Trade, Quality Upgrading and Wage Inequality in the Mexican Manufacturing Sector: Theory and Evidence from an Exchange-Rate Shock

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Abstract

This paper proposes a new model of the link between expanding trade and rising wage inequality in developing countries, and investigates its causal implications in a newly constructed panel of Mexican manufacturing establishments. In a theoretical setting with heterogeneous firms and quality differentiation, only the most productive firms in a developing country like Mexico enter the export market, and they produce a better-quality good for export than for the domestic market in order to appeal to richer developed-country consumers. Producing high-quality goods requires paying high wages both to white-collar and to blue-collar – but especially to white-collar – employees. An increase in the incentive for developing-country producers to export generates differential quality upgrading within industries, as more-productive firms increase exports and produce a greater share of high-quality goods, while less-productive firms remain focused on the domestic market. This process raises wage inequality both between firms and within the firms that upgrade. The empirical part of the paper uses a major exchange rate shock – the Mexican peso crisis of late 1994 – to test this causal mechanism. I find robust evidence that during the years of the crisis initially more-productive plants increased white-collar wages, blue-collar wages, and the relative wage of white-collar workers as compared to initially less-productive plants in the same industry. This pattern is absent in the periods before and after the crisis years. The results thus support the hypothesis that differential quality upgrading induced by the exchange rate shock contributed to the increase in wage inequality in Mexico in the mid-1990s.

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1 Introduction

Studies have found a coincidence between expanding trade and increasing wage inequality in many developing countries, including Argentina, Brazil, Chile, Colombia, Costa Rica, Malaysia, Morocco, Taiwan, and Uruguay.[1] Mexico is not an exception to the general pattern. Beginning with its unilateral trade liberalization in the mid-1980s, Mexico saw rapid increases both in the volume of trade and in the relative wage of skilled workers (Cragg and Epelbaum, 1996; Hanson and Harrison, 1999). The rising trend in inequality continued in the mid-1990s, with the implementation of the North American Free Trade Agreement (NAFTA) and a severe currency devaluation – the peso crisis – in December 1994. From 1994 to 1995, as the growth of manufacturing exports accelerated from 21% to 32% per year, the difference between wages of full-time male workers at the 90th percentile and 50th percentiles of the wage distribution rose by 6.7%.[2]

From the perspective of standard trade theory, the coincidence of expanding trade and rising wage inequality in Mexico is puzzling. The simplest version of the Hecksher-Ohlin model of trade predicts that wage inequality will fall in a country abundant in unskilled labor, as production shifts toward unskilled-labor-intensive industries, raising the demand for unskilled workers. More sophisticated Hecksher-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 focus exclusively on between-sector shifts as the mechanism through which trade affects labor markets[4], they can only explain a rise in inequality if trade causes a shift toward skilled-labor-intensive sectors. This condition is violated in the Mexican case. Figures 1a and 1b plot the change in employment over the period 1988-1998 by 4-digit manufacturing industries against measures of the level of skill- and capital-intensity. Both figures reveal a clear shift toward industries intensive in the use of unskilled labor, consistent with the simplest Hecksher-Ohlin story. The apparent inability of conventional trade theories to explain the rising relative wage of skilled workers in developing

[1] See Galiani and Sanguinetti (2003) on Argentina; Blom et al (2001), Green et al (2001), and Pavcnik et al (2002) on Brazil; Robbins (1994), Gindling and Robbins (2001) and Beyer et al (1999) on Chile; Robbins (1996b) and Attanasio et al (2002) on Colombia; Robbins and Gindling (1996) and Gindling and Robbins (2001) on Costa Rica. Robbins (1996a), Slaughter (2000), IADB (2002), Harrison and Hanson (1999), and Kremer and Maskin (2003) provide overviews.

[2] The calculations are based on the Encuesta Nacional de Empleo Urbano (ENEU), a household survey similar to the U.S. Current Population Survey. The calculation is for full-time male workers, ages 12-64, living in one of the 16 cities in the original ENEU sample. In these data, inequality declined slightly in the late 1990s, beginning in 1997. See Esquivel and Rodriguez-Lopez (2003), Hanson (2002) and Robertson (2000) for discussions of recent trends.

[3] We might expect a rise in the relative wage of skilled labor in a country like Mexico, for instance, if the country opens trade simultaneously with the U.S. and another country that is even more unskilled-labor-abundant (e.g. China) (Davis, 1996; Wood, 1997); if relatively unskilled-labor-intensive industries are more protected prior to liberalization (Revenga, 1997; Goldberg and Pavcnik, 2001; Feliciano, 2000); or if the production of maize is characterized by a factor intensity reversal (Larudee, 1995).

[4] An exception is the outsourcing model of Feenstra and Hanson (1996), to which I return below.

countries has led many observers to conclude that it must be due to factors unrelated to trade such as skill-biased technical change (Esquivel and Rodriguez-Lopez, 2003; Meza, 1999) or policy changes like deregulation and privatization that tend to accompany trade liberalization (Behrman et al, 2000).[5]

This paper proposes a new model of the link between trade and wage inequality in developing countries and tests its causal implications in a newly constructed panel of Mexican manufacturing plants. In the model, firms are heterogeneous in an underlying productivity parameter (which can be interpreted as technical know-how or entrepreneurial ability) and goods are differentiated in quality. Within each industry, only the most productive firms in a developing country like Mexico enter the export market, and they produce a better-quality good for export than for the domestic market in order to appeal to richer developed-country consumers. Producing high-quality goods in turn requires paying high wages to both white-collar and blue-collar employees, but especially to white-collar employees. An increase in the incentive to export leads to *differential quality upgrading* within industries: initially more-productive firms increase exports and shift toward greater production of higher-quality goods; initially less-productive firms remain solely in the domestic market and undertake no such upgrading. This process leads initially more-productive firms to raise wages across occupational categories, to raise the relative wage of white-collar workers, and to increase capital-intensity relative to initially less-productive firms within the same industry.

The empirical part of the paper uses the peso devaluation of December 1994 to test the prediction of differential quality upgrading within industries. My econometric strategy issues directly from the theoretical model. I generate a proxy for the unobserved know-how parameter using data from before the exchange rate shock, following one of two methods. In the first, I use the log of domestic sales deviated from industry means, which the model suggests will be proportional to the know-how parameter. In the second, I take the first principal component of a number of plant characteristics which the model suggests are correlated with the unobserved parameter. Once the proxy has been generated, I simply regress changes in plant behavior over the crisis period on the level of the proxy from before the shock. I find robust evidence that over the 1993-1997 period initially more productive plants increased the export share of sales, raised wages for both white-collar and blue-collar workers, raised the relative wage of white-collar workers, and increased the capital-labor ratio to a greater extent than initially less productive plants. Using an auxiliary dataset, I also find that over the 1994-1998 period initially more-productive plants were more

[5] Several papers have extended the technical-change argument to include the possibility that trade accelerates the process of skill-biased technical change, which then generates a rise in the skill premium (Acemoglu 1999; Pissarides 1997, Thoenig and Verdier 2002).

likely to acquire ISO 9000 certification, an international production standard commonly associated with high product quality. As a further test, I re-estimate the same model on periods before and after the peso crisis during which a currency devaluation did not intervene. I find essentially no evidence of quality upgrading in the 1989-1993 or the 1997-2001 periods. The only years in which I find similar (but weaker) results are 1986-1989, a period that itself was characterized by a significant depreciation of the peso. The results thus provide strong support for the hypothesis that differential quality upgrading due to the exchange rate shock contributed to the increase in wage inequality in Mexico in the mid-1990s.

In formalizing the mechanism of differential quality upgrading, the model draws on four elements from the existing theoretical literature.[6] The first element is monopolistic competition with heterogeneous producers, in the spirit of Melitz's (2003) extension of the seminal papers by Krugman (1979, 1980).[7] The second element is a micro-founded form of differentiation in product quality, drawn from Anderson et al's (1992) extension of the discrete-choice theory of McFadden (1978, 1981). The third element is an asymmetry in consumer demand between two countries, called North and South. In particular, consumers in North are assumed to be richer and hence more willing to pay for quality than consumers in South.[8] The fourth element is an O-ring production function from Kremer (1993) and Kremer and Maskin (1996), in which the production of high-quality goods requires highly skilled workers across occupational categories and is more sensitive to the skill of white-collar workers than to that of blue-collar workers.[9] The main contribution of the model is to synthesize these previously separate ideas and to elucidate a new mechanism through which trade-related shocks may affect outcomes at the plant level: shifts in the within-plant product mix between goods of different qualities destined for different markets.

The empirical part of the paper is related to a growing empirical literature on international trade and the behavior of individual plants. Studies in this literature have tended to find little evidence of within-plant changes in behavior in response to exposure to international markets.

[6] The idea that trade with developed countries induces quality upgrading in industrial firms in developing countries is present in, for instance, Morawetz (1981), Gereffi (1999) and Lopez (2003). The contribution of the theoretical part of this paper is to work out the logic of one rigorous version of the argument.

[7] Bernard, Eaton, Jensen and Kortum (2003) present an alternative model that also allows for plant-level heterogeneity. This paper is also related to a small literature on trade and wages under monopolistic competition, which assume symmetric countries and focus on scale effects as the mechanism through which trade affects relative wages (Dinopoulos and Segerstron, 1999; Dinopoulos et al, 2001; Epifani and Gancia, 2002; Eckholm and Midelfart Knarvik, 2001; Yeaple, 2003).

[8] This idea dates back at least to Linder (1961), and has been further elaborated upon by Shaked and Sutton (1982), Flam and Helpman (1987), Stokey (1991), Copeland and Kotwal (1996), Murphy and Shleifer (1997), and Brooks (2003). Hallak (2003), which I became aware of after the first draft of this paper had been circulated, has recently incorporated it into a model of trade under imperfect competition.

[9] Kremer and Maskin (2003) present a simple O-ring model to explain the coincidence of expanding trade and rising wage inequality in developing countries, but through matching of Southern and Northern workers in multinational firms, rather than through the quality-upgrading mechanism emphasized in this paper.

