-\left(1+ar_{t+\varepsilon}\right)prob_{Ct}}
\end{aligned} \tag{11}
$$

where ar_t are abnormal returns between periods t and $t+\varepsilon$, π_{iC} are per-period profits after a successful implementation of CUSFTA, and π_{iNC} per-period profits without CUSFTA. I_t denotes information available at time t, and $prob_{Ct}$ and $prob_{Ct+\varepsilon}$ the probability of a successful implementation of CUSFTA before and after the Conservative election victory, respectively. Intuitively, the size of the estimated abnormal returns is a function of profits under the free-trade regime and the alternative scenario without tariff cuts, as well as the change in the likelihood of CUSFTA's implementation brought about by the Conservative election victory (controlling for the ex-ante probability, $prob_{Ct}$).

The first line of the first column of Table 11 reports average predicted abnormal returns for exporters and non-exporters for the election event window on November 21 and 22. I first use my preferred binary measure of export status (see Table 4, column 2) to compute abnormal returns, since the 0-1 classification of firms into exporters and non-exporters used there makes the presentation of results straightforward. According to these estimates, exporters experienced average abnormal returns of around 0.9% and non-exporters of around -0.1%, yielding a predicted difference of around one percentage point.

As has been noted before, these abnormal returns are also likely to be influenced by the general impact of a Conservative election victory on stock markets, and possibly by a differential impact across smaller and larger firms (e.g., if the Conservatives were perceived to be "pro big business"). To strip out these two types of confounding impacts, columns 2 and 3 present average predicted abnormal returns based on (10) but disregard industry fixed effects (column 2) or industry fixed effects and the non-interacted export dummy (d_{ix}, column 3) in the return computation. Focusing on these parts of abnormal returns, which are more closely linked to the predictions of heterogenous firm models, yields a somewhat larger return difference between exporters and non-exporters of around 1.1 percentage points (column 2) and 2.7 percentage points (column 3).

Columns 4-6 compute the same statistics but use estimates based on my baseline measure of export status, the log of firm sales (see Table 3, column 2). For comparison with the previous binary measure, I classify all firms as exporters which have sales above the 30th percentile of their respective industry (but I do use their actual sales value to compute predicted abnormal returns). Results in columns 4-6 are very similar to columns 1-3, with estimated return differences of one percentage points for the full specification with industry fixed effects, 1.1 percentage points for the specification excluding industry fixed, and 3.1 percentage points for the specification excluding both industry fixed effects and the level term in log sales.

In lines 2-5 of Table 11, I present results for implied profit changes, using different sets of assumptions about ex-ante and ex-post implementation probabilities. Given the strong support for CUSFTA voiced by the Conservatives and the fact that their representatives had negotiated the agreement in the first place, it seems appropriate to set the ex-post implementation probability to 100% in all scenarios. The implied profit change is thus determined by assumptions about the ex-ante likelihood of implementation. In line 2, I use a value of 0% which is the most conservative assumption in the sense of yielding the smallest implied profit changes. The corresponding results thus provides a useful lower bound for the true profit impact of CUSFTA. Lines 3-5 make more realistic assumptions about the ex-ante probabilities. As discussed, the likelihood of a Conservative election victory was estimated by most observers to be not more than 50% prior to the publication of the opinion polls on November 19 (a Saturday). Thus, in lines 3-5 I choose ex-ante probabilites centered around 50% (30%, 50% and 70%, respectively).

As can be easily verified from (11), implied profit changes are equal to abnormal returns in the most conservative scenario of a 0%-100% change in the implementation probability of CUSFTA, and increase for higher ex-ante probabilities. Depending on the specific way of calculating predicted abnormal returns and the assumptions about ex-ante probabilities, the average implied difference in profit changes between exporters and non-exporters lies between 1 and 10 percentage points for my binary export proxy. The corresponding results for my log-sales measures span a slightly wider range, reaching from one percentage point to close to 14 percentage points in the least conservative scenario. In my view, these magnitudes are clearly economically significant but not implausibly large given the substantial effects of CUSFTA on the Canadian manufacturing sector found previously by authors such as Trefler (2004).

6 Conclusions

This paper presented new empirical evidence on key predictions of heterogeneous firm models. Using the uncertainty surrounding the negotiation and ratification of the Canada-United States Free Trade Agreement in 1987 and 1988, I showed that the pattern of abnormal returns of Canadian manufacturing firms was broadly consistent with the predictions of a class of models based on Melitz (2003).

Specifically, the election victory of the ruling Conservative party (a strong supporter of CUSFTA) led to significant stock market gains of exporting firms relative to non-exporters. Moreover, these relative gains were higher in sectors with larger U.S. tariff cuts. The same pattern was also found for earlier events which increased the likelihood of CUSFTA's implementation. In contrast, events which lowered the likelihood of implementation resulted in negative abnormal returns of exporters relative to non-exporters. Again, these losses were stronger in sectors with higher expected U.S. tariff cuts.

Results for Canadian tariff cuts were slightly less consistent across specifications. While most results indicate that exporting firms also gained relative to non-exporting firms in response to such tariff reductions, the corresponding coefficient estimates were generally small, sometimes insignificant and had the wrong sign in a few cases. I noted that this is not necessarily evidence against the relevance of heterogeneous firm models in general, given that many of the existing theoretical results on domestic tariff reductions seem to depend on assumptions about market entry and specific functional forms (e.g., linear demand or fixed costs) and need not carry over to more general settings.

To evaluate the quantitative importance and plausibility of the estimated return differences, I also calculated the CUSFTA-induced change in the expected future profits of active firms implied by my estimates. Based on assumptions about the change in the likelihood of CUSFTA's implementation brought about by the Conservative election victory, I estimated that CUSFTA increased exporters' per-period profits by 6%-7% relative to non-exporters in the most plausible scenarios, and up to 14% under more extreme assumptions.

