# Trade, Quality Upgrading and Wage Inequality in the Mexican Manufacturing Sector*

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Abstract

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.

Keywords: Trade and wage inequality, quality upgrading, heterogeneous firms, exchange-rate shock
JEL classification: F1, J3, O1, L1

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# 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, non-petroleum exports and total imports rose an average of 16.5 percent and 15.7 percent per year respectively over the 1985-2000 period. Figure I depicts the evolution of two measures of wage inequality: the log 90-10 ratio from a household 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 like Mexico when it integrates with a country like 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 like Mexico,[3] but because such models rely 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 et al. [1998]).

This paper proposes a new mechanism linking trade and wage inequality in developing countries — the quality-upgrading mechanism — and investigates its empirical implications

[1]Variable definitions are in Appendix I. The datasets are described in more detail in Section IV below and in Appendix II, posted online at http://www.columbia.edu/∼ev2124/.

[2]See Goldberg and Pavcnik [2007] for a review.

[3]Inequality may increase, for instance, if a unskilled-labor-abundant country like Mexico opens trade simultaneously with a skill-abundant country like the United States and an even more unskilled-labor-abundant country like China [Davis, 1996] or if relatively unskilled-labor-intensive industries are more protected prior to liberalization [Revenga, 1997].

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

[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 manufacturing censuses and schooling data from a household survey, Appendix Figures I and II in Appendix II (online) present evidence of shifts toward *less*-skill-intensive and *less*-capital-intensive sectors within manufacturing in Mexico over the 1988-1998 period, consistent with the simplest Heckscher-Ohlin model.

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. Since 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 1985-87, to investigate this mechanism. Using a number of different proxies for the underlying plant productivity parameter, I compare differential changes in outcomes between initially more-productive 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 dataset, 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 et al., 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.

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. There are many factors that 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 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 percent of the level of total variance and 34 percent of the change over the entire 1984-2001 period, but 52 percent and 51.5 percent of the changes in 1986-1988 and 1994-1995, respectively. While the increase in the between-industry component may itself reflect a differential shock to exporting across industries,

[6]Bernard et al. [2006] analyze product switching in the context of a model with symmetric goods.

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

[8]The dataset is described in more detail in Section IV below and in Appendix II (online).

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 datasets 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 (±.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 percent, 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 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 percent (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 $1.50 per hour to approximately $.90 per hour from 1994 to 1995,

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

rising back to only $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 non-tariff barriers had been removed, and approximately 95 percent of all imports into Mexico were covered by tariffs of 20 percent or lower. On the U.S. side, approximately 80 percent of imports were covered by tariffs of 5 percent or less. A majority of commodities were assigned phase-out 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 device 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. Between 1994 and 1995, exports as a share of total production rose sharply, due both to a decline in domestic sales and to 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 which had been in use in Germany since the 1950s. One consequence

[10] These figures are from the *Encuesta Nacional de Empleo Urbano (ENEU)* [National Urban Employment Survey]. For details, refer to Appendix II (online).

[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.

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 3-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 over the 1993-2001 period;[13] the shift toward exporting is evident.[14] More than 80 percent 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, anecdotal evidence

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

[13]The dataset is described in more detail in Section IV below and in Appendix II (online).

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

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. 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 which is assumed to buy one unit of a good from a continuum of goods indexed by $\omega$ 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 $\theta_d$ 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 $\theta_s$ and $\theta_n$ are constant within each country but that Northern consumers more willing to pay for quality than Southern ones: $\theta_n > \theta_s$. I treat $\theta_s$ and $\theta_n$ 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 $\delta_d$ be the ratio of the price level in country $d$ to the price level in South; that is, $\delta_s = 1$ and $\delta_n$ is the real exchange rate. The price of good $\omega$ relative to the Southern price level is then $p_d(\omega) = \delta_d \widetilde{p}_d(\omega)$.

The random consumer-product-match term, $\varepsilon$, is assumed to be independent and iden-

[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 $\kappa$ is the consumption of a non-differentiated 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 sub-utility 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).

tically distributed across consumer with a type 1 extreme-value distribution,[17] a standard multinomial-logit formulation [McFadden, 1974; Anderson et al., 1992]. A familiar derivation yields the following expected demand for each good [Anderson et al., 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 $\mu$ is a parameter of the distribution of $\varepsilon$ that captures the degree of differentiation between goods and $\Omega_d$ 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 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 $\lambda$, 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 fashion:[18]

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

[17] That is, $F(\varepsilon) = \exp\left(-\exp\left(-\frac{\varepsilon}{\mu} + \gamma\right)\right)$, where $\gamma = .5772$ (Euler's constant) by assumption to ensure that the expectation of $\varepsilon$ is zero.

