TRADE, QUALITY UPGRADING, AND WAGE INEQUALITY IN THE MEXICAN MANUFACTURING SECTOR*

ERIC A. VERHOOGEN

This paper proposes a new mechanism linking trade and wage inequality in developing countries—the quality-upgrading mechanism—and investigates its empirical implications in panel data on Mexican manufacturing plants. In a model with heterogeneous plants and quality differentiatiation, more productive plants produce higher-quality goods than less productive plants, and they pay higher wages to maintain a higher-quality workforce. Only the most productive plants enter the export market, and Southern exporters produce higher-quality goods for export than for the domestic market, to appeal to richer Northern consumers. An exchange-rate devaluation leads more-productive Southern plants to increase exports, upgrade quality, and raise wages relative to less-productive plants within the same industry, increasing within-industry wage dispersion. Using the late-1994 peso crisis as a source of variation and a variety of proxies for plant productivity, I find that initially more productive plants increased the export share of sales, white-collar wages, blue-collar wages, the relative wage of white-collar workers, and ISO 9000 certification more than initially less productive plants during the peso crisis period and that these differential changes were greater than in periods without devaluations before and after the crisis period. These findings support the hypothesis that quality upgrading induced by the exchange-rate shock increased within-industry wage inequality.

I. INTRODUCTION

Beginning with a series of liberalizing reforms in 1985 and 1986, the Mexican economy experienced more than a decade of both rapidly expanding trade and rising wage inequality. In current U.S. dollar terms, nonpetroleum exports and total imports rose an average of 16.5% and 15.7% per year, respectively, over the period 1985–2000. Figure I depicts the evolution of two measures of wage inequality: the log 90–10 ratio from a household

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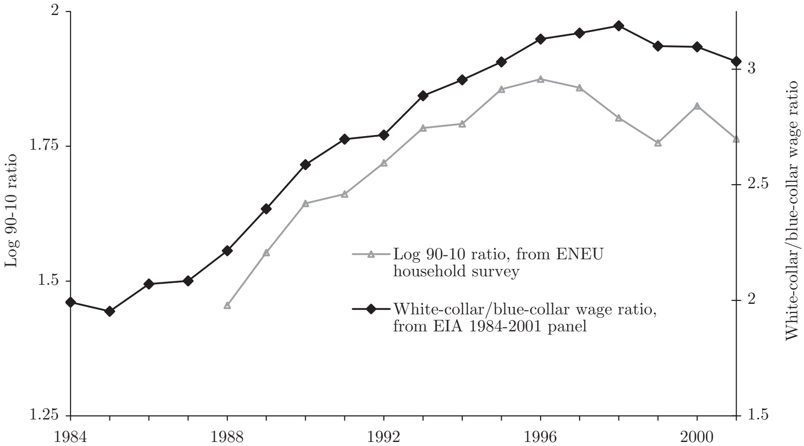

FIGURE I
Wage Inequality, 1984–2001

Notes: Log 90–10 ratio is for real hourly wages from the *Encuesta Nacional de Empleo Urbano* (ENEU) household survey. White-collar/blue-collar ratio is for hours-weighted averages of hourly wages for nonproduction workers and production workers in the *Encuesta Industrial Anual* [Annual Industrial Survey] (EIA) 1984–2001 panel of 1,114 plants. Variable definitions in Appendix I. Further details on data sets in Section IV of text and Appendix II (online).

employment survey, and the white-collar/blue-collar wage ratio from a balanced manufacturing plant panel.[1] Both illustrate a substantial increase, reaching a peak in 1996–1998. A similar coincidence of expanding trade and rising inequality has been observed in many other developing countries.[2]

From the perspective of standard Heckscher–Ohlin trade theory, this coincidence is puzzling. The simplest version of the Heckscher–Ohlin model predicts that wage inequality will *fall* in a country such as Mexico when it integrates with a country such as the United States, as production shifts toward unskilled-labor-intensive industries, raising the demand for unskilled labor. More sophisticated Heckscher–Ohlin-type models can account for a link between trade liberalization and wage inequality in a developing country such as Mexico,[3] but because such models rely

1. Variable definitions are in Appendix I. The data sets are described in more detail in Section IV below and in Appendix II, posted online on the QJE website.
2. See Goldberg and Pavcnik (2007) for a review.
3. Inequality may increase, for instance, if a unskilled-labor-abundant country such as Mexico opens trade simultaneously with a skill-abundant country such as the United States and an even more unskilled-labor-abundant country such as China (Davis 1996) or if relatively unskilled-labor-intensive industries are more protected prior to liberalization (Revenga 1997).

on between-sector shifts as the mechanism through which trade affects labor markets, they can only explain a rise in inequality if trade causes a shift in resources toward skilled-labor-intensive sectors.[4] Empirical studies have typically failed to find evidence of such shifts.[5] It is common in the literature to interpret increases in inequality that cannot be explained by between-sector shifts as evidence for non-trade-related factors such as skill-biased technical change (e.g., Berman, Bound, and Machin [1998]).

This paper proposes a new mechanism linking trade and wage inequality in developing countries—the quality-upgrading mechanism—and investigates its empirical implications in panel data on Mexican manufacturing plants. The theoretical framework combines three main elements. First, plants are heterogeneous in productivity and there is a fixed cost to entering the export market, such that only the most productive plants within each industry export, as in Melitz (2003). Second, goods are differentiated in quality and consumers differ in income and hence in willingness to pay for product quality across countries, such that a given poor-country exporting plant produces higher-quality goods for export than for the domestic market. Third, producing higher-quality goods requires higher-quality workers within each occupational category, and higher-quality workers must be paid higher wages, in the spirit of the O-ring model of Kremer (1993). In this context, an increase in the incentive to export in a developing country generates *differential quality upgrading*: initially more productive plants increase exports, produce a greater share of higher-quality goods, and raise wages relative to initially less productive plants in the same industry. Because initially more productive plants also tend to be initially higher-wage, this process increases within-industry wage dispersion.

The empirical part of the paper uses the peso devaluation of December 1994, as well as an earlier period of depreciation in 1985–1987, to investigate this mechanism. Using a number of different proxies for the underlying plant productivity parameter, I compare differential changes in outcomes between initially

4. The outsourcing model of Feenstra and Hanson (1996) is an exception; we return to it in Section VI.

5. Wacziarg and Wallack (2004) find little evidence of such shifts in response to trade liberalization in a large sample of countries. Using aggregate data from the Mexican *Censo Industrial* [Industrial Census] and schooling data from the ENEU household survey, Figures A1 and A2 in Appendix II (online) present evidence of shifts toward *less* skill-intensive and *less* capital-intensive sectors within manufacturing in Mexico over the period 1988–1998, consistent with the simplest Heckscher–Ohlin model.

more and initially less productive plants over the peso crisis period (1993–1997) to corresponding changes in periods without devaluations before (1989–1993) and after (1997–2001) the crisis period. I find greater differential changes in the export share of sales, white-collar wages, blue-collar wages, and the relative wage of white-collar workers in the peso crisis period. The results for 1986–1989, corresponding to the earlier period of depreciation, indicate patterns similar to those of 1993–1997. Using an auxiliary data set, I find greater differential changes during the peso crisis period in the likelihood of ISO 9000 certification, an international production standard commonly associated with product quality. These results support the hypothesis that differential quality upgrading induced by the exchange-rate shocks raised within-industry wage inequality.

In addition to the work cited above, this paper is related to a number of different strands of existing literature. In introducing quality-differentiated goods and asymmetric countries into a Melitz-type theoretical framework, it uses multinomial-logit microfoundations for consumer demand (McFadden 1974; Anderson, de Palma, and Thisse 1992), because they make transparent the dependence of willingness to pay for quality on consumer incomes. The paper is part of a growing literature using plant-level data from developing countries to examine plants' responses to exposure to international markets; for reviews, see Tybout (2000, 2003). The paper is also related to a number of papers examining the role of product quality in international trade, including Gabszewicz et al. (1982), Flam and Helpman (1987), Schott (2004), Hummels and Klenow (2005), Brooks (2006), and Hallak (2006). I am not aware of previous work that has focused on shifts in the within-plant product mix between goods of different qualities destined for different markets as a mechanism linking trade and labor-market outcomes.[6] Previous studies that have used exchange-rate shocks as a source of identification (without interacting with initial productivity) include Revenga (1992) and Abowd and Lemieux (1993). The sheer size of the late-1994 exchange-rate shock—along with the fact that it was largely unexpected, unlike most trade agreements—may explain why this paper finds stronger evidence of plant behavioral responses than is typical in the literature using tariff changes.

6. Bernard, Redding, and Schott (2006) analyze product switching in the context of a model with symmetric goods.

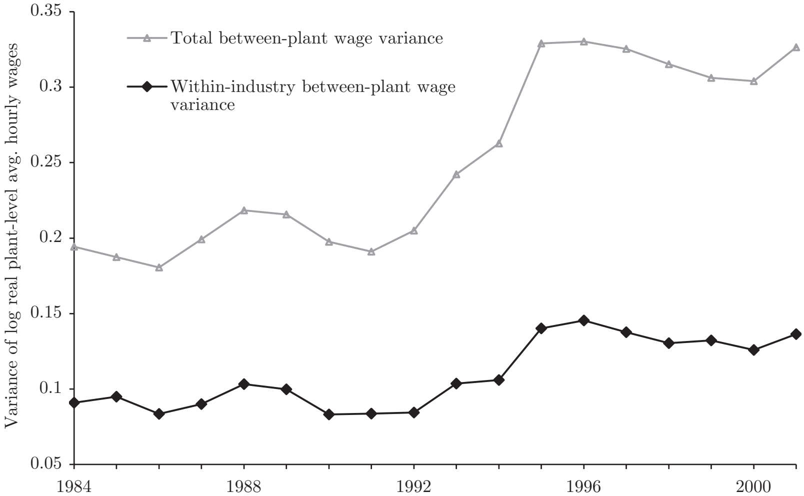

FIGURE II
Wage Variance, EIA 1984–2001 Panel

Notes: Total wage variance is hours-weighted variance of the log plant-average real hourly wage in the balanced EIA 1984–2001 panel of 1,114 plants. Within-industry-year variance is hours-weighted variance of residuals from a regression of the log plant-average real hourly wage on a full set of industry-year dummies (205 industries × 18 years) in EIA 1984–2001 panel. Variable definitions in Appendix I. Further details on data set in Section IV of text and Appendix II (online).

An important caveat is that the particular mechanism this paper focuses on—within-industry quality upgrading in response to exchange-rate devaluation—cannot explain the overall trend in Mexican wage inequality illustrated in Figure I. Many factors may have contributed to changes in aggregate inequality, among them exogenous technological change, migration (domestic and international), and other liberalization policies that have accompanied the opening to trade.[7] Figure II focuses more directly on a dimension of wage inequality that this paper may be able to explain: wage dispersion among manufacturing plants. The figure plots the variance of log plant-level average hourly wages in a balanced manufacturing plant panel, as well as the variance of the residuals from a regression of log plant-level average hourly wages on a full set of industry-year effects.[8] Total between-plant wage

7. For further discussion of wage trends in Mexico, see Cragg and Epelbaum (1996), Hanson and Harrison (1999), and Robertson (2004).

8. The data set is described in more detail in Section IV and in Appendix II (online).

dispersion rose sharply in 1994–1995 as well as in 1986–1988, during and immediately following the earlier period of depreciation. The within-industry component accounts for an especially large proportion of the increase in total (between-plant) variance in these periods: it accounts for 43.5% of the level of total variance and 34% of the change over the entire period 1984–2001, but 52% and 51.5% of the changes in 1986–1988 and 1994–1995, respectively. Although the increase in the between-industry component may itself reflect a differential shock to exporting across industries, this paper focuses on the within-industry component and the extent to which it can be attributed to the quality-upgrading mechanism.

The next section provides background on the peso crisis and presents a concrete example—a case study of the Volkswagen plant in Puebla, Mexico—to illustrate the process of quality upgrading. Section III develops the theoretical framework. Section IV describes the data sets and presents descriptive statistics. Section V presents the estimation strategy and results. Section VI addresses possible alternative explanations and Section VII concludes.

