IMPROVED ACCESS TO FOREIGN MARKETS RAISES PLANT-LEVEL PRODUCTIVITY. . . FOR SOME PLANTS*

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Market size matters for innovation and hence for productivity. Improved access to foreign markets will thus encourage firms to simultaneously export and invest in raising productivity. We examine this insight using the responses of Canadian plants to the elimination of U.S. tariffs. Unique "plant-specific" tariff cuts serve as an instrument for changes in exporting. We find that Canadian plants that were induced by the tariff cuts to start exporting or to export more (a) increased their labor productivity, (b) engaged in more product innovation, and (c) had higher adoption rates for advanced manufacturing technologies. Further, these responses were heterogeneous.

I. Introduction

Promoters of free trade tell us that improved access to foreign markets makes domestic firms stronger and more productive. Academic economists naturally wince at such boosterism, but the rhetoric actually has merit. At least as far back as Schmookler (1954), we knew that the larger the market, the more profitable it is for firms to invest in productivity-enhancing activities. Because improved access to foreign markets increases the effective size of the market, it should promote investments that raise firm-level productivity. An alternative way of stating this is that exporting and investing in productivity are complementary activities: doing one makes the other more profitable. This paper shows that Canadian plants that received preferential access to the U.S. market under the terms of the Canada–U.S. Free Trade Agreement

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(FTA) did in fact raise their labor productivity by investing in productivity-enhancing activities.

The complementarity between exporting and investing in productivity appears in Atkeson and Burstein (2007), Bustos (2007, 2008), and Constantini and Melitz (2008), all of whom provide conditions under which a reduction in the costs of exporting induces firms to simultaneously export and invest.[1] Verhoogen (2008) describes a related complementarity between exporting and investing in quality. These papers all allow for Melitz's (2003) heterogeneity in initial productivity. In practice, however, we also observe substantial heterogeneity in the returns to investing in productivity. Stories abound of firms that fail to implement new technologies as successfully as their competitors—one need only think of GM versus Toyota—and these stories are confirmed by careful analysis, such as Stalk and Hout (1990) in the management literature and Aw, Roberts, and Winston (2007, Table 6) in the trade literature. Once one allows two sources of heterogeneity, in initial productivity *and* in the productivity gains from investing, two prominent features of the Canadian data are easily explained.

First, many small and less-productive plants export.[2] This "unproductive exporters" paradox is inconsistent with the Melitz model. It is also inconsistent with the Bustos (2007, 2008) and Verhoogen (2008) models, in which only larger, more productive plants export and invest. Second, it is well known that new exporters have faster productivity growth than nonexporters (e.g., Bernard and Jensen [2004, Figure 1]). We find in addition that this productivity growth differential is declining in initial productivity: it is large for less productive plants and small for more productive plants. Restated, the exporter–nonexporter growth differential displays "negative selection."

To make sense of these two features of our data, we present a simple model of exporting and investing in productivity that features heterogeneity in initial productivity *and heterogeneity in*

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[1] Atkeson and Burstein (2007) focus on how the general equilibrium feedbacks of trade liberalization affect firms' innovation decisions. Aw, Roberts, and Winston (2007) and Aw, Roberts, and Xu (2008, 2009), using data from the Taiwanese electronics industry, emphasize that exporting and R&D are joint decisions. Their approach is very general and therefore subsumes our complementarity between exporting and investing in productivity. Ederington and McCalman (2008) explore the related question of how trade liberalization affects the diffusion rates of new technologies.

[2] This is not unique to Canada. It holds for U.S. data (Bernard et al. 2003, Figure 2B), for Spanish data (Delgado, Fariñas, and Ruano 2002, Figure 1), and for Colombian data (tabulations kindly provided to us by Bernardo Blum).

the productivity gains from investing. To see the role played by the latter, consider two firms with different initial productivities and suppose that both are just indifferent between (1) exporting and investing and (2) doing neither. The initially higher-productivity firm will do well in export markets, so its indifference must be due to low expected productivity gains from investing. The initially lower-productivity firm will do poorly in export markets, so its indifference must be due to large expected productivity gains from investing. Thus, productivity gains from investing are decreasing in initial productivity for the set of firms that are just indifferent between choices (1) and (2).

If the productivity benefits of improved access to foreign markets vary across firms, then we face a significant empirical problem: many of the moments of interest relating foreign market access to productivity are not identified. Imbens and Angrist (1994) showed that if there is a valid instrument for exporting then one can identify a weighted average of the productivity gains from exporting, *but only for those firms that are induced to export because of the instrument*. In terms of our theoretical model, these are the firms that are indifferent between choices (1) and (2).

Empirically we will be able to identify the increase in labor productivity, technology adoption, and innovation for those Canadian plants that were induced to export to the United States as a result of U.S. tariff cuts. We use the cuts associated with the FTA. Combining data on labor productivity, technology adoption, and innovation is an important feature of this paper. A novel feature is that our tariff-cut instrument is plant-specific. That is, we link the tariff-cut data to a plant's 6-digit Harmonized System (HS6) commodity data in order to compute the average tariff cut experienced by the plant. With this instrument in hand, we estimate a heterogeneous response model using the Angrist and Imbens (1995) variant of the local average treatment effect (LATE) estimator.

The role of the instrument turns out to be potentially very important for understanding the seemingly contradictory results reported in the related literature on the causal impact of exporting on productivity. A heterogeneous-response model can make sense of these divergent results. Consider the very different conclusions drawn by Bernard and Jensen (1999) for the United States and De Loecker (2007) for Slovenia. Slovenian firms likely started exporting because of improved access to the European Union and, as a prerequisite for joining European Union supply chains, Slovenian firms likely invested heavily in reducing product defect rates and lowering costs. The implicit instrument—entry into the European Union—picks off new exporters that were investing in new productivity-enhancing technologies. In contrast, most U.S. plants find themselves in a domestic market that is large enough to justify investing even without access to foreign markets. As Bernard and Jensen showed, plants in their U.S. sample likely started exporting because improved productivity from previous investing pushed them past the Melitz (2003) cutoff. These new exporters thus did not experience additional productivity gains from starting to export. The implicit instrument—past productivity growth—picks off new exporters that started investing before exporting. More generally, because (a) different instruments yield different predictions about who exports and (b) different exporters have different productivity gains from exporting, the choice of instrument will matter for conclusions about the productivity benefits of exporting.

Our main finding is that Canadian plants that were induced by the tariff cuts to start exporting or export more (a) increased their labor productivity, (b) engaged in more product innovation, and (c) had higher adoption rates of advanced manufacturing technologies. For plants that start to export (but not for plants that export more), the theory also predicts that there will be "negative selection"; that is, changes (a)–(c) will be largest for plants that are least productive. Our LATE estimates confirm this prediction.

The most significant of several weaknesses of our empirical work is that we measure productivity by value added per worker rather than total factor productivity (TFP). Data on capital stock are unavailable. To partially address this, in Section VI we show that the plants that were induced to raise their labor productivity were also the same plants that grabbed substantial *domestic* market share away from nonexporters. This suggests that these new exporters did indeed increase their TFP.

This paper is related to Bustos (2007, 2008). Using Argentinean data, Bustos (2007, Table 8) shows that firms that began exporting between 1992 and 1996 also increased their technology spending. Bustos (2008) shows that technology spending increased most in sectors that experienced improved access to Brazilian product markets (i.e., Mercursor tariff cuts). Our paper is also related to plant-level studies of the impact of the FTA. See Baldwin and Gu (2003), Trefler (2004), Baggs (2005), Baldwin, Caves, and Gu (2005), Baggs and Brander (2006), and Lileeva

(2008). Of particular interest here are two sets of papers. First, Baldwin, Beckstead, and Caves (2002) and Baldwin and Gu (2004, 2006) find that relative to nonexporters, exporters invest more in R&D and training, adopt more advanced manufacturing technologies, produce fewer products, and have longer production runs per product. Second, Feinberg and Keane (2006, 2009), and Keane and Feinberg (2007) find that the 1983–1996 increase in trade between U.S. multinationals and their Canadian affiliates was driven largely by improved logistics management, such as adoption of just-in-time production techniques. These studies point to a strong link between exporting and investing in productivity.

II. A MODEL OF SELECTION INTO INVESTING AND EXPORTING

Consider a model with two countries, home (Canada) and foreign (United States). Foreign values are denoted with an asterisk. Consumers have CES preferences and the market structure is monopolistic competition. A home firm producing variety i faces home demand $q(i) = p(i)^{-\sigma}A$ and foreign demand $q^*(i) = p^*(i)^{-\sigma}A^*$, where $\sigma > 1$ is the elasticity of substitution between varieties, A is a measure of domestic market size, A^* is a measure of foreign market size, $p(i)$ is the price charged at home, and $p^*(i)$ is the price (inclusive of tariff) charged abroad. Let $\tau(i) - 1$ be the *ad valorem* tariff the firm faces when selling to the foreign market. Turning to costs, a standardized bundle of inputs costs c and produces $\varphi_0'(i)$ units of output. $\varphi_0'(i)$ measures productivity. However, it is easier to work with a familiar transformation of productivity, namely, $\varphi_0 \equiv (\sigma-1)^{\sigma-1}\sigma^{-\sigma}(\varphi_0')^{\sigma-1}$. We are only interested in the firm's static optimization problem. We therefore treat the equilibrium outcomes A, A^*, and $c = 1$ as exogenous parameters. In what follows we drop all i indices.

Consider the standard Melitz (2003) problem as described in Helpman (2006). For a fixed cost F^E, the firm can export. Let $E = 1$ if the firm exports and $E = 0$ otherwise. Then the firm's maximum profits as a function of its exporting decision are

$$\pi_0(E) = \varphi_0\left[A + E\tau^{-\sigma}A^*\right] - EF^E \tag{1}$$

for $E = 0$, 1. See Helpman (2006, equations (1) and (2)). It follows that the firm exports when φ_0 exceeds the Melitz cutoff $F^E/(\tau^{-\sigma}A^*)$.

In addition to an exporting decision, we assume that for a fixed cost F^I the firm can raise its productivity from φ_0 to φ_1.[3] The firm's maximum profits when investing in productivity are

$$\pi_1(E) = \varphi_1\left[A + E\tau^{-\sigma}A^*\right] - EF^E - F^I. \tag{2}$$

The essence of the firm's problem is best understood by considering the difference between profits from (i) exporting and investing versus (ii) neither exporting nor investing. From equations (1) and (2), this difference is

$$\begin{aligned}\pi_1(1) - \pi_0(0) = {} & [\varphi_0\tau^{-\sigma}A^* - F^E] + [(\varphi_1 - \varphi_0)A - F^I] \\ & + [(\varphi_1 - \varphi_0)\tau^{-\sigma}A^*].\end{aligned} \tag{3}$$

The first term in brackets equals the increase in profits from exporting without investing in productivity. The second term in brackets equals the increase in profits from investing in productivity without exporting. The third term captures the complementarity between investing and exporting—it is the increase in variable profits that results from both exporting and investing as opposed to doing just one or the other. It is necessarily positive because productivity gains raise profits on all units sold, including foreign sales, and hence raise the profits from exporting. This complementarity can also be thought of as a familiar market-size effect that appears in many different models.

The firm's optimal choices are illustrated in Figure I, where initial productivity φ_0 is plotted against the productivity gains from investing, $\varphi_1 - \varphi_0$. When productivity gains are small the firm never invests and we are in a Melitz world: the firm exports if and only if initial productivity is above the Melitz threshold. The Melitz threshold is the vertical line in Figure I. Given that the firm is exporting, it will invest if and only if the productivity gains are above some threshold. This threshold is the horizontal line in Figure I.[4] The interesting region is where the first two terms in equation (3) are negative, so that the firm will not export without investing and will not invest without exporting. In this

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[3] It makes no difference to our conclusions if there are only marginal costs of investing or both marginal and fixed costs of investing. The key is that a switch from $E = 0$ to $E = 1$ makes investing in productivity more profitable. Restated, the key is that exporting and investing are complements in the sense of Milgrom and Roberts (1990).

[4] A firm that already exports will invest if $\pi_1(1) > \pi_0(1)$ or equivalently if $(\varphi_1 - \varphi_0)(A + \tau^{-\sigma}A^*) - F^I > 0$ or equivalently if $\varphi_1 - \varphi_0 > F^I/(A + \tau^{-\sigma}A^*)$. $\varphi_1 - \varphi_0 = F^I/(A + \tau^{-\sigma}A^*)$ defines the threshold and the horizontal line in Figure I.

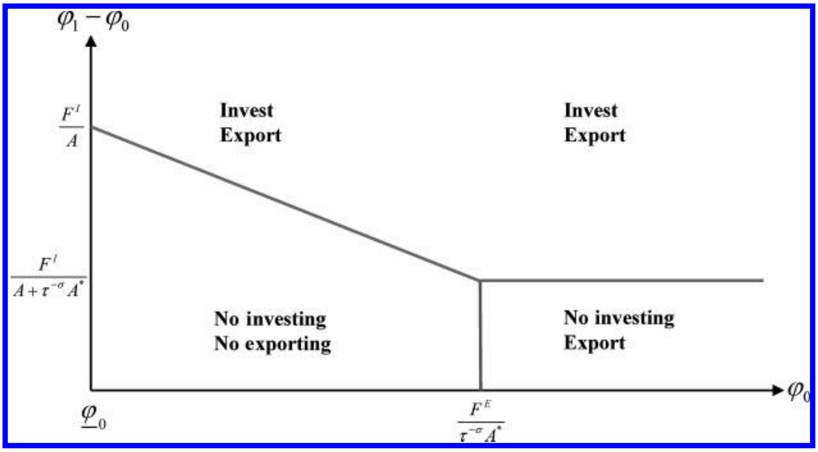

FIGURE I
The Optimal Choices of Exporting and Investing

region the complementarity between exporting and investing may nevertheless make it worthwhile for the firm to export and invest. To pin this down more precisely, suppose that in this region the firm must choose either (i) to export and invest or (ii) to do neither. The firm is indifferent between these two choices when $\pi_1(1) = \pi_0(0)$ or, from equation (3), when

$$\varphi_1 - \varphi_0 = -\varphi_0 \frac{\tau^{-\sigma}A^*}{A + \tau^{-\sigma}A^*} + \frac{F^I + F^E}{A + \tau^{-\sigma}A^*}. \tag{4}$$

Above this line the firm prefers to export and invest. Below it, the firm prefers to do neither. Equation (4) is the downward-sloping line in Figure I.