[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.

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 interests 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 to induce workers to supply it [Akerlof, 1982; Bowles, 1985; Shapiro and Stiglitz, 1984]; 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 $\rho$. 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

[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].

[21] The implications of the three interpretations are examined in matched employer-employee data in Kaplan and Verhoogen [2006].

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:

$$
\begin{aligned}
&\text{(5a)} \qquad & q_d^*(\lambda) &= (\eta\lambda\delta_d^\alpha\theta_d^\alpha)^{\frac{1}{1-\alpha}} \\
&\text{(5b)} \qquad & w_d^{h*}(\lambda) &= \underline{w}^h + \alpha^h\delta_d\theta_d q_d^*(\lambda) \\
&\text{(5c)} \qquad & w_d^{l*}(\lambda) &= \underline{w}^l + \alpha^l\delta_d\theta_d q_d^*(\lambda) \\
&\text{(5d)} \qquad & k_d^*(\lambda) &= \frac{\alpha^k}{\rho}\delta_d\theta_d q_d^*(\lambda) \\
&\text{(5e)} \qquad & p_d^*(\lambda) &= \mu\delta_d + \underline{w}^h + \underline{w}^l + \alpha\delta_d\theta_d q_d^*(\lambda)
\end{aligned}
$$

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

These equations carry several implications. First, all else equal, higher-$\lambda$ 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-$\lambda$ plants. It also follows directly that both output and profits are increasing in $\lambda$.

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

Third, all else equal, plant size and wages are positively correlated, since both are increasing in $\lambda$. 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, since all are increasing in $\lambda$.

[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].

Fifth, whether the ratio of the white-collar wage to the blue-collar wage is increasing or decreasing in quality, and hence in $\lambda$ and $\theta_d$, 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 $\lambda$ and $\theta_d$.[23]

Finally, the fact that profitability is increasing in $\lambda$ 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 have 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 $\lambda$. 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 $\lambda_n^{\min}$ enter the export market; the solid curve represents actual average quality as a function of $\lambda$, 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 that of 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, $\delta_n$, 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 $\delta_n$ or $N_s$ and hence is unaffected by the exchange-rate shock. Quality on the export line, $q_n(\lambda)$,

[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 $\frac{\partial}{\partial \theta_d}\left(\frac{w_d^{h*}(\lambda)}{w_d^{l*}(\lambda)}\right)$. If $\frac{\alpha^h}{\alpha^l} > \frac{\underline{w}^h}{\underline{w}^l}$ then the wage ratio, $\frac{w_d^h(\lambda)}{w_d^l(\lambda)}$, is increasing in $\lambda$ and $\theta_d$.

[24] Refer to the working paper [Verhoogen, 2007] for a discussion of the conditions.

increases to $q_n(\lambda)'$, since 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 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 $\lambda_n^{\min\prime}$ and $\lambda_n^{\min}$) see an especially large increase in average quality. Average quality is increasing in $\lambda$ within the category of switchers and within the category of always exporters (to the right of $\lambda_n^{\min}$). 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 *Encuesta Industrial Anual (EIA)* [Annual Industrial Survey], a yearly panel survey conducted by the *Instituto Nacional de 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 6-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

[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) below); the important point is that one would not expect such changes to generate differential changes between high-$\lambda$ and low-$\lambda$ plants.

[26] If product quality is sufficiently more sensitive to the quality of white-collar than 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.

[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 1993-2001 period, which I refer to as the EIA 1993-2001 Unbalanced Panel; refer to footnote 40 and the data construction details in Appendix II (online).

previous year. In 1995, 1999 and 2001 the ENESTyC elicited information on ISO 9000 certification, an international production standard. While 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 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 datasets 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 non-exporters 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 higher 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

[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 et al., 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 non-exporters and exporters are similar to those reported here.

[31] These differences are significant at the 5 percent 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.

peso crisis period than in other periods without an exchange-rate devaluation. A central econometric challenge in testing these implications is that the key variable, entrepreneurial ability $\lambda$, 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 $\lambda$, 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 $\lambda$ 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 $\lambda$, without the discontinuity at the cut-off 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, non-linear function depicted in Figure VIII is unlikely to hold empirically, 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-$\lambda$ and lower-$\lambda$ 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 $\lambda$ will be greater in periods with major devaluations than in periods without.