An emerging consensus in the literature on trade and productivity is that trade raises aggregate productivity by shifting production toward more-productive plants, rather than by improving productivity within plants (Clerides, Lach, and Tybout, 1998, Bernard and Jensen, 1999).[10] Studies that have examined the effects of industry-level changes in trade policy on plant-level changes in wage and employment decisions have found what many observers have described as puzzlingly small effects, in some cases despite large changes in tariffs or other trade policy measures.[11] In contrast, this paper finds strong, robust effects of a shock to the incentive to export on within-plant behavior. The strength of the results may be due to two advantages of using an exchange rate shock, rather than changes in trade policy, as the source of exogenous variation.[12] First, unlike most changes in trade policy, the shock was largely unexpected. Second, the shock was big. The peso lost approximately half of its value in a matter of days at the end of 1994, a change that dwarfs average tariff changes under NAFTA. This is especially important if we are interested in shocks to the incentive to export to a rich country. Tariff reductions by developed countries are typically small, in part because their tariffs tend already to be low. The challenge in making use of an exchange-rate shock is to identify a source of variation in its impact at the plant level. A main empirical contribution of this paper is to show how to use the interaction of the exchange-rate shock and pre-existing heterogeneity within industries to identify the heterogeneous effects of the shock at the plant level.

The main alternative theory of the link between expanding trade and rising wage inequality in Mexico is the outsourcing hypothesis of Feenstra and Hanson (1996, 1997). In their model, production in each industry is divided into phases of different skill intensities. Capital accumulation in Mexico leads to the outsourcing of progressively more skill-intensive phases to Mexico within each industry, raising the overall demand for skill in a way that does not show up in aggregate between-industry shifts like the ones illustrated by Figure 1. Their model is arguably best applied to the Mexican *maquiladoras*, plants legally committed to producing under subcontract for the export market.[13] Both the theory and the empirical work in this paper are primarily concerned with the non-maquiladora sector, and in this sense the two arguments are complementary. But

[10] For a review of the literature on trade and productivity in developing countries, see Tybout (2000).

[11] Levinsohn sums up his investigations in Chile with the statement: "Try as one might, it is difficult to find any differential employment response to the trade liberalization." See also Currie and Harrison (1997) and Harrison and Hanson (1999).

[12] This is not the first paper to use an exchange-rate shock identify an effect of international competition. Previous studies include Revenga (1992), Abowd and Lemieux (1993), and Bertrand (1999). What is new in this paper is the use of within-industry heterogeneity in the impact of such a shock to estimate its effects.

[13] Maquiladoras are plants participating in a government program that until recently required them to export nearly all of their output in exchange for relief from tariff duties on the value of imported inputs. In Mexico, the participants in this program are referred to as *maquiladoras de exportacion* (exporting maquiladoras). The word *maquiladora* (or *maquila* for short) is used more generally to apply to any plant producing under sub-contract. I will use the term only to refer to the former group.

I present two types of evidence that favor the quality-upgrading hypothesis over the outsourcing hypothesis as an explanation for increasing wage inequality in Mexico. First, plants in the non-maquiladora sector changed wages even in the absence of changes in the proportion of white- and blue-collar workers used in production, which suggests that the wage changes were not driven by shifts between activities of different skill-intensities, at least among non-maquiladora plants. Second, I present evidence (based on micro-data that were unavailable to Feenstra and Hanson) that maquiladoras are on average markedly *less* skill-intensive than the rest of the Mexican manufacturing sector. Although there may have been a shift toward more skill-intensive activities within the maquiladora sector, it appears that the first-order consequence of the expansion of the sector was an increase in the demand for less-skilled labor.

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 3 is the theoretical part of the paper. Section 3.1 develops the model for a closed economy, Section 3.2 sets out the two country version and relates the model to the observable variables available in the Mexican plant-level data, and Section 3.3 derives the comparative-static implications of an exchange rate shock for these observables. Section 4 is the empirical part. Section 4.1 describes the datasets and review broad patterns in the data, Section 4.2 discusses my econometric strategy, Section 4.3 presents the results, and Section 4.4 presents additional findings from the auxiliary dataset. Section 5 concludes.

2 Background and Brief Case Study

On Dec. 20, 1994, running short of reserves to defend its exchange-rate target, the new administration of Ernesto Zedillo announced that it would raise the ceiling on its exchange-rate band by 15%. This set off a speculative attack and investor flight from the peso, led by domestic Mexican investors. The currency promptly lost approximately 50% of its value, precipitating a major recession in Mexico. As Figure 2a illustrates, the aggregate price level in Mexico relative to the price level in U.S. dropped sharply with the devaluation and recovered only slowly thereafter.[14] Labor costs in dollar terms followed a similar pattern. Figure 2b plots the average wage level of full-time male workers with 9 years of education over the period 1993-1999, both in current U.S. dollars and in current pesos.[15] Interestingly, the nominal wages of manufacturing workers appear to have been almost entirely unaffected. In dollar terms, by contrast, the average wage

[14] It is interesting to note that there was a similar real depreciation of the peso in 1986. The 1986-1989 period will provide a corroborating test the differential quality upgrading hypothesis in the empirical section below.

[15] The data, again, are from the ENEU. See footnote 2.

for a male full-time worker with a junior high education fell from approximately \$1.50 per hour to approximately \$.90 per hour from 1994 to 1995, rising back only to \$1.10 per hour by 1999.

It is worth emphasizing that the peso crisis was a much larger shock than NAFTA, which had taken effect the previous January. Mexico's main round of trade liberalization came in the mid-1980s with its unilateral abandonment of import-substituting industrialization and entrance into the General Agreement on Tariffs and Trade. By 1993, almost all quotas and other non-tariff barriers had been removed, and approximately 95% of all imports into Mexico were covered by tariffs of 20% or less. On the U.S. side, tariffs were initially even lower: approximately 80% of imports into the U.S. were covered by tariffs of 5% or less. Moreover, the implementation of NAFTA did not represent a sudden shock. A majority of commodities were assigned phase-out schedules of five or more years. 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 1980s.[16]

How did the manufacturing sector respond to the crisis? To provide a concrete point of reference for the theoretical discussion, consider the example of one important plant in Mexico, the Volkswagen auto plant in Puebla, about two hours south of Mexico City. The Puebla plant is the sole producer in the world of one of Mexico's highest-profile exports to the U.S., the New Beetle. (It is also the sole producer for the U.S. market of another well-known model, the Jetta.) It comes as something of a surprise, then, that it is relatively rare to see a New Beetle in Mexico. The streets and highways of Mexico are dominated by the old model, the Original Beetle, known in Mexico as the *Sedan* (or, more affectionately, the *Vocho*), which was produced in the same plant until July 30, 2003.

The Original Beetle and the newer models, the New Beetle and the Jetta, represent a stark case of quality differentiation between goods produced primarily for the domestic market and goods produced primarily for export. 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 foam and partly of coconut fibers, a cheaper substitute. These and other quality differences are reflected in the prices of the models: the New Beetle and the Jetta sell for approximately US\$17,750 and US\$15,000 in Mexico, and roughly comparable prices in the U.S. The Original Beetle until recently sold for approximately US\$7,500 in Mexico.

The example of the Volkswagen plant is especially useful because it is possible to follow changes

[16] In unreported results, I find little effect of tariff changes at the industry level on plant level employment, sales, or wages, even when allowing for within-industry heterogeneity in plant-level responses, along the lines of the my empirical approach in this paper. These non-results reinforce the argument that the important shock in this period was the peso crisis.

in production by product line and see how the shock of the peso crisis affected the product mix within the plant. At the time of the crisis, the New Beetle had not yet been introduced. (It was introduced in 1998.) The plant was producing the Original Beetle, destined primarily for the domestic market, and the Jetta and the Golf (a model from which the New Beetle borrows the chassis and many underlying components), both destined primarily for export. Figure 3a plots output for the domestic market, output for the export market and total output over the period 1988-2002. The effect of the peso crisis is evident: output for the domestic market shrank precipitously and output for the export market rose in 1995. The net effect on output was a small decline. This shift was accompanied by a shift in the within-plant product mix away from the Original Beetle and toward the Jetta and Golf. As Figure 3b illustrates, the Original Beetle accounted for a significant majority of production for the domestic market, and the drop in domestic-oriented production primarily reflected a drop in production of the Original Beetle. Figure 3c illustrates that almost no Original Beetles were exported, either before or after the crisis. The increase in exports reflected an increase in production for export of Jettas and Golfs and, later, of New Beetles. Figure 3d capture the key facts: exports as a share of total production rose from 40% in 1994 to 80% in 1995, and this was accompanied by a shift in product mix from a cheaper, lower-quality model, the Original Beetle, to more expensive, higher-quality models, the Jetta and Golf.

What consequences did this shift in product mix have inside the plant? The most striking characteristic of the Puebla plant, until recently, was the juxtaposition of the production lines for the New Beetle and Jetta, which rely 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, which had been in use in Germany since the 1950s. The contrast was perhaps most evident in the welding area (*linea de soldadura*), which I visited in May 2003. The conveyor belt on the Original Beetle line had been in continuous operation since 1967. The welding was done by hand, with sparks flying, and line-workers banged irregularities into shape with hammers. Under the same roof, perhaps twenty yards away, the welding for the Jetta body was and continues to be performed entirely by robots; the labor requirements are limited to engineers to program the robots, and skilled maintenance workers to repair the machines in case of mechanical failure. One consequence of the shift in product mix, then, was 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 rather because of a shift toward greater reliance on technologies that were already in use in the plant.