The horizontal axis in Figure I starts at $\underline{\varphi}_0 \equiv F^E/(\tau^{-\sigma}A^*) - F^I/A$. For any firm with $\varphi_0 \geq \underline{\varphi}_0$, Figure I is *a complete description of the firm's optimal choices*. This is proved in the Appendix. The Appendix also fully characterizes the optimal choices of a firm with $\varphi_0 < \underline{\varphi}_0$, but such a firm is irrelevant for our subsequent empirical work on exporting because, in the terminology of LATE, it is a nonswitcher into exporting.

Consider now an improvement in access to the foreign market due to a fall in the foreign tariff τ. See Figure II. There are three effects. First, the downward-sloping equation (4) rotates clockwise around its fixed vertical intercept. Thus, some firms that previously neither exported nor invested now find themselves choosing

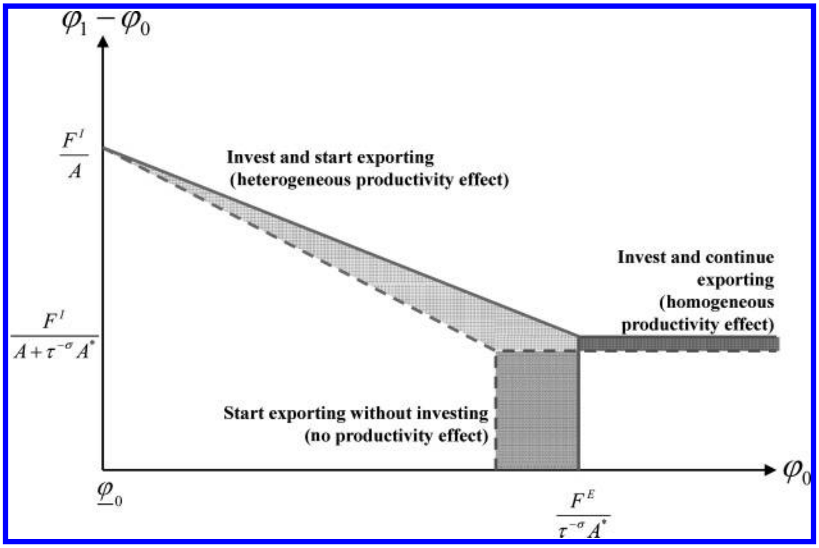

FIGURE II
Switching Behavior Induced by Improved Foreign Market Access

to both export and invest. For this group, the causal effect on productivity of improved market access is given by the downward-sloping line in Figures I and II, that is, by equation (4). Second, the fall in τ also causes a leftward shift of the Melitz cutoff. Thus, some firms that previously neither exported nor invested now find themselves exporting without investing. For these firms improved market access has no causal effect on productivity. Third, the fall in τ shifts the horizontal line in Figure II down. Thus, firms that already exported now invest. For this group, the causal effect on productivity of improved market access is given by the horizontal line in Figures I and II.

The primary result of this section is that improved access to foreign markets raises productivity for some firms but not all firms; that is, productivity responses are heterogeneous. This has important implications for empirical work. No researcher has ever adequately reported how productivity responses vary with initial productivity.

A much less important result of this section is that the complementarity between exporting and investing leads to the particular form of heterogeneity shown in Figure II. For firms that begin exporting, the effects are decreasing in φ_0, and for firms that already exported, the effects are the same for all firms. *A* *priori* there is no reason to think that our mechanism will be the dominant mechanism in a richer model that allows for other factors and other sources of heterogeneity. For now we simply note that the form of heterogeneity displayed in Figure II is what we find empirically. We also note that the empirical analysis to come imposes *none* of the theoretical structure developed in this section.

III. The Canada–U.S. Free Trade Agreement and the Data

III.A. A Brief History of the FTA

We are interested in the effects of improved market access on firms' decisions to export and invest. We use tariff reductions mandated under the Canada–U.S. Free Trade Agreement to examine these effects. Negotiations for the FTA began in September 1985. There was considerable uncertainty about whether there would be an agreement until after the November 1988 general election returned the Conservatives for a second term. The agreement went into effect on January 1, 1989. By 1996, the last year for which we have plant-level data, the tariff on each tariff-line item was down to less than one-fifth of its 1988 level, and by 1998 all tariffs were eliminated. See Brander (1991) for details.

Figure III plots real Canadian manufacturing exports to the United States. Data are from Trefler (2004). These exports changed little during the FTA negotiation period 1985–1988. They also changed little during the severe 1989–1991 recession, the worst recession in Canadian manufacturing since the 1930s. However, exports climbed spectacularly after 1991, increasing by 75% in just five years. Romalis (2007) shows a similar time profile for exports of goods that were subject to the largest tariff cuts. This 75% rise over five years was unprecedented in Canadian history and reflects the fact that a massive change occurred in Canadian manufacturing exports starting in 1991.

We have plant-level exports for 1984 and 1996. This means that we cannot examine the annual dynamics that are the focus of the literature spawned by the seminal papers of Roberts and Tybout (1997) and Clerides, Lach, and Tybout (1998). Instead, we will be interested in longer-run effects. This also means that we do not know the plant's first export date, information that is central to Bernard and Jensen's (1999, 2004) analysis of preentry versus postentry performance. However, we are not interested in

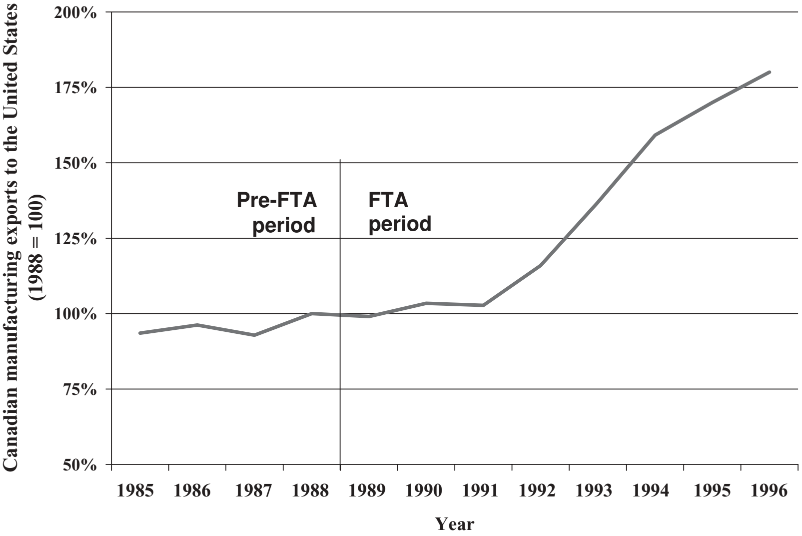

FIGURE III
Canadian Manufacturing Exports to the United States

the impact of exporting *per se*, but in the impact of improved foreign market access. Our identification comes not from *when* exporting occurs but from *why* it occurs. Specifically, identification comes from plant-specific tariff-cut instruments.[5]

There are a couple of outstanding data issues. First, as explained below, we will be looking only at plants that were alive from 1984 to 1996. When such long-lived plants enter export markets, they have a very strong tendency to remain in export markets. Statistics on this appear in Online Appendix IV. Here we simply note that in 1996 the average ratio of exports to sales was very high for exporters. It was 22% for those that started exporting

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[5] There is undoubtedly a small group of plants in our sample that started exporting during the pre-FTA 1985–1988 period for reasons that had nothing to do with the FTA tariff cuts. This would pose problems for a preentry versus postentry analysis, but poses no problem here. To see this, note that by definition these plants' export decisions were uncorrelated with the FTA tariff cuts. Presumably their productivity growth after 1988 was also uncorrelated with the FTA tariff cuts. Then, because the IV estimator purges out all data that are uncorrelated with the tariff-cut instrument, the IV estimator will not use the data associated with these plants. Their presence is irrelevant. (It is also possible that plants started exporting in anticipation of the FTA. Then their entry decision is correlated with the tariff-cut instrument. This will not matter provided their productivity gains began after 1988. If they began before 1988 then we will underestimate the impact of the tariff cuts.)

after 1984 and even higher (43%) for those that started earlier. These are committed exporters. Second, the 1984 export data do not indicate the destination of exports. However, 83% of Canadian manufacturing exports in 1984 went to the United States, and this number rose after implementation of the FTA. Thus, during the FTA period the vast majority of new entry into export markets and of increased exports likely involved the U.S. market.

III.B. Description of the Plant-Specific Tariff Variable

We will instrument changes in exporting with FTA-mandated tariff cuts. For each plant we have data on every 6-digit Harmonized System (HS6) good it produces. We use this information to construct the FTA-mandated average U.S. tariff cuts experienced by each Canadian plant. We are the first ever to construct plant-specific tariffs. We start with statutory tariff rates, which are legislated at the HS8 level. Let τ_{jt} be the U.S. tariff against Canadian imports of HS8 commodity j in year t. We compute the tariff change $\tau_{j,1988} - \tau_{j,1996}$ at the HS8 level and then aggregate up to the HS6 level in the standard way, using import weights (1996 U.S. imports from Canada). The resulting HS6 tariff changes are then matched to the HS6 plant-level commodity data. Additional information about construction of the tariff changes appears in Online Appendix I.

Having matched tariff changes to each HS6 commodity produced by the plant, we then average the tariff changes across those products produced by the plant. This is our plant-specific tariff change. We denote it by $\Delta\tau$. In the results reported, we use the unweighted average tariff change. However, as shown in Online Appendix II, it does not matter what weights are used in calculating the average tariff change.

We will be using the tariff changes as an instrument for exporting behavior. For reasons discussed in Angrist and Imbens (1995) and reviewed below, we will need to transform the tariff instrument into a set of mutually orthogonal binary variables. To this end, our 1984 nonexporters were divided into four groups, based on quartiles of the distribution of $\Delta\tau$. We define four mutually orthogonal binary variables $\Delta\tau_q$ $(q = 1, \ldots, 4)$ that indicate the quartile to which each plant belongs. Thus, if a plant's $\Delta\tau$ puts it in quartile q', then $\Delta\tau_{q'} = 1$ and $\Delta\tau_q = 0$ for $q \neq q'$. Many of the plants in the top quartile $(q = 4)$ have tariff cuts in excess of 10%. These are deep cuts, especially given that many of the affected plants produce "low-end," unskilled-intensive manufactures with mark-ups of less than 10%. All of the plants in the bottom quartile ($q = 1$) have zero tariff cuts.[6]

III.C. Sample Moments

In accordance with the theory (Figure II) we distinguish two subsamples of plants.

1984 nonexporters: These are the 5,233 Canadian plants that (a) did not export in 1984 and (b) survived until 1996. In terms of Figure II, these are the plants to the left of the Melitz cutoff. For these plants, we expect the heterogeneous productivity responses to be decreasing in initial productivity.

Old exporters: These are the 1,607 Canadian plants that exported in both 1984 and 1996. In terms of Figure II, these are the plants to the right of the Melitz cutoff. For these plants we expect the heterogeneous productivity responses to be independent of initial productivity.

Table I reports some basic sample statistics. Our 1984 nonexporters sample consists of 3,114 plants that did not export in either 1984 or 1996 (*nonexporters*) and 2,119 plants that did not export in 1984, but did export in 1996 (*new exporters*). Column (1) reports the difference between new exporters and nonexporters after controlling for four-digit SIC fixed effects (using OLS). The Canadian SIC has 208 four-digit industries. Table I makes it clear that new exporters and nonexporters were very different even before the FTA. New exporters employed more workers and had higher labor productivity than nonexporters. New exporters also had higher labor productivity growth both in the FTA period (1988–1996) *and in the pre-FTA period* (1984–1988).

Column (3) of Table I reports the difference between old exporters and new exporters after controlling for industry fixed effects. Old exporters were larger and more productive and had higher export-to-sales ratios.

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[6] There is a second reason for coding tariffs into quartiles. Some U.S. tariffs were so high as to be well above the level needed to choke off imports. This leads to a problem with using the continuous tariff data ($\Delta\tau$). For example, if the prohibitive level of tariffs is 10% and the actual level of tariffs varies between 10% and 20%, then we will estimate a zero impact of tariffs. Putting tariffs in quartile-based bins, especially for the highest-quartile bin, helps avoid this problem. See Online Appendix III for a more detailed discussion. Also see Online Appendix Table B.20 for an example of estimates using the continuous tariff data ($\Delta\tau$).

TABLE I
AVERAGE PLANT CHARACTERISTICS

	New exporters less nonexporters		Old exporters less new exporters	
	(1)	(2)	(3)	(4)
Log employment, 1984	0.582	(21.52)	0.739	(22.12)
Log labor productivity, 1984	0.074	(4.72)	0.058	(2.98)
Annual labor productivity growth, 1988–1996	0.023	(8.83)	−0.001	(−0.42)
Annual labor productivity growth, 1984–1988	0.030	(6.91)	0.011	(2.02)
Exports/sales, 1996	0.197	(37.24)	0.148	(22.62)
Proportion of plants with				
$\Delta\tau > 0$	0.265	(23.00)	−0.030	(−2.08)
$\Delta\tau >$ median	0.192	(18.80)	−0.069	(−5.59)
Canadian tariff cut on				
Final goods ($\Delta\tau^{\text{Output}}$)	0.020	(14.99)	0.002	(1.06)
Intermediate inputs ($\Delta\tau^{\text{Input}}$)	0.021	(22.97)	0.001	(0.46)

Notes. This table reports differences in means across groups of plants after controlling for industry fixed effects. There are three groups of plants: (1) plants that did not export in either 1984 or 1996 (nonexporters, $N = 3{,}114$); (2) plants that did not export in 1984, but exported in 1996 (new exporters, $N = 2{,}119$); and (3) plants that exported in both 1984 and 1996 (old exporters, $N = 1{,}607$). Each row reports the results of a regression which pools all three groups. The dependent variable is listed in the leftmost column. The independent variables are four-digit SIC fixed effects and three dummy variables indicating whether the plant is a nonexporter, new exporter, or old exporter. Column (1) reports the difference between the new exporter and nonexporter dummies. Column (3) reports the difference between the old exporter and new exporter dummies. t-statistics appear in parentheses.