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A Theoretical Predictions

In this appendix, I derive predictions for the impact of trade liberalization on profits using two recent and influential models of heterogeneous firms: Melitz and Ottaviano (2008) and Chaney (2008). Both allow for asymmetric country sizes and tariff barriers, which makes them particularly suitable for analyzing bilateral agreements such as CUSFTA, but still deliver closed-form solutions for relative profit changes. Below, I keep the authors' notation to facilitate the comparability of the analysis with the original contributions.

A.1 Chaney (2008)

I start with Chaney's extension of Melitz (2003) to asymmetric countries and trade barriers which is very similar to the model outlined in Section 2 of this paper. The main difference is that his upper-level utility function takes a Cobb-Douglas rather than a quasi-linear form, so that expenditure on the differentiated goods sector is not exogeneously fixed but depends on wages and profits. In equilibrium, however, total expenditure is a fixed multiple of wage income. Since, similar to my own analysis, wages are fixed through the presence of a freely traded numeraire good, expenditure is a function of exogeneous parameters only. The analysis of the impact of trade liberalization on profits is thus identical to my own, as I briefly demonstrate now.

Specifically, in Chaney's model firm-level profits of a firm in i with productivity φ from serving market j are (disregarding industry superscripts):

$$\pi_{ij}(\varphi) = \frac{p_{ij}(\varphi)\, q_{ij}(\varphi)}{\sigma} - f_{ij}$$

where $p_{ij}(\varphi)$ is the price charged in country j by a firm from country i with productivity φ, $q_{ij}(\varphi)$ is the local demand in j associated with this price, f_{ij} are fixed costs in units of the numeraire associated with entering a market j from i, and σ denotes the elasticity of substitution between varieties in the underlying CES subutility function.

Using equilibrium outcomes for prices, demand and productivity cutoffs (see equations 8 and 9 in Chaney), profits can be rewritten as:

$$\pi_{ij}(\varphi) = f_{ij}\left(\frac{\varphi}{\bar{\varphi}_{ij}}\right)^{\sigma-1} - f_{ij}$$

where $\bar{\varphi}_{ij}$ is the minimum productivity level of firms exporting from i to j, which in turn can be written as a function of income (Y_j), variable trade costs (τ_{ij}), the shape parameter γ of the underlying productivity distribution (also assumed to be Pareto) and sectoral expenditure shares μ:[31]

$$\bar{\varphi}_{ij} = \left[\frac{f_{ij}}{\mu}\frac{\gamma\sigma}{\gamma-\sigma+1}\frac{1}{(1+\lambda_5)}\right]^{1/\gamma}\left(\sum_{k=1}^{N}(Y_k/Y_j)\left(\frac{\tau_{kj}w_k}{\tau_{ij}w_i}\right)^{-\gamma}\left(\frac{f_{kj}}{f_{ij}}\right)^{\frac{\sigma-\gamma-1}{\sigma-1}}\right)^{1/\gamma}$$

Note the similarity to my earlier cutoff equation (4). In particular, the key parameters for the analysis of trade liberalizations (τ) enter in the exact same way. Thus, changes in cutoffs will be identical to my simplified version presented earlier. Together with the fact that profits can be written as a function of cutoffs as before, implies that all of my previous results carry through. In particular, new and existing exporters will see stronger relative profit increases compared to domestic Canadian firms as a reaction to lower U.S. tariffs; and lower Canadian tariffs will lead to proportionately higher losses for domestic firms than for exporters.

A.2 Melitz and Ottaviano (2008)

The model by Melitz and Ottaviano (2008) contains a number of differences compared to Chaney (2008) and the model of Section 2, such as linear demand and variable markups. Key for our analysis, however, is the distinction between the "short-run" version of their model, in which the number of potential entrants is fixed as before, and the "long-run" version in which the number of

[31] λ_5 collects constants in σ, μ, and γ (see footnote 11 in Chaney).

potential entrants is determined by a free entry condition. Similar to the model of Section 2 and Chaney (2008), wages are exogeneously fixed via the presence of a freely tradable numeraire good.

In the following, I focus on the two-country version of Melitz and Ottaviano (see Section 3 of their paper) and retain their original notation. The profits of a domestic firm can be split into profits derived from domestic sales and profits derived from export sales (π_D and π_X, respectively). Again, I focus on Canadian firms, denoted by a superscript H (for 'Home') in the following:

$$\pi_D^H(c) = \frac{L^H}{4\gamma}\left(c_D^H - c\right)^2$$

$$\pi_X^H(c) = \frac{L^F}{4\gamma}\left(\tau^F\right)^2\left(c_X^H - c\right)^2$$

where c denotes the marginal costs of a firm, and L^H and L^F the number of consumers in the home (Canadian) and foreign (U.S.) market, respectively. Iceberg-type trade costs associated with exporting to the U.S. are denoted by τ^F, and γ captures the degree of differentiation between products (see Melitz and Ottaviano, p. 297). Finally, c_D^H and c_X^H are the threshold levels of marginal costs above which Canadian firms do not enter their domestic and the U.S. market, respectively.

The change in profits of a Canadian firm in response to U.S. tariff reduction is thus:

$$\Delta\pi^H\left(c;\tau^F,\left(\tau^F\right)'\right) = \frac{L^H}{4\gamma}\left[\left(\left(c_D^H\right)' - c\right)^2 - \left(c_D^H - c\right)^2\right] + \frac{L^F}{4\gamma}\left[\left(\left(\tau^F\right)'\right)^2\left(\left(c_X^H\right)' - c\right)^2 - \left(\tau^F\right)^2\left(c_X^H - c\right)^2\right]$$

where τ^F denotes the initial tariff and $\left(\tau^F\right)'$ the new (lower) tariff, and $\left(c_D^H\right)'$ and $\left(c_X^H\right)'$ are the cutoffs associated with the new tariff. Likewise, for Canadian import tariff reductions (lower τ^H) we have:

$$\Delta\pi^H\left(c;\tau^H,\left(\tau^H\right)'\right) = \frac{L^H}{4\gamma}\left[\left(\left(c_D^H\right)' - c\right)^2 - \left(c_D^H - c\right)^2\right] + \frac{L^F}{4\gamma}\left(\tau^F\right)^2\left[\left(\left(c_X^H\right)' - c\right)^2 - \left(c_X^H - c\right)^2\right]$$

The change in profits is thus determined by the change in the cutoffs and (for U.S. tariff reductions) the direct impact of lower U.S. import tariffs (τ^F).