II. Background and Brief Case Study

For several years prior to the peso crisis, the Mexican government constrained the peso to vary within a narrow band (± 0.0004 pesos/day from Oct. 1992 to Dec. 1994). Over the same period, persistent trade and current-account deficits led to mounting pressure on government reserves. President Ernesto Zedillo took office on Dec. 1, 1994, and on Dec. 20 his new finance minister announced that the ceiling on the exchange-rate band would be raised 15.6%, from 3.46 to 4.00 pesos/dollar. This set off a speculative attack, and two days later the government was forced to allow the peso to float. The peso devalued immediately to 4.90 pesos/dollar and continued losing value. Figure III plots the real exchange rate, which reached a local maximum in March 1995 and recovered slowly thereafter. There are a variety of theories about what generated the extra pressure on the currency that prompted the government to devalue—a leading one points to a renewed offensive of Zapatista rebels in the Southern state of Chiapas (Economist Intelligence Unit 1995)—but whatever the precipitating event, it appears that the devaluation was largely unexpected. Both before and after the crisis, and in particular on

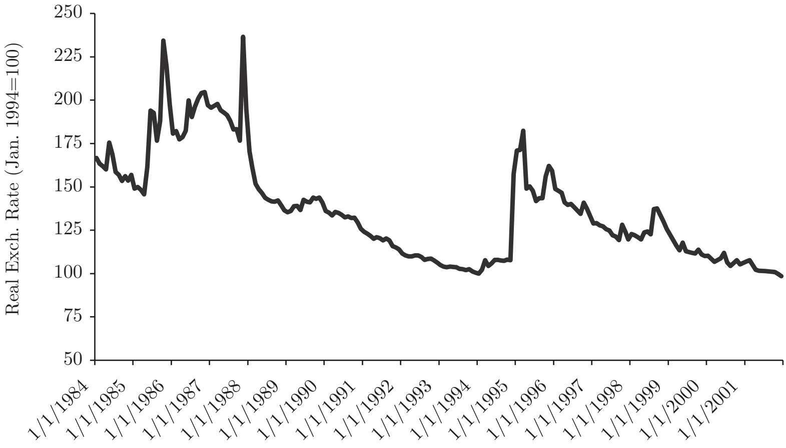

FIGURE III
Real Exchange Rate, 1984–2002

Notes. Real exchange rate calculated as RER $= e \times$ CPI(US)/CPI(Mex), where e is the peso/US$ nominal exchange rate. Data from IMF International Financial Statistics.

Dec. 1, 2004, the black market exchange rate (available monthly) and the official exchange rate coincided almost exactly.[9]

The devaluation led to a major economic contraction in Mexico, with GDP dropping by 6.2% (at constant prices) from 1994 to 1995. Nominal wages remained roughly constant through the crisis and labor costs for Mexican manufacturers fell in real peso or dollar terms. The average wage for a male full-time worker with nine years of education fell from approximately US$1.50 per hour to approximately US$0.90 per hour from 1994 to 1995, rising back to only US$1.10 per hour by 1999.[10]

It is worth emphasizing that the peso crisis was a much larger shock than the North American Free Trade Agreement (NAFTA), which had taken effect the previous January. By 1993, after eight years of liberalizing policies, almost all Mexican quotas and other nontariff barriers had been removed, and approximately 95% of all imports into Mexico were covered by tariffs of 20% or less. On the U.S. side, approximately 80% of imports were covered by tariffs of 5% or less. A majority of commodities were assigned phase-out

9. The black market rates are from Global Financial Data (http://www.globalfinancialdata.com/). See Figure A3 in Appendix II (online).

10. These figures are from the ENEU household survey. For details, refer to Appendix II (online).

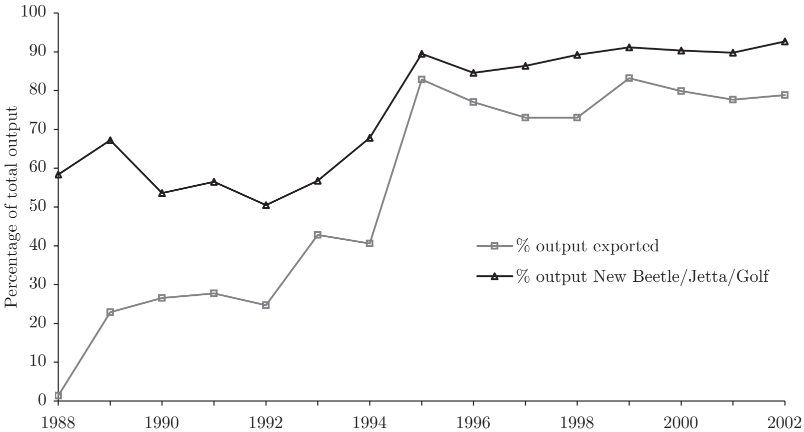

FIGURE IV
Exports, High-quality Models as Percentage of VW Output

Notes: Output measured in physical units. Omitted model from upper curve is the Original Beetle. Data from bulletins of the *Asociación Mexicana de la Industria Automotriz* (Mexican Automobile Industry Association).

schedules of five or more years. Tariff changes under NAFTA for the majority of commodities were thus typically on the order of a few percent per year or less. A common view among observers in Mexico is that NAFTA's main role was as a commitment to the general program of liberalization begun in the mid-1980s, rather than as a marked change in trade costs.

How did the manufacturing sector respond to the peso crisis? Consider the example of the Volkswagen plant in Puebla, Mexico. The Puebla plant is the sole world producer of the New Beetle and the sole North American producer of the Jetta. Until July 2003, the plant also produced the Original Beetle, almost all of which were sold in Mexico. There are clear quality differences between the Original Beetle and the newer models, the New Beetle, Jetta, and Golf (a model from which the New Beetle borrows many components).[11] These differences are reflected in prices: in July 2003, the New Beetle was selling for US$17,750, the Jetta for US$15,000, and the Original Beetle for US$7,500. Figure IV illustrates the effect of the peso crisis on the plant's product mix.

11. For example, the New Beetle and the Jetta have automatic window-raising mechanisms; the windows of the Original Beetle have to be cranked up by hand. The seats of the New Beetle and Jetta consist of polyurethane foam; the seats of the Original Beetle are made partly of lower-quality foam and partly of coconut fibers, a cheaper substitute.

Between 1994 and 1995, exports rose sharply as a share of total production because of both a decline in domestic sales and an increase in exports. The increase in the export share was accompanied by a sharp increase in production of the higher-quality varieties as a share of output.

What consequences did this shift in product mix have inside the plant? Until 2003, a striking characteristic of the plant was the juxtaposition of the production lines for the New Beetle and Jetta, which relied on state-of-the-art technology, and the production line for the Original Beetle, which employed essentially the same technology as when the plant opened in 1964, technology that had been in use in Germany since the 1950s. One consequence of the shift in product mix was thus a form of technological upgrading, an increase in the production-weighted average level of technological sophistication in the plant. This change occurred not because of an increase in the availability of new technologies but because of a greater reliance on technologies that were already in use in the plant.

As a consequence of the shift toward the high-quality varieties, demand appears to have fallen for frontline production workers (*técnicos*), who typically have a junior high school (*secundaria*) education and whose starting wage in 2002 was US$11.18/day. Demand appears to have risen for the skilled blue-collar workers who maintain robots and other automated machinery (*especialistas*), who are typically graduates of a three-year post-*secundaria* vocational school on the plant grounds and whose starting wage in 2002 was US$18/day.[12] I was unable to persuade the company to share detailed personnel data and hence am not able to make definitive statements about the changing skill composition, but conversations with both the former director of human resources and the president of the Volkswagen union confirm that the relative demand for *especialistas* rose with the shift in the product mix. At the white-collar level, it appears that the use of software engineers on the New Beetle and Jetta lines, highly skilled relative to the supervisors on the Original Beetle line, increased as well.

Does the example of Volkswagen generalize to the manufacturing sector as a whole? Figure V plots the export share of sales and the share of plants with positive exports for a balanced panel

12. The source for the wage figures is the 2002–2004 Volkswagen–Puebla collective bargaining agreement.

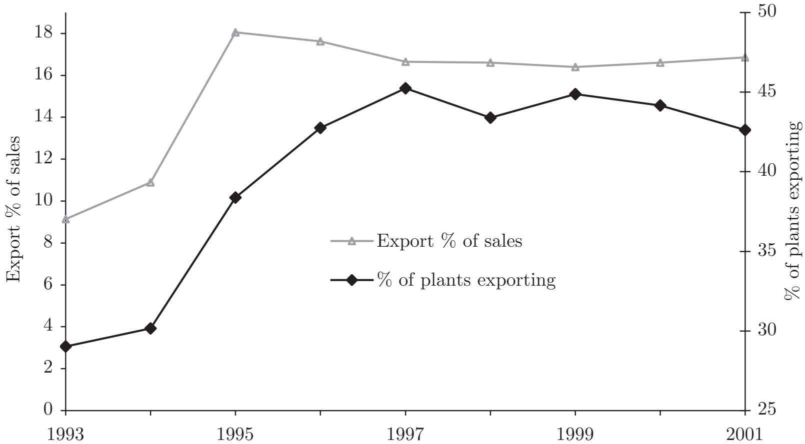

FIGURE V
Shift toward Exporting, 1993–2001

Notes: Data from EIA 1993–2001 panel. Export percentage of sales calculated as (total exports for all plants)/(total sales for all plants). Plants with exports greater than zero classified as exporting. Further details on data set in Section IV of the text and Appendix II (online).

over the period 1993–2001;[13] the shift toward exporting is evident.[14] More than 80% of Mexican exports over the period went to the United States; the increase in exports thus largely represents an increase in sales on the U.S. market. Generalizing from the Volkswagen example, it appears likely that the increase in exports to the United States was accompanied by an increase in the average quality of goods produced and an upgrading of the workforce in exporting plants.

III. Theory

To provide a framework for the empirical analysis, this section outlines a model that formalizes the quality-upgrading process as it played out at Volkswagen and, as anecdotal evidence suggests, across broad segments of the Mexican manufacturing sector. The model is partial-equilibrium, implicitly focused on a single industry that is small relative to the economy as a whole.

13. The data set is described in more detail in Section IV and in Appendix II (online).

14. It is a puzzle that the export share did not decline as the peso reappreciated after the peso crisis, but given this, we would not expect the quality-upgrading process to have reversed itself in the 1997–2001 period.

Readers interested in greater detail are referred to the working paper (Verhoogen 2007).[15]

III.A. Demand

There are two countries, North and South. In each market, indexed by $d = n, s$, there is a mass N_d of statistically identical consumers, each of whom is assumed to buy one unit of a good from a continuum of goods indexed by ω and to have the indirect utility function

$$V(\omega) = \theta_d q(\omega) - \widetilde{p}_d(\omega) + \varepsilon, \tag{1}$$

where q is product quality, assumed to be perfectly observable, and $\widetilde{p}$ is price relative to the price level in country d. The parameter θ_d captures consumers' willingness to pay for quality. It can be interpreted as a function of income: given identical direct utility functions, richer consumers have a lower marginal utility of income and are willing to pay more for a given level of quality.[16] I assume that θ_s and θ_n are constant within each country but that Northern consumers are more willing to pay for quality than Southern ones: $\theta_n > \theta_s$. I treat θ_s and θ_n as fixed parameters and abstract from changes in consumers' willingness to pay for quality arising from income changes due to the peso crisis. It will be convenient to keep track of prices relative to the price level in South, so let δ_d be the ratio of the price level in country d to the price level in South; that is, $\delta_s = 1$ and δ_n is the real exchange

15. The model is most closely related to that of Manasse and Turrini (2001), who also model heterogeneous firms producing quality-differentiated goods and frame their results in terms of wage inequality. Three features limit the usefulness of their model in this context: (1) it is not clear how to relate the utility function of their representative consumer to the choices of individual consumers, and hence not clear how to derive differences in aggregate quality demands from individual income differences; (2) product quality is a deterministic function of fixed firm characteristics, rather than a choice variable of the firm; and (3) each firm employs only one employee and the employee receives all the rents from production, with the result that it seems more natural to think of these individuals as entrepreneurs, and of dispersion of their payoffs as dispersion in profits rather than dispersion in wages.