The main estimating equation is the following:

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

where $i$, $j$ and $r$ index plants, industries and states, respectively; $\triangle y_{ijr}$ is a change in one of the outcome variables; $\alpha$ is an intercept term; $\widetilde{\lambda}_{ijr}$ is the value of the entrepreneurial-ability proxy in the initial year; $\psi_j$ is an industry fixed effect; $\xi_r$ 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 pre-treatment and post-treatment periods and understate the standard errors on the coefficient estimates, as discussed by Bertrand et al. [2004]. My strategy is to select just one year of data pre-crisis and one year post-crisis, i.e. 1993 and 1997 for the peso crisis period, 1997 and 2001 for the later placebo period without an exchange-rate shock. Bertrand et al. [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 appears in the denominator of the export share in the initial period and hence any measurement error in domestic sales will generate a mechanical positive bias in the estimate of $\beta$. Under the assumption that the measurement error in domestic sales is uncorrelated over time,

[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.

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 pre-devaluation 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 $\lambda$. It may be that the productivity improvements are correlated with $\lambda$ 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 $\beta$ 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, then 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-$\lambda$ 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 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 $\lambda$ 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

[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.

is instrumented with its lag, the coefficient is .005 and significant at the 5 percent 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 percent larger in log domestic sales terms than the other. Column 2 indicates that the larger plant had an approximately .7 percent greater wage increase for white-collar workers than the smaller plant during the 1993-1997 period, and a .16 percent 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 re-appreciated over the 1987-1993 period. Although the aggregate export response to the 1985-1987 devaluation was smaller than to the peso crisis, we would nonetheless expect to see greater differential trends between higher- and lower-$\lambda$ plants in the 1986-1989 and 1993-1997 periods than in the 1989-1993 period. 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 re-appreciation resembles the post-crisis re-appreciation period (1997-2001). This particular time pattern of changes

[36]Non-parametric 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 Appendix Figures IV-V 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 appears in the denominator.

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 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.

While 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 $\lambda$: log employment; predicted export share; predicted ISO 9000 certification; the first principal component of a number of variables hypothesized to be correlated with $\lambda$; 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; Appendix Table I in

[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.

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-$\lambda$ and low-$\lambda$ 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.

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 micro-data with balance-sheet information available in Mexico. (Summary statistics are in Appendix Table III.) In these data, exporting plants have a significantly larger share of their debt in dollar-denominated loans than do non-exporting plants. In manufacturing in 1993, the dollar-denominated shares of short-term debt were 38 percent and 20 percent for exporters and non-exporters respectively. For long-term debt, the corresponding figures were 36 percent and 13 percent. As a consequence, the balance sheets of exporting plants are likely to have been more *adversely* affected by the peso devaluation than those of non-exporters. However, exporters also face a lower cost of capital (defined as the ratio of total interest payments to total debt) than non-exporters; in manufacturing in 1993, the cost of capital was 12.5 percent for exporters and 17.8 percent for non-exporters. This suggests that increased exports may have differentially reduced the cost of capital for initially more-productive plants. To investigate

[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 non-robustness of the estimate for white-collar wages when using actual export share as the $\lambda$ 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 Appendix Table II. 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.

this, Table VI presents regressions of the form of (6) with the change in the cost of capital as the dependent variable. In order 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 1989-1993 or 1997-2000 periods. 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 export-promotion program.[42] When the *maquiladora* program began in 1965, *maquiladoras* were required to export 100 percent 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 since these are the variables that can be compared across

[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."

[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 vs. 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 percent of their output domestically. (Source: personal communication with Gerardo Durand, Director of Statistics of International Trade, Administrative Records, and Prices, INEGI.)

datasets. The results for the *maquiladora* sector are quite distinct from those for the non-*maquiladora* sector. The point estimates on initial log employment for the 1993-1997 period 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]

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 the greater differential wage changes between larger and smaller plants in the peso-crisis period in the *maquiladora* sector suggests the outsourcing hypothesis is unlikely to explain the differential wage changes we observe in 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]).

[44]This conclusion is reinforced by two additional findings. First, using data from the Mexican social security agency that covers a set of non-tradable sectors (construction, transportation, retail, and service) in addition to the manufacturing (tradable) sector, Appendix Table IV provides evidence of a significantly greater difference in differential trends in manufacturing than in the non-tradable sectors. For further details, see Appendix II and Kaplan and Verhoogen [2006]. Second, Kandilov [2005, 2007], 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 non-subsidized 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.

[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-$\lambda$ plants to be the most likely candidates. We have seen that there is no evidence that higher-$\lambda$ plants raised their white-collar employment share more than lower-$\lambda$ plants in the same industry. Hence it seems unlikely that shifts between activities of different skill intensities are driving the differential wage changes.

## 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 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 datasets, as well as all appendix tables and figures, are in Appendix II (online at http://www.columbia.edu/∼ev2124/).