The U.S. tariff cut instruments are highly correlated with exporting. First, 85% of new exporters received a tariff cut, as compared to only 56% of nonexporters. After controlling for industry fixed effects, there is a 26.5–percentage point difference ($t = 23.00$). See Table I. Second, 64% of new exporters received an above-median tariff cut, as compared to only 40% of nonexporters. After controlling for industry fixed effects, there is a 19.2-percentage point difference ($t = 18.80$). The tariff cuts mattered for exporting.[7,8]

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[7] The “Canadian tariff cut” rows of Table I are explained in Sections IV.F and IV.G. Additional sample statistics, including data on *levels* for nonexporters, new exporters and old exporters, appear in Online Appendix Table B.2.

[8] The 1984 survey was administered to plants that accounted for a remarkable 91% of total manufacturing output. Nevertheless, there are two types of selection issues, both of which are dealt with in an earlier version of this paper (Lileeva and Trefler 2007). These earlier results are reviewed briefly in Online Appendix IV and very briefly here. First, there are plants that start exporting and then stop. This is much less common in our data than one might surmise from Eaton et al. (2008, Table 8.6) because we are looking at long-lived plants, that is, plants that were alive at least from 1984 to 1996. Second, there are plant deaths. Modeling

TABLE II

LABOR PRODUCTIVITY GROWTH, 1988–1996: NEW EXPORTERS LESS NONEXPORTERS

	Labor productivity quartiles in 1988			
	(1)	(2)	(3)	(4)
New exporters less nonexporters	0.053 (9.34)	0.036 (8.21)	0.030 (6.93)	0.027 (5.45)
Average number of employees per plant	53	74	95	93

Notes. The first row of this table reports the difference between new exporters and nonexporters in average annual labor productivity growth ($\Delta\varphi$). t-statistics are in parentheses. The first (last) column deals with plants in the lowest (highest) labor–productivity quartile. The sample consists of 1984 nonexporters.

IV. STARTING TO EXPORT AND LABOR PRODUCTIVITY GROWTH

In Sections IV–VI we study our sample of 1984 nonexporters. In Section VII we return to the old exporters sample.

IV.A. *Preliminary Results on Heterogeneous Labor Productivity Responses*

Let LP_t be value added per worker (labor productivity) in year t and let

$$\Delta\varphi \equiv \ln(\mathrm{LP}_{1996}/\mathrm{LP}_{1988})/8$$

denote the average annual log change in labor productivity. This is our dependent variable throughout this section. We begin with simple data displays that point to the presence of positive, heterogeneous labor productivity responses to improved U.S. market access. To examine heterogeneity, we assign plants to one of four bins based on quartiles of the distribution of initial labor productivity (LP_{1988}). To control for industrial structure, we calculate within-industry quartiles. This ensures that each quartile contains plants from all industries.[9]

Table II reports the mean of $\Delta\varphi$ for new exporters less the mean of $\Delta\varphi$ for nonexporters. The first column reports this difference for plants that were in the first or lowest quartile of the labor-productivity distribution. Within this quartile, labor productivity

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deaths increases our estimates of the impacts of improved foreign market access because dying plants tend to be nonexporters with negative productivity growth. See Online Appendix IV and especially Lileeva and Trefler (2007) for details.

[9] Specifically, choose a four-digit SIC industry and, for plants in the industry, calculate the quartiles of the distribution of 1988 labor productivity. Then assign each plant a number from 1 to 4 based on which quartile it is in. Repeat this for each industry so that all plants are assigned to quartiles. This is what we mean by “within-industry” quartiles.

Employment quartiles in 1988	Labor productivity quartiles in 1988: 1	2	3	4
1	**0.061**	**0.052**	**0.050**	**0.041**
2	**0.045**	**0.039**	**0.030**	0.023
3	**0.057**	**0.033**	**0.021**	0.013
4	0.022	0.015	0.008	0.005

FIGURE IV

Labor Productivity Growth, 1988–1996: New Exporters Less Nonexporters by Productivity and Size

As in Table II, each cell is the difference between new exporters and nonexporters in average annual labor productivity growth ($\Delta\varphi$). Boldface indicates statistically significant differences at the 1% level. t-statistics appear in Online Appendix Table B.3. The sample consists of 1984 nonexporters.

growth was 0.053 log points higher for new exporters than for nonexporters ($t = 9.34$). As one moves to higher quartiles, the difference shrinks monotonically to 0.027 log points. The monotonicity is consistent with the downward-sloping new exporters effect in Figure II.

In the Melitz model and, by extension, in our model, productivity is perfectly correlated with size. Productivity matters because productive firms are large and large firms can profitably incur the fixed costs of exporting. Restated, the economic insight about exporting is as much about size as it is about productivity. Yet the largest plants in Canada are not always the most productive plants. The bottom row of Table II shows that the average number of employees per plant is about the same in both the third and fourth quartiles. To control for size, we additionally assign each plant to one of four bins based on within-industry quartiles of the distribution of initial (1988) employment size. We then form the 16 bins that appear in Figure IV. Each cell in the figure reports the difference between new exporters and nonexporters in average annual labor productivity growth ($\Delta\varphi$).

The dominant feature of Figure IV is the large differences to the upper left (small, less productive plants) and the small differences to the lower right (large, more productive plants). Boldface indicates statistically significant differences at the 1% level, so that the differences to the lower right are not significant. It is remarkable that whether moving across any row, down any column, or down along any diagonal there are almost always declining productivity growth differences between new exporters and nonexporters. We thus conclude that heterogeneity is pervasive.[10]

In what follows we will be conducting our econometric analysis as nonparametrically as possible, that is, separately for different bins. From the theory it is natural to divide up the sample into four bins based on quartiles of the productivity distribution, as in Table II. As we will show, this yields very nice results. However, Figure IV shows that such a procedure leads one to overstate the labor productivity benefits for the most productive plants: these benefits are statistically significant in Table II ($t = 5.45$), but insignificant in Figure IV. It is therefore of interest to create bins based on both initial productivity and size.

In our subsequent econometric work we will be estimating everything separately within bins. Because we only have 5,233 plants, we consolidate the 16 bins in Figure IV into five bins, giving us approximately 1,000 observations per bin. The five bins are demarcated by the diagonal bands in the Figure IV matrix. For example, the first bin consists of elements (1, 1), (1, 2), and (2, 1), whereas the second bin consists of elements (3, 1), (2, 2), and (1, 3). *This will be our baseline definition of bins*. A parametric alternative is to estimate a probit of export status on initial productivity and size, use this to create a propensity score for each plant, and then divide the sample into five groups based on propensity scores. Yet another alternative is to use the four productivity-based bins of Table II. As we shall see, the way in which bins are defined will not matter beyond what is already evident from Table II and Figure IV.

IV.B. More Preliminary Results on Heterogeneous Labor Productivity Responses

The analysis of Table II and Figure IV is suggestive but not rigorous. In this section we provide a textbook IV approach before turning in the next section to the Angrist and Imbens (1995) approach. Our dependent variable is labor productivity growth $\Delta\varphi$. Letting EXP_{1996} be plant exports in 1996, we define our "treatment" T as $T = \ln \mathrm{EXP}_{1996}$ for new exporters and $T = 0$ for nonexporters. Below, we also consider alternative definitions of the treatment, but this does not alter our results. We instrument for

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[10] t-statistics appear in Online Appendix Table B.3. Table B.3 also repeats the Figure IV exercise but using OLS with four-digit SIC industry fixed effects. The results are very similar.

T using $\Delta\tau_2$, $\Delta\tau_3$, and $\Delta\tau_4$, defined in Section III.B. Our controls, denoted by the vector X, are (1) log employment in 1984, (2) log labor productivity in 1984, and (3) the average annual log change in labor productivity during 1984–1988. With this notation in hand, we estimate the following equations:

$$\Delta\varphi = \beta T + \gamma X + \varepsilon \tag{5}$$

$$T = \Sigma_{i=2}^{4}\delta_i \Delta\tau_i + \gamma' X + \eta. \tag{6}$$

The estimates appear in Table III. This table is important. The bottom panel presents OLS estimates of the first-stage equation (6). The first row of the panel reports estimates for the pooled sample. The remaining rows report estimates separately for each of the five bins defined by the bands in Figure IV. From columns (10)–(15) of the bottom panel, the tariff-cut instruments are all statistically significant and have the expected positive signs.

The top and middle panels of Table III present OLS and IV estimates of equation (5), respectively. Column (2) presents the coefficient on T, our focus variable. In the pooled sample the coefficient is positive and statistically significant both for OLS ($t = 13.01$) and for IV ($t = 15.92$). The effect of starting to export not only is positive, it also is heterogeneous. As one moves from bin 1 (small, less productive plants) to bin 5 (large, more productive plants), the estimated effects decline. *This is exactly as predicted by the model.* The hypothesis that the coefficients are equal across bins is rejected both for OLS ($F = 14.48$, $p < .001$) and for IV ($F = 12.26$, $p < .001$). Further, Online Appendix Table B.4 shows that bin 5 is statistically different from bins 1–3 and that bin 4 is statistically different from bins 1 and 2.

Before moving on, we briefly demonstrate the insensitivity of these results to the way in which bins are defined. At the end of Section IV.A, we described a parametric way of constructing the five bins. Pool the 1984 nonexporter and old exporter samples and estimate a probit for 1996 export status. The regressors are 1988 log labor productivity, 1988 log employment, and four-digit SIC industry fixed effects.[11] Next, create an estimated probability of exporting in 1996 for each plant and divide the full sample into quintiles based on this probability. This assigns each plant to one of five bins. Using this new bin definition, reestimate equations (5)

[11] The labor productivity coefficient (0.41) and the employment size coefficient (0.59) are both statistically significant. See Online Appendix Table B.5.

TABLE III
Labor Productivity Growth, 1988–1996: Standard IV Estimation

Bin	Treatment: $T = \ln(\text{EXP}_{1996})$		Labor productivity		Labor prod. growth 1984–1988		Employment size		Tariff cut instruments						Alternative bins $T = \ln(\text{EXP}_{1996})$	
									$\Delta\tau_2$		$\Delta\tau_3$		$\Delta\tau_4$			
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(2′)	(3′)
OLS, dependent variable: Productivity growth																
All	0.0022	(13.01)	−0.04	(−22.58)	−0.28	(−35.47)	0.01	(5.55)								
1	0.0042	(8.12)	−0.05	(−7.46)	−0.25	(−10.27)	0.01	(2.04)							0.0040	(8.94)
2	0.0041	(9.64)	−0.05	(−9.40)	−0.33	(−15.74)	0.00	(−0.58)							0.0036	(9.45)
3	0.0027	(6.67)	−0.05	(−10.44)	−0.32	(−18.03)	0.00	(1.39)							0.0023	(6.73)
4	0.0013	(3.48)	−0.04	(−8.45)	−0.25	(−12.45)	0.01	(3.36)							0.0008	(2.42)
5	0.0008	(2.79)	−0.03	(−7.19)	−0.20	(−9.16)	0.01	(3.64)							0.0002	(0.43)
Second-stage IV, dependent variable: Productivity growth																
All	0.010	(15.92)	−0.05	(−25.33)	−0.31	(−37.77)	−0.01	(−6.01)								
1	0.017	(9.87)	−0.05	(−7.88)	−0.24	(−10.06)	0.00	(−0.70)							0.016	(10.11)
2	0.015	(10.30)	−0.06	(−10.52)	−0.32	(−15.35)	−0.02	(−4.98)							0.014	(10.07)
3	0.012	(7.72)	−0.05	(−10.45)	−0.30	(−16.99)	−0.01	(−3.43)							0.010	(7.68)
4	0.008	(4.57)	−0.03	(−7.14)	−0.24	(−11.90)	0.00	(−0.19)							0.006	(4.00)
5	0.003	(2.44)	−0.03	(−7.25)	−0.20	(−9.28)	0.01	(2.79)							0.001	(0.85)
First-Stage IV, dependent variable: $T = 0, \ln(\text{EXP}_{1996})$																
All			1.68	(20.48)	2.50	(4.03)	0.78	(5.03)	3.1	(12.04)	4.4	(18.82)	3.1	(13.01)		
1			0.13	(0.35)	−0.50	(−0.34)	0.50	(2.61)	1.8	(4.21)	3.1	(7.64)	2.9	(7.33)		
2			0.75	(1.80)	−0.70	(−0.46)	1.36	(5.99)	3.4	(6.10)	4.1	(8.33)	2.9	(5.92)		
3			−0.04	(−0.13)	−1.39	(−1.12)	1.63	(7.94)	3.4	(6.02)	4.1	(8.19)	3.2	(6.47)		
4			−0.45	(−1.14)	−0.37	(−0.21)	1.36	(6.13)	2.2	(3.37)	4.0	(6.54)	2.5	(4.05)		
5			0.33	(0.76)	1.75	(0.79)	0.64	(2.86)	3.2	(4.93)	4.6	(7.37)	2.4	(3.67)		

Notes. The top panel of this table reports OLS estimates of the productivity-growth equation (5). The dependent variable $\Delta\varphi$ is the average annual log change in labor productivity over the period 1988–1996. The key regressor is the treatment variable T: $T = \ln(\text{EXP}_{1996})$ for new exporters and $T = 0$ for nonexporters. The middle panel reports the IV estimates of the productivity-growth-equation (5). The bottom panel reports estimates of the first-stage equation (6) (dependent variable is T). The first row of each panel reports the results for the full sample $N = 5{,}233$. Each of the following five rows reports results by bin, where bins are defined in Figure IV. Bins sample sizes are $N_1 = 980$, $N_2 = 953$, $N_3 = 1{,}208$, $N_4 = 1{,}020$, and $N_5 = 1{,}072$. Column (2′) reports the coefficient on the treatment T for the case where the bins are redefined using the probit–based method. The sample consists of 1984 nonexporters. t-statistics are in parentheses.

and (6) by bin for those plants in the 1984 nonexporters sample.[12] The estimates of the treatment coefficient β when probit-based bins are used appear in column (2′) of Table III. As is apparent, the two bin definitions yield almost identical results.[13]

One can also define bins either in terms of quartiles of the 1988 labor productivity distribution or in terms of quartiles of the 1988 employment size distribution. Estimates of equations (5) and (6) for these two bin definitions appear in Online Appendix Tables B.6 and B.7. They display the same pattern of heterogeneity as in Table III.