16. If consumers have direct utility $U(\omega) = u(\kappa) + q(\omega) + \widetilde{\varepsilon}$, where κ is the consumption of an undifferentiated numeraire good, then optimization yields the indirect utility function $\widetilde{V}(\omega) = u(y_d - \tilde{p}_d(\omega)) + q(\omega) + \widetilde{\varepsilon}$. If $\tilde{p}_d(\omega)$ is small relative to the consumer's income, y_d, then a first-order expansion of the subutility function $u(\cdot)$ gives $\widetilde{V}(\omega) = u(y_d) - \tilde{p}_d(\omega)u'(y_d) + q(\omega) + \widetilde{\varepsilon}$. Let $\theta_d \equiv 1/u'(y_d)$, $V \equiv [\widetilde{V}/u'(y_d)] - [u(y_d)/u'(y_d)]$, $\varepsilon \equiv \widetilde{\varepsilon}/u'(y_d)$. We then have (1). Note that the $u(y_d)/u'(y_d)$ term in V does not affect the choice probabilities and drops out of the expression for demand, (2).

rate. The price of good ω relative to the Southern price level is then $p_d(\omega) = \delta_d \widetilde{p}_d(\omega)$.

The random consumer–product-match term, ε, is assumed to be independent and identically distributed across consumers with a type 1 extreme-value distribution,[17] a standard multinomial-logit formulation (McFadden 1974; Anderson, de Palma, and Thisse 1992). A familiar derivation yields the following expected demand for each good (Anderson, de Palma, and Thisse 1992, Theorem 2.2, p. 39):

$$x_d(\omega) = \frac{N_d \exp\left[\frac{1}{\mu}\left(\theta_d q(\omega) - \frac{p_d(\omega)}{\delta_d}\right)\right]}{\int_{\Omega_d} \exp\left[\frac{1}{\mu}\left(\theta_d q(\omega) - \frac{p_d(\omega)}{\delta_d}\right)\right] d\omega}, \tag{2}$$

where μ is a parameter of the distribution of ε that captures the degree of differentiation between goods and Ω_d is the set of goods available in consumer market d. I assume throughout that plants are risk-neutral and write demand without the expectation operator. Note that this specification of demand combines horizontal differentiation, in the sense that if the prices of all goods are equal, each will be purchased with positive probability, with vertical differentiation, in the sense that if the prices of all goods are equal, higher-quality goods will be purchased with higher probability.

III.B. Production

In each country, there is a continuum of potential entrepreneurs of mass 1, heterogeneous in an exogenously fixed productivity parameter λ, which can be interpreted as entrepreneurial ability or technical know-how. To streamline the exposition, consider only the decisions of Southern plants; the analysis for Northern plants is similar. It is convenient to think of a plant that enters both the domestic and export markets as producing on different production lines, indexed by $d = n, s$.

I assume that each unit of output carries fixed factor requirements: one white-collar worker, one blue-collar worker, and one machine. Product quality is assumed to depend on the "quality" of the two workers, the technical sophistication of the machine, and the ability of the entrepreneur, combining in Cobb–Douglas

17. That is, $F(\varepsilon) = \exp(-\exp(\gamma - \varepsilon/\mu))$, where $\gamma = 0.5772$ (Euler's constant) by assumption, to ensure that the expectation of ε is zero.

fashion,[18]

$$q_d\left(k_d, e_d^h, e_d^h; \lambda\right) = \lambda (k_d)^{\alpha^k} \left(e_d^h\right)^{\alpha^h} \left(e_d^l\right)^{\alpha^l}, \tag{3}$$

where k represents the amount of capital embodied in the machine, and e^h and e^l represent the quality of the white-collar and blue-collar worker, respectively.[19] As in the O-ring production functions of Kremer (1993) and Kremer and Maskin (1996), which this function emulates, many low-skilled workers cannot substitute for one high-skilled one, and the qualities of different workers are complementary.[20] Let $\alpha \equiv \alpha^k + \alpha^h + \alpha^l$ and assume that $\alpha < 1$. This will ensure an interior solution in the choice of product quality.

Plants face worker quality–wage schedules that are assumed to be upward-sloping and, in the interest of simplicity, linear:

$$e_d^h = z^h \left(w_d^h - \underline{w}^h\right) \tag{4a}$$

$$e_d^l = z^l \left(w_d^l - \underline{w}^l\right), \tag{4b}$$

where w_d^h and w_d^l are the wages of white-collar and blue-collar workers on a particular production line and z^h and z^l are positive constants. The variables $\underline{w}^h$ and $\underline{w}^l$ represent the average wages of white-collar and blue-collar workers in the outside labor market and are taken to be exogenous. These worker quality–wage schedules can be justified on the basis of a number of different models: a model in which worker quality represents general skill, workers are heterogeneous in skill levels within each occupational category, and plants must pay high wages to attract high-skill workers, as in Kremer (1993); a model in which worker quality represents effort and plants must offer efficiency wages in order

18. This function can be interpreted as a reduced-form relationship between factor inputs and resulting product quality. It may arise, for instance, because producing high quality requires sophisticated machinery, which in turn requires high-quality workers.

19. It is valid to interpret (3) as indicating that product quality is skill-intensive, as long as it is understood that this means intensive in worker skill or effort for given factor shares of employment, rather than a high white-collar share of employment.

20. The ideal experiment to test the assumption that high-quality products require high-quality workers would randomly assign quality requirements to different plants and then examine how plants adjust skill levels, skill shares, wages, and technology. Although this assumption is the basis for a parsimonious account of a number of patterns presented in the empirical part of this paper, in the absence of such an experiment the ultimate validation of the assumption will depend on how well the model predicts out of sample, and hence must await future empirical work. For evidence on the related point that the production of newly invented goods requires highly skilled workers, see Xiang (2005).

to induce workers to supply it (Akerlof 1982; Shapiro and Stiglitz 1984; Bowles 1985); or a model in which worker quality represents plant-specific skills and workers bargain for a share of the gains to investments in those skills (Hashimoto 1981). For present purposes, the important point is that worker quality improves product quality and is costly to the plant to acquire.[21]

The rental cost of capital is ρ. There is a fixed cost to each plant of entering its domestic market and an additional fixed cost of entering the export market. The combination of constant marginal cost and the fixed cost of entry generates increasing returns to scale. There is no cost to differentiation and plants are constrained to offer just one variety. As a consequence, all plants differentiate and have a monopoly in the market for their particular variety.

III.C. Plants' Optimization

Each plant chooses the white-collar wage, the blue-collar wage, capital intensity, and output price to maximize profits, separately for each production line. The input decisions determine quality; quality and price pin down demand and hence output. As is standard in monopolistic competition models, each plant thinks of itself as small relative to the market as a whole and treats the denominator in (2) as unaffected by its own choices. Given this assumption, optimization yields the following:

$$q_d^*(\lambda) = \left(\eta\lambda\delta_d^\alpha\theta_d^\alpha\right)^{1/(1-\alpha)} \tag{5a}$$

$$w_d^{h*}(\lambda) = \underline{w}^h + \alpha^h\delta_d\theta_d q_d^*(\lambda) \tag{5b}$$

$$w_d^{l*}(\lambda) = \underline{w}^l + \alpha^l\delta_d\theta_d q_d^*(\lambda) \tag{5c}$$

$$k_d^*(\lambda) = \frac{\alpha^k}{\rho}\delta_d\theta_d q_d^*(\lambda) \tag{5d}$$

$$p_d^*(\lambda) = \mu\delta_d + \underline{w}^h + \underline{w}^l + \alpha\delta_d\theta_d q_d^*(\lambda), \tag{5e}$$

where $\eta \equiv (z^h\alpha^h)^{\alpha^h}(z^l\alpha^l)^{\alpha^l}(\alpha^k/r)^{\alpha^k}$ is a constant.

These equations carry several implications. First, all else equal, higher-λ plants produce higher-quality goods, pay higher wages to both white-collar and blue-collar workers, are more capital-intensive, and charge higher prices than lower-λ plants.

21. The implications of the three interpretations are examined in matched employer–employee data in Kaplan and Verhoogen (2006).

It also follows directly that both output and profits are increasing in λ.

Second, if a plant enters both markets, then it chooses greater quality, prices, wages, and capital intensity for goods sold in the North than for goods sold in the South because $\theta_n > \theta_s$.[22]

Third, all else equal, plant size and wages are positively correlated because both are increasing in λ. The model thus provides a natural explanation for the employer size-wage effect, documented by Brown and Medoff (1989) and others.

Fourth, all else equal, prices and quality are positively correlated with plant size because all are increasing in λ.

Fifth, whether the ratio of the white-collar wage to the blue-collar wage is increasing or decreasing in quality, and hence in λ and θ_d, depends on the sensitivity of product quality to the quality of each type of worker in (3). Kremer and Maskin (1996) hypothesize that production is more sensitive to the skill of white-collar workers. In the current context, if product quality is sufficiently more sensitive to the skill of white-collar workers, then the wage ratio will be increasing in both λ and θ_d.[23]

Finally, the fact that profitability is increasing in λ implies that in equilibrium there will be a cut-off value for each destination, $\lambda_d^{\min}$, above which all plants will enter and earn positive profits, and below which no plants will enter. The cut-off is determined by the condition that the marginal plant has zero profits after paying the fixed cost of entry to the destination market, as in Melitz (2003).

Figure VI summarizes the cross-sectional relationship between quality and λ. The dotted $\overline{q}^*(\lambda)$ curve represents average quality if all plants were to enter both markets, a weighted average of quality on the domestic production line, $q_s^*(\lambda)$, and quality on the export production line, $q_n^*(\lambda)$, with the weights given by the export share of output of each plant. In fact, only plants above

22. A single plant will produce different qualities for different markets even in the absence of the quality bias due to trade quotas or per-unit trade costs explored by Feenstra (1988) and Hummels and Skiba (2004).

23. From (5b) and (5c),

$$\frac{\partial}{\partial \lambda}\left(\frac{w_d^{h*}(\lambda)}{w_d^{l*}(\lambda)}\right) = \frac{\theta_d \delta_d}{\left(w_d^{l*}(\lambda)\right)^2}\left(\alpha^h \underline{w}^l - \alpha^l \underline{w}^h\right)\frac{\partial q_d^*(\lambda)}{\partial \lambda}.$$

A similar result holds for $\partial/\partial\theta_d(w_d^{h*}(\lambda)/w_d^{l*}(\lambda))$. If $\alpha^h/\alpha^l > \underline{w}^h/\underline{w}^l$ then the wage ratio, $w_d^h(\lambda)/w_d^l(\lambda)$, is increasing in λ and θ_d.

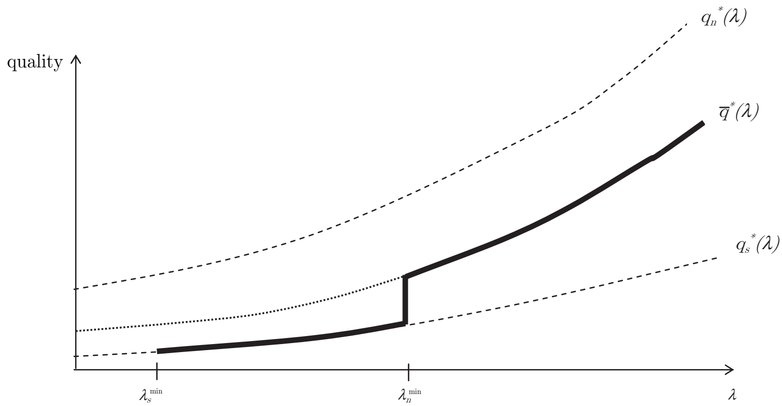

FIGURE VI
Average Quality

$\lambda_n^{\min}$ enter the export market; the solid curve represents actual average quality as a function of λ, taking into account entry patterns. In practice it is rare to have data by production line within plants. In the Mexican data, we observe plant-level averages, analogous to the solid curve. As equations (5a)–(5e) indicate, the model predicts cross-sectional patterns similar to the solid curve for observed white-collar wages, blue-collar wages, and capital intensity.