The employees of the Puebla plant are members of an activist union with a history of militancy,[17] and a collective-bargaining agreement has constrained the ability of management to adjust its labor practices in response to the changing product mix. The shift of production toward the Jetta and Golf/New Beetle lines nevertheless appears to have had consequences for skill demands and wages. Demand has increased for *especialistas* (specialists), the skilled production workers who maintain the automated machines such as the robots in the welding area. A typical production worker (*técnico*) in the plant has a junior high school (*secundaria*) education. The especialistas, by contrast, are graduates of a 3-year post-secundaria vocational school that the company administers on the plant grounds. The starting wage for a técnico under the 2002-2004 collective bargaining agreement is 122.87 pesos (US$11.18)[18] per day. The starting wage for an especialista is 195.06 pesos (US$17.74) per day. I was unable to persuade the company to share detailed data on the number of workers of each type and each salary level working on each production line, and hence am not able to make definitive statements about the changing skill composition and earnings of the workforce, but conversations with both the former director of Human Resources and the president of the Volkswagen union suggest that the relative demand for especialistas has risen with the share of production making intensive use of automated technology. At the white-collar level, it appears that the use of software engineers, highly skilled relative to the supervisors on the Original Beetle line, has increased as well.

Does the example of Volkswagen generalize to the manufacturing sector as a whole? Figures 4a-4c show that, as in the Volkswagen example, the peso crisis induced a significant shift toward production for export. Figure 4a plots total exports, total domestic sales, and total sales from the EIA 1993-2001 balanced panel. The increase in exports and the decline in domestic sales are evident. It is notable that the two effects largely appear to have offset each other, in this sample of large plants. Figure 4b plots exports as a percentage of total sales; the shift toward exporting appears in even sharper focus. Figure 4c shows that the increase in the volume of exports was accompanied by an increase in the fraction of establishments exporting, from 30% in 1993 to approximately 45% in 1997. The U.S. has long been the overwhelmingly important destination market for Mexican exports. In 1992, the U.S. was the recipient of 80.6% of Mexican exports; by 2000, the percentage had risen to 88.7%. Thus the increase in exports largely represents an increase in sales of Mexican plants on the U.S. market. The Volkswagen example suggests by extension that the increase in exports to the U.S. was likely to have been accompanied by an increase in the average quality of goods produced and an upgrading of the industrial workforce in

[17] The VW union, the Sindicato Independiente Volkswagen, is independent of the main Mexican labor confederation, the Confederacion Mexicana de Trabajadores (CTM), and has a long history of strikes, the most recent in 2001. For a concise history, see Hanson and Shapiro (1995).

[18] At the Oct. 1, 2003 exchange rate of 10.99 pesos/dollar.

exporting plants.

3 Theory

This section develops a model of trade, quality upgrading and wage inequality that formalizes the salient features of the upgrading process as it has played out at Volkswagen and, anecdotal evidence suggests, across broad segments of the Mexican manufacturing sector. For expositional purposes, I begin with the less notation-intensive case of a single, closed economy and then move on to the more notation-heavy two-country version. I begin by deriving the demand curve facing each firm and the solution to firms' optimization problems taking entry decisions as given, and then solve for entry, given these optimizing decisions.

It is worth emphasizing at the outset that the model has a number of special features. I use specific functional forms and ignore dynamic issues. Incomes are assumed to be homogeneous within countries and heterogeneous only across countries. The model is partial-equilibrium, implicitly focused on a single differentiated-goods industry that is small relative to the economy as a whole in either country. Nonetheless, my hope is that these sacrifices of mathematical elegance on one hand and realism on the other are justified by the extent to which the model achieves two goals: to bring the insights of new trade theory – which, in assuming symmetric countries, has implicitly been focused on integration among developed countries – to bear on the distinctive experience of developing countries; and to tie the model directly to an empirical approach capable of estimating its causal implications in real data on Mexican plants.

3.1 One-Country Model

3.1.1 Demand

There are N statistically identical consumers, indexed by i. Each is assumed to buy one unit of some variety, with varieties indexed by j, for $j = 1..J$. Each has the indirect utility function:

$$V_{ij} = \theta q_j - p_j + \varepsilon_{ij} \tag{1}$$

The variables q_j and p_j are the quality (observable to the consumer) and price of variety j, respectively. The parameter θ captures consumers' willingness to pay for quality. It can be interpreted as a function of income. If richer consumers have identical utility functions as poorer consumers but the marginal utility of income is declining, they will be willing to pay more for a given level of quality.[19] I assume that θ is constant across consumers within a country, and differs only across countries. I treat it as a fixed parameter, and abstract from changes in consumers' willingness to pay for quality arising from income changes due to the peso crisis.

The individual-specific random-utility terms, ε_{ij}, are assumed to have identical double-exponential distributions and to be independent across goods and consumers. Except for the quality term, the set-up is a standard multinomial-logit model of consumer demand (McFadden 1978, 1981). This particular specification appears in Anderson et al (1992).[20] Appendix A.1 shows that this specification yields the following expected demand for each good j:

$$E(x_j) = \frac{N \exp\left[\left(\theta q_j - p_j\right)/\mu\right]}{\sum_{t=1}^{J} \exp\left[\left(\theta q_t - p_t\right)/\mu\right]} \tag{2}$$

The parameter μ here captures the degree of differentiation between goods. As $\mu \to 0$, any quality or price difference between goods will be magnified, such that the most attractive good will capture all the demand, and the model approaches perfect competition.[21]

As Anderson et al (1992) point out, the resulting model 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 a higher probability. The advantage of this approach is that it allows us to consider large numbers of heterogeneous firms in a tractable way. This feature distinguishes the monopolistic-competition approach from standard models of differentiation along a single, vertical quality dimension (Gabszewicz and Thisse, 1979, 1980, Gabszewicz et al, 1981; Shaked and Sutton, 1982; Gabszewicz and Turrini, 2000) which quickly become intractable with more than two firms; from patent-race models in which a single firm captures the market for an entire industry (Grossman and Helpman, 1991a, 1991b); and from Ricardian-type vertical-differentiation models based on constant returns to scale and perfect competition (Falvey and Kierzkowski, 1987; Flam and Helpman, 1987; and Stokey, 1991), which do not specify the boundaries between individual firms and hence do not carry implications for firm- or plant-level

[19] Following Tirole (2000, p. 97 fn 1), if consumers have direct utility $U_{ij} = u(x_0) + q_j + \widetilde{\varepsilon}_{ij}$ where x_0 is the consumption of a non-differentiated numeraire good, then optimization yields the indirect utility function $\widetilde{V}_{ij} = u(y - p_j) + q_j + \widetilde{\varepsilon}_{ij}$. If p_j is small relative to the consumer's income, y, then a first-order expansion of the sub-utility function $u(\cdot)$ gives: $\widetilde{V}_{ij} = u(y) - p_j u'(y) + q_j + \widetilde{\varepsilon}_{ij}$. Let $\theta \equiv 1/u'(y), V_{ij} \equiv u'(y)\widetilde{V}_{ij}, \varepsilon_{ij} \equiv u'(y)\widetilde{\varepsilon}_{ij}$. The $u(y)$ term will drop out of the expression for aggregate demand, and is for that reason immaterial. We thus have (1).

[20] Anderson and de Palma (2001) extend the framework to the case of asymmetric firms choosing different quality levels, but do not apply it to the analysis of international trade. Hallak (2003) presents a similar specification in a trade context, but assumes goods have the same quality level within industries and uses the model to analyze the pattern of trade at the industry level, rather than within-industry heterogeneity.

[21] In this sense, the parameter μ is analogous to the constant elasticity of substitution parameter in the better-known Dixit-Stiglitz (1977) approach.

behavior.[22]

3.1.2 Production

There is a mass M of firms, heterogeneous in a productivity parameter λ distributed continuously over the interval $[0, \lambda^{\max}]$, with probability distribution $g(\lambda)$. As mentioned above, we can think of λ as representing technical know-how or entrepreneurial ability in the firm.[23] (I will use the terms know-how and productivity interchangeably to refer to this parameter.) I assume that know-how is a fixed characteristic of the firm,[24] and is not available (or is prohibitively costly) on the open market. The parameter λ uniquely identifies firms within the country.

I assume that production is governed by an O-ring production function. In Kremer's original O-ring paper (Kremer, 1993), the production process is divided into a fixed number of tasks, and output is a multiplicative function of the skill-levels of the workers employed in each task.[25] Kremer and Maskin (1996) extend this framework to the case where there are two categories of workers, which I will refer to as white-collar and blue-collar, and output is given by a Cobb-Douglas function with a larger coefficient on the skill of the white-collar worker. I adopt the Kremer-Maskin framework, with four modifications.

First, I separate the quality and the quantity of output. I assume that quantity is determined by a simple fixed-coefficient production function, with one white-collar and one blue-collar worker producing one unit of output.[26] The empirical part of the paper will present evidence that the assumption of fixed proportions of white-collar and blue-collar workers is not unreasonable. I assume that quality is given by a Cobb-Douglas function of the skill-levels of the two workers.

Second, I allow for three alternative interpretations of the "skill" of workers that enters the quality equation: (1) Workers are heterogeneous in skill levels and firms must pay higher wages to attract high-skill workers to the firm, as in Kremer's original paper. (2) Workers are homogeneous

[22] Manasse and Turrini (2001) also combine quality differentiation with a model of trade under monopolistic competition and frame their results in terms of wage inequality. Three issues limit the usefulness of their model in this context, however. First, 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. Second, product quality in their model is a deterministic function of fixed firm characteristics, rather than a choice variable of the firm. Third, each firm employs only one employee and the employee receives all the rents from production. It thus seems more natural to think of these individuals as entrepreneurs rather than employees, and of dispersion of their payoffs as dispersion in profits rather than dispersion in wages.

[23] Lucas (1978) posits a similar underlying characteristic of firms and describes it as "managerial talent."

[24] An interesting direction for future work is to relax this assumption and allow know-how to evolve at different rates over time depending on the type of goods the firm produces, in the spirit of the learning-by-doing model of Young (1991).

[25] The idea is that in sophisticated products, mistakes in any aspect of the production process – for instance, the O-rings in the Space Shuttle Challenger – can drastically reduce the value of the product.