Turning to other specifications issues, adding four-digit industry fixed effects makes no difference to the results. See Online Appendix Table B.8. In all subsequent sections of this paper, these fixed effects will be included in the analysis. In addition, the three covariates X play little role. Omitting them makes no difference to the results, as shown in Online Appendix Table B.9. This completes the discussion of the sensitivity of the Table III results.

Table IV assesses the size of the Table III coefficients. This is another important table. Column (1) of Table IV reports ΔT, the change in exporting induced by the U.S. tariff cuts (from equation (6)). The mean value of T is 13.7, so the ΔT of around two is reasonable. Column (2) reports the impact on labor productivity of the change in exporting induced by the U.S. tariff cuts. It is $\widehat{\beta} \times \Delta T \times 8$, where (a) $\widehat{\beta}$ is the IV coefficient on T from column (2) of Table III, (b) ΔT is the induced change in exporting from column (1) of Table IV, and (c) 8 converts annual labor productivity changes to a 1988–1996 change. *The numbers in column (2) of Table IV are one of our most important results*. To summarize them, the last row of column (2) reports their employment-weighted average

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[12] An alternative is to estimate the probit only for 1984 nonexporters rather than for the full sample (1984 nonexporters plus old exporters). However, as pointed out by a referee, using only 1984 nonexporters can lead to thorny selection issues. As an empirical matter, it does not matter whether the probit is estimated with the full sample or only with the 1984 nonexporters sample.

[13] The slight difference in results is attributable to differences in bin sample sizes. The sample sizes in the probit-based bins (column (2′)) are 1,198, 1,221, 1,080, 957, and 777. Thus relative to the Figure IV–based bins, the probit-based bins have more plants in the low bins (e.g., bin 1) and fewer plants in the high bins (e.g., bin 5). This puts more low-response plants in the low bins and fewer high-response plants in the high bins. Restated, the probit-based results will have smaller estimates of β in all bins. That the sample sizes are smaller for the higher probit-based bins is a combined consequence of two facts: (a) the probit was estimated using the full sample and (b) the highest bins have many old exporters and few 1984 nonexporters.

TABLE IV
SPECIFICATION TESTS AND COEFFICIENT MAGNITUDES

	Coefficient magnitudes			Hausman test		Over-id test		First stage F-tests			
	ΔT	$\beta \times \Delta T \times 8$	Emp. wgt.					3 tariffs		All variables	
Bin	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
1	1.43	0.196	0.06	34.62	.00	3.56	.03	25.50	.00	18.28	.00
2	2.20	0.264	0.10	34.43	.00	1.53	.22	29.10	.00	32.68	.00
3	2.71	0.267	0.18	25.36	.00	0.05	.95	26.20	.00	43.88	.00
4	2.25	0.146	0.25	11.22	.02	1.40	.25	14.35	.00	22.88	.00
5	2.77	0.071	0.40	3.26	.52	2.47	.09	18.89	.00	18.46	.00
Total		0.153									

Notes. Column (1) is the estimated impact of the U.S. tariff cut on the treatment T, that is, on exporting. It is based on the estimates of equation (6) reported in the bottom panel of Table III. Column (2) is the impact of exporting on labor productivity for those plants that were induced to export as a result of the U.S. tariff cuts. The "Total" row reports the average of column (2), averaged using the column (3) employment weights. The table also provides specification tests for the IV specification that appears in the middle panel of Table III. Column (4) reports the Hausman χ^2_4 statistic for the difference between the OLS and IV estimates of four parameters (coefficients on T and three covariates). Column (6) reports the Basmann (1960) χ^2_3 statistic for overidentification. Column (8) reports the F^3_N statistic for the joint null that all three tariff coefficients in the first stage equal 0. (Bin sample sizes N are given in the notes to Table III.) Column (10) reports the F^6_N statistics for the joint null that all six coefficients in the first stage equal 0. p-values appear in columns (5), (7), (9), and (11).

across all five bins. We average using 1996 employment weights. These weights appear in column (3).[14]

The average effect is 0.153, which indicates that improved access to the U.S. market raised the productivity of 1984 nonexporters by 0.153 log points. More precisely, labor productivity rose on average by 0.153 log points for those plants that were induced to export as a result of improved access to the U.S. market. The result 0.153 log points strikes us as a very large number. Because these plants accounted for 23% of manufacturing employment in 1996, improved market access raised manufacturing productivity by 0.035 log points ($= 0.153 \times 0.23$). This 3.5% effect is one of our paper's major take-home results.

Columns (4)–(11) of Table IV report some standard specification tests. Columns (5), (7), (9), and (11) are p-values. The Hausman (1978) tests imply that the OLS and IV coefficients are different for bins 1–3, but not for bins 4 and 5. Column (6) reports overidentification tests. All the p-values exceed .01, which supports the exogeneity of the tariff instruments. The F-tests in

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[14] If this paragraph is unclear, see Online Appendix V for a lengthier description.

column (8) reject the null that the three tariffs have 0 coefficients in the first stage. The F-tests in column (10) reject the null that all the first-stage coefficients are 0. The large size of these F-statistics is important in light of research on weak instruments, such as Staiger and Stock (1997).

IV.C. Clarifications

At this point in seminars two questions are often posed. First, can't the standard Melitz model explain our finding? Specifically, is it not possible that the tariff cuts shifted the Melitz cutoff so far to the left that even plants in bins 1 and 2 moved above the cutoff? The answer is no! By defining bins in terms of quartiles of labor productivity and size we have ensured that plants in bins 1 and 2 are the very smallest and least productive. The Melitz model simply cannot explain why so many of these plants started exporting.[15]

The second question we receive is about our choice of focus. Why do we focus on the within-plant labor productivity gains from the FTA when there are other sources of labor productivity gains? The answer is that we have documented these other sources elsewhere. Putting together the available literature, three sources of productivity gain emerge as being important empirically. First, Trefler (2004) and Lileeva (2008) showed that the fall in the Canadian tariff against U.S. plants resulted in a substantial amount of contraction and exit of import-competing plants. Because these are the least productive plants, their contraction and exit raised average productivity. In particular, it raised average manufacturing productivity by 4.3%. Second, exporters are more productive than nonexporters, so that when Canadian exporters grow by exporting to the United States, average productivity rises. A simple share-shift analysis (e.g., Bernard and Jensen [2004]) shows that this raised average manufacturing productivity by 4.1%. These two effects involve raising average productivity by increasing the market share of more productive plants. The third effect, which is the 3.5% effect we just documented, operates via within-plant changes in productivity. Below we will find two additional within-plant effects. The reduction in Canadian tariffs on intermediate inputs purchased

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[15] The percentages of plants that started exporting in each of bins $1, \ldots, 5$ are 16%, 31%, 42%, 50%, and 61%, respectively. Thus even in bin 1, 16% of plants started exporting.

from the United States (a form of improved foreign market access) raised productivity by a further 0.5%. See Section IV.F. In addition, we will also find a 1.4% effect on within-plant labor productivity for old exporters. See Section VII. The sum of these effects is 13.8%. The idea that a single government policy could raise labor productivity by so much is indeed remarkable.[16]

IV.D. Review of Econometrics with Unobserved Heterogeneous Responses

Given that we have documented heterogeneous responses based on observables, it seems likely that there are heterogenous responses based on unobservables. If so, then one cannot interpret the estimated coefficients on exporting in the way we have been doing (e.g., Imbens and Angrist [1994]). To understand this, suppose there is a binary instrument (plants either receive a tariff cut or not) and let ΔT be the impact of the tariff cut on exports. Let $\Delta\varphi^c$ be the causal effect of exporting on productivity growth: $\Delta\varphi^c \equiv (\beta + U)\Delta T$, where β is the same for all plants and U is the plant-specific or heterogeneous causal response. Also assume that there are no covariates. Then IV consistently estimates

$$\text{LATE} = \beta + \frac{\mathbf{E}\left[U \cdot \Delta T\right]}{\mathbf{E}\left[\Delta T\right]}. \tag{7}$$

See Card (2001, pp. 1141–1142) for a simple exposition of this point. This has two implications. First, LATE is a weighted average of the heterogeneous responses $\beta + U$, where the weights are ΔT. Second, in our empirical context, we expect LATE $> \beta$ because firms that expect large gains from exporting and investing (large U) will also be firms that are likely to switch their behavior as a result of improved market access (large ΔT).

The interpretation of the IV estimator of β as a weighted average of the heterogeneous responses no longer holds when there are covariates and/or continuous instruments. Indeed, no simple interpretation holds. To restore the weighted-average interpretation, Angrist and Imbens (1995, Theorem 3) recommend the following.

1. Convert the instrument into a set of mutually orthogonal binary variables, as we have already done.

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[16] Because this is a long paper, a reader who is not interested in econometric details may want to jump straight to Section V. Table VII and Panel A of Table IX are the other crucial results of the paper.

2. Convert the covariates into binary variables. Recall that we have three covariates (1984 labor productivity, 1984 employment size, 1984–1988 labor productivity growth). Let $k = 1, 2, 3$ index covariates and to fix ideas, consider just one of these covariates (1984 labor productivity). We calculate the quartiles of the 1984 distribution of labor productivity and then create four dummy variables, one for each quartile.[17] Because the dummies sum to one, we drop one dummy and denote the remaining ones by X_{kq}, where $q = 2, 3, 4$, indexes the quartile. This procedure creates three dummy variables for each of three covariates, for a total of nine covariates.
3. Expand the instrument set to include instrument–covariate interactions. That is, interact the nine covariates with the three tariffs to create an additional twenty-seven instruments $X_{kq}\Delta\tau_{q'}$. This is a large instrument set. One of our aims in showing Table III with its small number of instruments (three instruments) was to show that our results will not be sensitive to having many instruments.

With these changes in place, we estimate the following Angrist–Imbens equations:

$$\Delta\varphi = \beta T + \Sigma_{k=1}^{3}\Sigma_{q=2}^{4}\gamma_{kq}X_{kq} + \theta_{\text{SIC}} + \varepsilon, \tag{8}$$

where θ_{SIC} is an industry fixed effect and the treatment T is instrumented by the first-stage equation

$$\begin{aligned} T =\ & \Sigma_{q=2}^{4}\delta_q\Delta\tau_q + \Sigma_{k=1}^{3}\Sigma_{q=2}^{4}\gamma'_{kq}X_{kq} \\ & + \Sigma_{k=1}^{3}\Sigma_{q,q'=2}^{4}\lambda_{kqq'}X_{kq}\Delta\tau_{q'} + \theta'_{\text{SIC}} + \eta. \end{aligned} \tag{9}$$

Note that equations (8) and (9) will be estimated separately by bin; however, for notational simplicity we have suppressed bin subscripts.[18]

IV.E. Angrist–Imbens Results

Table V presents the estimates of the Angrist–Imbens equations (8) and (9). Four-digit SIC industry fixed effects are used. Our baseline specification appears in Panel A of Table V, so we review it in some detail. Estimates of the treatment effect β in

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[17] These are within-industry quartiles, as described in footnote 9.

[18] Finally, the Angrist–Imbens method requires us to round $\ln \text{EXP}_{1996}$ to the nearest integer, a procedure that makes no difference to our results.