## I.A EIA 1993-2001 and 1984-2001 Panels

Employment (white-collar, blue-collar, total) = average yearly employment for non-production 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 is has a formal training program, = 0 otherwise.

Turnover rate = 100*(2*(separations + new hires over previous 6 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 4-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 previous week, in 16 cities in original (1987) ENEU sample.[46] The wage and average schooling calculations use the sampling weights reported by INEGI.

### 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 $\geq 10$ percent foreign ownership, and 4-digit industry effects. I then recovered the predicted values $x'\hat{\beta}$, deviated from 6-digit industry means, and standardized the variable to have variance 1.

[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.

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, then deviated from 6-digit industry means and standardized to have variance 1.

First principal component: I deviated the co-variates used for the ISO 9000 index from 6-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 6-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.

Since 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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[47] I am indebted to Matthias Schuendeln 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.

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**Table I**
**Summary Statistics by Export Status**

| A. EIA 1993-2001 Panel, 1993 | Non-exporters | Exporters | All |
|---|---|---|---|
| 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) | |
| share with foreign ownership | 0.09 | 0.30 | 0.15 |
| | (0.01) | (0.01) | (0.01) |
| N | 2316 | 747 | 3263 |

| B. EIA-ENESTyC Panel | Non-exporters | Exporters | All |
|---|---|---|---|
| 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 parentheses. Exporter means export sales > 0. Has foreign ownership means FDI share > 0. Data on FDI from 1994; other data from EIA Balanced panel from 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/US$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 non-exporters (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 datasets in Section IV of text and Appendix II (online).

**Table II**
**Baseline Estimates, EIA 1993-2001 Panel**

| A. Cross-sectional regressions, 1993 | | export share of sales | log(white-collar wage) | log(blue-collar wage) | log(wage ratio) | log(K/L ratio) | white-collar emp. share |
|---|---|---|---|---|---|---|---|
| | | (1) | (2) | (3) | (4) | (5) | (6) |
| | 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] |
| | R2 | 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) | Δ log(white-collar wage) | Δ log(blue-collar wage) | Δ log(wage ratio) | Δ log(K/L ratio) | Δ (white-coll. emp. share) |
|---|---|---|---|---|---|---|---|
| | | (1) | (2) | (3) | (4) | (5) | (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] |
| | R2 | 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] |
| | R2 | 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] |
| | R2 | 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] |
| | R2 | 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. (Co-variate at left; dependent variables at top, with changes in Panel B over period at left.) All regressions include 205 industry (6-digit) and 32 state dummies. IV regressions instrument log domestic sales in 1994 and 1998 with values from previous year. N=3263 for all regressions. Variable definitions in Appendix I. Further details on dataset in Section IV of 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.

**Table III**
**Estimates from EIA 1984-2001 Panel**

| | | Δ (export share of sales) | Δ log(white-collar wage) | Δ log(blue-collar wage) | Δ log(wage ratio) | Δ log(K/L ratio) | Δ (white-coll. emp. share) |
|---|---|---|---|---|---|---|---|
| | | (1) | (2) | (3) | (4) | (5) | (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] |
| | R2 | 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] |
| | R2 | 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] |
| | R2 | 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] |
| | R2 | 0.200 | 0.193 | 0.217 | 0.163 | 0.188 | 0.146 |

Notes: Table reports coefficients on log domestic sales for 24 separate regressions. (Co-variate at left; dependent variables at top, with changes over period at left.) All regressions include 205 industry (6-digit) and 32 state dummies (coefficients omitted). N=1114 for all regressions. Variable definitions in Appendix I. Further details on dataset in Section IV of text and Appendix II (online). Robust standard errors in brackets. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

**Table IV**
**Estimates from EIA-ENESTyC Panel**

| **A. Cross-sectional Regressions, 1993** | | ISO 9000 certification | white-collar avg. schooling | blue-collar avg. schooling | has formal training | turnover rate | accident rate | absentee rate |
|---|---|---|---|---|---|---|---|---|
| | | (1) | (2) | (3) | (4) | (5) | (6) | (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 |
| | R2 | 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 | Δ white-collar avg. schooling | Δ blue-collar avg. schooling | Δ has formal training | Δ turnover rate | Δ accident rate | Δ absentee rate |
|---|---|---|---|---|---|---|---|---|
| | | (1) | (2) | (3) | (4) | (5) | (6) | (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] |
| | R2 | 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] |
| | R2 | 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. (Co-variate at left; dependent variables at top, with changes in Panel B over period at left.) All regressions include dummies for 50 industries (4-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 dataset in Section IV of 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.