III.D. Discussion of Effects of Exchange-Rate Devaluation

In the context of this model, we can think of the devaluation and the ensuing recession as having two effects: an increase in the real exchange rate, δ_n, and a decline in the number of domestic consumers, N_s. Under plausible conditions, the shock has the effects illustrated by Figure VII.[24] Quality on the domestic production line, $q_s^*(\lambda)$ does not depend on either δ_n or N_s and hence is unaffected by the exchange-rate shock. Quality on the export line, $q_n^*(\lambda)$, increases to $q_n^*(\lambda)'$, because the peso devaluation reduces Southern plants' cost of producing quality relative to Northern demand. The dotted average quality line, $\overline{q}^*(\lambda)$, shifts up to $\overline{q}^*(\lambda)'$ with the crisis, both because the $q_n^*(\lambda)$ curve shifts up and because

24. Refer to the working paper (Verhoogen 2007) for a discussion of the conditions.

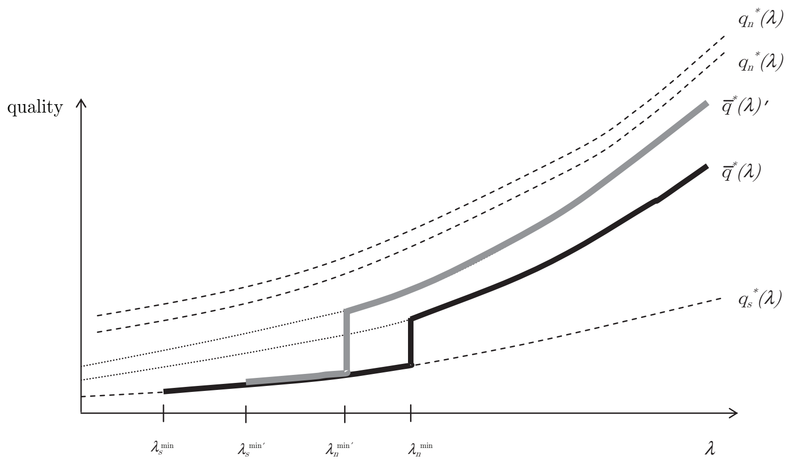

FIGURE VII
Response to Exchange-Rate Devaluation

the export share of output increases. The cut-off value for entry into the export market shifts to the left.[25]

Figure VIII depicts the difference between the two solid lines in Figure VII, the change in observable quality taking into account entry patterns. The plants that switch into exporting (between $\lambda_n^{min'}$ and λ_n^{min}) see an especially large increase in average quality. Average quality is increasing in λ within the category of switchers and within the category of always exporters (to the right of λ_n^{min}). As suggested by equations (5a)–(5e), the model predicts a similar pattern for white-collar wages, blue-collar wages, and capital intensity.[26] These are the implications to be taken to data in the remainder of the paper.

IV. DATA

The results in this paper are primarily based on the EIA, a yearly panel survey conducted by the *Instituto Nacional de*

25. One might also imagine that the devaluation would have an effect on the outside wage terms, $\underline{w}^h$ and $\underline{w}^l$. In the empirical section, such changes will be absorbed by industry–year effects (or industry effects when the model is written in changes, as in (6)); the important point is that one would not expect such changes to generate differential changes between high-λ and low-λ plants.

26. If product quality is sufficiently more sensitive to the quality of white-collar than of blue-collar workers (refer to footnote 23), then the wage ratio will follow a similar pattern. By assumption, there is no such prediction for the white-collar share of employment.

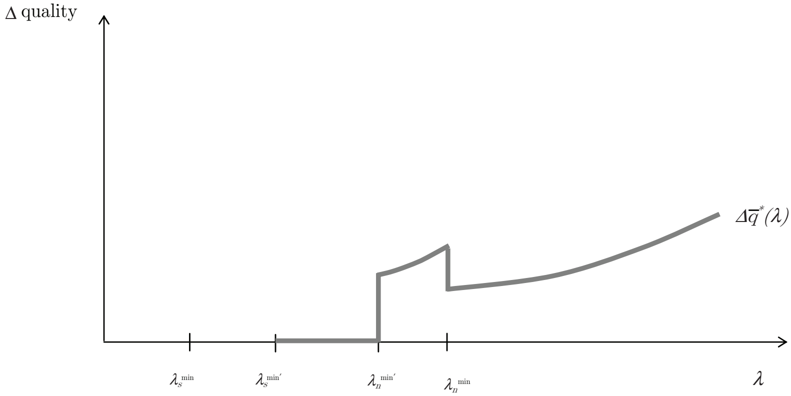

FIGURE VIII
Change in Average Quality in Response to Devaluation

Estadísticas, Geografía, e Información (INEGI), the Mexican government statistical agency. The EIA is based on a deterministic sample of the largest plants in 205 of the 309 six-digit manufacturing industries in the Mexican industrial classification system. The EIA survey excludes *maquiladoras*, assembly plants that participate in a Mexican government export-promotion program.[27] I constructed two balanced panels from the EIA: the EIA 1993–2001 panel, which contains 3,263 plants, and the EIA 1984–2001 panel, which contains fewer plants—1,114—but over a longer period.[28]

INEGI carried out a more qualitative plant survey, the *Encuesta Nacional de Empleo, Salarios, Tecnología y Capacitacíon* (ENESTyC) [National Survey of Employment, Wages, Technology and Training] in 1992, 1995, 1999, and 2001, with questions referring to the previous year. In 1995, 1999, and 2001 the ENESTyC elicited information on ISO 9000 certification, an international production standard. Although ISO 9000 is mainly a procedural standard, the common view among Mexican managers is that ISO 9000 is a signal of high product quality, and I take it as such for the

27. In Mexico, the participants in this program are referred to as *maquiladoras de exportación* (exporting *maquiladoras*). The word *maquiladora* is used more generally to apply to any plant producing under subcontract. I use the term *maquiladora* only to refer to the former group.

28. I also constructed an unbalanced panel for the period 1993–2001, which I refer to as the EIA 1993–2001 Unbalanced Panel; refer to footnote 40 and the data construction details in Appendix II (online).

purposes of this paper.[29] I refer to the 844 plants that appear in both the EIA and the ENESTyC and that have data on ISO 9000 certification in 1995, 1999, and 2001 as the EIA-ENESTyC panel.

Variable definitions are included in Appendix I. Details on the processing of these data as well as of the other data sets used in this paper, including the *Encuesta Nacional de Empleo Urbano* (ENEU) [National Urban Employment Survey] and the *Estadísticas Mensuales de la Industria Maquiladora de Exportación* (EMIME) [Monthly Statistics on Maquiladora Export Industry], are included in Appendix II (online).

Panel A of Table I reports summary statistics for the EIA 1993–2001 panel for the initial year, 1993, separately by export status.[30] As observed by Bernard and Jensen (1999) in U.S. data, there are systematic differences between exporters and nonexporters in cross section: exporters are larger in terms of employment and total sales, are more capital-intensive, and pay higher wages. As observed by Aw and Batra (1999) in Taiwanese data, exporters have a higher white-collar/blue-collar wage ratio. Exporters have greater domestic sales, consistent with the theoretical model. They also have a larger share of imported inputs. Panel B of Table I reports means by export status for relevant variables from the EIA-ENESTyC panel. We see further that exporters are more likely to have ISO 9000 certification, hire white-collar and blue-collar workers with more years of schooling, and have lower rates of absenteeism, accidents, and turnover.[31]

V. Estimation

V.A. *Econometric Strategy*

The main implication of the theoretical model is that the pattern illustrated by Figure VIII should hold for a number of observable variables—export share, white-collar wage, blue-collar wage, capital intensity, and ISO 9000 certification—to a greater extent during the peso crisis period than in periods without

29. ISO 9000 certification is not cheap talk. Obtaining certification typically takes between nine months and two years and costs $187,000 (1996 U.S. dollars) on average (Guler, Guillen, and Macpherson 2002).

30. The EIA 1984–2001 panel contains a greater share of large plants than the EIA 1993–2001 panel, but the qualitative differences between nonexporters and exporters are similar to those reported here.

31. These differences are significant at the 5% level for training, white-collar schooling, and absentee rate but not for ISO 9000 certification, blue-collar schooling, the accident rate, or the turnover rate.

TABLE I
SUMMARY STATISTICS BY EXPORT STATUS

	Nonexporters	Exporters	All
A. EIA 1993–2001 Panel, 1993			
Employment	182.39	333.76	226.32
	[4.94]	[11.98]	[5.08]
Revenues	42.24	89.84	56.05
	[1.75]	[4.59]	[1.86]
Domestic sales	41.17	70.16	49.58
	[1.71]	[3.59]	[1.61]
K/L ratio	42.58	55.59	46.36
	[1.40]	[2.60]	[1.25]
White-collar hourly wage	20.53	28.40	22.81
	[0.27]	[0.50]	[0.25]
Blue-collar hourly wage	8.04	9.64	8.50
	[0.08]	[0.15]	[0.07]
White-collar employment share	0.30	0.33	0.31
	[0.004]	[0.005]	[0.003]
Export share of sales		0.16	
		[0.01]	
Import share of input expenditures	0.14	0.30	0.19
	[0.01]	[0.01]	[0.005]
Share with foreign ownership	0.09	0.30	0.15
	[0.01]	[0.01]	[0.01]
N	2,316	947	3,263
B. EIA-ENESTyC Panel			
Employment	308.37	414.29	348.91
	[12.18]	[21.08]	[11.16]
Share with ISO 9000 certification	0.09	0.12	0.10
	[0.01]	[0.02]	[0.01]
Share with formal training program	0.69	0.79	0.73
	[0.02]	[0.02]	[0.02]
Avg. schooling, white-collar	12.04	12.47	12.20
	[0.08]	[0.10]	[0.06]
Avg. schooling, blue-collar	7.26	7.38	7.30
	[0.07]	[0.09]	[0.05]
Absentee rate	1.41	1.20	1.33
	[0.07]	[0.08]	[0.05]
Turnover rate	74.70	66.62	71.56
	[3.66]	[4.44]	[2.83]
Accident rate	5.01	4.63	4.87
	[0.29]	[0.31]	[0.21]

Notes. Standard errors of means in brackets. Exporter means export sales >0. Has foreign ownership means FDI share >0. Data on FDI from 1994; other data in Panel A from EIA 1993–2001 panel for 1993. Revenues, domestic sales measured in millions of 1994 pesos, K/L ratio in thousands of 1994 pesos, wages in 1994 pesos per hour, employment in number of workers, employment share in hours. Average 1994 exchange rate: 3.38 pesos/U.S.$1. Data on ISO 9000, training, accidents from ENESTyC 1995, reporting for 1994; data on schooling from ENESTyC 1992, reporting for 1991. Numbers of observations for nonexporters (exporters) are 521 (323) for employment and ISO 9000, 366 (224) for white-collar schooling, 367 (223) for blue-collar schooling, 510 (318) for accident rate, 317 (198) for absentee rate, 459 (292) for turnover rate; otherwise as reported above. Further variable definitions in Appendix I. Further details on data sets in Section IV of the text and Appendix II (online).

an exchange-rate devaluation. A central econometric challenge in testing these implications is that a key variable, entrepreneurial ability λ, is unobserved. This section presents an econometric strategy that uses simple observable proxies for this parameter. I have also estimated a factor-analytic model that estimates the key parameters for all periods simultaneously by maximum likelihood, without the need to construct a proxy for λ, at the cost of stronger distributional assumptions. The results are consistent with those presented here; interested readers are referred to the working paper for details (Verhoogen 2007).

The theoretical model suggests a number of variables that are correlated with λ in cross section and hence that are candidates to be proxies. My preferred proxy is log domestic sales, deviated from industry means. The main argument for this proxy is that sales is the only variable that is observed separately by production line. In cross section, domestic sales thus bears a smooth, continuous relationship to λ, without the discontinuity at the cutoff for entry into the export market. The domestic sales variable has the additional advantage that it is relatively well measured. I present results for a variety of alternative proxies below.

The discontinuous, nonlinear function depicted in Figure VIII is unlikely to hold exactly in the data, both because of the noise in domestic sales (and the other proxies) and because the costs of entering the export market are likely to be heterogeneous across plants and industries. Rather than attempting to estimate that precise curve, I approximate it with a linear function. In the absence of background trends between higher-λ and lower-λ plants, we would expect a positively sloped line during the peso-crisis period and a horizontal line in other periods. In practice such background trends are likely to exist, however, so the applicable prediction is that the slope of changes in the observable variables—export share, white-collar wage, blue-collar wage, capital intensity, and product quality—against λ will be greater in periods with major devaluations than in periods without.