For most of their analysis, Melitz and Ottaviano assume a Pareto parameterization of the cost draws c, i.e. $G(c) = \left(c/c_M\right)^k$ with $c \epsilon [0, c_M]$. They also distinguish between short-run and long-run effects as discussed. In the short run, the number of incumbent firms in Canada and the U.S. is fixed at $\bar{N}_D^H$ and $\bar{N}_D^F$, respectively. Incumbents observe their cost draw c and decide whether to produce or not. The domestic Canadian cutoff is then implicitly defined by (see equation 28, p. 308):

$$\frac{\alpha - c_D^H}{\left(c_D^H\right)^{k+1}} = \frac{\eta}{2(k+1)\gamma}\left[\frac{\bar{N}_D^H}{\left(\bar{c}_M^H\right)^k} + \left(\tau^H\right)^{-k}\frac{\bar{N}_D^F}{\left(\bar{c}_M^F\right)^k}\right]$$

where $\bar{c}_M^F$ and $\bar{c}_M^H$ denote the upper bound of the distribution of marginal costs of incumbent firms in the two countries. In the long run, with the number of incumbent firms determined by a zero profit condition, this becomes (see equation 23, p. 305):

$$c_D^H = \left[\frac{\gamma\phi}{L^H} \frac{1 - \left(\tau^F\right)^{-k}}{1 - \left(\tau^H \tau^F\right)^{-k}} \right]^{1/(k+2)}$$

Export cutoffs are simply the other country's domestic cutoff, divided by the trade costs of accessing the foreign market (for Canada, $c_X^H = c_D^F / \tau^F$). Thus, in the short-run c_X^H is implicitly defined by:

$$\frac{\alpha - c_X^H \left(\tau^F\right)}{\left(c_X^H\right)^{k+1} \left(\tau^F\right)^{k+1}} = \frac{\eta}{2\left(k+1\right)\gamma} \left[\frac{N_D^F}{\left(\bar{c}_M^F\right)^k} + \left(\tau^F\right)^{-k} \frac{N_D^H}{\left(\bar{c}_M^H\right)^k} \right]$$

In the long-run, c_X^H becomes:

$$c_X^H = \left(\tau^F\right)^{-1} \left[\frac{\gamma\phi}{L^F} \frac{1 - \left(\tau^H\right)^{-k}}{1 - \left(\tau^F \tau^H\right)^{-k}} \right]^{1/(k+2)}$$

A.2.1 Short-run effects

Taking partial derivates of the domestic cutoff with respect to the two tariffs, I obtain $\partial c_D^H / \partial \tau^F = 0$ and $\partial c_D^H / \partial \tau^H > 0$. That is, unilateral domestic liberalization lowers the domestic cost cutoff (i.e., the least efficient firms exit), whereas lower U.S. tariffs have no impact. Likewise, $\partial c_X^H / \partial \tau^F < 0$ and $\partial c_X^H / \partial \tau^H = 0$, i.e., U.S. tariff reductions raise the cost cutoff (less efficient firms start exporting) but Canadian tariff reductions have no impact.

Starting with the response to U.S. tariff reductions, I again compare the profit changes for continuing exporters, new exporters and domestic firms:[32]

$$\frac{\Delta\pi_{D,X}^H \left(c; \tau^F, \left(\tau^F\right)'\right)}{\pi_{D,X}^H \left(c; \tau^F\right)} = \frac{\frac{L^F}{4\gamma} \left[\left(\left(\tau^F\right)'\right)^2 \left(\left(c_X^H\right)' - c\right)^2 - \left(\tau^F\right)^2 \left(c_X^H - c\right)^2 \right]}{\frac{L^H}{4\gamma} \left(c_D^H - c\right)^2 + \frac{L^F}{4\gamma} \left(\tau^F\right)^2 \left(c_X^H - c\right)^2} > 0$$

$$\frac{\Delta\pi_{D,S}^H \left(c; \tau^F, \left(\tau^F\right)'\right)}{\pi_{D,S}^H \left(c; \tau^F\right)} = \frac{\frac{L^F}{4\gamma} \left[\left(\left(\tau^F\right)'\right)^2 \left(\left(c_X^H\right)' - c\right)^2 - \left(\tau^F\right)^2 \left(c_X^H - c\right)^2 \right]}{\frac{L^H}{4\gamma} \left(c_D^H - c\right)^2} > 0$$

$$\frac{\Delta\pi_{D,DOM}^H \left(c; \tau^F, \left(\tau^F\right)'\right)}{\pi_{D,DOM}^H \left(c; \tau^F\right)} = \frac{\left[\frac{L^H}{4\gamma} \left(\left(c_D^H\right)' - c\right)^2 - \frac{L^H}{4\gamma} \left(c_D^H - c\right)^2 \right]}{\frac{L^H}{4\gamma} \left(c_D^H - c\right)^2} = 0$$

That is, as in Section 2 new and continuing exporters gain relative to domestic firms in response to U.S. tariff reductions.