**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 | 1997-2001 | Difference | 1993-1997 | 1997-2001 | Difference | 1993-1997 | 1997-2001 | Difference | 1993-1997 | 1997-2001 | Difference |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (11) | (12) |
| Proxy for $\lambda$: | | | | | | | | | | | | |
| 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] |
| R2 | 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] |
| R2 | 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] |
| R2 | 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] |
| R2 | 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] |
| R2 | 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] |
| | 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] |
| R2 | | | | 0.129 | 0.086 | | 0.121 | 0.091 | | 0.085 | 0.075 | |

Notes: Table reports coefficients on log domestic sales for 54 separate regressions. (Co-variate at left; dependent variables at top, with changes over period at top.) All regressions include 205 industry (6-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 dataset in Section IV of 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.

**Table VI**
**Estimates for Cost of Capital, Publicly Listed Mexican Firms, 1989-2000**

| | | Dep. Var.: Δ cost of capital | | | |
|---|---|---|---|---|---|
| | | All Sectors | | Manufacturing only | |
| | | OLS | IV | OLS | IV |
| | | (1) | (2) | (3) | (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 |
| | R2 | 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 |
| | R2 | 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 |
| | R2 | 0.050 | 0.083 | 0.002 | 0.088 |

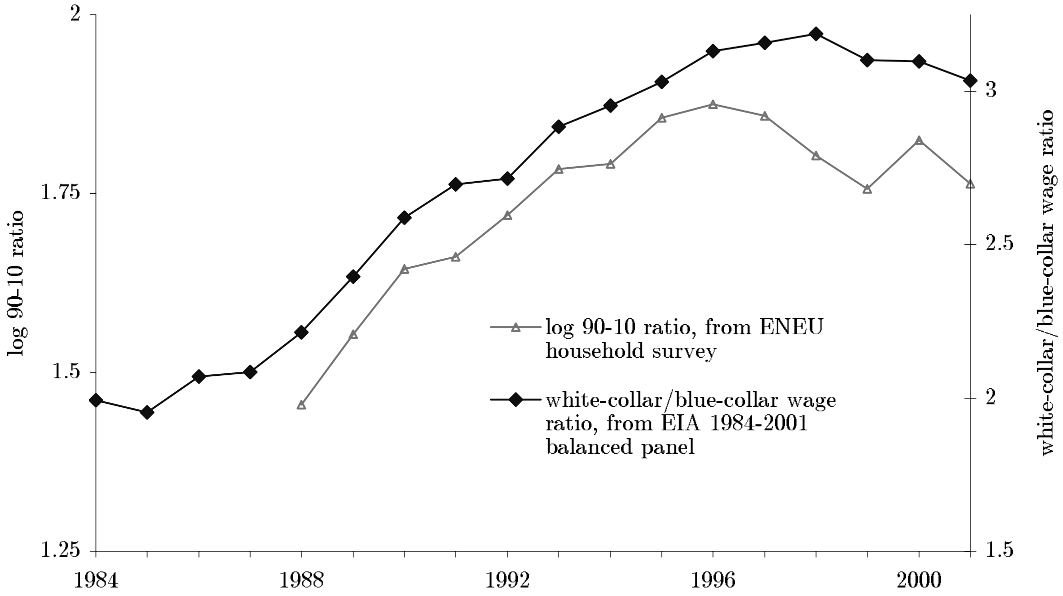
Notes: Table reports coefficients on initial log domestic sales for 12 separate regressions. (Co-variate at left; dependent variable at top, with changes over period at left.) Regressions for all sectors together include industry effects (6 coarse industry categories). IV regressions instrument log domestic 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). Dataset is unbalanced panel of publicly listed firms on Mexican stock market (Bolsa Mexicana de Valores) for 1989-2000. For further details on dataset, see Pratap and Urrutia (2004) and Aguiar (2005). Robust standard errors in brackets. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

**Table VII**
**Comparing Non-Maquiladora and Maquiladora Sectors**

| | | EIA 1993-2001 Panel (Non-Maquiladora Sector) | | | EMIME Panel (Maquiladora Sector) | | |
|---|---|---|---|---|---|---|---|
| | | Δ log(avg. hourly wage) | Δ log(white-collar yearly earnings) | Δ log(blue-collar yearly earnings) | Δ log(avg. hourly wage) | Δ log(white-collar yearly earnings) | Δ log(blue-collar yearly earnings) |
| | | (1) | (2) | (3) | (4) | (5) | (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] |
| | R2 | 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] |
| | R2 | 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 | | 3263 | 3263 | 3263 | 1088 | 1088 | 1088 |