[26] In this sense, the model is distinct from existing Hecksher-Ohlin and Ricardian trade models, which rely on differences in the proportions of different types of workers across industries (or production phases within industries in the Feenstra-Hanson model).

and skills are firm-specific; acquiring skill is costly because firms must pay to train workers. (3) Workers are homogeneous and "skill" represents the degree of effort or care exercised in production, rather than a particular capability or body of knowledge, as in the efficiency-wage theories of Akerlof (1982), Bowles (1985) or Shapiro and Stiglitz (1984).[27] A realistic model would probably combine all three of these interpretations. For present purposes, the important point is simply that "skill" improves quality and is costly to the firm to acquire.

Third, I include an additional input, machines. Because we typically only know the total value of machinery, rather than the number of machines and their level of sophistication separately, it is not obvious how machines should enter the quantity and quality equations. Here I assume that one white-collar and one blue-collar worker always combine with one machine, and that the capital-labor ratio, given by k, enters the quality equation as a proxy for the technical sophistication of the machine.[28]

Fourth, I assume that the know-how parameter λ enters as a multiplicative coefficient in the production function for quality. For the sake of simplicity, I exclude it from the function for quantity, although this is not crucial.[29]

Let h and l index white-collar and blue-collar workers. Let e^h and e^l represent the skill or effort of white-collar workers and blue-collar worker, respectively. Let w^h and w^l represent the wages paid by the firm, and $\underline{w}^h$ and $\underline{w}^l$ the outside market wages for each type of worker, which are taken to be exogenous to the industry being modeled.[30] In the interests of simplicity, I assume that skill/effort level is a linear function of the difference between the wage in the firm and the outside wage:

$$\begin{aligned} e^h(\lambda) &= z^h \left[w^h(\lambda) - \underline{w}^h\right] \\ e^l(\lambda) &= z^l \left[w^l(\lambda) - \underline{w}^l\right] \end{aligned} \tag{3}$$

where z^h and z^l are positive constants. The production function for quality is given by:

$$q(\lambda) = \lambda \left[k(\lambda)\right]^{\alpha^k} \left[e^h(\lambda)\right]^{\alpha^h} \left[e^l(\lambda)\right]^{\alpha^l} \tag{4}$$

[27] See also Dalmazzo (2002), who integrates this efficiency-wage idea explicitly into the model of Kremer (1993).

[28] None of the qualitative results in the paper depend on this assumption. Capital is included in the model mainly to generate implications for a variable on which we will have data, and could easily be dropped.

[29] A model in which know-how does not affect quality directly but instead reduces input requirements in the quantity equation for the two types of labor (which affect quality) but not for other costs such as transport or raw materials that do not affect quality would generate similar results, at the cost of messier algebra.

[30] In interpretations (2) and (3) above, this market wage is the wage that a worker in the firm would actually receive in the outside market. In (1), we can think of it as an average wage among the heterogeneous workers in the outside market, rather than the wage that a worker in the firm would actually receive if she left the firm.

Let $\alpha \equiv \alpha^k + \alpha^h + \alpha^l$. Assume that improvements to quality from a given increase in the skill and sophistication of inputs are diminishing: $\alpha < 1$. This ensures an interior solution in the choice of quality. Let r be the rental cost of capital. The marginal cost of producing one unit of output, which by assumption is independent of the quantity produced, is $mc(\lambda) = w^h(\lambda) + w^l(\lambda) + rk(\lambda)$. There is a fixed cost of entry, f.

The combination of constant marginal cost (conditional on quality) and the fixed cost of entry generates increasing returns to scale. There is no cost to differentiation and firms are constrained to offer just one variety. As a consequence, all firms differentiate and have a monopoly in the market for their particular variety. The fixed parameter λ thus indexes goods as well as firms. Since λ is distributed continuously, we can think of the number of goods J in equation (2) as going to infinity. Then assuming firms are risk-neutral and dropping the expectation, the demand function for each good (2) can be rewritten:

$$x(\lambda) = \frac{N}{D} \exp\left[\frac{\theta q(\lambda) - p(\lambda)}{\mu}\right] \tag{5}$$

with

$$D \equiv \int\limits_{\lambda \in \Lambda} \exp\left[\frac{\theta q(\lambda) - p(\lambda)}{\mu}\right] Mg(\lambda)\, d\lambda \tag{6}$$

where Λ is the set of firms that enter the market.

3.1.3 Firms' Optimization

Each firm seeks to maximize its profit: $\pi(\lambda) = [p(\lambda) - mc(\lambda)]\, x(\lambda) - f$. The firm's choice variables are w^h, w^l, k (which together determine quality, q) and p. As is standard in monopolistic competition models, I assume that each firm thinks of itself as small relative to the market as a whole, and treats the aggregate quantity in the denominator of the expression for output, D, as unaffected by its own choices. Given this assumption, optimization yields:

$$\begin{aligned} w^{h*}(\lambda) - \underline{w}^h &= \alpha^h [\eta\lambda\theta]^{\frac{1}{1-\alpha}} \\ w^{l*}(\lambda) - \underline{w}^l &= \alpha^l [\eta\lambda\theta]^{\frac{1}{1-\alpha}} \\ k^*(\lambda) &= \frac{\alpha^k}{r} [\eta\lambda\theta]^{\frac{1}{1-\alpha}} \end{aligned} \tag{7}$$

where $\eta \equiv \left(\frac{\alpha^k}{r}\right)^{\alpha^k} \left(z^h\alpha^h\right)^{\alpha^h} \left(z^l\alpha^l\right)^{\alpha^l}$. These choices for wages and the capital-labor ratio determine the quality level:

$$q^*(\lambda) = [\eta\lambda\theta^{\alpha}]^{\frac{1}{1-\alpha}} \tag{8}$$

It is convenient to write the firm's choice variables in terms of the resulting value of quality:

$$\begin{aligned} w^{h*}(\lambda) = \underline{w}^h &+ \alpha^h \theta q^*(\lambda) \\ w^{l*}(\lambda) = \underline{w}^l &+ \alpha^l \theta q^*(\lambda) \\ k^*(\lambda) &= \frac{\alpha^k}{r} \theta q^*(\lambda) \end{aligned} \tag{9}$$

The quality of goods produced by the firm is a summary indicator that captures the dependence of wages and the capital-labor ratio on the firm's underlying know-how. Equations (8) and (9) give us our first key implications of the model: in cross-section, higher-λ firms produce higher-quality goods, pay higher wages to both white-collar and blue-collar workers, and are more capital-intensive than lower-λ firms. Note also that quality, wages and capital intensity are greater, the greater the willingness of consumers to pay for quality, θ.

Appendix A.2 shows that the wage ratio: $\omega^*(\lambda) \equiv \frac{w^{h*}(\lambda)}{w^{l*}(\lambda)}$ is increasing in λ and θ if (and only if) $\alpha^h/\alpha^l > \underline{w}^h/\underline{w}^l$, that is, if the production of quality is sufficiently more sensitive to the skill of white-collar workers than to that of blue-collar workers.

Marginal cost and price at the optimum are given by $mc^*(\lambda) = \underline{w}^h + \underline{w}^l + \alpha\theta q^*(\lambda)$ and $p^*(\lambda) = \mu + \underline{w}^h + \underline{w}^l + \alpha\theta q^*(\lambda)$. Note that both are increasing in λ and θ. As is standard in logit demand models, the mark-up is constant: $p^*(\lambda) - mc^*(\lambda) = \mu$.

Define $I^*(\lambda) \equiv [\theta q^*(\lambda) - p^*(\lambda)]/\mu = \left[(1-\alpha)(\eta\theta\lambda)^{\frac{1}{1-\alpha}} - \mu - \underline{w}^h - \underline{w}^l\right]/\mu$ and note that $I(\cdot)$ is increasing in λ and θ as well. Output and profit can be written:

$$x^* = \frac{N}{D^*} \exp\{I^*(\lambda)\} \tag{10}$$

$$\pi^*(\lambda) = \frac{\mu N}{D^*} \exp\{I^*(\lambda)\} - f \tag{11}$$

where

$$D^* = \int_{\lambda \in \Lambda} \exp\{I^*(\lambda)\} M g(\lambda)\, d\lambda \tag{12}$$

Equations (10) and (11) complete the set of cross-sectional implications in the one-country version of the model: taking as given the set Λ of firms that enter, output and profit are greater for firms with higher λ.

A corollary of equations (7) and (10) is that the size of a firm will be correlated with the wages it pays to both types of workers, since both are increasing in λ. The model thus provides a natural explanation for the employer size-wage effect, documented by Brown and Medoff (1989) for the U.S. and by Velenchik (1997) and Schaffner (1998), among others, for developing countries.

3.1.4 Entry

The fact that profitability is increasing in λ implies that in equilibrium there will be a single cut-off value of productivity, call it $\lambda^{\min}$, above which all firms will enter and earn positive profits, and below which no firms will enter. The cut-off is defined implicitly by the requirement that the marginal firm have zero profits:

$$\pi^*\left(\lambda^{\min}\right) = \frac{\mu N}{D^*\left(\lambda^{\min}\right)} \exp\left\{I^*\left(\lambda^{\min}\right)\right\} - f = 0 \tag{13}$$

where now we can specify the limits of integration for D^* and write it as a function of the variable lower cut-off, $\lambda^{\min}$:

$$D^*\left(\lambda^{\min}\right) = \int_{\lambda^{\min}}^{\lambda^{\max}} \exp\left\{I^*\left(\lambda\right)\right\} Mg\left(\lambda\right) d\lambda \tag{14}$$

Note that $D^*\left(\cdot\right)$ is a decreasing function of $\lambda^{\min}$: the higher is the cut-off, the smaller the mass of firms over which we integrate.

3.2 Two-Country Model

3.2.1 Modifications to Basic Set-up

Now suppose that there are two countries in the model, North and South, indexed by n and s. I assume that the two markets are segmented, in the sense that firms can sell a different quality product and charge a different price in each market.[31] The key difference between countries, as mentioned above, is that Northern consumers are more willing to pay for quality than Southern consumers: $\theta_n > \theta_s$.