Looking next at Canadian tariff reductions, profit changes for exporters, and for continuing and exiting domestic firms are as follows:

$$\frac{\Delta\pi_{D,X}^H \left(c; \tau^H, \left(\tau^H\right)'\right)}{\pi_{D,X}^H \left(c; \tau^H\right)} = \frac{\frac{L^H}{4\gamma} \left[\left(\left(c_D^H\right)' - c\right)^2 - \left(c_D^H - c\right)^2 \right]}{\frac{L^H}{4\gamma} \left(c_D^H - c\right)^2 + \frac{L^F}{4\gamma} \left(\tau^F\right)^2 \left(c_X^H - c\right)^2} < 0$$

[32] Details of the derivations of these and the following results are available from the author upon request.

$$\frac{\Delta\pi^H_{D,DOM}\left(c;\tau^H,\left(\tau^H\right)'\right)}{\pi^H_{D,DOM}\left(c;\tau^H\right)} = \frac{\frac{L^H}{4\gamma}\left[\left(\left(c^H_D\right)'-c\right)^2-\left(c^H_D-c\right)^2\right]}{\frac{L^H}{4\gamma}\left(c^H_D-c\right)^2} < \frac{\Delta\pi_X\left(\gamma,\tau_{ji},\tau'_{ji}\right)}{\pi_X\left(\gamma\right)}$$

$$\frac{\Delta\pi^H_{D,EXIT}\left(c;\tau^H,\left(\tau^H\right)'\right)}{\pi^H_{D,EXIT}\left(c;\tau^H\right)} = \frac{0-\frac{L^H}{4\gamma}\left(c^H_D-c\right)^2}{\frac{L^H}{4\gamma}\left(c^H_D-c\right)^2} = -1 < \frac{\Delta\pi_{DOM}\left(\gamma,\tau_{ji},\tau'_{ji}\right)}{\pi_{DOM}\left(\gamma\right)}$$

Thus, Canadian tariff reductions will reduce profits of all Canadian firms but exporters will be less affected than both continuing and exiting domestic firms. Intuitively, the part of exporters' total profit derived from the U.S. market is not affected by Canadian tariff cuts, so that the relative decline in total profits is smaller. Secondly, linear demand implies that the percentage loss in domestic profits is smaller for more productive and thus bigger firms. To summarize, the short-run predictions of Melitz and Ottaviano with respect to profits (and thus stock prices) are identical to the model from Section 2.

A.2.2 Long-run effects[33]

From the above long-run cutoffs, it is easy to see that $\frac{\partial c^H_D}{\partial \tau^F} > 0$, $\frac{\partial c^H_D}{\partial \tau^H} < 0$, $\frac{\partial c^H_X}{\partial \tau^F} < 0$ and $\frac{\partial c^H_X}{\partial \tau^H} > 0$. The corresponding changes in profits in response to U.S. tariff reductions for continuing domestic firms, and for new and existing exporters are:

$$\frac{\Delta\pi^H_{D,DOM}\left(c;\tau^F,\left(\tau^F\right)'\right)}{\pi^H_{D,DOM}\left(c;\tau^F\right)} = \frac{\frac{L^H}{4\gamma}\left[\left(\left(c^H_D\right)'-c\right)^2-\left(c^H_D-c\right)^2\right]}{\frac{L^H}{4\gamma}\left(c^H_D-c\right)^2} < 0$$

$$\begin{aligned}\frac{\Delta\pi^H_{D,S}\left(c;\tau^F,\left(\tau^F\right)'\right)}{\pi^H_{D,S}\left(c;\tau^F\right)} &= \frac{\frac{L^H}{4\gamma}\left[\left(\left(c^H_D\right)'-c\right)^2-\left(c^H_D-c\right)^2\right]+\frac{L^F}{4\gamma}\left[\left(\left(\tau^F\right)'\right)^2\left(\left(c^H_X\right)'-c\right)^2-\left(\tau^F\right)^2\left(c^H_X-c\right)^2\right]}{\frac{L^H}{4\gamma}\left(c^H_D-c\right)^2}\\ &> \frac{\Delta\pi_{DOM}\left(c;\tau^F,\left(\tau^F\right)'\right)}{\pi_{DOM}\left(c;\tau^F\right)}\end{aligned}$$

$$\begin{aligned}\frac{\Delta\pi^H_{D,X}\left(c;\tau^F,\left(\tau^F\right)'\right)}{\pi^H_{D,X}\left(c;\tau^F\right)} &= \frac{\frac{L^H}{4\gamma}\left[\left(\left(c^H_D\right)'-c\right)^2-\left(c^H_D-c\right)^2\right]+\frac{L^F}{4\gamma}\left[\left(\left(\tau^F\right)'\right)^2\left(\left(c^H_X\right)'-c\right)^2-\left(\tau^F\right)^2\left(c^H_X-c\right)^2\right]}{\frac{L^H}{4\gamma}\left(c^H_D-c\right)^2+\frac{L^F}{4\gamma}\left(\tau^F\right)^2\left(c^H_X-c\right)^2}\\ &> \frac{\Delta\pi_{DOM}\left(c;\tau^F,\left(\tau^F\right)'\right)}{\pi_{DOM}\left(c;\tau^F\right)}\end{aligned}$$

Intuitively, U.S. tariff reductions increase export profits which raises profits of exporters relative to non-exporters. Domestic profits are reduced for all firms but the percentage profit decline is again stronger for non-exporters because of linear demand.

[33] Note that all results in this subsection are comparisons of two long-run equilibria. They are thus best understood as applying only to those firms active in both equilibria.