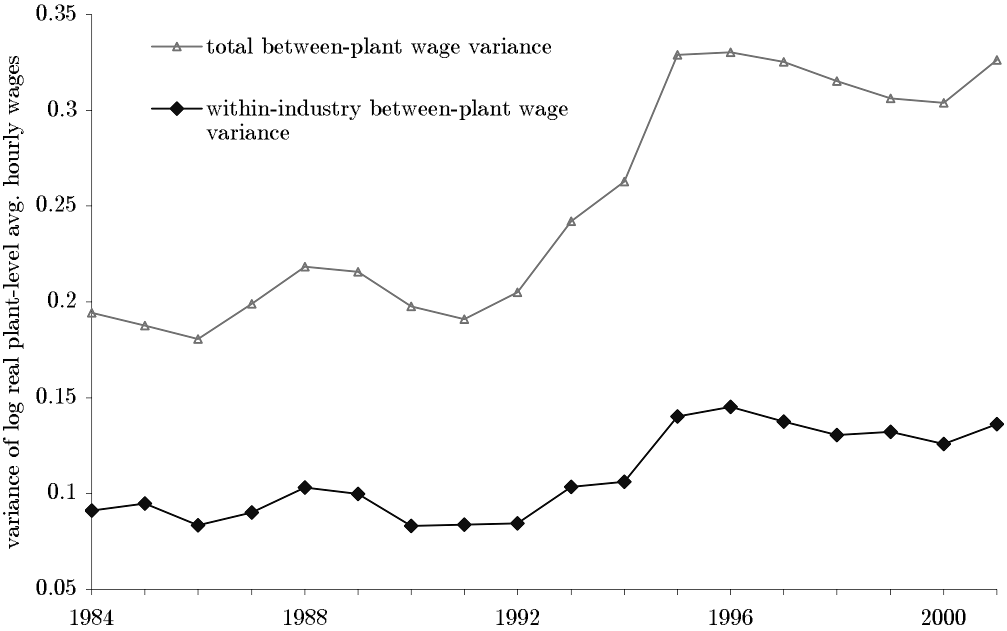
Notes: Table reports coefficients on initial log employment for 12 separate regressions. (Co-variate at left; dependent variables at top, with changes over period at left.) Columns (1)-(3) use the EIA 1993-2001 Panel and include 205 industry (6-digit) and 32 state dummies. Columns (4)-(6) use the EMIME Panel over same period and include 12 industry and 32 state dummies. Variable definitions in Appendix I. Further details on datasets in Section IV of 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.

**Figure I**
**Wage Inequality, 1984-2001**



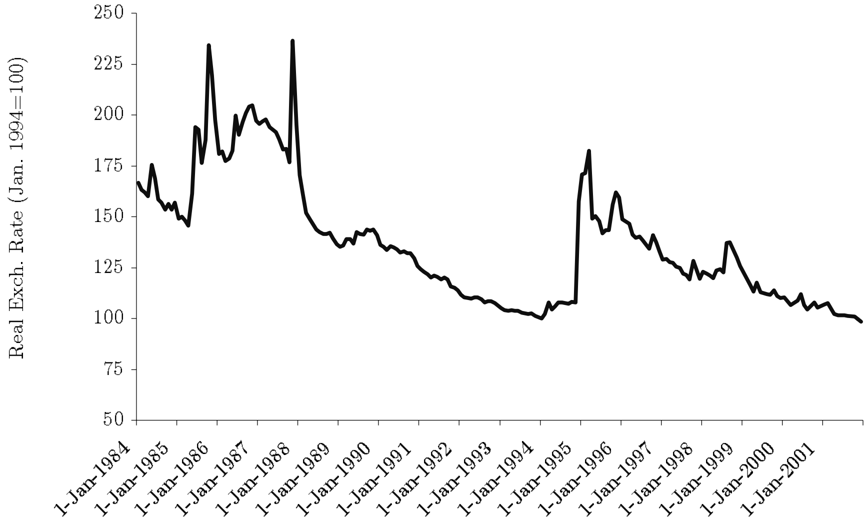

Notes: Log 90-10 ratio is for real hourly wages from ENEU household survey. White-collar/blue-collar ratio is for hours-weighted averages of hourly wages for non-production workers and production workers in EIA 1984-2001 balanced panel of 1114 plants. Variable definitions in Appendix I. Further details on datasets in Section IV of text and Appendix II (online).