TABLE V
LABOR PRODUCTIVITY GROWTH 1988–1996, ANGRIST–IMBENS IV ESTIMATOR

	IV						OLS			Alternative bins	
	β	t	ΔT	$\beta \times \Delta T \times 8$	Overidentification		β	t	R^2	β	t
Bin	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(1′)	(2′)
	A. Baseline: Five bins based on productivity and size; $T = 0$, $\ln(\text{EXP}_{1996})$										
1	0.012	(7.34)	1.58	0.147	1.20	(.22)	0.0041	(7.32)	.19	0.013	(9.65)
2	0.010	(7.77)	2.85	0.237	1.53	(.04)	0.0041	(8.29)	.20	0.012	(9.45)
3	0.009	(6.21)	3.46	0.241	1.28	(.15)	0.0031	(6.71)	.25	0.009	(8.18)
4	0.005	(2.54)	2.02	0.085	0.80	(.77)	0.0011	(2.76)	.20	0.002	(1.19)
5	0.002	(1.01)	1.51	0.022	1.13	(.29)	0.0007	(1.97)	.16	0.002	(1.65)
Total				0.107							
	B. Five bins based on productivity and size; binary treatment ($T = 0, 1$)										
1	0.154	(7.34)	0.12	0.149	1.18	(.24)	0.0519	(6.99)	.18	0.183	(9.76)
2	0.140	(7.71)	0.21	0.234	1.54	(.04)	0.0534	(7.86)	.20	0.159	(9.43)
3	0.117	(6.06)	0.24	0.228	1.32	(.12)	0.0388	(6.08)	.24	0.129	(8.14)
4	0.084	(2.93)	0.14	0.091	0.70	(.88)	0.0144	(2.46)	.20	0.021	(1.21)
5	0.021	(0.88)	0.11	0.018	1.15	(.27)	0.0081	(1.59)	.16	0.032	(1.47)
Total				0.105							

TABLE V

CONTINUED

	IV						OLS			Alternative bins	
	β	t	ΔT	$\beta \times \Delta T \times 8$	Overidentification		β	t	R^2	β	t
Bin	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(1′)	(2′)
C. Four bins based on productivity; $T = 0$, $\ln(\text{EXP}_{1996})$											
1	0.010	(8.27)	2.88	0.240	1.67	(.02)	0.0036	(7.86)	.20	0.011	(10.48)
2	0.009	(7.66)	3.34	0.251	1.71	(.01)	0.0028	(8.10)	.13	0.009	(9.13)
3	0.006	(4.65)	3.55	0.162	2.42	(.00)	0.0021	(6.22)	.13	0.005	(4.33)
4	0.005	(3.38)	1.58	0.067	1.58	(.03)	0.0016	(4.04)	.14	0.005	(3.88)
Total				0.168							
D. Baseline, but without the twenty-seven covariate–tariff interaction instruments											
1	0.015	(7.91)	1.50	0.183	1.20	(.31)	0.0041	(7.32)	.19	0.016	(9.71)
2	0.013	(8.72)	2.65	0.282	1.53	(.21)	0.0041	(8.29)	.20	0.014	(10.47)
3	0.012	(7.32)	3.42	0.317	1.28	(.28)	0.0031	(6.71)	.25	0.012	(9.08)
4	0.011	(2.95)	1.55	0.139	0.80	(.49)	0.0011	(2.76)	.20	0.007	(4.02)
5	0.004	(1.15)	1.47	0.042	1.13	(.34)	0.0007	(1.97)	.16	0.003	(1.82)
Total				0.150							

Notes. This table reports the results of estimating the Angrist–Imbens equations (8) and (9). Column (1) is the IV estimate of the treatment effect (β). Column (3) is the estimated change in exporting induced by the U.S. tariff cuts. Column (4) is the effect of exporting on labor productivity for those Canadian plants that were induced to export by the U.S. tariff cuts. Column (5) is the overidentification test (a χ^2 statistic) and column (6) is its p-value. A p-value below .01 indicates that the exclusion restriction is rejected. Column (7) is the OLS estimate of the treatment effect (β) and column (8) is its t-statistic. "Total" rows are the employment-weighted averages of the $\beta \times \Delta T \times 8$ across bins. The remaining estimated coefficients in equations (8) and (9) appear in Online Appendix Tables B.10–B.13. Column (1′) reports the IV coefficient on the treatment T for the case where the bins are redefined using the probit-based method. Sample sizes for Panels A, B, and D are the same as in the notes to Table III. Sample sizes for the four bins in Panel C are between 1,228 and 1,394. The sample consists of 1984 nonexporters.

equation (8) appear in column (1) for IV and column (7) for OLS. All the remaining coefficients in equations (8) and (9) appear in Online Appendix Table B.10. The baseline specification is estimated separately for each of our five bins (from Figure IV). The main finding is that the estimates of β are positive and decline monotonically as one moves from bin 1 (less productive, small plants) to bin 5 (more productive, large plants). This is exactly what we saw in Table III. An F-test rejects the hypothesis that all five coefficients are equal ($p < .001$ for both OLS and IV). Further, as shown in Online Appendix Table B.4, the coefficient for bin 5 is statistically larger than those for bins 1–3 and the coefficient for bin 4 is statistically larger than those for bins 1 and 2. Hence there is coefficient heterogeneity. (Coefficient homogeneity is rejected for all the specifications in Table V.) The impact of tariffs on exporting appears in column (3), which gives the average predicted change in $T = \ln(\mathrm{EXP}_{1996})$ for plants in each bin. Column (4) reports the impact on labor productivity growth over our eight-year period for those plants that were induced to export as a result of the U.S. tariff cuts. The "Total" row reports the employment-weighted average of these impacts.

Panel B of Table V alters the specification using a binary definition of the treatment: $T = 0$ for nonexporters and $T = 1$ for new exporters. Our main findings are unaltered: the estimates of β are positive and decline monotonically. Although we cannot compare coefficient magnitudes across different definitions of the treatment, we can compare magnitudes of the treatment effects. These appear in column (4) and are virtually identical in Panels A and B of the table. We conclude from this that the exact definition of the treatment does not matter.

In Panel C of Table V, we define bins as in Table II, that is, based on quartiles of the 1988 labor productivity distribution. There are now only four bins. Once again the estimates of β are positive and decline monotonically.[19]

One surprise is that our baseline total treatment effect of 0.107 is so much smaller than our 0.153 estimate from Table IV. To investigate the source of this difference, we modified the Angrist–Imbens first-stage equation (9) by dropping the twenty-seven instruments that come from interacting the nine covariate dummies with the three tariff-cut dummies. The results appear in Panel D

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[19] As expected from a comparison of Table II with Figure IV, the total effect is larger using these four bins.

of Table V. This leads to larger estimates of β and to a total treatment effect of 0.150. It appears that including these interactions leads to smaller treatment effects.[20]

IV.F. Improved Access to U.S. Intermediate Inputs

The FTA mandated reductions in U.S. tariffs against products sold by Canadian plants in the U.S. market. It also mandated reductions in *Canadian* tariffs against intermediate inputs and capital equipment purchased by Canadian plants from the United States. We denote these Canadian tariff cuts by $\Delta\tau^{\text{Input}}$. In terms of our model, these cuts lower F^I, the cost of investing in productivity. A fall in F^I leads to a parallel shift down of the lines in Figures I and II. $\Delta\tau^{\text{Input}}$ thus has effects similar to those of $\Delta\tau$. Restated, $\Delta\tau^{\text{Input}}$ also captures improved foreign market access, in the sense of improved access by Canadian plants to U.S. intermediate inputs.

To investigate, we construct $\Delta\tau^{\text{Input}}$ as follows. We have HS6 Canadian imports from the United States and the corresponding duties collected. This allows us to calculate the Canadian tariff reductions against the United States by HS6 code. For each plant we also have its HS6 purchases of intermediate inputs. We can therefore match up each intermediate input purchase with a Canadian tariff cut. As with the plant-level U.S. tariff cuts, we aggregate these intermediate input tariff cuts up to the plant level using the unweighted average of Canadian tariff cuts on intermediate inputs purchased by the plant. Table I shows that $\Delta\tau^{\text{Input}}$ was 2.1 percentage points larger for new exporters than for nonexporters ($t = 22.97$), suggesting that $\Delta\tau^{\text{Input}}$ mattered for exporting.

Turning to IV estimation, we begin by introducing $\Delta\tau^{\text{Input}}$ as an additional regressor in the first stage (equation (9)). Consider Panel B of results in Table VI. Column (1) reports the estimated coefficients on $\Delta\tau^{\text{Input}}$. (Recall that the first-stage dependent variable is T. None of the other first-stage coefficients are reported in the table.) The coefficients on $\Delta\tau^{\text{Input}}$ are all positive and most are statistically significant, indicating that Canadian export decisions were correlated with access to U.S. intermediate inputs.

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[20] To see whether we could drive the treatment effect to zero by adding more covariate–$\Delta\tau_q$ interactions, we considered a much larger set of covariates. This made no difference. As should be apparent from Table V, we have chosen as our baseline a specification that yields the smallest estimates of the impact of improved market access.

TABLE VI
LABOR PRODUCTIVITY GROWTH, 1988–1996: CANADIAN TARIFF CUTS

	First stage				Second stage						Alternative second stage				Tests	
	$\Delta\tau^{\text{Input}}$		$\Delta\tau^{\text{Output}}$, $\Delta\ln M$		$T=\ln(\text{EXP}_{1996})$		$\Delta\tau^{\text{Input}}$		$\Delta\tau^{\text{Output}}$, $\Delta\ln M$		$T=\ln(\text{EXP}_{1996})$					
	Coeff.	t	Coeff.	t	β	t	Coeff.	t	Coeff.	t	β	t	ΔT	$\beta\times\Delta T\times 8$	Hausman	Overidentification
Bin	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
					A. Baseline (from Table V)								A. Baseline (from Table V)			
1					0.012	(7.34)					0.012	(7.34)	1.58	0.147	20.6	1.20
2					0.010	(7.77)					0.010	(7.77)	2.85	0.237	20.9	1.53
3					0.009	(6.21)					0.009	(6.21)	3.46	0.241	15.6	1.28
4					0.005	(2.54)					0.005	(2.54)	2.02	0.085	3.7	0.80
5					0.002	(1.01)					0.002	(1.01)	1.51	0.022	0.4	1.13
Total														0.107		
					B. $\Delta\tau^{\text{Input}}$ in both stages								B. $\Delta\tau^{\text{Input}}$ as an instrument			
1	39.5	(7.09)			0.008	(3.22)	0.30	(2.03)			0.012	(8.82)	1.96	0.180	30.3	1.16
2	26.4	(4.28)			0.010	(5.65)	0.00	(−0.01)			0.010	(8.22)	3.24	0.267	23.3	1.50
3	36.5	(6.35)			0.007	(3.85)	0.12	(1.01)			0.008	(6.70)	4.04	0.262	17.8	1.28
4	22.9	(2.86)			0.004	(1.96)	0.15	(1.41)			0.006	(3.27)	2.71	0.132	6.2	0.76
5	1.9	(0.25)			0.002	(0.93)	0.05	(0.61)			0.002	(1.02)	1.59	0.023	0.4	1.10
Total														0.128		

TABLE VI
CONTINUED

	First stage				Second stage						Alternative second stage				Tests	
	$\Delta\tau^{\text{Input}}$		$\Delta\tau^{\text{Output}}$, $\Delta \ln M$		$T = \ln(\text{EXP}_{1996})$		$\Delta\tau^{\text{Input}}$		$\Delta\tau^{\text{Output}}$, $\Delta \ln M$		$T = \ln(\text{EXP}_{1996})$					
	Coeff.	t	Coeff.	t	β	t	Coeff.	t	Coeff.	t	β	t	ΔT	$\beta \times \Delta T \times 8$	Hausman	Overidentification
Bin	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
	C. $\Delta\tau^{\text{Input}}$ and $\Delta\tau^{\text{Output}}$ in both stages										C. $\Delta\tau^{\text{Input}}$ and $\Delta\tau^{\text{Output}}$ as instruments					
1	43.3	(7.60)	−11.54	(−2.87)	0.007	(3.25)	0.22	(1.48)	0.15	(2.52)	0.011	(8.67)	1.93	0.172	29.0	1.21
2	27.2	(4.33)	−3.30	(−0.68)	0.009	(4.56)	−0.02	(−0.17)	0.18	(2.72)	0.010	(8.12)	3.21	0.259	22.6	1.56
3	37.2	(6.25)	−2.17	(−0.45)	0.006	(3.39)	0.07	(0.62)	0.15	(2.19)	0.008	(6.66)	4.01	0.257	17.5	1.29
4	23.0	(2.84)	−0.68	(−0.11)	0.004	(1.92)	0.15	(1.38)	0.01	(0.17)	0.006	(3.27)	2.70	0.132	6.3	0.73
5	−0.4	(−0.05)	3.80	(0.78)	0.001	(0.79)	0.02	(0.21)	0.05	(0.98)	0.002	(1.09)	1.71	0.026	0.5	1.07
Total														0.127		
	D. $\Delta\tau^{\text{Input}}$ and $\Delta \ln M$ in both stages										D. $\Delta\tau^{\text{Input}}$ and $\Delta \ln M$ as instruments					
1	39.7	(7.14)	−3.97	(−1.71)	0.008	(3.35)	0.29	(1.97)	−0.04	(−0.96)	0.012	(9.04)	1.59	0.148	31.6	1.13
2	26.5	(4.31)	−2.35	(−0.76)	0.010	(5.66)	0.00	(−0.01)	0.01	(0.21)	0.010	(8.22)	3.01	0.246	23.2	1.45
3	36.7	(6.36)	−1.15	(−0.42)	0.007	(3.79)	0.13	(1.07)	−0.03	(−0.72)	0.008	(6.72)	3.92	0.254	17.8	1.24
4	22.7	(2.82)	0.47	(0.12)	0.004	(1.93)	0.15	(1.37)	0.02	(0.42)	0.006	(3.28)	2.73	0.133	6.4	0.73
5	1.3	(0.17)	−2.64	(−0.75)	0.002	(0.95)	0.05	(0.70)	0.04	(1.13)	0.002	(0.92)	1.30	0.017	0.3	1.09
Total														0.121		

Notes. Each row of this table reports estimates of the second-stage equation (8) and the first-stage equation (9). The second-stage dependent variable is labor productivity growth $\Delta\varphi$. The first-stage dependent variable is the treatment $T = 0, \ln(\text{EXP}_{1996})$. Panel A is the baseline specification carried over from Table V. Panels B–D augment this specification with additional regressors. This is most clearly explained with reference to Panel C. It introduces two more regressors into the first stage. These are the plant-specific Canadian tariff cuts both on intermediate inputs purchased by the plant ($\Delta\tau^{\text{Input}}$) and on commodities produced by the plant ($\Delta\tau^{\text{Output}}$). Columns (1) and (3) report the estimated first-stage coefficients on $\Delta\tau^{\text{Input}}$ and $\Delta\tau^{\text{Output}}$, respectively. Columns (5)–(10) report second-stage coefficients when $\Delta\tau^{\text{Input}}$ and $\Delta\tau^{\text{Output}}$ are included in the second stage. Columns (11)–(14) report second-stage coefficients when $\Delta\tau^{\text{Input}}$ and $\Delta\tau^{\text{Output}}$ are omitted from the second stage. Columns (15) and (16) report Hausman and Basmann overidentification statistics, respectively, for the alternative second-stage specification (columns (11)–(14)). The Basmann statistics are tiny, indicating that the Canadian tariff cuts are exogenous. Turning to differences across panels, Panel B omits $\Delta\tau^{\text{Output}}$, but keeps $\Delta\tau^{\text{Input}}$. Panel D replaces $\Delta\tau^{\text{Output}}$ with $\Delta \ln M$, the growth in Canadian imports from the United States on commodities produced by the plant. The sample consists of 1984 nonexporters. t-statistics are in parentheses.