On the other hand Canadian tariff reductions make continuing never-exporters better off relative to continuing exporters and firms which leave the export market and only sell domestically. This is because export profits decrease and the increase in domestic profits is stronger for the (smaller) never-exporters, again because of linear demand:

$$\frac{\Delta\pi^H_{D,X}\left(c;\tau^H,\left(\tau^H\right)'\right)}{\pi^H_{D,X}\left(c;\tau^H\right)}=\frac{\frac{L^H}{4\gamma}\left(\left(c^H_D\right)'-c\right)^2+\frac{L^F}{4\gamma}\left(\tau^F\right)^2\left(\left(c^H_X\right)'-c\right)^2}{\frac{L^H}{4\gamma}\left(c^H_D-c\right)^2+\frac{L^F}{4\gamma}\left(\tau^F\right)^2\left(c^H_X-c\right)^2}-1$$

$$\frac{\Delta\pi^H_{D,XEXIT}\left(c;\tau^H,\left(\tau^H\right)'\right)}{\pi^H_{D,XEXIT}\left(c;\tau^H\right)}=\frac{\frac{L^H}{4\gamma}\left(\left(c^H_D\right)'-c\right)^2}{\frac{L^H}{4\gamma}\left(c^H_D-c\right)^2+\frac{L^F}{4\gamma}\left(\tau^F\right)^2\left(c^H_X-c\right)^2}-1<\frac{\Delta\pi^H_{D,X}\left(c;\tau^H,\left(\tau^H\right)'\right)}{\pi^H_{D,X}\left(c;\tau^H\right)}$$

$$\frac{\Delta\pi^H_{D,DOM}\left(c;\tau^H,\left(\tau^H\right)'\right)}{\pi^H_{D,DOM}\left(c;\tau^H\right)}=\frac{\frac{L^H}{4\gamma}\left(\left(c^H_D\right)'-c\right)^2}{\frac{L^H}{4\gamma}\left(c^H_D-c\right)^2}-1>\frac{\Delta\pi^H_{D,X}\left(c;\tau^H,\left(\tau^H\right)'\right)}{\pi^H_{D,X}\left(c;\tau^H\right)}.$$

Table 1: Summary of Events

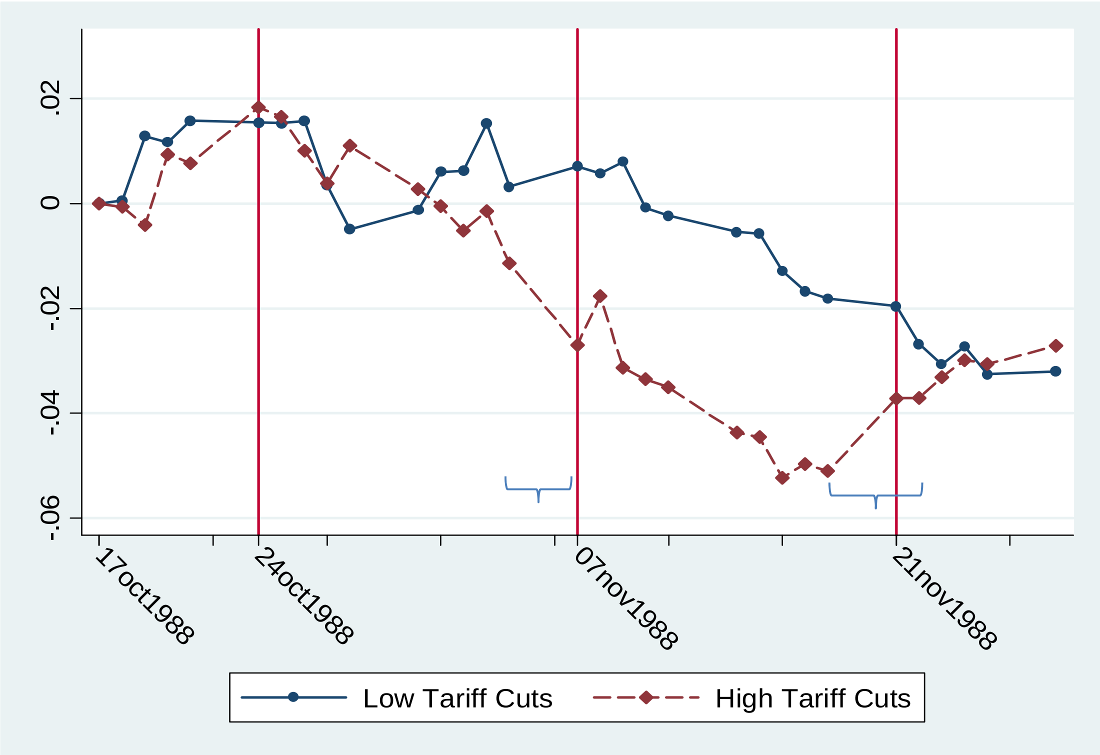
Event Description	Event Date	Likelihood of CUSFTA's implementation
1. Three nationwide opinion polls put the Conservative Party ahead of the opposition on Saturday, November 19. The Conservatives win the election on November 21.	November 21 and 22, 1988	Strongly increased
2. The United States and Canada reach an agreement on CUSFTA on Saturday, October 3, 1987.	October 5, 1987	Increased
3. John Turner instructs the Liberal majority in the Canadian Senate to block the ratification of CUSFTA until after a general election.	July 20, 1988	Decreased
4. A Gallup poll published on the morning of November 7 shows a twelve percentage point lead for the oppositional Liberal Party.	November 7, 1988	Decreased