**Figure II**
**Wage Variance, EIA 1984-2001 Balanced Panel**



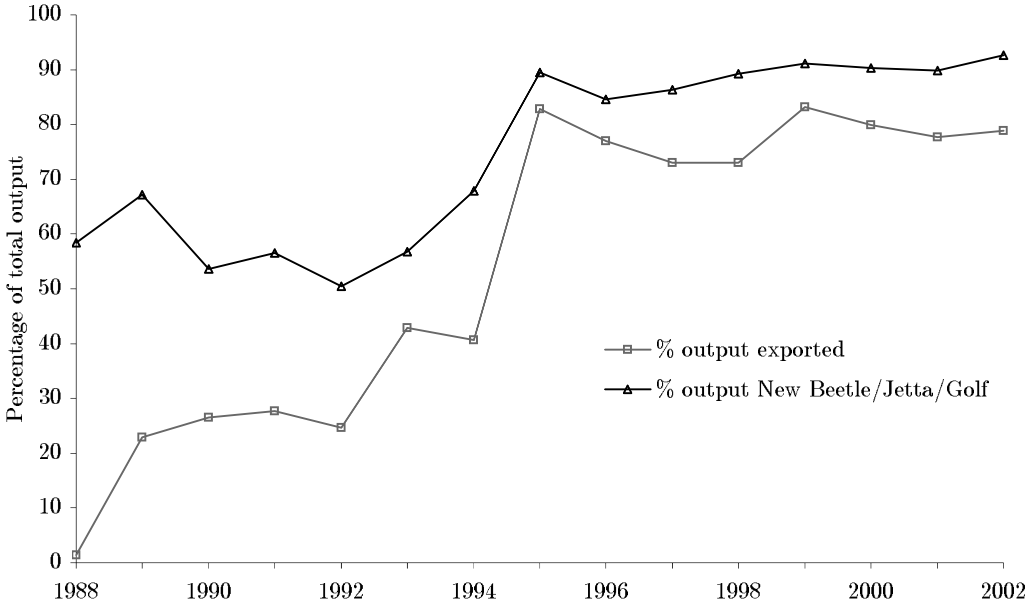

Notes: Total wage variance is hours-weighted variance of the log plant-average real hourly wage in EIA 1984-2001 balanced panel of 1114 plants. Within-industry-year variance is hours-weighted variance of residual from regression of the log plant-average real hourly wage on a full set of industry-year dummies (205 industries * 18 years) in EIA 1984-2001 balanced panel. Variable definitions in Appendix I. Further details on dataset in Section IV of text and Appendix II (online).

**Fig. III**
**Real Exchange Rate, 1984-2002**



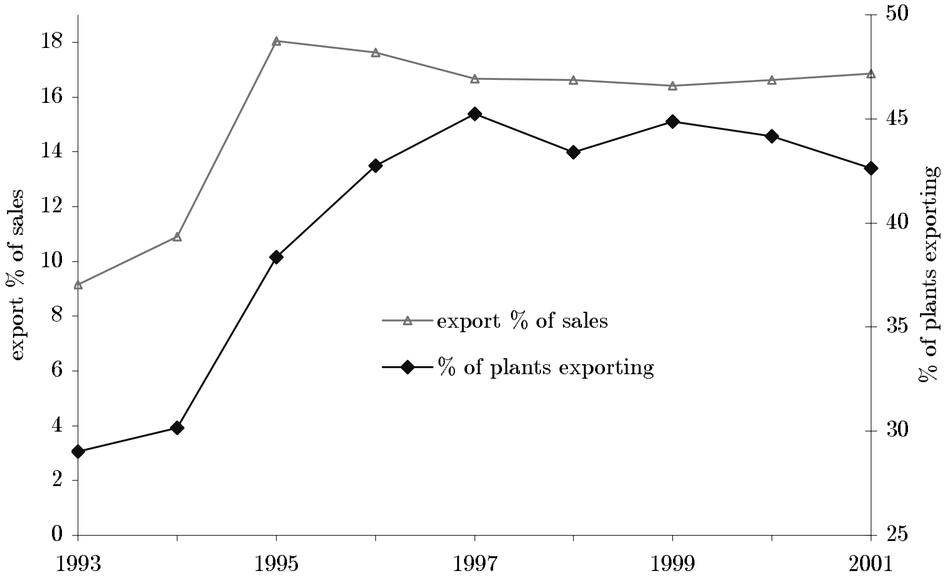

Notes: Real exchange rate calculated as RER = $e$ * CPI(US)/CPI(Mex), where $e$ is the peso/US$ nominal exchange rate. Data from IMF International Financial Statistics.