The main estimating equation is

$$\triangle y_{ijr} = \alpha + \widetilde{\lambda}_{ijr}\beta + \psi_j + \xi_r + u_{ijr}, \tag{6}$$

where i, j, and r index plants, industries, and states, respectively; $\triangle y_{ijr}$ is a change in one of the outcome variables; α is an intercept term; $\widetilde{\lambda}_{ijr}$ is the value of the entrepreneurial ability proxy in the initial year; ψ_j is an industry fixed effect; ξ_r is a state fixed effect;

and u_{ijr} is a mean-zero disturbance. I estimate this equation separately by period and compare the coefficient estimates $\widehat{\beta}$ across periods. When using the EIA 1993–2001 panel, I estimate for the periods 1993–1997 and 1997–2001. When using the EIA 1984–2001 panel (for which exports are available beginning in 1986), I use the periods 1986–1989, 1989–1993, 1993–1997, and 1997–2001. The theoretical predictions are that $\widehat{\beta}_{1993-1997} > \widehat{\beta}_{1997-2001}$, $\widehat{\beta}_{1993-1997} > \widehat{\beta}_{1989-1993}$, and $\widehat{\beta}_{1986-1989} > \widehat{\beta}_{1989-1993}$ when the dependent variable is the change in the export share, white-collar wage, blue-collar wage, capital intensity, or ISO 9000 certification. Note that if there were just two groups of plants, high-$\tilde{\lambda}$ and low-$\tilde{\lambda}$, then this strategy would amount to a familiar triple-differences strategy: $\widehat{\beta}_{1993-1997}$ would reflect the difference in differences between high-$\tilde{\lambda}$ and low-$\tilde{\lambda}$ plants from 1993 to 1997, $\widehat{\beta}_{1997-2001}$ the difference in differences from 1997 to 2001, and $\widehat{\beta}_{1993-1997} - \widehat{\beta}_{1997-2001}$ the difference in difference in differences.[32]

The analogy with triple-difference designs highlights a potential pitfall in the estimation. If we use information from many years of pre- or post-crisis data without taking into account serial correlation across periods, then we may overstate the amount of independent variation in the pretreatment and posttreatment periods and understate the standard errors on the coefficient estimates, as discussed by Bertrand, Duflo, and Mullainathan (2004). My strategy is to select just one year of data precrisis and one year postcrisis, i.e., 1993 and 1997 for the peso crisis period, 1997 and 2001 for the later placebo period without an exchange-rate shock. Bertrand, Duflo, and Mullainathan (2004) refer to this strategy as "ignoring time-series information" and find that it performs reasonably well in a Monte Carlo study.[33]

A potential concern with regression (6) when the change in export share appears on the left-hand side and initial log domestic sales on the right is that domestic sales appear in the denominator of the export share in the initial period and hence any

32. There is also an instrumental-variables interpretation of (6). One could think of $\widetilde{\lambda}_{ijr}$ as an instrument for the change in exports, which could then be used to estimate the relationship between exporting and plant-level behavior. The danger with this interpretation is that the relationship cannot be considered causal: a plant's decision to enter the export market and its decision to upgrade quality are simultaneous outcomes of the plant's solution to a single underlying optimization problem. If we think of the estimation strategy as investigating the reduced-form relationship between initial level of productivity and both the change in exports and the average quality of goods produced, then this danger is avoided and the interpretation is clear.

33. I have varied the years used as initial and final years of each period; the basic results are not sensitive to these changes.

measurement error in domestic sales will generate a mechanical positive bias in the estimate of β. Under the assumption that the measurement error in domestic sales is uncorrelated over time, an IV strategy instrumenting log domestic sales with its lag will correct this bias.[34] A disadvantage of this IV strategy is that it requires two years of predevaluation data. In the case of the EIA 1993–2001 panel, this means using 1993 as the lagged value and looking at changes over the 1994–1997 period. It may be that in 1994 plants were already responding to NAFTA, which had been implemented at the beginning of the year. For this reason, I present both the OLS and the IV results below.

It is important to consider the possibility that plants are subject to time-varying productivity shocks in addition to the time-invariant level of productivity represented by λ. It may be that the productivity improvements are correlated with λ if, for instance, new technologies periodically become available to all plants but more technologically advanced plants have stronger incentives to adopt them, as in Aghion et al. (2005). If so, and if the productivity improvements also affect the dependent variable, $\triangle y$, as one would expect, then the OLS estimate of β will be inconsistent. This problem is not addressed by instrumenting with lagged domestic sales. Note, however, that as long as the distribution of the productivity shocks and the effect of the shocks on the dependent variable are constant over time, the *difference* in OLS coefficients (e.g., $\widehat{\beta}_{1993-1997} - \widehat{\beta}_{1997-2001}$) will be consistent. This underlines the need to have a "control" period without a devaluation against which to compare the "treatment" period of the peso crisis. The treatment period–control period comparison can difference out any such differential background trends between higher- and lower-λ plants.[35]

V.B. Results

Panel A of Table II reports cross-sectional regressions of six different outcome variables on log domestic sales in 1993 (and

34. The assumption that the measurement error in domestic sales is uncorrelated over time is admittedly quite strong, and for this reason the IV results should be treated with caution.

35. A related point is that caution is warranted in drawing causal inferences from regressions of, for instance, changes in wages on changes in the export share or other measures of export status, as in Bernard and Jensen (1997) or Bustos (2005). The theoretical model suggests that a positive productivity shock will generate both an increase in exports and an increase in wages, giving rise to a positive omitted-variables bias in the coefficient on the change in exports.

TABLE II
BASELINE ESTIMATES, EIA 1993–2001 PANEL

		Export share of sales (1)	log(white-collar wage) (2)	log(blue-collar wage) (3)	log(wage ratio) (4)	log(K/L ratio) (5)	White-collar emp. share (6)
A. Cross-sectional regressions, 1993							
	log domestic sales, 1993	−0.001	0.209***	0.133***	0.075***	0.343***	0.010***
		[0.003]	[0.008]	[0.006]	[0.008]	[0.017]	[0.002]
	R^2	0.220	0.391	0.358	0.185	0.370	0.343
B. Differential changes, 1993–1997 and 1997–2001		Δ (export share of sales) (1)	Δ log(white-collar wage) (2)	Δ log(blue-collar wage) (3)	Δ log(wage ratio) (4)	Δ log(K/L ratio) (5)	Δ (white-coll. emp. share) (6)
OLS regressions							
1993–1997	log domestic sales, 1993	0.020***	0.072***	0.036***	0.036***	0.083***	−0.002
		[0.002]	[0.008]	[0.006]	[0.009]	[0.011]	[0.002]
	R^2	0.173	0.15	0.129	0.09	0.134	0.111
1997–2001	log domestic sales, 1997	0.007***	0.016**	0.008	0.008	0.026***	−0.001
		[0.002]	[0.007]	[0.005]	[0.007]	[0.009]	[0.001]
	R^2	0.123	0.088	0.092	0.075	0.107	0.102
Difference (1993–1997 vs. 1997–2001)		0.014***	0.056***	0.028***	0.028**	0.057***	−0.002
		[0.003]	[0.010]	[0.007]	[0.011]	[0.014]	[0.002]
IV regressions							
1994–1997	log domestic sales, 1994	0.014***	0.058***	0.033***	0.026***	0.058***	0.000
		[0.002]	[0.007]	[0.005]	[0.008]	[0.009]	[0.002]
	R^2	0.161	0.148	0.118	0.093	0.119	0.092
1998–2001	log domestic sales, 1998	0.004**	0.005	0.004	0.001	0.016**	−0.001
		[0.002]	[0.006]	[0.004]	[0.007]	[0.008]	[0.001]
	R^2	0.111	0.082	0.097	0.077	0.102	0.099
Difference (1993–1997 vs. 1997–2001)		0.010***	0.053***	0.029***	0.024**	0.042***	0.001
		[0.003]	[0.009]	[0.007]	[0.010]	[0.012]	0.002

Notes. Table reports coefficients on log domestic sales for 30 separate regressions. (Covariate at left; dependent variables at top, with changes in Panel B over period at left.) All regressions include 205 industry (six-digit) and 32 state dummies. IV regressions instrument log domestic sales in 1994 and 1998 with values from previous year. $N = 3,263$ for all regressions. Variable definitions in Appendix I. Further details on data set in Section IV of the text and Appendix II (online). Robust standard errors in brackets. Standard errors on differences allow for cross-equation correlation. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

industry and region effects), using plants in the EIA 1993–2001 panel. As expected, white-collar wages, blue-collar wages, and capital intensity are positively correlated with the λ proxy within industries. The coefficient in the export share equation is negative, but this appears to be due to a mechanical bias generated by measurement error in domestic sales, negative in this case; when log domestic sales in 1994 are instrumented with their lag, the coefficient is 0.005 and significant at the 5% level.[36]

Panel B of Table II reports estimates of (6) for the same six outcome variables, estimated separately for 1993–1997 and 1997–2001. In the OLS results, the coefficient on log domestic sales is significantly larger in 1993–1997 than in 1997–2001 for the export share, white-collar wages, blue-collar wages, the wage ratio, and capital intensity. The magnitudes indicate economically significant effects. Consider two plants, one of which is 10% larger in log domestic sales terms than the other. Column (2) indicates that the larger plant had an approximately 0.72% greater wage increase for white-collar workers than the smaller plant during the 1993–1997 period, and a 0.16% greater increase during 1997–2001. Note that there is a significant difference in differential changes in the log wage ratio of white-collar to blue-collar workers, consistent with the hypothesis that product quality is more sensitive to the skill of white-collar than blue-collar workers. There is no evidence of differential trends in the ratio of white-collar hours to total hours; the coefficients are precisely estimated and essentially zero in both periods. The IV estimates, instrumenting initial log domestic sales with its lag, are of smaller magnitude than the OLS estimates, in part a consequence of the fact that the dependent variables are changes over a shorter period,[37] but the differences in slopes between the two periods for the wage variables and capital intensity are robust.

Table III presents OLS results for the EIA 1984–2001 panel, which allows us to look before the peso crisis period as well as after. As was illustrated by Figure III, the peso depreciated sharply in 1985–1987 and subsequently reappreciated over the period

36. Nonparametric cross-sectional regressions of these outcome variables against log domestic sales (deviated from industry means) suggest that the assumption of linearity in (6) is not unreasonable, at least for the observable wage variables that are the outcomes of primary interest. See Figures A4–A5 in Appendix II (online).

37. In the case of the export share, the IV procedure also removes the mechanical positive bias arising from the fact that domestic sales appear in the denominator.

TABLE III
ESTIMATES FROM EIA 1984–2001 PANEL

		Δ (export share of sales) (1)	Δ log(white-collar wage) (2)	Δ log(blue-collar wage) (3)	Δ log(wage ratio) (4)	Δ log(K/L ratio) (5)	Δ (white-coll. emp. share) (6)
1986–1989	log domestic sales, 1986	0.011***	0.047***	0.020***	0.023**	0.040***	−0.006**
		[0.002]	[0.010]	[0.008]	[0.011]	[0.015]	[0.003]
	R^2	0.185	0.182	0.254	0.194	0.208	0.207
1989–1993	log domestic sales, 1989	0.007**	0.004	0.006	0.002	0.076***	−0.001
		[0.003]	[0.011]	[0.009]	[0.013]	[0.027]	[0.003]
	R^2	0.219	0.173	0.211	0.167	0.179	0.195
1993–1997	log domestic sales, 1993	0.013***	0.067***	0.021**	0.045***	0.092***	0.002
		[0.003]	[0.012]	[0.008]	[0.012]	[0.017]	[0.003]
	R^2	0.277	0.272	0.236	0.188	0.266	0.219
1997–2001	log domestic sales, 1997	0.007**	0.002	0.009	−0.009	0.014	−0.001
		[0.003]	[0.009]	[0.008]	[0.011]	[0.014]	[0.002]
	R^2	0.200	0.193	0.217	0.163	0.188	0.146

Notes. Table reports coefficients on log domestic sales for 24 separate regressions. (Covariate at left; dependent variables at top, with changes over period at left.) All regressions include 205 industry (six-digit) and 32 state dummies (coefficients omitted). $N = 1{,}114$ for all regressions. Variable definitions in Appendix I. Further details on data set in Section IV of the text and Appendix II (online). Robust standard errors in brackets. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

1987–1993. Although the aggregate export response to the 1985–1987 devaluation was smaller than that to the peso crisis, we would nonetheless expect to see greater differential trends between higher- and lower-λ plants in the periods 1986–1989 and 1993–1997 than in the period 1989–1993. The coefficients in the export share equations are just marginally distinguishable, and there is an anomalous result for capital intensity, but the important message of the table is that, for the key wage outcomes, the earlier devaluation period (1986–1989) resembles the peso-crisis period (1993–1997), and the 1989–1993 period of reappreciation resembles the post-crisis reappreciation period (1997–2001). This particular time pattern of changes points to exchange-rate movements as a driving force of the differential wage changes.