I also assume production is segmented, in the sense that firms can make different optimization decisions for production for the Northern market than for production for the Southern market. It is convenient to think of each firm as producing (or potentially producing) on two different production lines, one for the domestic market and one for export.[32] We thus potentially have two

[31] Firms are still constrained to sell one variety in each market.

[32] A firm is allowed pay different wages to workers of the same type producing on the different lines. There is thus no internal equity constraint in this model, although it is likely that in real firms fairness considerations between workers producing on the different lines would constrain firms' ability to make completely separable decisions on the two lines.

sets of optimization equations for each firm in each country, one for each production line. To keep track of decisions on each line, I introduce two indices: $c = s, n$ indicates the country in which the plant is located; and $d = s, n$ indicates the destination market. For variables that pertain either only to the production location or only to the destination market, I use just one subscript.

The equations governing each firm's decisions in the one-country model map almost directly into the equations governing each firm's decisions on a particular production line, with one main difference. The ratio of the price levels between the two countries – the real exchange rate – is allowed to vary. To keep track of these movements, it is convenient to define δ_{cd} to be the ratio of the price level in country d to the price level in country c. Given this definition, we have that $\delta_{ss} = \delta_{nn} = 1$ and $\delta_{ns} = 1/\delta_{sn}$. Assume that purchasing power parity holds initially, so that $\delta_{sn} = \delta_{ns} = 1$; below I will assume that a currency shock in South lowers the price level in South and raises δ_{sn}. Consider the price charged by the firm with know-how λ in South for goods sold on the Northern market. Let $p_{sn}(\lambda)$ be the price charged by the firm in terms of Southern goods. For Northern consumers the relevant price is the price in terms of Northern goods, which is to say $\frac{p_{sn}(\lambda)}{\delta_{sn}}$. In general, the consumer demand function (equation (5) in the one-country version) should be rewritten as:

$$x_{cd}(\lambda) = \frac{N_d}{D_d} \exp\left[\frac{\theta_d q_{cd}(\lambda) - \frac{p_{cd}(\lambda)}{\delta_{cd}}}{\mu}\right] \tag{15}$$

In the two country case, the term in the denominator, D_d, now involves an aggregation in each consumer market over two sets of firms, domestic firms selling in their home market and foreign firms exporting to it. Let $\lambda_{cd}^{\min}$ represent the cut-off for a firm located in country c to enter destination market d. Then we can write:

$$D_d\left(\lambda_{sd}^{\min}, \lambda_{nd}^{\min}\right) = D_{sd}\left(\lambda_{sd}^{\min}\right) + D_{nd}\left(\lambda_{nd}^{\min}\right) \tag{16}$$

where, as in (6),

$$D_{cd}\left(\lambda_{cd}^{\min}\right) = \int_{\lambda_{cd}^{\min}}^{\lambda_c^{\max}} \exp\left\{\frac{\theta_d q_{cd}(\lambda) - \frac{p_{cd}(\lambda)}{\delta_{cd}}}{\mu}\right\} M_c g_c(\lambda)\, d\lambda \tag{17}$$

Finally, I assume that firms must pay a fixed cost to export, f_e, in addition to the fixed cost to enter the domestic market. It will be convenient to indicate fixed costs using the subscript notation: $f_{sn} = f_{ss} + f_e$ and $f_{ns} = f_{nn} + f_e$. The fact that exporters must pay to enter the domestic market before paying to enter the export market means that there will be no firms that only enter the export market. The number of consumers in each market, N_n and N_s, the total masses of firms, M_n and M_s, the distributions of firms, $g_n(\lambda)$ and $g_s(\lambda)$, and the maximum values of entrepreneurial ability, $\lambda_n^{\max}$ and $\lambda_s^{\max}$, are allowed to differ across countries, although these differences are of little consequence. The extent-of-differentiation parameter, μ, the cost of capital, r, the slopes of the skill-wage schedules, z^h and z^l, and the exponents of the Cobb-Douglas production function for quality, α^h, α^l and α^k and hence the term η in (7) are assumed not to differ across countries.

3.2.2 Firms' Optimization for Each Production Line

Given the two-country set-up, firms' optimization yields the following (refer to equation (7)):

$$
\begin{aligned}
q_{cd}^*(\lambda) &= \left[\eta\lambda\delta_{cd}^{\alpha}\theta_d^{\alpha}\right]^{\frac{1}{1-\alpha}} \\
w_{cd}^{h*}(\lambda) - \underline{w}_c^h &= \alpha^h\delta_{cd}\theta_d q_{cd}^*(\lambda) \\
w_{cd}^{l*}(\lambda) - \underline{w}_c^l &= \alpha^l\delta_{cd}\theta_d q_{cd}^*(\lambda) \\
k_{cd}^*(\lambda) &= \frac{\alpha^k}{r}\delta_{cd}\theta_d q_{cd}^*(\lambda) \\
\omega_{cd}^*(\lambda) &= \frac{\underline{w}_c^h + \alpha^h\delta_{cd}\theta_d q_{cd}^*(\lambda)}{\underline{w}_c^l + \alpha^l\delta_{cd}\theta_d q_{cd}^*(\lambda)} \\
mc_{cd}^*(\lambda) &= \underline{w}_c^h + \underline{w}_c^l + \alpha\delta_{cd}\theta_d q_{cd}^*(\lambda) \\
p_{cd}^*(\lambda) &= \mu + \underline{w}_c^h + \underline{w}_c^l + \alpha\delta_{cd}\theta_d q_{cd}^*(\lambda) \\
x_{cd}^*(\lambda) &= \frac{N_d}{D_d^*}\exp\left\{I_{cd}^*(\lambda)\right\} \\
\pi_{cd}^*(\lambda) &= \frac{\mu N_d}{D_d^*}\exp\left\{I_{cd}^*(\lambda)\right\} - f_{cd}
\end{aligned}
\tag{18}
$$

where

$$
\begin{aligned}
I_{cd}^*(\lambda) &\equiv \left[\theta_d q_{cd}^* - \frac{p_{cd}^*}{\delta_{cd}}\right]/\mu \\
&= \left[(1-\alpha)\left(\eta\lambda\delta_{cd}^{\alpha}\theta_d\right)^{\frac{1}{1-\alpha}} - \frac{\mu + \underline{w}_c^h + \underline{w}_c^l}{\delta_{cd}}\right]/\mu
\end{aligned}
\tag{19}
$$

and D_d^* is given by (16) and (17) above.

For each location-destination pair, we have the same set of cross-sectional relationships as in the one-country case. Among firms in a given country producing for a given market, firms with greater know-how λ will produce a higher-quality good; pay higher wages to both white-collar and blue-collar workers; pay a higher relative wage to white-collar workers (conditional on the assumption that $\alpha^h/\alpha^l > \underline{w}^h/\underline{w}^l$); use more capital per worker; charge a higher price; produce more output and earn more profits.

From the expression for quality in (18), we see that if a given firm enters both markets, then (assuming $\delta_{ss} = \delta_{sn} = 1$ initially) it will produce a higher-quality good for the Northern market than for the Southern one.[33] Within each firm, the production of higher-quality goods on the export line is accompanied by higher wages, higher relative wages of white-collar workers, higher capital intensity, higher costs, and higher prices than on the production line for the domestic market.

3.2.3 Entry

In the two-country model, there will be four entry cut-offs, $\lambda_{cd}^{\min}$, one for each location-destination pair ($c = n, s$, $d = n, s$). The cut-offs are determined by four zero-profit conditions:

$$\pi_{cd}^{*}\left(\lambda_{cd}^{\min}\right) = \frac{\mu N_d}{D_d^{*}} \exp\left\{I_{cd}^{*}\left(\lambda_{cd}^{\min}\right)\right\} - f_{cd} = 0 \tag{20}$$

The fact that firms must first pay the fixed cost to enter the domestic market before paying the additional fixed cost to export means that $\lambda_{sn}^{\min} > \lambda_{ss}^{\min}$ and $\lambda_{ns}^{\min} > \lambda_{nn}^{\min}$. Within the cross-section of firms in each country at a given time, we have three distinct groups of firms: firms that do not enter (call them non-entrants), firms that enter only the domestic market (call them non-exporters), and firms that enter both markets (call them exporters). In South, non-entrants correspond to the interval $\lambda \in [0, \lambda_{ss}^{\min})$, non-exporters to the interval $\lambda \in [\lambda_{ss}^{\min}, \lambda_{sn}^{\min})$, and exporters to the interval $\lambda \in [\lambda_{sn}^{\min}, \lambda_{s}^{\max})$.

3.2.4 Observables at the Firm Level

The notion of two production lines per firm is conceptually straightforward, but in practical terms it is rare to have production data on firms or plants by production line, except in particular industries such as automobiles. This section relates the variables defined above to the variables that are observable in the Mexican establishment panels and similar datasets. For ease of exposition, I focus on Southern firms, but the analysis for Northern firms would be analogous. Before deriving the implications for the observables, it is convenient first to define two variables which are not directly observable but which will be useful in what follows: the export share of output[34] and the

[33] A single firm will produce different qualities for different markets even in the absence of the quality bias due to per-unit trade costs first hypothesized by Alchian and Allen (1964). The effect is due solely to differences in quality demands by consumers arising from income differences between countries. This phenomenon cannot be captured in models based on perfect competition in the product market, such as the Ricardian-type quality upgrading model of Stokey (1991). Under perfect competition, firms face identical product market conditions across countries, and have no incentive to produce different goods for different markets.