We next introduce $\Delta\tau^{\text{Input}}$ into the second stage (equation (8)). That is, $\Delta\tau^{\text{Input}}$ appears in both stages. The second-stage IV coefficients on $\Delta\tau^{\text{Input}}$ appear in column (7). They are never statistically significant. As a result, we prefer an alternative second stage that excludes $\Delta\tau^{\text{Input}}$. This appears in columns (11)–(14).[21] The two second-stage specifications yield very similar estimates of β (compare columns (5) and (11)). The exception is bin 1; however, this bin 1 sensitivity completely disappears when we use probit-based bins or simple IV (equations (5) and (6)). See Online Appendix Tables B.14 and B.21. This and the fact that $\Delta\tau^{\text{Input}}$ is not statistically significant in the second stage leads us to focus on the alternative specification (columns (11)–(14)). For this specification the hypothesis that the β are equal across bins is rejected ($F = 4.78$, $p < .001$).

How does the inclusion of $\Delta\tau^{\text{Input}}$ change our conclusions? Our baseline results (i.e., without $\Delta\tau^{\text{Input}}$ in either stage) appeared in Table V and are repeated as Panel A of Table VI. Comparing across Panels A and B of Table VI, three results are apparent. First, column (11) shows that the estimated β are virtually identical across the two panels of results. Adding $\Delta\tau^{\text{Input}}$ does not affect the estimated β. Second, column (13) shows the impact of improved U.S. market access on exports both for plants' final goods and for plants' intermediate input purchases, that is, it is the induced change in ΔT due to both the $\Delta\tau_q$ ($q = 2, 3, 4$) *and* $\Delta\tau^{\text{Input}}$. This joint impact has grown as a result of including $\Delta\tau^{\text{Input}}$. Third, column (14) shows the impact on labor productivity of being induced to export because of improved U.S. market access. The impact has grown because ΔT has grown. Consequently, the estimated total effect across all bins is larger, 0.128 log points as compared to 0.107 log points for the baseline specification. Because these plants accounted for 23% of manufacturing employment in 1996, improved access to U.S. intermediate inputs raised labor productivity in manufacturing by 0.005 log points ($= 0.23 \times [0.128 - 0.107]$). In short, there were additional gains to Canadian plants from improved access to U.S. markets for intermediate inputs.

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[21] In this alternative specification, the first stage remains as before. $\Delta\tau^{\text{Input}}$ is thus an instrument for T and the exogeneity of $\Delta\tau^{\text{Input}}$ is supported by the small overidentification test statistics in column (16).

IV.G. Bilateral Trade Liberalization—The Effect of Canadian Tariff Cuts on Final Goods

The final goods produced by Canadian plants were subject not just to U.S. tariff reductions ($\Delta\tau$), but also to Canadian tariff reductions ($\Delta\tau^{\text{Output}}$). In terms of our model, a fall in Canadian tariffs raises U.S. exports to Canada, which reduces A and leads to a market-shrinking effect, that is, makes it less likely that a plant will export or invest in productivity. The Canadian tariff cuts also raise an econometric issue. Because the two tariff cuts $\Delta\tau$ and $\Delta\tau^{\text{Output}}$ are correlated, $\Delta\tau$ may not be a clean instrument.

To examine these issues, we add $\Delta\tau^{\text{Output}}$ to the analysis in exactly the same way that we added $\Delta\tau^{\text{Input}}$ in the preceding section. The results appear in Panel C of Table VI. The first stage (equation (9)) is now augmented by the inclusion of $\Delta\tau^{\text{Output}}$ and $\Delta\tau^{\text{Input}}$. Their estimated coefficients appear in columns (3) and (1), respectively. None of the five coefficients on $\Delta\tau^{\text{Output}}$ are economically large and only one is statistically significant ($t = -2.87$). Columns (5)–(10) show estimates of the second stage (equation (8)) when augmented by $\Delta\tau^{\text{Output}}$ and $\Delta\tau^{\text{Input}}$. Only one of the coefficients on $\Delta\tau^{\text{Output}}$ is significant ($t = 2.72$) and all have the wrong sign. We therefore also consider a specification with $\Delta\tau^{\text{Output}}$ and $\Delta\tau^{\text{Input}}$ omitted from the second stage. See columns (11)–(14). Comparing these with column (11)–(14) of Panel B, it is apparent that adding $\Delta\tau^{\text{Output}}$ to the analysis does not affect our results.

To examine further we turn to import data rather than tariff data. We calculated the log change in HS6 Canadian imports from the United States between 1988 and 1996. We then matched these HS6 changes to the HS6 commodities produced by Canadian plants and calculated the average change for each plant, averaged across HS6 products produced by the plant. As always, we use simple, unweighted averages. Let $\Delta \ln M$ denote the resulting plant-level measure of import competition. Panel D of Table VI introduces $\Delta \ln M$ into the analysis in place of $\Delta\tau^{\text{Output}}$. The estimated coefficients on $\Delta \ln M$ are economically and statistically small in both stages (columns (3) and (9)). We therefore focus on the alternative second stage. The column (11) estimated β continue to display heterogeneity. (Equality of β across bins is rejected, $F = 5.11$, $p < .001$.) Further, from column (14) the total effect barely changes: from 0.128 in Panel B to 0.121 in Panel D.

Summarizing, adding Canadian imports of final goods or Canadian tariff cuts on final goods to the analysis does not alter our conclusions.[22]

V. STARTING TO EXPORT AND INVESTING IN PRODUCTIVITY

This is a long paper, so it is useful at this point to flag the importance of this section. We have now accomplished the first of two major goals of this paper dealing with new exporters: we have shown that there is a reduced-form correlation between labor productivity gains and exporting for the low- and medium-productivity plants that were induced to export as a result of improved access to U.S. markets. Our second goal is to link these labor productivity gains to active investments in productivity. We will show in this section that the same plants that benefited from being induced to export—plants in bins 1, 2, and 3—were also the plants that engaged in product innovation and adopted advanced manufacturing technologies.

Data are from the 1993 Survey of Innovation and Advanced Technologies (SIAT). See Baldwin and Hanel (2003) for a description of the survey. The surveyed plants include 512 plants that are in our group of 5,233 plants. We start with the survey's technology-adoption questions. The survey asks plants about their current use of various types of technologies and the year of initial adoption. The most important of these technologies is manufacturing information systems (MIS), which deals with computer-based production management and scheduling systems for orders, inventory, and finished goods. MIS also deals with computer-based management of machine loading, production scheduling, inventory control, and material handling. These systems are necessary for a variety of productivity-enhancing production techniques such as just-in-time inventory and lean manufacturing. Investments in MIS are thus a central component of any productivity-enhancing change in production techniques.

The first pair of rows in Table VII deal with MIS adoption rates over the period 1989–1993. We start in 1989 because the FTA came into effect on January 1, 1989. With only 512 plants we must be careful about degrees of freedom. We thus group plants in bins 1, 2, and 3 into a single "low" bin and group plants in bins 4

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[22] This is not to say that the Canadian tariff cuts were unimportant in other ways. See Section IV.C.

TABLE VII
TECHNOLOGY ADOPTION AND PRODUCT INNOVATION

	Raw adoption and innovation rates				OLS		IV			
	New exporter	Nonexporter	New–Non							
			Difference	%	β	t	β	t	Difference t	$\beta\Delta T$
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)
Adoption of advanced manufacturing technologies, 1989–1993										
1. Manufacturing information systems										
Low bin	0.16	0.06	0.10	183	0.0077	(3.21)	0.018	(3.36)	(3.55)	0.07
High bin	0.16	0.17	−0.01	−5	−0.0008	(−0.18)	−0.018	(−1.92)		−0.06
2. Inspection and communications										
Low bin	0.18	0.10	0.07	72	0.0068	(2.55)	0.021	(3.61)	(3.17)	0.08
High bin	0.14	0.20	−0.06	−30	−0.0064	(−1.31)	−0.013	(−1.39)		−0.05
Engagement in innovative activities, 1989–1991										
3. Any product or process innovation										
Low bin	0.30	0.20	0.09	46	0.0073	(2.17)	0.021	(2.77)	(2.78)	0.08
High bin	0.53	0.57	−0.03	−6	−0.0011	(−0.18)	−0.018	(−1.50)		−0.06
4. Any product innovation										
Low bin	0.26	0.14	0.12	82	0.0083	(2.75)	0.019	(2.77)	(2.31)	0.07
High bin	0.43	0.47	−0.05	−10	−0.0022	(−0.34)	−0.011	(−0.90)		−0.04
Labor productivity growth, 1988–1996										
5. Labor productivity growth										
Low bin	0.030	0.005	0.024		0.0025	(3.92)	0.005	(3.37)	(2.50)	0.018
High bin	−0.005	−0.007	0.002		−0.0001	(−0.16)	−0.002	(−0.81)		−0.005

Notes. Plants are grouped into four types depending on whether they are (i) new exporters or nonexporters and (ii) in low bins (bins 1, 2, and 3 of Figure IV) or high bins (bins 4 and 5 of Figure IV). As always, low-bin plants are smaller and less productive. Columns (1)–(5) deal with raw adoption and innovation rates. For concreteness, the first number in the table states that 16% of low-bin new exporters adopted manufacturing information systems during the 1989–1993 period. Column (3) reports the difference between columns (1) and (2). Column (4) reports the percentage difference between columns (1) and (2): 100 · (column (1) − column (2))/(column (2)). Columns (6)–(11) provide estimates of the treatment coefficient β using the Angrist–Imbens estimating equation (8), but with a binary dependent variable (adopt or not, engage in innovation or not). Column (10) reports the t-statistic for the IV difference between the low and high bins ($\beta_{\text{Low}} - \beta_{\text{High}}$). Column (11) reports the impact of improved U.S. market access on adoption rates and engagement in innovation. The specifications in columns (6)–(10) are given by equations (8) and (9), but with two-digit SIC fixed effects in both equations. The sample consists of those 1984 nonexporters that appear in the SIAT survey ($N = 512$). t-statistics are in parentheses.

and 5 into a single "high" bin. Columns (1)–(5) of Table VII provide summaries of the raw adoption rates for the two groups. Within each of these two groups, columns (2) and (3) provide the adoption rates of new exporters and nonexporters, respectively. Among low-bin plants, 16% of new exporters adopted MIS between 1989 and 1993, whereas only 6% of nonexporters did. Column (4) reports the difference, 10 percentage points. Restated, new exporters were 183% ($\cong (16-6)/6$) more likely than nonexporters to have adopted at least one advanced manufacturing technology by 1993. See column (5). Among high-bin plants, 16% of new exporters and 17% of nonexporters had adopted at least one technology by 1993. Their adoption rates were virtually identical. Putting the low- and high-bin results together, among the group of plants with labor productivity gains for new exporters (i.e., low-bin plants), new exporters were adopting advanced technologies more frequently than nonexporters. In contrast, among the group of plants with no labor productivity gains for new exporters (i.e., high-bin plants), new exporters were adopting advanced technologies about as frequently as nonexporters. This is exactly what the model predicts: productivity gains are the result of the *joint* decision to export and invest.[23]

Columns (6)–(11) of Table VII report the estimates of β in equation (8). This is the Angrist–Imbens specification that we have repeatedly used above, except that (a) the dependent variable is now a binary indicator of whether or not the plant adopted an MIS technology during the period 1989–1993 and (b) we use two-digit rather than four-digit SIC fixed effects in order to conserve on degrees of freedom. (Results using four-digit fixed effects are very similar.) Let $\widehat{\beta}_{\text{Low}}$ and $\widehat{\beta}_{\text{High}}$ be estimates of β for the low and high bins, respectively. Columns (6) and (8) report the OLS and IV estimates, respectively. The IV estimate $\widehat{\beta}_{\text{Low}}$ equals 0.018 ($t = 3.36$). To gauge its magnitude, column (11) reports $\widehat{\beta}_{\text{Low}}\Delta T_{\text{Low}}$, the increase in adoption rates for those plants that were induced to export as a result of improved access to the U.S. market. The increase is 7%, which accounts for much of the 10%

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[23] Note that adoption rates of advanced technologies were very low in 1988 for all four types of plants (low or high bin, new exporter or nonexporter), so our results cannot be explained as technology catch-up by laggards. Also note that high-bin plants tend to invest more than low-bin plants in advanced manufacturing technologies and in innovation. This is not surprising, given that high-bin plants are larger and there are other factors outside our model that drive these decisions. The main point is that *within bins* there are systematic differences between new exporters and nonexporters.

difference (column (4)) between new exporters and nonexporters. In contrast, $\widehat{\beta}_{\text{High}}$ is not statistically significant. Column (10) reports the t-statistic on $\widehat{\beta}_{\text{Low}} - \widehat{\beta}_{\text{High}}$. The difference is significant ($t = 3.55$).[24] Thus, improved access to the U.S. market was important for MIS adoption rates, but only for the same low-bin group that experienced labor productivity gains.

The second panel of results in Table VII shows a similar pattern for inspection and communications technologies.[25] Raw adoption rates were higher for new exporters than for nonexporters, but only in the low bin. See column (5). Further, the IV estimates of the productivity gains from improved U.S. market access are positive in the low bin ($t = 3.61$) and statistically insignificant in the high bin ($t = -1.39$). The difference $\widehat{\beta}_{\text{Low}} - \widehat{\beta}_{\text{High}}$ is also statistically significant ($t = 3.17$).