Table 2: Summary Statistics

Industry	#	Sales			$d\tau_{CAN}$	$d\tau_{US}$
		Median	Min	Max		
Aerospace & Defense	10	238.7	39.5	1456.4	-2.7%	-2.6%
Automobiles & Parts	6	412.0	113.2	15943.3	-0.4%	-0.2%
Beverages	9	57.1	4.7	4611.0	-26.6%	-1.8%
Chemicals	7	158.0	32.8	1385.4	-5.2%	-4.5%
Construction & Materials	21	206.5	0.7	4715.0	-6.0%	-2.9%
Electronic & Electrical Equipment	14	72.3	0.1	1797.7	-3.3%	-2.7%
Food Producers	19	354.5	3.2	3804.0	-4.3%	-2.2%
Forestry & Paper	22	526.1	43.1	5819.1	-3.3%	-0.6%
General Industrials	8	467.5	1.5	6499.8	-7.5%	-2.8%
Healthcare Equipment & Services	4	33.0	0.3	205.9	-4.3%	-2.8%
Household Goods	12	101.8	10.4	450.5	-8.2%	-3.0%
Industrial Engineering	18	97.2	2.7	1737.5	-0.8%	-0.4%
Industrial Metals	24	408.6	0.1	10175.0	-2.8%	-2.0%
Leisure Goods	6	308.9	93.7	1110.5	-4.6%	-3.0%
Media	27	159.2	0.2	4467.9	0.0%	0.0%
Oil Equipment & Services	20	14.5	0.7	3941.0	-2.3%	-1.5%
Personal Goods	3	157.1	8.7	1217.2	-12.7%	-8.7%
Pharmaceuticals & Biotechnology	6	0.9	0.1	156.3	-4.7%	-2.3%
Technology Hardware & Equipment	9	28.5	2.7	6451.3	-1.6%	-1.9%
Tobacco	2	2629.2	413.9	4844.5	-1.4%	0.0%
Total	247	178.3	0.1	15943.3	-5.1%	-2.3%

Notes: Table shows descriptive statistics on the number of firms per industry, firm-level sales (in mill. $CND), and average tariff cuts implemented under CUSFTA. See text for details.

Figure 1: Cumulative average returns during the Canadian election campaign of 1988



Notes: Figure shows differences in cumulative average returns (CARs) between firms above and below the 50^{th} sales percentile in each industry for two groups: the 50% of industries with the largest U.S. tariff cuts and the 50% of industries with the smallest U.S. tariff cuts. All CARs are normalised to zero on Oct. 17 and calculated at the end of each day (i.e., CAR_t-CAR_{t-1} measures the market reaction on day t).

Table 3: Baseline Results

	(1) Return	(2) Return
$d_e * d_x$	0.003	-0.006
	(9.936)**	(12.661)**
$d_e * d_x * d\tau_{US}$		-0.420
		(18.832)**
$d_e * d_x * d\tau_{CAN}$		-0.015
		(3.745)**
Export Status Definition	log(sales)	log(sales)
Firms	247	247
Event Window	Nov. 21-22	Nov. 21-22
Length Event Window	2 days	2 days
Observations Event Window	494	494

Notes: Table shows cumulative average abnormal returns from market-model OLS regressions (t-stats in brackets, based on standard errors clustered per trading day). The dependent variable is daily stock returns. See text for specification details (equations 7 and 8). The independent variables are event dummies (d_e) interacted with export status (d_x), and triple interactions between event dummies, export status and Canadian tariff cuts ($d\tau_{CAN}$) or US tariff cuts ($d\tau_{US}$). Both columns use a continuous definition of export status (log sales). All specifications include industry fixed effects interacted with the event dummy. +, *, and ** denote statistical significance at the 10%, 5% and 1% level, respectively.

Table 4: Alternative Proxies for Export Status based on Firm Sales

	(1) Return	(2) Return	(3) Return	(4) Return	(5) Return	(6) Return	(7) Return	(8) Return	(9) Return	(10) Return	(11) Return	(12) Return
$d_e * d_x$	0.009	-0.015	0.019	-0.009	0.007	-0.024	0.008	-0.020	0.004	-0.013	0.014	-0.012
	(7.923)**	(8.461)**	(12.325)**	(3.548)**	(6.661)**	(12.428)**	(8.815)**	(14.059)**	(4.996)**	(12.861)**	(11.799)**	(7.512)**
$d_e * d_x * d\tau_{US}$		-0.928		-1.792		-1.311		-1.351		-0.755		-1.140
		(7.991)**		(11.689)**		(13.568)**		(17.165)**		(13.939)**		(14.625)**
$d_e * d_x * d\tau_{CAN}$		-0.208		0.084		-0.166		-0.061		-0.065		-0.097
		(6.784)**		(1.915)+		(6.797)**		(2.639)**		(3.869)**		(3.855)**
Export status definition	>30th percent.	>30th percent.	>20th percent.	>20th percent.	>40th percent.	>40th percent.	>60th percent.	>60th percent.	>80th percent.	>80th percent.	Sectoral variation	Sectoral variation
Firms	247	247	247	247	247	247	247	247	247	247	247	247
Event Window	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22
Event Window Length	2 days	2 days	2 days	2 days	2 days	2 days	2 days	2 days	2 days	2 days	2 days	2 days
Observations Event Window	494	494	494	494	494	494	494	494	494	494	494	494

Notes: Table shows cumulative average abnormal returns from market-model OLS regressions (figures in brackets are t-stats based on standard errors clustered per trading day). The dependent variable is daily stock returns. See text for specification details (equations 7 and 8). The independent variables shown in the table are an event dummy (d_e) interacted with export status (d_x), and triple interactions between the event dummy, export status and Canadian tariff cuts ($d\tau_{CAN}$) or US tariff cuts ($d\tau_{US}$), respectively. Firms are classified as exporters if their sales are bigger than the percentile of their industry's sales distribution indicated in row 5. All specifications include industry fixed effects interacted with the event dummy. See text for details. +, *, and ** denote statistical significance at the 10%, 5% and 1% level, respectively.