[34] The export share of output is not observable because although we observe sales in each market we do not observe the prices of goods sold on each market, and hence cannot make inferences about the quantities of goods sold.

average quality of goods produced.[35]

Define the export share of output for Southern firms as: $\chi_s^*(\lambda) \equiv \frac{x_{sn}^*(\lambda)}{x_{sn}^*(\lambda)+x_{ss}^*(\lambda)}$. Suppose that all firms enter both markets. Appendix A.3 shows that under this counterfactual assumption output on each production line is increasing in λ and output on the export line increases more steeply in λ. Consequently, the export share of output is also increasing in λ. But in general not all firms enter both markets. For non-exporters, $\chi_s^*(\lambda) = 0$. (For non-entrants, the export share of output is obviously undefined.) The result that the export share of output is an increasing function of λ holds only for exporters. Note that there will be a discontinuity in $\chi_s^*(\lambda)$ at the cut-off for entry into the export market. Firms just to the left of the cut-off have zero exports. Firms just to the right of the cut-off must sell enough on the export market to recoup the indivisible fixed cost, and have an export share strictly greater than zero.

The average quality of goods produced by a given firm can be expressed as a weighted average of the quality levels produced on the firm's two production lines, where the weights are given by the export share of output: $\overline{q}_s^*(\lambda) = \chi_s^*(\lambda)\, q_{sn}^*(\lambda) + (1-\chi_s^*(\lambda))\, q_{ss}^*(\lambda)$. Figure 5a summarizes the relationship between $\overline{q}_s^*(\lambda)$ and λ in the cross-section of Southern firms. The dashed (as opposed to dotted) curves represent $q_{ss}^*(\lambda)$ and $q_{sn}^*(\lambda)$, the quality levels on each production line. From (18), we know that $q_{ss}^*(\lambda)$ and $q_{sn}^*(\lambda)$ are increasing in λ. Since $\theta_n > \theta_s$, we also know that the $q_{sn}^*(\lambda)$ curve will lie above and have a greater slope than the $q_{ss}^*(\lambda)$ curve. The dotted curve represents the counterfactual average quality that would obtain if all firms entered both markets. The solid curve represents actual average quality as a function of λ. For non-exporters, average quality is equal to the quality level on the domestic production line, $\overline{q}_s^*(\lambda) = q_{ss}^*(\lambda)$. For exporters, the average quality is given by the dotted curve. Note that the average product quality is increasing in λ for two reasons: first because product quality is increasing in λ on each production line; and second because, for the exporters, the export share of output $\chi_s^*(\lambda)$ is increasing in λ as well. There is a discontinuity in average quality at the cut-off for entry into the export market, a consequence of the discontinuity in the export share of output discussed above.

Wages for each type of worker and capital intensity follow a pattern qualitatively similar to average quality. The firm-level averages for these variables are given by:

$$\begin{aligned}
\overline{w}_s^{h*}(\lambda) &= \chi_s^*(\lambda)\, w_{sn}^{h*}(\lambda) + (1-\chi_s^*(\lambda))\, w_{ss}^{h*}(\lambda) \\
\overline{w}_s^{l*}(\lambda) &= \chi_s^* w_{sn}^{l*}(\lambda) + (1-\chi_s^*(\lambda))\, w_{ss}^{l*}(\lambda) \\
\overline{k}_s^*(\lambda) &= \chi_s^*(\lambda)\, k_{sn}^*(\lambda) + (1-\chi_s^*(\lambda))\, k_{ss}^*(\lambda)
\end{aligned} \tag{21}$$

[35] The data on ISO 9000 certification discussed in section 4.4 provide a partial but clearly incomplete measure of product quality.

Like average quality, these variables will be increasing in λ in cross-section, both because wages and the capital-labor ratio on each production line are increasing in λ and also because the export share of output is increasing. They will also have a discontinuity at the cut-off for entry into the export market.

Define the average wage ratio at the firm level as the ratio of the average white-collar wage to the average blue-collar wage: $\overline{\omega}_s^*(\lambda) \equiv \overline{w}_s^{h*}(\lambda) / \overline{w}_s^{l*}(\lambda)$. It can then be rewritten as a weighted average of the wage ratios on each production line: $\overline{\omega}_s^*(\lambda) = \widetilde{\chi}_s^*(\lambda)\,\omega_{sn}^*(\lambda) + (1 - \widetilde{\chi}_s^*(\lambda))\,\omega_{ss}^*(\lambda)$ where $\widetilde{\chi}_s^*(\lambda) \equiv \frac{\chi_s^*(\lambda) w_{sn}^{l*}(\lambda)}{\chi_s^*(\lambda) w_{sn}^{l*}(\lambda) + (1-\chi_s^*(\lambda)) w_{ss}^{l*}(\lambda)}$. Like $\chi_s^*(\lambda)$, $\widetilde{\chi}_s^*(\lambda)$ will be zero for non-exporters and increasing in λ for exporters. Conditional on the assumption that $\alpha^h/\alpha^l > w^h/w^l$, the wage ratio will follow the same qualitative pattern as wage levels.

Under the assumption that one white-collar worker and one blue-collar worker are required to produce one unit of output, employment of each type of worker is simply equal to total output. Let E represent total employment. For non-exporters, $E_s^*(\lambda) = 2x_{ss}^*(\lambda)$ and is increasing in λ. For exporters, $E_s^*(\lambda) = 2\left(x_{ss}^*(\lambda) + x_{sn}^*(\lambda)\right)$ and is increasing more steeply in λ. There will also be a discontinuity in employment at the cut-off for the export market.

The observable variables domestic sales and export sales are distinguished from the other observables by the fact that they correspond to a single production line. Let $S_{ss}^*(\lambda) = p_{ss}^*(\lambda)\,x_{ss}^*(\lambda)$ and $S_{sn}^*(\lambda) = p_{sn}^*(\lambda)\,x_{sn}^*(\lambda)$ represent domestic sales and exports, respectively. For all firms that enter the domestic market, non-exporters and exporters, the underlying variables $p_{ss}^*(\lambda)$ and $x_{ss}^*(\lambda)$ – and hence domestic sales, $S_{sn}^*(\lambda)$ – are continuous, monotonically increasing functions of λ. I will rely on this fact below, when using domestic sales as a proxy for the underlying productivity parameter. Export sales, $S_{sn}^*(\lambda)$, are smooth and increasing within the set of exporters, but are zero for non-exporters, which will make up the majority of plants in our data. Finally, the export share of sales is given by $\sigma_s^*(\lambda) = \frac{S_{sn}^*(\lambda)}{S_{ss}^*(\lambda) + S_{sn}^*(\lambda)}$. Appendix A.4 shows that, conditional on firms entering both markets, the export share of sales is increasing in λ.

3.3 Comparative-Static Effects of an Exchange-Rate Shock

I model the shock as having two (exogenous) effects. First, the real exchange rate, δ_{sn}, defined as the Northern price level over the Southern price level, rises (δ_{ns} falls). Second, the shock reduces total consumer demand in South, manifested in our single differentiated goods industry as a decline in the number of Southern consumers, N_s. In addition, to restrain the proliferation of algebra, I assume that North is large relative to South, such that the increased entry of Southern firms into the Northern market in response to the peso crisis does not appreciably affect the profitability of other firms selling in the Northern market. I focus on the comparative-static effects for Southern firms, since they will be the focus of the empirical work.

3.3.1 Effects on Entry of Southern Firms

The effect of the devaluation of the peso is to make Southern goods more competitive relative to Northern goods in both markets. Appendix A.5 shows that $\partial\lambda_{sn}^{\min}/\partial\delta_{sn} < 0$ and $\partial\lambda_{ss}^{\min}/\partial\delta_{sn} < 0$. That is, as the real exchange rate rises, more Southern firms will enter both markets. The effect of the decline in consumer demand in South is to reduce the profitability of all firms selling in South. Given the segmentation of the consumer markets, it has no effect on the profitability of selling in North. Formally, Appendix A.5 shows that $\partial\lambda_{ss}^{\min}/\partial N_s < 0$ and $\partial\lambda_{sn}^{\min}/\partial N_s = 0$.

The net effect on Southern firms' entry into the export market is clear: more Southern firms enter the export market ($\lambda_{sn}^{\min}$ falls). The net effect on entry of Southern firms into the Southern market is ambiguous: the increased competitiveness of Southern firms and the reduction in Southern consumer demand work in opposite directions. In fact, the peso crisis led to an increase in the number of bankruptcies of manufacturing firms in Mexico. It appears that the empirically relevant case is the one in which $\lambda_{ss}^{\min}$ rises. The remainder of the paper will focus on this case. Appendix A.5 gives a precise statement of the condition required to obtain it.

Given our assumptions, the entire continuum of potential firms in South can be classified into five categories (let *pre* indicate the pre-crisis period and *post* the post-crisis period.): (1) Never entrants ($\lambda < \lambda_{ss,pre}^{\min}$) do not enter either market in either period. (2) Exiters from the domestic market ($\lambda_{ss,pre}^{\min} \leq \lambda < \lambda_{ss,post}^{\min}$) enter only the domestic market in the first period and go out of business in the second period. (3) Always non-exporters ($\lambda_{ss,post}^{\min} \leq \lambda < \lambda_{sn,post}^{\min}$) enter only the domestic market, in both periods. (4) Switchers into exporting ($\lambda_{sn,post}^{\min} \leq \lambda < \lambda_{sn,pre}^{\min}$) enter only the domestic market in the first period, but enter both markets in the second period. (5) Always exporters ($\lambda_{sn,pre}^{\min} \leq \lambda$) enter both markets in both periods.