Turning from technology adoption rates to innovation, the third panel of results in Table VII is from the 1989–1991 innovation component of the SIAT survey. The survey asks plants whether they were engaged in product and process innovation during the 1989–1991 period. For low-bin plants, new exporters were 46% more likely than nonexporters to have engaged in any innovation. As expected, this difference disappears for high-bin plants. Most of the innovation effect is associated with product rather than process innovation. Low-bin exporters were 82% more likely than nonexporters to engage in product innovation, a difference that disappears for high-bin plants. Columns (6)–(11) of the table show that these raw-data results carry over to our IV framework. The low-bin IV estimates have the expected signs and are economically and statistically significant. As expected, the high-bin IV estimates are not statistically significant.

Table VIII examines the effects of adding additional plant-specific instruments: the Canadian tariff cuts on intermediate inputs ($\Delta\tau^{\text{Input}}$), the Canadian tariff cuts on final goods ($\Delta\tau^{\text{Output}}$), and the log change in Canadian imports from the United States ($\Delta \ln M$). The four panels in the table correspond exactly to those in Table VI (alternative second stage). For the sake of space, we

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[24] See Online Appendix VII for a description of how this t-statistic is calculated.

[25] Inspection and communications includes (a) automated sensor-based equipment used for inspection/or testing of incoming materials, in-process materials and final products (e.g., tests of failure rates); (b) local area networks for technical data and factory use and inter-company computer networks linking the plant to subcontractors, suppliers, and/or customers; (c) programmable controllers; and (d) computers used for control on the factory floor.

TABLE VIII
TECHNOLOGY ADOPTION AND PRODUCT INNOVATION: ROLE OF CANADIAN TARIFF CUTS

Adoption of advanced manufacturing technologies, 1989–1993						Engagement in innovative activities, 1989–1991								
Manufacturing information systems (MIS)			Inspection and communication			Any innovation			Any product innovation			Labor productivity growth 1988–1996		
(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)	(14)	(15)
A. Baseline (from Table VIII)														
0.018	(3.36)	(3.55)	0.021	(3.61)	(3.17)	0.021	(2.77)	(2.78)	0.019	(2.77)	(2.31)	0.0048	(3.37)	(2.50)
B. $\Delta\tau^{\text{Input}}$ as an instrument														
0.018	(3.49)	(3.21)	0.018	(3.13)	(2.99)	0.019	(2.73)	(3.20)	0.015	(2.42)	(2.64)	0.0055	(4.11)	(3.11)
C. $\Delta\tau^{\text{Input}}$ and $\Delta\tau^{\text{Output}}$ as instruments														
0.019	(3.78)	(3.64)	0.017	(3.10)	(3.14)	0.018	(2.65)	(3.25)	0.015	(2.41)	(3.11)	0.0054	(4.14)	(3.11)
D. $\Delta\tau^{\text{Input}}$ and $\Delta \ln M$ as instruments														
0.017	(3.45)	(3.15)	0.017	(3.14)	(3.12)	0.019	(2.73)	(3.31)	0.015	(2.49)	(2.82)	0.0054	(4.16)	(3.20)

Notes. This table reports IV estimates of β_{Low}, the low-bin coefficient on the treatment T in the second-stage equation (8). The dependent variable is the binary variable listed in the column heading. t-statistics are in parentheses and two t-statistics are reported. The first is for $H_0 : \beta_{\text{Low}} = 0$. The second is for $H_0 : \beta_{\text{Low}} = \beta_{\text{High}}$. Panel A of results is carried over from columns (8)–(10) of Table VII. Panels A–D of this table correspond in structure to Panels A–D of Table VI (alternative second stage). The sample consists of those 1984 nonexporters that also appear in the SIAT survey ($N = 512$).

only report the IV estimates of β_{Low}. The IV estimates of β_{High} are never significant. We report two t-statistics, the first for $H_0 : \beta_{\text{Low}} = 0$ and the second for $H_0 : \beta_{\text{Low}} = \beta_{\text{High}}$. For each technology the results are very stable across the four specifications and imply the same conclusions as implied by Table VII. Online Appendix Table B.15 reports the IV estimates of β_{Low} when $\Delta\tau^{\text{Input}}$, $\Delta\tau^{\text{Output}}$, and/or $\Delta \ln M$ are added to both the first and second stages. In Table B.15 the second-stage coefficients on $\Delta\tau^{\text{Input}}$, $\Delta\tau^{\text{Output}}$, and $\Delta \ln M$ are never statistically significant and the estimates of β_{Low} are very similar to those in Table VIII.[26]

To conclude this important section, we have shown that those plants that have a positive correlation of exporting with labor productivity growth (i.e., small, less productive plants) are the same plants that have positive correlations of exporting with both technology adoption and product innovation. This is consistent with a model featuring a complementarity between exporting and investing.

VI. Problems with Labor Productivity

We have shown that for plants that were induced by U.S. tariff cuts to export, those that were initially smaller and less productive experienced (a) high rates of investment in advanced technology adoption and product innovation and (b) high rates of labor productivity growth. It is possible that the labor productivity growth does not reflect any TFP growth, but instead reflects either growing mark-ups or high rates of investment. We do not have the capital stock data needed to rule out this possibility. However, there are three indirect pieces of evidence that bear on the issue.

The first and most important was suggested to us by Kala Krishna. Suppose that our result is an artifact of higher mark-ups: improved access to the U.S. market led Canadian firms to charge higher prices and thus have higher value added and higher labor productivity. If there were no underlying difference in the TFP performance of new exporters relative to nonexporters, then the higher mark-ups and prices charged by new exporters would cause them to lose sales relative to nonexporters in the *domestic* (Canadian) market. Yet exactly the opposite happened. New

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[26] Online Appendix Tables B.16 and B.17 repeat Tables VII and VIII, respectively, using the probit-based alternative definition of bins. The results are less significant for MIS and more significant for product innovation.

exporters increased their sales in the Canadian market relative to nonexporters. It follows that new exporters must have been offering lower prices in the Canadian market. This is inconsistent with rising mark-ups, but consistent with rising TFP.

To show that new exporters increased their Canadian sales relative to nonexporters we reestimate equations (8) and (9) just as in specification A of Table V, but with a single change: the dependent variable is now the average annual log change in domestic (Canadian) sales, 1988–1996. The results appear in Panel A of Table IX. The IV estimates of the impact of treatment T on domestic sales appear in column (1). As expected, the pattern of heterogeneity is similar to that for labor productivity growth. Thus, the pattern of domestic sales growth mirrors the pattern of labor productivity growth.

We turn to a second argument in favor of interpreting our labor productivity results in terms of TFP. We have seen that the same firms that experienced rising labor productivity were also rapid adopters of MIS. We know that these new systems—which include just-in-time inventory management and lean manufacturing—are associated both with higher TFP (e.g., Brynjolfsson and Hitt [2003], Bloom and Reneen [2007]) and with reductions in inventories of intermediate inputs (Feinberg and Keane 2006; Keane and Feinberg 2007). If our correlation of labor productivity with exporting at least partly reflects the impact of MIS adoption on TFP, then this impact should also be reflected in inventory reductions.

To examine this, we again reestimated equations (8) and (9), this time using the 1988–1996 average annual log change in raw material inventories as our dependent variable in equation (8). Note that this is an eight-year change, so that it is unlikely we are capturing short-run business cycle phenomena. The results appear in Panel B of Table IX. As expected, we see that inventories fell significantly, but only in the lower bins, that is, for initially smaller, less productive plants. See column (5).

The third argument in favor of interpreting our labor productivity results in terms of TFP is weaker, but still informative. If TFP rose, one might expect that it would reduce inputs per unit of output. To examine the argument, we again reestimate equations (8) and (9), this time with the dependent variable in equation (8) redefined either as the 1988–1996 average annual log change in intermediate input purchases divided by sales or as the 1988–1996 average annual log change in energy purchases divided by

TABLE IX

Evidence That Labor–Productivity Gains Reflect TFP Gains

Bin	β (1)	t (2)	$\beta\times\Delta T\times 8$ (3)	N (4)	Bin	β (5)	t (6)	$\beta\times\Delta T\times 8$ (7)	N (8)
	A. Domestic (Canadian) sales					B. Inventories of intermediate inputs			
1	0.0064	(3.18)	0.08	979	1	−0.0022	(−7.80)	−0.03	979
2	0.0060	(4.50)	0.14	953	2	−0.0010	(−4.44)	−0.02	953
3	0.0031	(2.32)	0.09	1,208	3	−0.0010	(−5.62)	−0.03	1,208
4	0.0029	(1.72)	0.05	1,020	4	−0.0007	(−3.64)	−0.01	1,020
5	0.0022	(1.00)	0.03	1,072	5	−0.0002	(−1.07)	0.00	1,072
Total			0.057		Total			−0.013	
	C. Intermediate input purchases/shipments					D. Energy purchases/shipments			
1	−0.0074	(−6.45)	−0.09	980	1	−0.0086	(−4.75)	−0.11	980
2	−0.0049	(−4.56)	−0.11	951	2	−0.0109	(−7.94)	−0.26	951
3	−0.0044	(−3.97)	−0.12	1,205	3	−0.0092	(−6.64)	−0.26	1,205
4	−0.0034	(−1.77)	−0.05	1,020	4	−0.0056	(−2.63)	−0.09	1,020
5	0.0002	(0.14)	0.00	1,071	5	0.0025	(1.27)	0.03	1,071
Total			−0.052		Total			−0.092	

Notes. This table reports IV estimates of equation (8), except that the dependent variable is now the 1988–1996 average annual log change in the variable indicated in the heading of the panel, such as, the 1988–1996 average annual log change in domestic (Canadian) sales. β in columns (1) and (5) is the coefficient on the treatment variable ($T = 0, \ln \text{EXP}_{1996}$). Aside from the dependent variable, the specification is identical to that in the baseline specification of Table V. The predicted treatment used for IV (and for the change in treatment ΔT used in columns (3) and (7)) comes from equation (9) and uses the same specification as in the baseline specification of Table V. The sample consists of 1984 nonexporters. t-statistics are in parentheses.

sales. From Table IX, we see economically significant drops in input usage per unit of output for both intermediates and energy, but as expected, only for initially smaller, less productive plants. (A minor exception is the first bin of the energy results.)

To summarize, the heterogeneous pattern of labor productivity growth that we estimated is also to be found in other outcomes, such as increased domestic sales and improved inventory management, that are likely to be correlated with TFP growth.

VII. OLD EXPORTERS

So far we have examined plants that did not export in 1984. We now examine the old exporters sample, that is, plants that were already exporting in 1984 (before the FTA negotiations began) and that continued to export in 1996. The theory makes two predictions about such plants. First, U.S. tariff cuts will induce some of them to start investing in productivity. Second, those that invest will all experience *the same* log change in productivity. This can be seen from Figure II, where to the right of the Melitz cutoff, the U.S. tariff cut causes a parallel shift downward of the horizontal line. *Thus, the productivity gains are predicted to be independent of initial productivity* φ_0. Intuitively, initial productivity only matters when there is a *joint* decision about exporting and investing in productivity. For old exporters, the exporting fixed costs have already been incurred, so that the joint decision has been replaced by a single decision about whether or not to invest.

To investigate these predictions, we consider the set of 1,607 plants that exported in both 1984 and 1996. Sample statistics appear in Table I. Other details of the sample appear in Online Appendix VI. Because old exporters exported in both 1984 and 1996, we can define the treatment as the average annual log change in exporting, $T \equiv \ln(\text{EXP}_{1996}/\text{EXP}_{1984})/12$. The median value of T is 0.075. We begin by computing the average annual log change in labor productivity over the period 1988–1996 ($\Delta\varphi$) for two groups of old exporters, those with $T \geq 0.075$ and those with $T < 0.075$. The difference in labor productivity growth between these two groups is a very substantial 0.018 log points per year ($t = 4.30$). See column (1) of Table X.

The remaining rows of Table X report the same difference in labor productivity growth, but for three subsamples of plants that were selected based on initial (1988) labor productivity and initial (1988) employment size. The first bin contains plants that have

TABLE X

LABOR PRODUCTIVITY GROWTH 1988–1996, OLD EXPORTERS SAMPLE: OLS

	Labor productivity growth differences		OLS			
	Mean (1)	t (2)	β (3)	t (4)	R^2 (5)	N (6)
All plants	0.018	(4.30)	0.060	(4.04)	.15	1,607
1. Productivity *and* employment below median	0.021	(2.22)	0.061	(1.70)	.11	351
2. Other	0.011	(1.87)	0.060	(2.84)	.15	802
3. Productivity *and* employment above median	0.023	(3.38)	0.059	(2.23)	.09	454

Notes. This table reports results for the sample of old exporters, plants that exported both in 1984 and in 1996. The treatment is defined as the average annual log change in exports, $T = \ln(\text{EXP}_{1996}/\text{EXP}_{1984})/12$. Columns (1) and (2): Divide plants into two groups based on whether the plant has a T that is above or below 0.075 (the median of T). Column (1) reports the *difference* between the two groups in labor productivity growth $\Delta\varphi$ (above median minus below median). Columns (3)–(5): These report OLS estimates of equation (8). The coefficient on T is reported in column (3). The specification of equation (8) differs from that reported in our baseline specification (Panel A of Table V) in three ways: (i) the sample consists of old exporters, (ii) the treatment is $T = \ln(\text{EXP}_{1996}/\text{EXP}_{1984})/12$, and (iii) industry fixed effects are defined at the two-digit SIC level.

below-median productivity *and* below-median employment. The third bin contains plants that have above-median productivity *and* above-median employment. Restated, these two bins contain the extremes of smaller, less productive plants and larger, more productive plants. Bin 2 ("Other") contains the remaining "middle" plants.[27] For all three bins, column (1) of Table X shows that plants with above-median export growth experienced more rapid labor productivity growth. However, this is not always statistically significant.