Table IV presents results for additional variables available only in the EIA-ENESTyC panel. Panel A presents cross-sectional regressions similar to those of Panel A of Table II. There is strong evidence of cross-sectional correlation between log domestic sales and indicators that are plausibly associated with high-quality production. Controlling for industry and region effects, larger plants are significantly more likely to have ISO 9000 certification, to have a formal worker training program, to employ white-collar and blue-collar workers with more schooling, and to have lower turnover, accident, and absentee rates than smaller plants. Panel B reports estimates of (6) for these outcome variables. For ISO 9000 certification there is strong evidence of differential trends between initially smaller and initially larger plants, and the difference in differential trends across periods is marginally significant. There is also evidence of a greater differential increase in the average schooling of blue-collar workers in the earlier period.[38] There is no strong evidence of differential changes in the other variables. One interpretation of these results is that quality upgrading requires changing the composition of the blue-collar workforce but not other workplace practices. Another plausible explanation, however, is simply that the other workplace-practice variables are too poorly measured to pick up subtle changes in plant behavior. The results using the ENESTyC variables should be treated with caution—ISO 9000 is not an ideal measure of product quality, sample sizes are relatively small, and the intervals between waves of the ENESTyC are of different lengths, among other reasons—but

38. It is worth emphasizing, however, that the schooling measures are based on managers' estimates of the number of workers in broad schooling categories, and are likely to be quite noisy.

TABLE IV
ESTIMATES FROM EIA-ENESTyC PANEL

		A. Cross-sectional regressions, 1993						
		ISO 9000 certification (1)	White-collar avg. schooling (2)	Blue-collar avg. schooling (3)	Has formal training (4)	Turnover rate (5)	Accident rate (6)	Absentee rate (7)
	Log domestic sales, 1993	0.023**	0.286***	0.156***	0.049***	−20.239***	−0.802***	−0.250***
		[0.011]	[0.067]	[0.058]	[0.017]	[2.995]	[0.216]	[0.044]
	N	844	590	590	843	751	828	515
	R^2	0.154	0.258	0.240	0.117	0.168	0.*206	0.245
		B. Differential Responses, 1993–1997 and 1997–2001						
		Δ ISO 9000 certification (1)	Δ white-collar avg. schooling (2)	Δ blue-collar avg. schooling (3)	Δ has formal training (4)	Δ turnover rate (5)	Δ accident rate (6)	Δ absentee rate (7)
1993–1997	Log domestic sales, 1993	0.079***	−0.105	0.204***	0.008	1.067	0.219	−0.025
		[0.018]	[0.104]	[0.078]	[0.020]	[4.224]	[0.247]	[0.093]
	R^2	0.171	0.164	0.194	0.1	0.184	0.141	0.243
1997–2001	Log domestic sales, 1997	0.036***	0.058	−0.023	−0.024	−4.294	0.045	−0.140
		[0.015]	[0.088]	[0.075]	[0.017]	[4.655]	[0.222]	[0.093]
	R^2	0.127	0.151	0.173	0.082	0.161	0.134	0.138
Difference (1993–1997 vs. 1997–2001)		0.042*	−0.163	0.228**	0.032	5.361	0.174	0.115
		[0.024]	[0.136]	[0.109]	[0.026]	[6.286]	[0.332]	[0.131]
	N	844	484	484	836	513	713	354

Notes. Table reports coefficients on log domestic sales for 21 separate regressions. (Covariate at left; dependent variables at top, with changes in Panel B over period at left.) All regressions include dummies for 50 industries (four-digit) and 32 states. Data on ISO 9000, training, turnover rate, accident rate, absentee rate from 1994, 1998, 2000; on schooling from 1991, 1998, 2000. Since requiring plants to have complete data on all variables would have reduced the panel prohibitively, I allow the sample size to change across columns. Variable definitions in Appendix I. Further details on data set in Section IV of the text and Appendix II (online). Robust standard errors in brackets. Standard errors on differences allow for cross-equation correlation. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

the cross-sectional patterns and the difference in differential trends in ISO 9000 certification and blue-collar schooling are corroborative evidence for the quality-upgrading hypothesis.

Although the domestic sales proxy has much to recommend it, it can be criticized on the grounds that plant size may reflect a number of factors that are unrelated to entrepreneurial ability or productivity. Table V reports estimates of (6) for the export share, white-collar wages, blue-collar wages, and the wage ratio using a variety of alternative proxies that theory predicts will be correlated with λ: log employment; predicted export share; predicted ISO 9000 certification; the first principal component of a number of variables hypothesized to be correlated with λ; total factor productivity; log domestic sales per worker; and actual export share. Details on the construction of each of these proxies are in Appendix I. Unsurprisingly, they tend to be highly correlated within industries; Table A1 in Appendix II (online) reports bivariate correlations among them. The results for the export share, white-collar wages, and blue-collar wages are generally quite robust to the choice of proxy: the difference in coefficients is at least marginally significant in all but one case.[39] The results for the wage ratio are somewhat less robust but are generally consistent with those using the domestic-sales proxy in Table II.[40]

VI. Alternative Hypotheses

The identification strategy of this paper, like other triple-difference-type research designs, relies on the assumption that there are no other factors generating a difference in differential trends, in this case between high-λ and low-λ plants. But the peso crisis had many effects on the Mexican economy, and one might argue that some of them generated such a difference in differential trends. The leading candidate is the banking crisis and general contraction of credit that followed the exchange-rate devaluation.

39. Note that in any given year more than half of plants have zero exports, and there is variation in actual export share only among initially exporting plants. This may explain the nonrobustness of the estimate for white-collar wages when using actual export share as the λ proxy.

40. Using an unbalanced panel from the EIA for 1993–2001, I have also estimated a selection–correction model to correct for endogenous exit; see Table A2 in Appendix II (online). The model is identified on the basis of functional form assumptions, rather than an excluded instrument, and the estimates should be treated with caution, but it is nonetheless reassuring that they correspond closely to the results using the EIA 1993–2001 balanced panel.

TABLE V
ESTIMATES USING ALTERNATIVE PROXIES, EIA 1993–2001 PANEL

	Δ (export share of sales)			Δ log(white-collar wage)			Δ log(blue-collar wage)			Δ log(wage ratio)		
	1993–1997 (1)	1997–2001 (2)	Difference (3)	1993–1997 (4)	1997–2001 (5)	Difference (6)	1993–1997 (7)	1997–2001 (8)	Difference (9)	1993–1997 (10)	1997–2001 (11)	Difference (12)
Proxy for λ:												
Log employment	0.017***	0.004*	0.013***	0.086***	0.013	0.073***	0.050***	0.018***	0.033***	0.036***	−0.005	0.040***
	[0.003]	[0.002]	[0.004]	[0.010]	[0.009]	[0.014]	[0.007]	[0.006]	[0.010]	[0.012]	[0.009]	[0.015]
R^2	0.156	0.12		0.147	0.087		0.131	0.093		0.088	0.075	
Predicted exp. share index	0.017***	0.003	0.014***	0.080***	0.015**	0.066***	0.048***	0.015***	0.033***	0.031***	0.000	0.031***
	[0.002]	[0.002]	[0.003]	[0.008]	[0.007]	[0.011]	[0.006]	[0.005]	[0.008]	[0.009]	[0.008]	[0.012]
R^2	0.162	0.119		0.154	0.087		0.137	0.093		0.088	0.075	
Predicted ISO 9000 index	0.008***	−0.005**	0.013***	0.088***	0.011	0.078***	0.024***	0.003	0.021**	0.061***	0.008	0.053***
	[0.003]	[0.002]	[0.003]	[0.009]	[0.008]	[0.012]	[0.007]	[0.006]	[0.009]	[0.010]	[0.008]	[0.013]
R^2	0.148	0.121		0.151	0.087		0.122	0.091		0.094	0.075	
First principal component	0.013***	−0.001	0.015***	0.084***	0.020***	0.064***	0.044***	0.010*	0.034***	0.039***	0.009	0.030**
	[0.002]	[0.002]	[0.003]	[0.008]	[0.007]	[0.011]	[0.006]	[0.005]	[0.008]	[0.009]	[0.008]	[0.012]
R^2	0.156	0.119		0.156	0.088		0.134	0.092		0.09	0.075	
TFP (fixed effects)	0.013***	0.002	0.011***	0.060***	0.013*	0.047***	0.029***	0.002	0.027***	0.031***	0.01	0.021
	[0.002]	[0.002]	[0.003]	[0.009]	[0.008]	[0.012]	[0.006]	[0.006]	[0.009]	[0.010]	[0.009]	[0.013]
R^2	0.156	0.119		0.144	0.087		0.125	0.091		0.088	0.075	
Log (dom sales/worker)	0.029***	0.012***	0.017***	0.059***	0.022**	0.036**	0.017*	−0.006	0.023*	0.041***	0.027**	0.013
	[0.004]	[0.003]	[0.005]	[0.014]	[0.011]	[0.017]	[0.010]	[0.008]	[0.012]	[0.015]	[0.012]	[0.019]
R^2	0.166	0.125		0.134	0.088		0.119	0.091		0.087	0.076	
Export share				0.144**	0.029	0.115	0.173***	0.000	0.173***	−0.026	0.033	−0.06
				[0.067]	[0.048]	[0.079]	[0.053]	[0.036]	[0.062]	[0.081]	[0.051]	[0.092]
R^2				0.129	0.086		0.121	0.091		0.085	0.075	

Notes. Table reports coefficients on log domestic sales for 54 separate regressions. (Covariate at left; dependent variables at top, with changes over period at top.) All regressions include 205 industry (six-digit) and 32 state dummies. Coefficients when initial export share is proxy and change in export share is outcome omitted because measurement error generates severe biases. Variable definitions in Appendix I. Further details on data set in Section IV of the text and Appendix II (online). Robust standard errors in brackets. Standard errors on differences allow for cross-equation correlation. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

To address this concern, I draw on balance sheet data from publicly listed firms on the Mexican stock market over the 1989–2000 period. These are the data used by Pratap and Urrutia (2004) and Aguiar (2005) and are the most detailed microdata with balance-sheet information available in Mexico. (Summary statistics are in Table A3 in Appendix II (online).) In these data, exporting plants have a significantly larger share of their debt in dollar-denominated loans than do nonexporting plants. In manufacturing in 1993, the dollar-denominated shares of short-term debt were 38% and 20% for exporters and nonexporters, respectively. For long-term debt, the corresponding figures were 36% and 13%. As a consequence, the balance sheets of exporting plants are likely to have been more *adversely* affected by the peso devaluation than those of nonexporters. However, exporters also face a lower cost of capital (defined as the ratio of total interest payments to total debt) than nonexporters; in manufacturing in 1993, the cost of capital was 12.5% for exporters and 17.8% for nonexporters. This suggests that increased exports may have differentially reduced the cost of capital for initially more productive plants. To investigate this, Table VI presents regressions of the form of (6) with the change in the cost of capital as the dependent variable. For the lower cost of capital for exporters to explain the differential wage changes, it would have to be the case that the cost of capital declined more for initially larger plants than initially smaller ones during the peso crisis period and that this relative decline was greater in magnitude during the peso crisis period than during the periods 1989–1993 or 1997–2000. In fact, the point estimates for the 1993–1997 period are *positive* in all specifications and not significantly smaller than the coefficients for the other periods. There is no evidence that differential changes in the cost of capital can explain the differential wage changes of Table II.[41]

One might still be worried that some factor unrelated to exports generated the difference in differential trends. One way to investigate this possibility is to examine whether the same differential trends are displayed even within a set of plants that saw no differential change in exports over the period—the *maquiladora* plants participating in the government's

41. This is consistent with the conclusion of Aguiar (2005, p. 106): "Given that export propensity and foreign debt composition are significantly correlated ... the balance sheet effect is offsetting much of the benefits of the real devaluation."