3.3.2 Effects on Production Decisions of Southern Firms

Now consider firm-level production decisions conditional on λ and a given set of entry cut-offs. Consider first the effect of the crisis on the change in the export share of output. Following the same approach as in Section 3.3 above, we first calculate the counterfactual export shares of output assuming that all firms enter both market, both before and after the crisis. Let the change in the export share be given by $d\chi_s(\lambda) \equiv \chi_{s,post}(\lambda) - \chi_{s,pre}(\lambda)$. Appendix A.6 shows that under the counterfactual that all firms enter both markets, the change in export share of output is given by:

$$d\chi_s(\lambda) = \Sigma_s \chi_{s,pre}(\lambda)\left[1 - \chi_{s,pre}(\lambda)\right] \tag{22}$$

where $\Sigma_s > 0$ and is increasing in λ. Thus $d\chi_s(\lambda) > 0$; the export share of output increases as a result of the shock. In addition, the facts that $\chi_s(\lambda)$ and Σ_s are increasing in λ imply that a sufficient condition for $d\chi_s(\lambda)$ to be increasing in λ is that $\chi_s(\lambda) < 1/2$.[36] We will see in the empirical section that among the minority of Mexican plants that export, the average fraction of output exported is 20-25%.[37] Exports made up more than half of total sales in only 3-7% of plants over the 1993-2001 period. It thus appears reasonable to focus on the case where $\chi_s(\lambda) < 1/2$, and I do so hereafter. Consider the five cases from above. The never entrants and exiters (Cases (1) and (2)) are not observed after the crisis. For the always non-exporters (case (3)): $d\chi_s(\lambda) = 0$. For the switchers into exporting (case (4)), the export share is zero before the crisis and positive after the crisis; the change is $d\chi_s(\lambda) = \chi_{s,post}(\lambda)$. For the always exporters, the change in the export share is given by (22). Note that $d\chi_s(\lambda)$ is increasing in λ both within the category of switchers into exporting and within the category of always exporters, and that there are two discontinuities in the relationship between $d\chi_s(\lambda)$ and λ, a jump up at $\lambda_{sn,pre}^{\min}$ and a jump down at $\lambda_{sn,post}^{\min}$.

Now consider the effect on the average quality of goods produced. Suppose again that all firms enter both markets. Quality on the domestic production line, $q_{ss}(\lambda)$ does not depend on either δ_{sn} or N_s and hence is unaffected by the exchange-rate shock. Quality on the export line, $q_{sn}(\lambda)$ increases with the increase in the real exchange rate δ_{sn}. Intuitively, the peso devaluation reduces Southern firms' cost of producing quality relative to what Northern consumers are willing to pay for it, and this induces them to supply more of it. Algebraically, we have:

$$dq_{sn}^*(\lambda) = \left(\frac{\alpha}{1-\alpha}\right)\left[\eta\lambda\left(\theta_n\right)^{\alpha}\right]^{\frac{1}{1-\alpha}}\left(\delta_{sn}\right)^{\frac{2\alpha-1}{1-\alpha}} d\delta_{sn}$$

which is positive and increasing in λ. The increase in the average level of quality in the firm is then a combination of the increase in the quality on the export line and the increase in the export share of output:

$$d\overline{q}_s(\lambda) = \chi_s(\lambda)\, dq_{sn}(\lambda) + \left(q_{sn}^*(\lambda) - q_{ss}^*(\lambda)\right) d\chi_s(\lambda) \tag{23}$$

Under the counterfactual that firms enter both markets, $\chi_s(\lambda)$, $d\chi_s(\lambda)$, $dq_{sn}(\lambda)$, and $q_{sn}^*(\lambda) - q_{ss}^*(\lambda)$ are all positive and increasing in λ; hence $d\overline{q}_s(\lambda)$ is positive and increasing in λ as well. Figures 5b depicts the effect of the exchange rate shock on the level average quality. The $q_{sn}^*(\lambda)$

[36] Intuitively, if a firm is already exporting a large share of output before the shock, then there is little room to increase the export share further.

[37] Moreover, the model suggests that the export share of sales is an upper bound on the export share of output, since the price of output sold in the Northern market is higher than the price of goods sold in South.

curve shifts up in response to the shock, and the dotted counterfactual $\overline{q}_s(\lambda)$ curve also shifts up, reflecting both the increase in quality on the export line and the increase in the export share of output. The thinner, dark solid line represents the actual average quality curve pre-crisis. The thicker, gray solid line represents the actual average quality curve post-crisis. The always non-exporters continue to produce exclusively for the domestic market and their average quality level does not change. The switchers into exporting see an especially large increase in average quality. The always exporters see the change in quality given by equation (23). Figure 5c depicts the change in average quality as a function of λ. Average quality is increasing in λ within the category of switchers and within the category of always exporters. There is a positive discontinuity at the post-crisis cut-off for entry into the export market, $\lambda_{sn,post}^{\min}$, and a negative discontinuity at the pre-crisis cut-off for entry into the export market, $\lambda_{sn,pre}^{\min}$, both consequences of the discontinuities in the export share of output.

Recalling (21), the changes in wages for each type of worker and the capital-labor ratio are:

$$\begin{aligned} d\overline{w}_s^{h*}(\lambda) &= \chi_s(\lambda)\,\alpha^h\theta_n\delta_{sn}dq_{sn}^*(\lambda) + \alpha^h\left(\theta_n\delta_{sn}q_{sn}^*(\lambda) - \theta_s q_{ss}^*(\lambda)\right)d\chi_s(\lambda) \\ d\overline{w}_s^{l*}(\lambda) &= \chi_s(\lambda)\,\alpha^l\theta_n\delta_{sn}dq_{sn}^*(\lambda) + \alpha^l\left(\theta_n\delta_{sn}q_{sn}^*(\lambda) - \theta_s q_{ss}^*(\lambda)\right)d\chi_s(\lambda) \\ d\overline{k}_s &= \chi_s(\lambda)\left(\frac{\alpha^k}{r}\right)\theta_n\delta_{sn}dq_{sn}^*(\lambda) + \frac{\alpha^k}{r}\left(\theta_n\delta_{sn}q_{sn}^*(\lambda) - \theta_s q_{ss}^*(\lambda)\right)d\chi_s(\lambda) \end{aligned} \tag{24}$$

The pattern of changes in these variables are qualitatively similar to the pattern for average quality. Note that although I have assumed that outside market wages for each type of worker are exogenously fixed, it would be straightforward to allow them to vary with the shock. The important point is that the change in the outside market wage would affect wages in all firms equally; we would still observe a larger differential increase in higher-λ firms. Appendix A.7 shows that the change in the average wage ratio $\overline{\omega}_s$ in response to the shock also follows the same qualitative pattern as average quality depicted in Figure 5c. Appendix A.8 shows that, for the case where the export share of sales is less than 1/2, the model carries similar implications for changes in the export share of sales, $\sigma_s(\lambda)$. Finally, Appendix A.9 shows why the model does not carry a strong prediction for the effect of the exchange-rate shock on total employment and total sales. The intuition is that although higher-λ firms see a larger increase in exports, they also have higher levels of domestic sales prior to the shock and hence a greater drop in domestic sales in response to the shock.

To sum up, the testable implication I will take to the data is that the pattern illustrated by Figure 5c will hold for a number of observable variables: wages of each type of worker, the wage ratio, the capital-labor ratio, the export share of sales, and, to the extent that it is observable, product quality. Unfortunately, our estimates of λ will be sufficiently imprecise and our data on the these outcome variables will be sufficiently noisy that it will not be possible to estimate the specific predictions of discontinuities at the post-crisis and pre-crisis cut-offs for entry into exporting. We will have to content ourselves with estimating roughly linear relationships between these changes in these observables and estimates of the underlying know-how parameter λ. In this sense, the empirical work should be seen as testing the first-order qualitative implications of the model, rather than its specific functional form.

4 Empirics

4.1 Data

The analysis in this paper is based primarily 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 a panel of the largest establishments in Mexico in 205 6-digit manufacturing industries, excluding maquiladoras. The panel on which the estimates in this paper are primarily based covers the year 1993-2001. The sample is non-random: In 1993, plants with 100 or more employees were included with certainty, and smaller plants were included in descending order of size until 85% of the production of the industry was covered. These plants were then followed over time. Because of this design, the survey is not appropriate for analyzing shifts in production across industries or turnover of plants, but it is well-suited to a study of the evolution of the behavior of individual establishments over time. After cleaning, I am left with 3,003 plants that appear with complete data in all nine years of the panel; I refer to this panel as the EIA 1993-2001 Balanced Panel. An advantage of the 1993-2001 EIA data is that it reports not just whether or not an establishment exited, but also why it exited. I construct a second panel that includes plants from the original 1993 sample that went out of business over the 1994-2001 period, discarding plants that disappeared from the dataset for reasons unrelated to their economic performance. The resulting panel has 3,605 plants; I refer to it as the EIA 1993-2001 Unbalanced Panel.

There is also an earlier EIA panel, covering the period 1984-1994. The sampling design for this panel was similar to that of the 1993-2001 panel: the sample was drawn in 1984, and then followed over time. The advantage of including this earlier panel is that it is possible to follow a subset of plants (706 total) over the entire 1984-2001 period. I refer to this dataset as the EIA 1984-2001 Balanced Panel. Further details on the sample and cleaning procedure are in the data appendix.

Table 1 reports summary statistics for the EIA 1993-2001 Balanced panel in the years 1993, 1997 and 2001, for non-exporters (plants with no exports), exporters (plants with positive exports), and all plants together. The first important point is that, consistent with the theoretical model, there are systematic differences between exporters and non-exporters in cross-section. Exporters tend to be larger in employment and in revenues; to be more capital-intensive; to pay higher wages, especially to white-collar workers; and to receive more foreign investment.[38] Note further that, again consistent with the theoretical model, exporters tend to have greater sales on the domestic market than non-exporters. The second important point is that the employment share of white-collar workers does *not* differ significantly between exporter and non-exporters, or across time. This suggests that the assumption of the theoretical model that the proportion of each type of worker in total employment is fixed is not grossly unrealistic. The third important point is that, among exporters, the average share of sales that derives from exports is on the order of 20%. This suggests that focusing the model on the case where the export share of sales is less than 50% is appropriate.

Figure 6 shows time paths of log real revenues, log employment, and log white-collar and blue-collar wages, aggregated separately for always non-exporters, switchers into exporting, and always exporters.[39] It reveals that the peso crisis had quite different effects on manufacturing establishments depending on the extent of their integration into export markets. The first point to notice is that, consistent with the theoretical model, the levels of employment, revenues, and wages of both types of workers for switchers into exporting are intermediate between the low levels of the always non-exporters and the high levels of the always exporters. The second point is that exporting plants weathered the crisis better than non-exporting plants. Always exporters saw no dip in revenues and a small dip in employment during the crisis. Switchers into exporting saw larger dips, but rebounded more quickly than the always non-exporters. The third point is that real wages fell sharply during the crisis, for both types of workers, and that the decline appears to have been slightly steeper for always non-exporters than for always exporters. The wages of switchers into exporting moved subtly away from the always non-exporters and toward the always exporters. This tendency appears to be stronger for white-collar wages than for blue-collar.