Table 5: Actual Export Status and Comparison with Proxies based on Sales

	(1) Return	(2) Return	(3) Return	(4) Return	(5) Return	(6) Return	(7) Return	(8) Return	(9) Return	(10) Return
$d_e * d_x$	0.016	0.013	0.015	0.012	0.014	0.010	0.015	0.013	0.002	0.002
	(14.266)**	(12.098)**	(14.726)**	(10.993)**	(16.220)**	(11.151)**	(15.854)**	(11.511)**	(14.798)**	(10.639)**
$d_e * d_x * d\tau_{US}$		-0.338		-0.380		-0.386		-0.299		-0.051
		(8.122)**		(9.551)**		(10.615)**		(8.430)**		(10.062)**
$d_e * d_x * d\tau_{CAN}$		0.061		0.064		0.057		0.061		0.009
		(3.056)**		(4.989)**		(7.814)**		(8.137)**		(9.785)**
Export status definition	Actual (1988 only)	Actual (1988 only)	Actual (extended)	Actual (extended)	>30th percent.	>30th percent.	>20th percent.	>20th percent.	log(sales)	log(sales)
Firms	54	54	54	54	54	54	54	54	54	54
Event Window	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22
Event Window Length	2 days	2 days	2 days	2 days	2 days	2 days	2 days	2 days	2 days	2 days
Observations Event Window	108	108	108	108	108	108	108	108	108	108

Notes: Table shows cumulative average abnormal returns from market-model OLS regressions (figures in brackets are t-stats based on standard errors clustered per trading day). The dependent variable is daily stock returns. See text for specification details (equations 7 and 8). The independent variables shown in the table are an event dummy (d_e) interacted with export status (d_x), and triple interactions between the event dummy, export status and Canadian tariff cuts ($d\tau_{CAN}$) or US tariff cuts ($d\tau_{US}$), respectively. In columns 1-4, I use actual export status. In columns 5-6 and 7-8, firms are classified as exporters if their sales are larger than the 30th and 20th percentile of their industry's sales distribution, respectively. Columns 9 and 10 use a continuous definition of export status (log sales). See text for details. +, *, and ** denote statistical significance at the 10%, 5% and 1% level, respectively.

Table 6: Longer Event Period, Fama-French Portfolios, Log Returns, Influential Sectors, Adjusted Tariffs

	(1) Return	(2) Return	(3) Log Return	(4) Return	(5) Return	(6) Return	(7) Return
$d_e * d_x$	-0.012	-0.007	-0.006	-0.004	-0.005	-0.004	-0.001
	(7.089)**	(7.357)**	(13.795)**	(9.626)**	(11.090)**	(6.660)**	(2.480)*
$d_e * d_x * d\tau_{US}$	-0.539	-0.425	-0.438	-0.333	-0.185	-0.369	-0.345
	(6.605)**	(10.007)**	(18.687)**	(13.870)**	(3.695)**	(16.889)**	(11.787)**
$d_e * d_x * d\tau_{CAN}$	-0.011	-0.018	-0.014	-0.016	-0.134	-0.011	-0.014
	(0.759)	(2.651)**	(3.706)**	(3.864)**	(4.274)**	(2.551)*	(2.060)*
Abnormal Returns Model	Market Model	Fama-French	Market Model	Market Model	Market Model	Market Model	Market Model
Sectors excluded?	None	None	None	Personal Goods	Beverages	Mixed Sectors	None
Tariffs	Unadjusted	Unadjusted	Unadjusted	Unadjusted	Unadjusted	Unadjusted	Minus RoW
Export Status Definition	log(sales)	log(sales)	log(sales)	log(sales)	log(sales)	log(sales)	log(sales)
Firms	247	247	247	244	238	196	247
Event Window	Nov. 14-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22	Nov. 21-22
Length Event Window	7 days	2 days	2 days	2 days	2 days	2 days	2 days
Observations Event Window	1729	494	494	488	476	392	494

Notes: Table shows cumulative average abnormal returns from OLS regressions (figures in brackets are t-stats based on standard errors clustered per trading day). The dependent variable is daily stock returns in columns 1-2 and 4-7, and log returns in columns 3. See text for specification details (equations 7 and 8). The independent variables shown in the table are an event dummy (d_e) interacted with export status (d_x), and triple interactions between the event dummy, export status and Canadian tariff cuts ($d\tau_{CAN}$) or US tariff cuts ($d\tau_{US}$), respectively. All columns use a continuous definition of export status (log sales). All specifications include industry fixed effects interacted with the event dummy. See text for details. +, *, and ** denote statistical significance at the 10%, 5% and 1% level, respectively.

Table 7: Input Tariffs and MNE Status as Controls

	(1) Return	(2) Return	(3) Return
$d_e * d_x$	-0.011	-0.009	-0.016
	(12.645)**	(13.394)**	(8.249)**
$d_e * d_x * d\tau_{US}$	-0.395	-0.555	-0.480
	(17.511)**	(15.780)**	(10.915)**
$d_e * d_x * d\tau_{CAN}$	-0.011	0.003	0.011
	(2.581)*	(0.487)	(2.148)*
$d_e * d_x * d\tau_{INPUT}$	-0.118		-0.192
	(6.162)**		(3.819)**
$d_e * d_{MNE}$		0.018	0.017
		(10.840)**	(10.509)**
$d_e * d_{MNE} * d\tau_{US}$		0.681	0.649
		(6.760)**	(6.252)**
$d_e * d_{MNE} * d\tau_{CAN}$		0.045	0.042
		(0.987)	(0.928)
Export Status Definition	log(sales)	log(sales)	log(sales)
Firms	247	194	194
Event Window	Nov. 21-22	Nov. 21-22	Nov. 21-22
Length Event Window	2 days	2 days	2 days
Observations Event Window	494	388	388

Notes: Table shows cumulative average abnormal returns from market-model OLS regressions (figures in brackets are t-stats based on standard errors clustered per trading day). The dependent variable is daily stock returns. See text for specification details (equations 7 and 8). The independent variables shown in the table are event dummies (d_e) interacted with export status (d_x), and triple interactions between the event dummy, export status and Canadian tariff cuts ($d\tau_{CAN}$) or US tariff cuts ($d\tau_{US}$), respectively. Columns 1 and 3 also include triple interactions between event dummies, export status and intermediate input tariffs ($d\tau_{INPUT}$), and columns 2 and 3 include triple interactions of event dummies, MNE status (d_{MNE}) and U.S. or Canadian tariff cuts. All columns use a continuous definition of export status (log sales). All specifications include industry fixed effects interacted with the event dummy. +, *, and ** denote statistical significance at the 10%, 5% and 1% level, respectively.