To control for plant characteristics, we reestimated equation (8) for the old exporters sample ($N = 1{,}607$). The OLS estimates of the treatment coefficient β appear in column (3) of Table X. Even with controls, the correlation of increased exporting with labor productivity growth is positive. *More importantly, this labor productivity effect is the same for all three bins. There is no negative selection.*[28] *This is exactly as predicted by the theory and is in marked contrast to our previous predictions and results about*

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[27] In deciding whether a plant is below the median, we use the median of all plants in the same two-digit SIC industry. That is, we use within-industry rankings, as elsewhere in the paper. See footnote 9.

[28] The hypothesis of coefficient equality across the three bins is accepted ($F = 0.01$, $p = 1.00$).

new exporters. The fact that the treatment effect β is decreasing across bins for new exporters and constant across bins for old exporters makes it more difficult (but not impossible) to argue that our results are driven by unobservables that are correlated with initial productivity.[29]

The IV estimates of equations (8) and (9) appear in the first row of Table XI. Columns (1) and (2) show that the estimate of the treatment effect β is positive and statistically significant ($t = 3.46$). However, this must be viewed with caution, because for many old exporters increased exports were due less to tariff cuts and more to technology-enabled integration of North American operations (Feinberg and Keane 2006). Also, the Hausman test statistic is small (column (9)) and the F-test for the joint significance of the tariff variables in the first stage is only 5.73, below the Staiger and Stock (1997) threshold of 10.

With these caveats in mind, column (4) reports the impact on labor productivity growth of improved foreign market access, $\beta \times \Delta T \times 8 = 0.067 (= 0.33 \times 0.026 \times 8)$. That is, for plants that were induced by the U.S. tariff cuts to increase their exports, increased exports raised labor productivity over the 1988–1996 period by 6.7%. Because these plants accounted for 21% of manufacturing employment in 1996, improved market access raised manufacturing productivity by 0.014 log points ($= 0.067 \times 0.21$). *This 1.4% effect is an important message of this paper.*

The remainder of Table XI introduces the Canadian tariff cuts into the first stage. The four panels of the table coincide with the four panels in Table VI. The main results are similar to what we saw for 1984 nonexporters. The reduction in Canadian tariffs on intermediate inputs purchased from the United States is important for predicting an increase in exporting (columns (5) and (6)). Whereas adding $\Delta\tau^{\text{Input}}$ reduced the estimated treatment coefficient β (column (1)), it raises the impact of improved market access on exporting (ΔT in column (3)) because access is improved not only for plants' final goods, but also for plants' purchases of U.S. intermediate inputs. This in turn raises the impact of improved market access on labor productivity to 8.8%. See column (4).

Panels C and D of Table XI report the results of adding as regressors either the Canadian tariff cuts on final goods

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[29] Online Appendix Table B.18 repeats Table X using the probit-based definition of bins. The conclusions are the same.

TABLE XI
LABOR PRODUCTIVITY GROWTH 1988–1996, OLD EXPORTERS SAMPLE: IV

Second stage				First stage				Tests	
Treatment T				$\Delta\tau^{\text{Input}}$		$\Delta\tau^{\text{Output}}$, $\Delta\ln M$			
β (1)	t (2)	ΔT (3)	$\beta\times\Delta T\times 8$ (4)	Coeff. (5)	t (6)	Coeff. (7)	t (8)	Hausman (9)	Overidentification (10)
A. Baseline ($\Delta\tau^{\text{Input}}$, $\Delta\tau^{\text{Output}}$, and $\Delta\ln M$ omitted from both stages)									
0.33	(3.46)	0.026	0.067					6.80	0.72
B. $\Delta\tau^{\text{Input}}$ as an instrument									
0.27	(3.45)	0.040	0.088	0.53	(4.08)			6.64	0.77
C. $\Delta\tau^{\text{Input}}$ and $\Delta\tau^{\text{Output}}$ as instruments									
0.23	(2.92)	0.042	0.075	0.46	(3.42)	0.17	(1.80)	4.41	1.00
D. $\Delta\tau^{\text{Input}}$ and $\Delta\ln M$ as instruments									
0.29	(3.74)	0.044	0.102	0.49	(3.75)	0.06	(1.34)	7.90	0.78

Notes. This table is similar in structure to Table VI. (The Alternative second stage in Table VI corresponds to the Second stage in this table.) The table reports estimates of equations (8) and (9). The dependent variable is 1988–1996 labor productivity growth ($\Delta\varphi$). The treatment is $T = (\ln \text{EXP}_{1996}/\text{EXP}_{1984})/12$, the average annual log change in exports. The overidentification χ^2 statistics in column (10) are all small, indicating that U.S. and Canadian tariff cuts are exogenous. See the notes to Table VI for an explanation of the remaining columns and rows. The sample consists of the set of plants that exported in both 1984 and 1996 ($N = 1{,}607$). t-statistics appear in parentheses.

(Panel C) or the change in Canadian import competition (Panel D). The first-stage coefficients have the wrong sign and are statistically insignificant. See columns (7) and (8). They thus add nothing to the analysis.[30]

Finally, we have not reported IV results by bin because with twenty-seven Angrist–Imbens "interaction" instruments and twenty two-digit SIC industry fixed effects, our bins of 351 and 454 plants are too small. We thus consider a more parsimonious specification that (a) divides the sample into only two bins (based on 1988 labor productivity) and (b) eliminates the interaction terms from the first stage.[31] The IV estimates of the treatment coefficients β are 0.171 ($t = 1.66$) and 0.195 ($t = 1.53$) for the low- and high-productivity bins, respectively. Thus, there is cross-bin coefficient homogeneity, exactly as predicted by the theory.

VIII. CONCLUSIONS

This paper presented three core empirical results for the 1984 nonexporters sample, that is, for plants that did not export in 1984 and survived to 1996:

1. Table V showed that there were labor productivity gains for those Canadian manufacturing plants that were induced to export because of improved access to the U.S. market. Table VI showed that two distinct elements of market access matter: selling final products into the United States ($\Delta\tau$) and buying intermediate inputs from the United States ($\Delta\tau^{\text{Input}}$). The former is more important.[32]
2. Table VII showed that the labor productivity gainers also had high post-Agreement adoption rates of advanced manufacturing technologies and high post-Agreement levels of product innovation. That is, the new exporters who experienced labor productivity gains were investing in productivity.

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[30] In Table XI, the Canadian tariffs were added to the first stage (equation (9)), but not to the second stage (equation (8)). This is because, as shown in Online Appendix Table B.19, neither $\Delta\tau^{\text{Input}}$ nor $\Delta \ln M$ is statistically significant when added to the second stage. Interestingly, $\Delta\tau^{\text{Output}}$ is statistically significant in the second stage ($t = -2.86$), and when added to the second stage, it raises our estimate of $\beta \times \Delta T \times 8$ by a lot, to 0.137.

[31] That is, $\lambda_{kqq'} = 0$ in equation (9).

[32] The labor productivity gains were 10.7% in our Angrist–Imbens baseline specification with only $\Delta\tau$ (Table V), 12.8% when both $\Delta\tau$ and $\Delta\tau^{\text{Input}}$ were included (Table VI), and 15.3% in a standard IV approach with only $\Delta\tau$ (Table III).

3. Table IX showed that the labor productivity gainers also increased their domestic (Canadian) sales relative to nonexporters. This is exactly what one would expect if the labor productivity gains reflected underlying TFP gains.

Each of these outcomes was heterogeneous and displayed "negative selection" in initial productivity. For example, the labor productivity gains were largest for the initially least productive plants and fell to zero for the initially most productive plants. We argued that such negative selection is consistent with a model featuring two-dimensional heterogeneity. One dimension of heterogeneity is in initial productivity, as in Melitz (2003), and the second is in the productivity gains from investing. With two-dimensional heterogeneity, negative selection is an immediate consequence of the fundamental complementarity between exporting and investing in productivity. Restated, market size matters for innovation and improved foreign market access induces innovation. This insight explains all of our results.

There are additional insights to be had from old exporters, that is, from plants that exported both in 1984 and in 1996. Old exporters that were induced to export because of improved access to the U.S. market also experienced labor productivity gains (Table XI). Further, and as predicted by the model, the old-exporter labor productivity gains did not display negative selection or any other form of heterogeneity (Table X).[33] The fact that the treatment effect β was decreasing across bins for new exporters and constant across bins for old exporters makes it more difficult (but not impossible) to argue that our results are driven by unobservables that are correlated with initial productivity.

An important goal of our work was to assess the impact of the Canada–U.S. Free Trade Agreement. We estimated that *within-plant* increases in labor productivity were large enough to raise labor productivity in Canadian manufacturing as a whole by between 4.8% (using equations (8) and (9)) and 5.6% (using equations (5) and (6)). In addition, the FTA led to *between-plant* increases in labor productivity of 4.3% due to plant exit and 4.1% due to expansion of high–productivity plants. (See Sections IV.C and VII for sample calculations.) Summing these figures, the FTA increased Canadian manufacturing productivity by between 13.2%

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[33] Negative selection is due to the complementarity of the *joint* decision to export and invest. For old exporters, the fixed costs of exporting have already been incurred, so there is no joint decision and no negative selection.

and 14.0%. The fact that a single government policy can be so important is truly remarkable.

Finally, we argued in the Introduction that in the presence of heterogeneous responses we should expect results to differ across studies depending on the choice of instrument. This neatly explained why results in the exporting-and-productivity literature (e.g., Clerides, Lach, and Tybout [1998] and Bernard and Jensen [1999]) vary across studies. The productivity effect depends on who exports and why. These in turn are framed by the instrument.

APPENDIX

Let I be a binary indicator of whether the firm invests ($I = 1$) or not ($I = 0$). Let $\pi_I(E)$ be profits as in equations (1) and (2). The firm chooses one of four alternatives, $(E, I) \in \{(0, 0), (0, 1), (1, 0), (1, 1)\}$. Each line in Figure A.1 corresponds to an indifference condition between two alternatives. For example, the comparison $\pi_1(1) = \pi_0(1)$ is the horizontal line to the right of the Melitz cutoff $F^E/\tau^{-\sigma}A^*$. The label is always above the line and indicates the region for which the inequality holds. For example, $\pi_1(1) > \pi_0(1)$ holds above the line and $\pi_1(1) < \pi_0(1)$ holds below the line. It is trivial to verify that the lines are correctly drawn.

Consider the region to the right of the Melitz cutoff. We know from equation (3)—see the first term and the discussion following the equation—that the firm always exports in this region. We therefore only have to consider alternatives $(E, I) = (1, 1)$ and $(E, I) = (1, 0)$, that is, we only have to consider the horizontal line. Thus, the firm exports and invests above the horizontal line and exports without investing below the horizontal line. This completes the proof for the region to the right of the Melitz cutoff.

Now consider the region to the left of the Melitz cutoff, but to the right of $\underline{\varphi}_0 \equiv [F^E/(\tau^{-\sigma}A^*)] - (F^I/A)$. Because we are to the left of the Melitz cutoff, the firm will never export without investing; that is, we can ignore the choice $(E, I) = (1, 0)$. Above the solid line we have $\pi_1(1) > \pi_0(0)$ and $\pi_1(1) > \pi_1(0)$; that is, $(E, I) = (1, 1)$ is preferred to (0, 0) and (0, 1). Hence, the firm exports and invests. Below the solid line we have $\pi_1(1) < \pi_0(0)$ and $\pi_1(0) < \pi_0(0)$. Hence the firm neither exports nor invests. This completes the proof of the theory in the main text, which assumed that $\varphi_0 > \underline{\varphi}_0$.

Finally, consider the region to the left of $\underline{\varphi}_0 \equiv [F^E/(\tau^{-\sigma}A^*)] - (F^I/A)$. As in the preceding paragraph, we need not consider

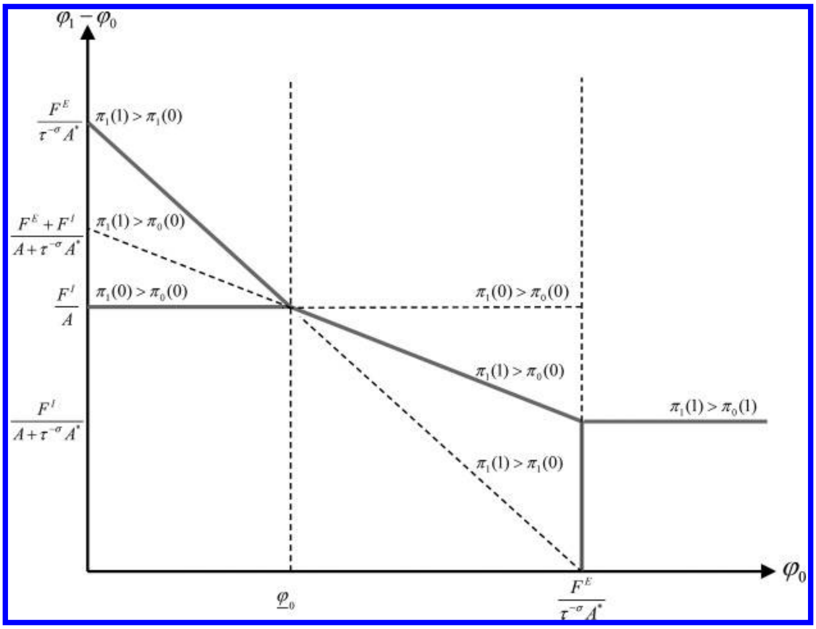

FIGURE A.1
Proof of the Theory

exporting without investing. Above the top solid line we have $\pi_1(1) > \pi_1(0)$ and $\pi_1(1) > \pi_0(0)$. Hence, the firm exports and invests. Below the bottom solid line we have $\pi_1(0) < \pi_0(0)$ and $\pi_1(1) < \pi_0(0)$. Hence the firm neither exports nor invests. Between the two solid lines we have $\pi_1(1) < \pi_1(0)$ and $\pi_1(0) > \pi_0(0)$. Hence the firm invests without exporting.

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