TABLE VI
ESTIMATES FOR COST OF CAPITAL, PUBLICLY LISTED MEXICAN FIRMS, 1989–2000

		Dep. var.: Δ cost of capital			
		All sectors		Manufacturing only	
		OLS (1)	IV (2)	OLS (3)	IV (4)
1989–1993	Log domestic sales, 1989	1.545	−1.13	−1.483	−1.724
		[1.850]	[1.170]	[2.353]	[1.649]
	N	68	68	37	37
	R^2	0.130	0.042	0.011	0.005
1993–1997	Log domestic sales, 1993	0.22	0.379	0.088	0.397
		[0.671]	[0.414]	[1.059]	[0.554]
	N	122	122	57	57
	R^2	0.086	0.128	0.001	0.010
1997–2000	Log domestic sales, 1997	0.642	1.473	0.155	0.948*
		[0.411]	[0.981]	[0.471]	[0.515]
	N	113	113	52	52
	R^2	0.050	0.083	0.002	0.088

Notes. Table reports coefficients on initial log domestic sales for 12 separate regressions. (Covariate at left; dependent variable at top, with changes over period at left.) Regressions for all sectors together include industry effects (six coarse industry categories). IV regressions instrument log domestic sales in 1990, 1994, and 1998 with value from previous year; second-stage dependent variables are changes over 1990–1993, 1994–1997, and 1998–2001, respectively. Cost of capital defined as 100 $*$ (interest paid)/(total debt). Data set is unbalanced panel of publicly listed firms on Mexican stock market (*Bolsa Mexicana de Valores*) for 1989–2000. For further details on data set, see Pratap and Urrutia (2004) and Aguiar (2005). Robust standard errors in brackets.*** indicates significance at 1% level,** at 5% level,* at 10% level.

export-promotion program.[42] When the *maquiladora* program began in 1965, *maquiladoras* were required to export 100% of their output. Although this requirement has gradually been loosened since 1989, *maquiladora* plants continue to export nearly all of their output.[43] Thus the peso devaluation, although it was a boon to the sector as a whole, did not generate a differential change in the extent of exporting between larger and smaller, or more productive and less productive, *maquiladora* plants. Table VII presents a comparison of results for the non-*maquiladora* sector (from the EIA 1993–2001 panel) and the *maquiladora* sector (from a balanced panel for the same period created from the EMIME data described in Appendix II [online]). The proxy for initial productivity is log employment and the wage variables are average hourly wages, white-collar yearly earnings, and blue-collar yearly earnings, because these are the variables that can be compared across data sets. The results for the *maquiladora* sector are quite distinct from those for the non-*maquiladora* sector. The point estimates of initial log employment for the period 1993–1997 are *negative*. In the later period the coefficients are significantly *larger* than in the earlier period. It does not appear that the differential wage changes in the non-*maquiladora* sector are due to a generalized differential trend between larger and smaller, or more productive and less productive plants.[44]

42. An alternative approach would be to focus on initially small plants that are likely to be below the export margin even after the devaluation. Estimates of (6) for plants initially below median domestic sales in each industry yield differential wage trends similar to those in Table II, which ostensibly casts doubt on the quality-upgrading hypothesis. But this test is not as clean as the comparison with *maquiladoras*, because even plants that report zero domestic sales may be producing intermediate goods for plants that export or may be selling on the export market through distributors; export-driven quality upgrading may be present even among plants that report zero direct exports.

43. Although the fraction of revenues from export v. domestic sales is not reported in the published statistics, INEGI has carried out unpublished surveys, and has consistently found that *maquiladoras* sell less than 5% of their output domestically. (Source: personal communication with Gerardo Durand, Director of Statistics of International Trade, Administrative Records, and Prices, INEGI.)

44. This conclusion is reinforced by two additional findings. First, using data from the Mexican social security agency that cover a set of nontradable sectors (construction, transportation, retail, and service) in addition to the manufacturing (tradable) sector, Table A4 in Appendix II (online), provides evidence of a significantly greater difference in differential trends in manufacturing than in the nontradable sectors. For further details, see Appendix II and Kaplan and Verhoogen (2006). Second, Kandilov (2005), building explicitly on an earlier incarnation of this paper (Verhoogen 2004), uses exogenous variation in the incentive to export from an export-subsidy program in Chile, rather than a macroeconomic shock, and finds that industries offered the subsidy displayed greater wage growth for white-collar workers than nonsubsidized industries, and that within subsidized industries initially larger and more productive plants saw greater wage increases for both white-collar and blue-collar workers than initially smaller or less productive plants.

TABLE VII
COMPARING NON-MAQUILADORA AND MAQUILADORA SECTORS

		EIA 1993–2001 panel (non-*maquiladora* sector)			EMIME panel (*maquiladora* sector)		
		Δ log(avg. hourly wage) (1)	Δ log(white-collar yearly earnings) (2)	Δ log(blue-collar yearly earnings) (3)	Δ log(avg. hourly wage) (4)	Δ log(white-collar yearly earnings) (5)	Δ log(blue-collar yearly earnings) (6)
1993–1997	Log employment 1993	0.072***	0.055***	0.050***	−0.019	−0.023	−0.029**
		[0.007]	[0.010]	[0.007]	[0.016]	[0.016]	[0.014]
	R^2	0.157	0.124	0.140	0.055	0.044	0.041
1997–2001	Log employment 1997	0.025***	0.013	0.017***	0.032**	0.050***	0.038***
		[0.006]	[0.008]	[0.006]	[0.014]	[0.014]	[0.013]
	R^2	0.109	0.086	0.104	0.047	0.049	0.061
Difference (1993–1997 vs. 1997–2001)		0.048***	0.042***	0.033***	−0.050**	−0.073***	−0.067***
		[0.009]	[0.013]	[0.009]	[0.021]	[0.021]	[0.019]
N		3,263	3,263	3,263	1,088	1,088	1,088

Notes. Table reports coefficients on initial log employment for 12 separate regressions. (Covariate at left; dependent variables at top, with changes over period at left.) Columns (1)–(3) use the EIA 1993–2001 panel and include 205 industry (six-digit) and 32 state dummies. Columns (4)–(6) use the EMIME panel over same period and include 12 industry and 32 state dummies. Variable definitions in Appendix I. Further details on data sets in Section IV of the text and Appendix II (online). Robust standard errors in brackets. Standard errors on differences allow for cross-equation correlation.*** indicates significance at 1% level,** at 5% level,* at 10% level.

These results for the *maquiladora* sector also argue against the outsourcing hypothesis of Feenstra and Hanson (1996) as an explanation for the differential wage trends in Table II. Feenstra and Hanson hypothesize that U.S. firms outsource production activities that are low-skill for the United States but high-skill for Mexico, raising skill demands in both places. One would expect this mechanism to be strongest in the *maquiladoras*, which are explicitly dedicated to outsourcing from Northern firms. The outsourcing hypothesis provides a compelling account of the increase in skill and capital intensity of the *maquiladora* sector over time, but the fact that we do not see greater differential wage changes between larger and smaller plants in the peso crisis period in the *maquiladora* sector suggests that the outsourcing hypothesis is unlikely to explain the differential wage changes we observe in the non-*maquiladora* sector.[45]

Interested readers are referred to the working paper (Verhoogen 2007) for a discussion of two additional alternative hypotheses: the rent-sharing hypothesis and the scale-economies hypothesis of Yeaple (2005) (further developed by Bustos [2005]).

VII. Conclusion

This paper has proposed a new mechanism linking trade and wage inequality in Mexico: quality upgrading due to increased exports. It has offered robust evidence that initially larger, more productive plants were more likely to increase exports, white-collar wages, blue-collar wages, and ISO 9000 certification than initially smaller, less productive plants during the peso-crisis period (1993–1997) and that this differential change was greater than during adjacent periods without devaluations (1989–1993 and 1997–2001). The pattern in 1986–1989, also a period of devaluation of the peso, is similar to that of 1993–1997. These findings are robust to the choice of proxy for plant productivity and do not appear to be explained by a variety of alternative hypotheses. There is also evidence that the differential wage changes were greater for white-collar than for blue-collar workers, suggesting

45. It is also worth noting that the Feenstra–Hanson mechanism relies on shifts between activities of different skill intensity within industries to explain wage changes. Although the model does not carry predictions for which particular plants will undertake the new skill-intensive activities, one would expect the higher-λ plants to be the most likely candidates. We have seen that there is no evidence that higher-λ plants raised their white-collar employment share more than lower-λ plants in the same industry. Hence it seems unlikely that shifts between activities of different skill intensities are driving the differential wage changes.

that quality upgrading leads to increases in within-plant wage inequality as well as to increased wage dispersion across plants. These results support the hypothesis that devaluation-induced quality upgrading contributed to rising wage inequality within industries in the Mexican manufacturing sector.

An important question that remains unanswered is the extent to which the quality-upgrading mechanism can be generalized to other contexts. On the basis of the theoretical model, we would expect any bilateral reduction in trade costs—due to a trade agreement or declining transport costs, for instance—to have effects similar to those of an exchange-rate devaluation. The incentive to export would increase and domestic sales would decrease—in this case because of an increase in import competition rather than a contraction of domestic demand—generating a differential increase in the export share and hence differential quality upgrading within industries. Case-study evidence (Gereffi 1999; Nadvi 1999) also suggests that quality upgrading may be a more general phenomenon. If so, the quality-upgrading hypothesis may help to explain the link between trade liberalization and wage inequality more broadly.

The main relevance of the quality-upgrading hypothesis to policy debates may lie in its identification of a new dimension of relative winners and losers from trade in developing countries. In the simplest Heckscher–Ohlin model, the relative winners are unskilled workers and the relative losers are skilled workers and owners of capital. From this perspective, it is a mystery why many unskilled workers in developing countries protest against globalization and why many of the foremost proponents of globalization are educated, urban elites. In the quality-upgrading view, by contrast, the relative winners are the entrepreneurs and employees, especially the most skilled, with either the qualifications or the good fortune to be employed in the most modern, export-oriented plants within each industry, and the relative losers are the entrepreneurs and employees, especially the unskilled, in less productive, domestically oriented plants. In this view, both the enthusiasm for globalization of the relatively better off and the pessimism of the relatively worse off may make economic sense.

Appendix I: Variable Definitions

Details on the processing of the various data sets, as well as all appendix tables and figures, are in Appendix II (online).

I.A. EIA 1993–2001 and 1984–2001 Panels

Employment (white-collar, blue-collar, total) = average yearly employment for nonproduction workers (*empleados*), production workers (*obreros*), and all workers, respectively.

White-collar hourly wage = total white-collar wage bill/total hours worked by white-collar workers, deflated to 1994 pesos using the main consumer price index (INPC) from *Banco de Mexico*, the Mexican central bank.

Blue-collar hourly wage = total blue-collar wage bill/total hours worked by blue-collar workers, deflated to 1994 pesos by consumer price index.

Wage ratio = white-collar hourly wage/blue-collar real hourly wage.

White-collar employment share = white-collar hours worked/ total hours worked

Domestic sales = domestic sales as reported, measured in thousands of 1994 pesos, deflated by the producer price index (INPP) from *Banco de Mexico*.

Export sales = export sales as reported, measured in thousands of 1994 pesos, deflated by the producer price index.

Export share of sales = export sales/(domestic sales + export sales).

Import share of input expenditures = imported input expenditures/(domestic input expenditures + imported input expenditures).

Exporter = 1 if export percentage of sales > 0, = 0 otherwise.

Capital–labor ratio = real capital stock/total employment. Capital stock was constructed using a perpetual-inventory method; see Appendix II (online).

Foreign ownership indicator = 1 if plant had ≥ 10 percent foreign ownership in 1994 (the only year for which data were available during 1993–2001), = 0 otherwise.

I.B. EIA-ENESTyC Panel

ISO 9000 indicator = 1 if the establishment has ISO 9000 or similar quality certification, = 0 otherwise.

Average schooling of blue-collar workers = (3*(# with less than primary school) + 6*(# with primary school) + 9*(# with junior high school) + 12*(# with high school) + 16*(# with college/professional degree) + 18*(# with postgraduate education))/employment of blue-collar workers at time of survey

Average schooling of white-collar workers: similar to blue-collar average schooling.

Formal training indicator = 1 if establishment reports that it has a formal training program, = 0 otherwise.

Turnover rate = 100*(2*(separations + new hires over previous six months))/total employment at time of survey.

Accident rate = 100*(number of accidents over previous calendar year)/average employment for previous calendar year.

Absentee rate = average number of daily absences/employment at time of survey.

I.C. EMIME Panel

Average hourly wage = total wage bill (all workers)/total hours. Note that hours by occupation are not available prior to 1997.