Table 8: Parameter Estimates for Non-Event Dates

Coefficient estimate	Mean (sd)	Test mean≠0 (t-stat)	Percentiles 1st	5th	10th	50th	90th	95th	99th
β_{1e}, log sales export proxy	0.000 (0.003)	1.10	-0.009	-0.006	-0.004	0.000	0.004	0.005	0.008
β_{2e}, Can. tariff-export status interaction, log sales export proxy	-0.001 (0.032)	1.48	-0.077	-0.053	-0.047	-0.001	0.039	0.048	0.064
β_{3e}, U.S. tariff-export status interaction, log sales export proxy	0.001 (0.192)	0.09	-0.478	-0.312	-0.251	-0.012	0.258	0.305	0.438
Number of draws					1,000				
Number of firms					247				
Length Event Window					2 days				
Obs. Event Window					494				

Notes: Table shows means, standard deviation and percentiles for the distributions of coefficient estimates shown in the left column. Also shown is the t-stat of a regression of the coefficient estimates on a constant (column "Test mean≠0"). The coefficient estimates were obtained by estimating equation (8) in the main text for randomly chosen pairs of consecutive days in the period 1 November 1987 to 30 June 1988. Results are based on 1,000 repetitions. See text and Table 3 for further details.

Table 9: Absolute Price Changes

	(1) $p_t - p_{t-1}$	(2) $p_t - p_{t-1}$
$d_e * d_x$	-0.012	-0.065
	(1.437)	(1.394)
$d_e * d_x * d\tau_{US}$	-1.392	-16.549
	(3.480)**	(5.481)**
$d_e * d_x * d\tau_{CAN}$	0.252	2.909
	(1.456)	(2.227)*
Export Status Definition	log(sales)	binary (sales>30^{th} percentile)
Firms	247	247
Event Window	Nov. 21-22	Nov. 21-22
Length Event Window	2 days	2 days
Observations Event Window	494	494

Notes: Table shows cumulative average abnormal price changes from market-model OLS regressions (figures in brackets are t-stats based on standard errors clustered per trading day). The dependent variable is daily stock returns. See text for specification details (equation 9). The independent variables shown in the table are event dummies (d_e) interacted with export status (d_x), and triple interactions between the event dummy, export status and Canadian tariff cuts ($d\tau_{CAN}$) or US tariff cuts ($d\tau_{US}$), respectively. All specifications include industry fixed effects interacted with the event dummy. +, *, and ** denote statistical significance at the 10%, 5% and 1% level, respectively.

Table 10: Additional Events

	(1) Return	(2) Return	(3) Return
$d_e * d_x$	-0.000	0.000	-0.000
	(2.197)*	(1.883)+	(0.085)
$d_e * d_x * d\tau_{US}$	-0.062	0.058	0.099
	(7.491)**	(5.997)**	(6.405)**
$d_e * d_x * d\tau_{CAN}$	-0.005	0.005	0.028
	(3.525)**	(2.374)*	(8.703)**
Export Status Definition	log(sales)	log(sales)	log(sales)
Firms	247	247	247
Event Window	Oct.5, 1987	July 20, 1988	Nov. 7, 1988
Length Event Window	1 day	1 day	1 day
Observations Event Window	247	247	247

Notes: Table shows cumulative average abnormal returns from market-model OLS regressions (figures in brackets are t-stats based on standard errors clustered per trading day). The dependent variable is daily stock returns. See text for specification details (equations 7 and 8). The independent variables shown in the table are event dummies (d_e) interacted with export status (d_x), and triple interactions between the event dummy, export status and Canadian tariff cuts ($d\tau_{CAN}$) or US tariff cuts ($d\tau_{US}$), respectively. All columns use a continuous definition of export status (log sales). All specifications include industry fixed effects interacted with the event dummy. +, *, and ** denote statistical significance at the 10%, 5% and 1% level, respectively.

Table 11: Quantification

	(1)	(2)	(3)	(4)	(5)	(6)
Predicted Abnormal Returns						
- Non-Exporters	-0.1%	0.0%	0.0%	-0.1%	0.3%	1.6%
- Exporters	0.9%	1.1%	2.7%	0.9%	1.4%	4.7%
Implied Profits Changes (0-100%)						
- Non-Exporters	-0.1%	0.0%	0.0%	-0.1%	0.3%	1.6%
- Exporters	0.9%	1.1%	2.7%	0.9%	1.4%	4.7%
Implied Profits Changes (30-100%)						
- Non-Exporters	-0.0%	0.0%	0.0%	-0.0%	0.4%	2.4%
- Exporters	1.3%	1.6%	3.9%	1.3%	2.1%	6.9%
Implied Profits Changes (50-100%)						
- Non-Exporters	0.0%	0.0%	0.0%	0.0%	0.7%	3.5%
- Exporters	1.8%	2.3%	5.5%	1.8%	3.1%	10.2%
Implied Profits Changes (70-100%)						
- Non-Exporters	0.3%	0.0%	0.0%	0.4%	1.3%	6.2%
- Exporters	3.2%	4.1%	9.8%	3.2%	6.2%	19.8%
Export Status Definition	>30th percent.	>30th percent.	>30th percent.	log(sales)	log(sales)	log(sales)
Components used in computation of Predicted Abnormal Returns	All	No industry FE	Inter-actions only	All	No industry FE	Inter-actions only

Notes: Table shows predicted average abnormal returns and implied per-period profit changes for exporters and non-exporters. Columns 1-3 use a binary sales-proxy for export status and columns 4-6 use log-sales. See equations (8) and (10) for the underlying specification and Tables 3 and 4, column 2, for the coefficient estimates used. The implied profit changes in rows 2-5 are based on the assumptions about the pre-post change in the likelihood of CUSFTA's implementation indicated in the table. See text for details.