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




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

**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 dataset in Section IV of text and Appendix II (online).

**Figure VI**
**Average Quality**

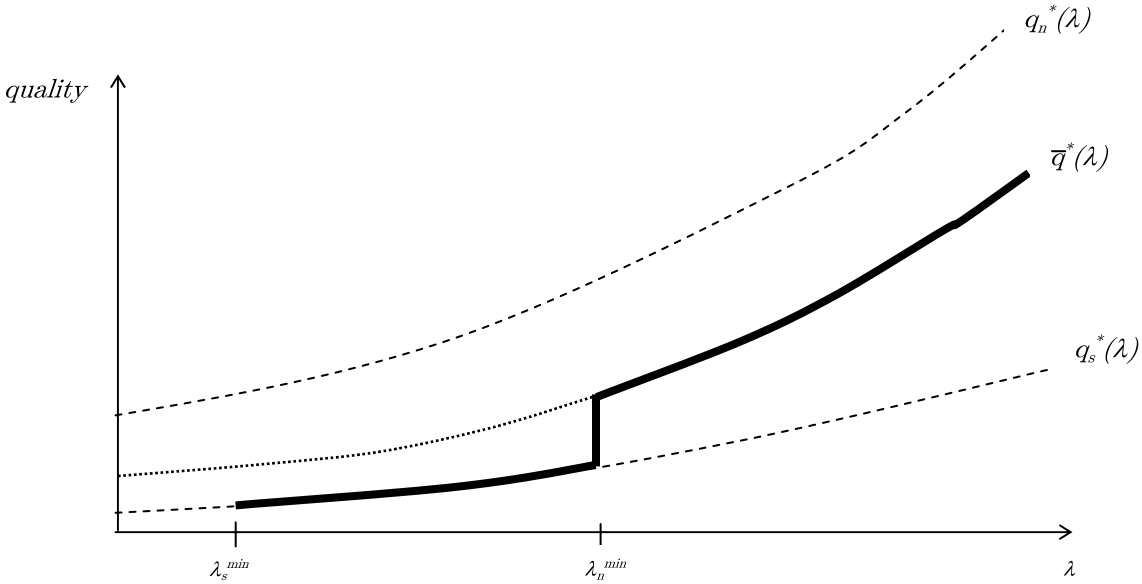


**Figure VII**
**Response to Exchange-Rate Devaluation**

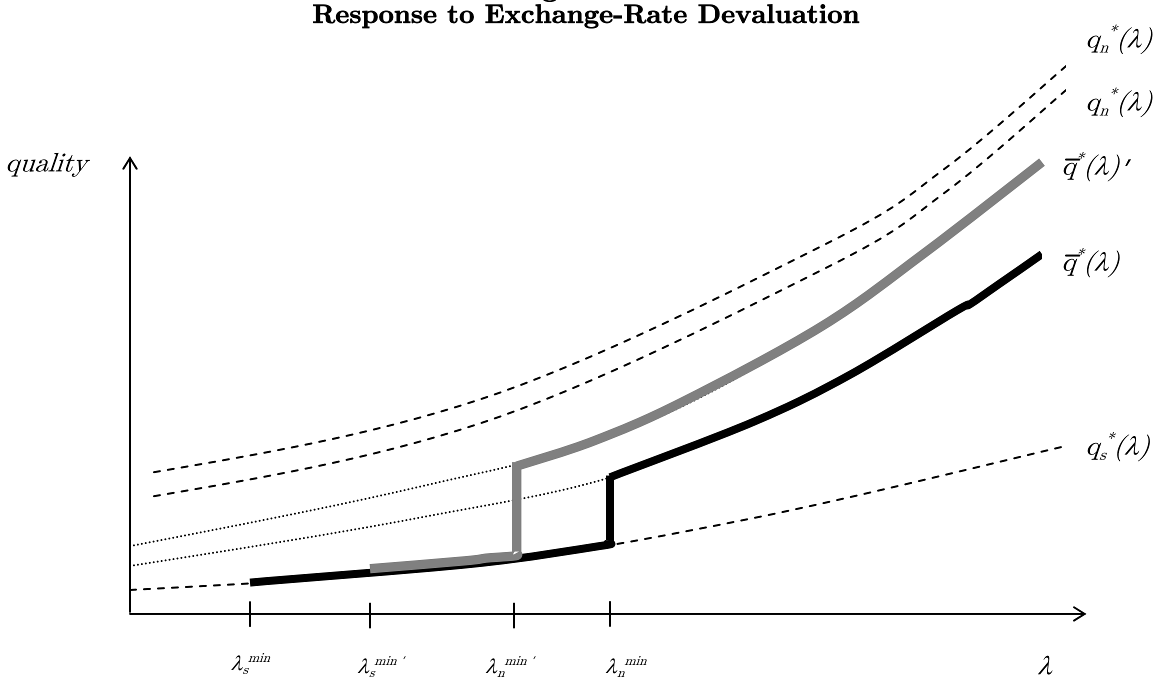


**Figure VIII**
**Change in Average Quality in Response to Devaluation**