Blue-collar yearly earnings = total blue-collar (*obrero*) wage bill/average blue-collar employment over 12 months, deflated to 1994 pesos by consumer price index.

White-collar yearly earnings = total white-collar (*empleados* + *técnicos*) wage bill/average white-collar employment over 12 months, deflated to 1994 pesos by consumer price index.

I.D. ENEU Household Data

Hourly wage (Figure I) = (wage in job worked last week, converted to monthly basis)/(hours worked in job last week, converted to monthly basis), deflated to 1994 pesos using consumer price index.

Average schooling, by four-digit industry (Figure I) = average of years of schooling of workers in industry.

The ENEU sample used in this paper consists of men, ages 16–64, who worked 35 or more hours in the previous week, in 16 cities in original (1987) ENEU sample.[46] The wage and average schooling calculations use the sampling weights reported by INEGI.

46. The processing of the ENEU data in Figure 1b of Robertson (2004) differs from this sample in that it includes women and only 8 of the original 16 cities. These differences do not explain the discrepancy between that figure and Figure I of this paper, and I have not been able to replicate his figure to determine the source of the discrepancy.

I.E. Alternative Proxies

The alternative proxy variables used in Table V were constructed as follows:

Predicted export share index: I ran a tobit of the export share of sales on log hours worked for each occupation, log total sales, log capital–labor ratio, log electricity intensity (kilowatt hours used per hour worked), an indicator for whether the plant has $\geq$10% foreign ownership, and four-digit industry effects. I then recovered the predicted values $x'\hat{\beta}$, deviated from six-digit industry means, and standardized the variable to have variance 1.

Predicted ISO 9000 index: Using the EIA-ENESTyC panel, I ran a probit of ISO 9000 certification on the same covariates as for the export share index plus the export share itself. With the estimate of $\hat{\beta}$ from this probit, I calculated $x'\hat{\beta}$ for all plants in the EIA 1993–2001 panel, and then deviated from six-digit industry means and standardized to have variance 1.

First principal component: I deviated the covariates used for the ISO 9000 index from six-digit industry means, took the first principal component (the linear combination capturing the maximum variance of the joint distribution of the variables), and standardized the variable to have variance 1.

Total factor productivity (TFP): I pooled data for two years (i.e., 1993 and 1994, or 1997 and 1998) and regressed log revenues on log hours worked for each occupation, log capital–labor ratio, log materials costs, log electricity costs, and plant fixed effects. I recovered the coefficients on the plant fixed effects, deviated them from six-digit industry means, and standardized.[47] Note that this approach to estimating TFP is attractive because, unlike other standard TFP measures, it does not mechanically include year-specific measurement error in revenues.[48]

Domestic sales per worker = total real domestic sales/average employment for year.

47. I am indebted to Matthias Schündeln for suggesting this approach.

48. It is important to note that standard methods for estimating TFP impose assumptions on the production function and market structure that are inconsistent with the model presented in this paper. I include the TFP proxy for the sake of completeness.

Because wages are the primary outcome variables of interest, they are omitted from the construction of these proxies, so that mean reversion in wages due to measurement error does not bias the estimates. An FDI indicator is included; although this variable does not enter into the theoretical model, it seems plausible that foreign-owned plants have greater access to the knowledge and technologies of their parent companies, and hence that foreign ownership is a predictor for unobserved know-how or productivity.

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REFERENCES

Abowd, John, and Thomas Lemieux, "The Effects of Product Market Competition on Collective Bargaining Agreements: The Case of Foreign Competition in Canada," *Quarterly Journal of Economics*, 108 (1993), 983–1014.

Aghion, P., R. Burgess, S. Redding, and F. Zilibotti, "Entry Liberalization and Inequality in Industrial Performance," *Journal of the European Economic Association*, 3 (2005), 291–302.

Aguiar, Mark, "Investment, Devaluation, and Foreign Currency Exposure: The Case of Mexico," *Journal of Development Economics*, 78 (2005), 95–113.

Akerlof, George, "Labor Contracts as Partial Gift Exchange," *Quarterly Journal of Economics*, 97 (1982), 543–569.

Anderson, Simon, Andre de Palma, and Jacques-Francois Thisse, *Discrete-Choice Theory of Product Differentiation* (Cambridge, MA: MIT Press, 1992).

Aw, Bee Yan, and Geeta Batra, "Wages, Firm Size, and Wage Inequality: How Much Do Exports Matter?" in *Innovation, Industry Evolution, and Employment*, D. B. Audretsch and A. R. Thurik, eds. (Cambridge: Cambridge University Press, 1999).

Berman, Eli, John Bound, and Stephen Machin, "Implications of Skill-Biased Technological Change: International Evidence," *Quarterly Journal of Economics*, 113 (1998), 1245–1279.

Bernard, Andrew B., and J. Bradford Jensen, "Exporters, Skill Upgrading and the Wage Gap," *Journal of International Economics*, 42 (1997), 3–31.

——, "Exceptional Exporter Performance: Cause, Effect, or Both?" *Journal of International Economics*, 47 (1999), 1–25.

Bernard, Andrew B., Stephen J. Redding, and Peter K. Schott, "Multi-Product Firms and Trade Liberalization," NBER Working Paper No. 12782, Dec. 2006.

Bertrand, Marianne, Esther Duflo, and Sendhil Mullainathan, "How Much Should We Trust Difference-in-Differences Estimates?" *Quarterly Journal of Economics*, 119 (2004), 249–276.

Bowles, Samuel, "The Production Process in a Competitive Economy: Walrasian, Neo-Hobbesian and Marxian Models," *American Economic Review*, 75 (1985), 16–36.

Brooks, Eileen L., "Why Don't Firms Export More? Product Quality and Colombian Plants," *Journal of Development Economics*, 80 (2006), 160–178.

Brown, Charles, and James Medoff, "The Employer Size–Wage Effect," *Journal of Political Economy*, 97 (1989), 1027–1059.

Bustos, Paula, "The Impact of Trade on Technology and Skill Upgrading: Evidence from Argentina," Unpub. Paper, Universitat Pompeu Fabra, 2005. Available online at http://www.crei.cat/people/bustos/wp/Trade,%20Technology%20and%20skill.pdf

Cragg, Michael, and Mario Epelbaum, "Why Has Wage Dispersion Grown in Mexico? Is It the Incidence of Reforms or the Growing Demand for Skills?" *Journal of Development Economics*, 51 (1996), 99–116.

Davis, Donald R., "Trade Liberalization and Income Distribution," NBER Working Paper No. 5693, 1996.
Economist Intelligence Unit, *Mexico: Country Report, 1st quarter* (London: Economist Intelligence Unit, 1995).
Feenstra, Robert C., "Quality Change under Trade Restraints in Japanese Autos," *Quarterly Journal of Economics*, 103 (1988), 131–146.
Feenstra, Robert, and Gordon Hanson, "Foreign Investment, Outsourcing and Relative Wages," in *Political Economy of Trade Policy: Essays in Honor of Jagdish Bhagwati*, R. Feenstra and G. Grossman, eds. (Cambridge, MA: MIT Press, 1996).
Flam, Harry, and Elhanan Helpman, "Vertical Product Differentiation and North–South Trade," *American Economic Review*, 77 (1987), 810–822.
Gabszewicz, J. Jaskold, Avner Shaked, John Sutton, and Jacques-Francois Thisse, "International Trade in Differentiated Products," *International Economic Review*, 22 (1982), 527–535.
Gereffi, Gary, "International Trade and Industrial Upgrading in the Apparel Commodity Chain," *Journal of International Economics*, 48 (1999), 37–70.
Goldberg, Pinelopi Koujianou, and Nina Pavcnik, "Distributional Effects of Globalization in Developing Countries," *Journal of Economic Literature*, 45 (2007), 39–82.
Guler, Isin, Mauro Guillen, and John Muir Macpherson, "Global Competition, Institutions and the Diffusion of Organizational Practices: The International Spread of ISO 9000 Quality Certificates," *Administrative Science Quarterly*, 47 (2002), 207–232.
Hallak, Juan Carlos, "Product Quality and the Direction of Trade," *Journal of International Economics*, 68 (2006), 238–265.
Hanson, Gordon, and Ann Harrison, "Trade Liberalization and Wage Inequality in Mexico," *Industrial and Labor Relations Review*, 52 (1999), 271–288.
Hashimoto, Masanori, "Firm-Specific Human Capital as a Shared Investment," *American Economic Review*, 71 (1981), 475–482.
Hummels, David, and Peter J. Klenow, "The Variety and Quality of a Nation's Exports," *American Economic Review*, 95 (2005), 704–723.
Hummels, David, and Alexandre Skiba, "Shipping the Good Apples Out? An Empirical Confirmation of the Alchian–Allen Conjecture," *Journal of Political Economy*, 112 (2004), 1384–1402.
Kandilov, Ivan, "How International Trade Affects Wages and Employment," Ph.D. Dissertation, University of Michigan, 2005.
Kaplan, David S., and Eric A. Verhoogen, "Exporting and Individual Wage Premia: Evidence from Mexican Employer-Employee Data," Unpub. paper, ITAM and Columbia University, 2006. Available online at http://www.columbia.edu/~ev2124/.
Kremer, Michael, "The O-Ring Theory of Economic Development," *Quarterly Journal of Economics*, 108 (1993), 551–575.
Kremer, Michael, and Eric Maskin, "Wage Inequality and Segregation by Skill," NBER Working Paper No. 5718, 1996.
Manasse, Paolo, and Alessandro Turrini, "Trade, Wages and Superstars," *Journal of International Economics*, 54 (2001), 97–117.
McFadden, Daniel, "Conditional Logit Analysis of Qualitative Choice Behavior," in *Frontiers in Econometrics*, P. Zarembka, ed. (New York: Academic Press, 1974).
Melitz, Marc, "The Impact of Trade on Intra-industry Reallocations and Aggregate Industry Productivity," *Econometrica*, 71 (2003), 1695–1725.
Nadvi, Khalid, "Collective Efficiency and Collective Failure: The Response of the Sialkot Surgical Instrument Cluster to Global Quality Pressures," *World Development*, 27 (1999), 1605–1626.
Pratap, Sangeeta, and Carlos Urrutia, "Firm Dynamics, Investment and Debt Portfolio: Balance Sheet Effects of the Mexican Crisis of 1994," *Journal of Development Economics*, 75 (2004), 535–563.
Revenga, Ana, "Exporting Jobs? The Impact of Import Competition on Employment and Wages in U.S. Manufacturing," *Quarterly Journal of Economics*, 107 (1992), 255–284.

——, “Employment and Wage Effects of Trade Liberalization: The Case of Mexican Manufacturing,” *Journal of Labor Economics*, 15 (1997), s20–s43.
Robertson, Raymond, “Relative Prices and Wage Inequality: Evidence from Mexico,” *Journal of International Economics*, 64 (2004), 387–409.
Schott, Peter, “Across-Product versus Within-Product Specialization in International Trade,” *Quarterly Journal of Economics*, 119 (2004), 647–678.
Shapiro, Carl, and Joseph Stiglitz, “Equilibrium Unemployment as a Worker Discipline Device,” *American Economic Review*, 74 (1984), 433–444.
Tybout, James, “Manufacturing Firms in Developing Countries: How Well Do They Do, and Why?” *Journal of Economic Literature*, 38 (2000), 11–44.
——, “Plant- and Firm-Level Evidence on the ‘New’ Trade Theories,” in *Handbook of International Trade*, E. K. Choi and J. Harrigan, eds. (Oxford: Basil Blackwell, 2003).
Verhoogen, Eric A., “Trade, Quality Upgrading and Wage Inequality in the Mexican Manufacturing Sector: Theory and Evidence from an Exchange-Rate Shock,” Center for Labor Economics, UC Berkeley, Working Paper No. 67, Jan. 2004.
——, “Trade, Quality Upgrading and Wage Inequality in the Mexican Manufacturing Sector,” CEPR Discussion Paper No. 6385, July 2007.
Wacziarg, Romain, and Jessica Seddon Wallack, “Trade Liberalization and Intersectoral Labor Movements,” *Journal of International Economics*, 64 (2004), 411–439.
Xiang, Chong, “New Goods and the Relative Demand for Skilled Labor,” *Review of Economics and Statistics*, 87 (2005), 285–298.
Yeaple, Stephen Ross, “A Simple Model of Firm Heterogeneity, International Trade, and Wages,” *Journal of International Economics*, 65 (2005), 1–20.