[38] These patterns are consistent with the stylized facts for manufacturing plants in the U.S. documented by Bernard and Jensen (1999).

[39] Plants were classified as always exporters if they had positive exports in both 1993 and 1997, as switchers into exporting if they had zero exports in 1993 and positive exports in 1997, and as always non-exporters if they had zero exports in both years. The switchers out of exporting were few (less than 5% of the sample) and I omit them from the figures.

4.2 Econometric Strategy

Although such descriptive comparisons of means across plants with different export status are suggestive, they can also be misleading. Differences between exporters and non-exporters in cross-section may simply be due to selection of more-productive (higher-λ) plants into the export market, and changes between always non-exporters, switchers, and always exporters over time may reflect changing patterns of selection based on shocks to productivity over time. This section presents an econometric strategy that is not confounded by such endogenous selection. The strategy is based on a comparison of within-plant changes over time between plants distinguished by their underlying productivity *prior* to the crisis, rather than by the choices they make in response to the crisis.

The first econometric task is to find a proxy for the unobserved parameter, λ. The theoretical model suggests that a number of different observable variables – wages, the wage ratio, the capital-labor ratio, total employment, total sales, the export share of sales – are related to λ. But the problem of inferring λ is complicated by the fact that the observable variables that are weighted averages of the two production lines depend on λ in two ways: through the plant's decisions on each line and through the export share of output. My primary strategy for avoiding this difficulty is to use as a proxy the one observable variable that is both non-zero for the all plants in the dataset and observable at the production-line level: domestic sales. The model suggests that in cross-section domestic sales will be a smooth, continuous, monotonically increasing function of the know-how parameter λ. Appendix A.10 shows that, using a first-order approximation, $\lambda - \overline{\lambda} = \frac{1}{B}\left(\ln(S_{ss}) - \overline{\ln(S_{ss})}\right)$ where $\overline{\lambda}$ and $\overline{\ln(S_{ss})}$ are industry means, and B is a positive constant that depends on the industry-average productivity but not on the productivity of a particular plant. Up to a scaling factor, the deviated log of domestic sales can thus be taken directly as a proxy for (deviated within industry) know-how.

Alternatively, a simple and transparent method for combining the information from the several variables that the model suggests are correlated with λ into a single index is to take their first principal component, the single linear combination of the several variables that best accounts for the variance in their joint distribution.[40] As a robustness check, I estimate the same econometric model using this proxy for λ. The variables included in the principal component calculation are

[40] Another approach would be to use the multiple-indicator multiple-cause framework of Goldberger (1972, 1974), and use factor-analytic techniques to make more efficient use of the information available in the various plant characteristics. This approach is unattractive for present purposes because it depends on a number of restrictive assumptions on the disturbance terms in the equations relating the indicators to the unobserved factor, assumptions that are unlikely to hold in a dataset as rife with measurement error as a plant-level panel from a developing country. The factor-analysis approach may still be promising, however, and it remains a subject for future work.

total employment, total revenues, the export share of sales, the capital labor ratio, wage level for each type of worker, and the employment ratio (the ratio of white-collar to blue-collar employment). I also include an indicator for whether the plant has received foreign direct investment. While 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 an indicator for unobserved know-how.

Once I have estimated the proxy, the remainder of the estimation procedure is straightforward. I simply regress within-plant changes in plant characteristics on the proxy dummies for the 205 6-digit industries and 32 states in Mexico. Let $\widehat{\lambda}$ be the estimate for λ from before the crisis. The model can be written:

$$dy_j = \widehat{\lambda}_j \cdot \beta + D_j \cdot \gamma + \varepsilon_j \tag{25}$$

where j indexes plants, D_j is a vector of indicators for industry and state within Mexico, ε_j is a mean-zero disturbance, assumed to be uncorrelated with $\widehat{\lambda}_j$, and dy_j is a within-plant change in one of the set of plant characteristics that the model suggests should be affected by quality upgrading. The parameter of interest in this equation is β. Identification of β is based implicitly on a comparison of within-plant changes over the crisis years between initially high-$\widehat{\lambda}$ and low-$\widehat{\lambda}$ plants.[41] A significant positive coefficient can be interpreted as evidence of differential upgrading within industries.

This specification presumes that no omitted variables are correlated with both $\widehat{\lambda}_j$ and dy_j. A possible concern is that plants' underlying know-how itself changes over time, and that the changes are correlated with its initial level. This would be the case, for instance, if plants near the technological frontier are better able to learn about new technological advancements. I address this possibility by re-estimating the same econometric model in different periods. If a positive coefficient on $\widehat{\lambda}_j$ for the peso crisis years were due to a process of non-uniform technical diffusion rather than the exchange rate shock, then we would expect to find a similar coefficient on $\widehat{\lambda}_j$ in periods during which a major exchange-rate shock did not intervene.

[41] This estimation strategy can also be interpreted in an instrumental-variables framework. The know-how proxy can be thought of as an instrument for the change in exports during the crisis, which might then be used to estimate the effect of exporting on plant-level behavior. The danger with this interpretation is that it is not clear what it means to talk about a causal effect of exporting on plant behavior: a plant's decision to enter the export market and its decision to upgrade quality can both be interpreted as manifestations of the plant's solution to a single underlying optimization problem. But 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. This interpretation underlines the similarity of this approach to that of Chay and Greenstone (2001), who use initial levels of air pollution as an instrument for changes in pollution at the country level in response to the 1981-1982 recession in the U.S.

A second concern with this specification is that it regresses changes on levels, and mean reversion due to measurement error may bias the coefficient estimates. This issue is not relevant to the main estimates of the paper, which regress changes in wages and the wage ratio on the initial level of domestic sales. But it may be important for regressions involving changes in sales, changes in the export share of sales (in which domestic sales enters in the denominator), or possibly changes in total employment (for reasons explained below), as well as for regressions using the first principal component. To deal with this issue, when using log domestic sales, I present both OLS results and IV results, instrumenting the initial level of domestic sales with sales from the previous year. When using the first principal component, I leave out of the principal component estimation the variable that appears in changes on the left hand side of the subsequent regression.

4.3 Basic Results

Let us begin with a set of graphs that illustrate the main results. Figures 7a-7d plots locally smoothed non-parametric regression lines of the levels of the export share, white-collar and blue-collar wages, and the wage ratio in pre- and post-crisis years against log domestic sales in 1993. The dark, bold curves are for the pre-crisis year, the gray curves for the post-crisis year. All variables have been deviated from year-specific industry means; hence the relevant information is contained in the slopes of the curves, not the levels. To avoid the issue of mean reversion, Figure 7a plots the export percentage of sales in 1994 instead of 1993, as well as in 1997. The figure reveals (1) that in cross-section higher-$\widehat{\lambda}$ have higher export shares, and (2) that there was a larger increase in the export share in plants with higher domestic sales in 1993.[42] Figures 7b-7d plot similar graphs for white-collar wages, blue-collar wages, and the wage ratio in 1993 and 1997. In cross-section, wages for both types of workers and the wage ratio are increasing in $\widehat{\lambda}$ (log domestic sales). The changes in the slopes of the curves between 1993 and 1997 are subtler than for the export share, but still evident: higher-$\widehat{\lambda}$ plants saw greater increases in these wage variables than lower-$\widehat{\lambda}$ plants from 1993-1997. It appears that the increase in slope was greater for white-collar than for blue-collar wages, and the fact that the slope of the wage ratio curve also increased confirms this.

As argued above, if the relationship between initial productivity and the changes in wages and capital intensity were due to an omitted variable unrelated to the peso crisis, then we would expect to see the same relationship showing up in other periods in which there was no such shock. To

[42] The spikes at the left ends of both curves are due to a small number of plants with very low domestic sales and high exports. These plants appear to be maquiladoras that were included in the EIA dataset by mistake (and that my cleaning procedures have failed to filter out because they report non-zero domestic sales.)

investigate this, Figures 8a-8d compare changes in our key variables over 1993-1997 period (dark, bold curves) to changes over the 1997-2001 period (gray curves).[43] The changes over the 1993-1997 period are visibly larger for higher-$\widehat{\lambda}$ plants. The wage and wage-ratio changes over the 1997-2001 display no gradient in $\widehat{\lambda}$ whatsoever.[44] To get a better sense of the statistical significance of these patterns, we now turn to the regression results.

Table 2 reports results of linear regressions of the model given by (25), for six different dependent variables (24 regressions in all) using the EIA 1993-2001 Balanced Panel. The coefficients on the industry and state dummies are omitted. Column (1) reports regressions of changes in our key variables over 1993-1997 on log domestic sales in 1993. To address the mean reversion issue, column (3) reports regressions of the changes over 1994-1997 on log domestic sales in 1994 instrumented by its value in the previous year. The point estimate for the export share declines in the IV specification – consistent with the idea that there is measurement error in the domestic sales term in the denominator of the export share – but remains highly significant. The differential increases in the white-collar wage, the blue-collar wage, the wage ratio we observed in Figures 7-8 for the 1993-1997 period are highly significant, as is the differential increase in the capital-labor ratio. It is also notable that there is no evidence of differential within-plant changes in the ratio of white-collar employment to blue-collar employment.

Columns (5)-(8) report results for the same specification re-estimated for the 1997-2001 period (using 1997 as the initial year). As in Figure 8, the contrast between these results and the results for the 1993-1997 period is striking. Although the estimates for the change in the export share are still significant, the magnitude of the estimate declines by more than 50 percent. The estimates for the changes in wages and the wage ratio drop essentially to zero. The estimate for the log capital-labor ratio in column (5) is significant at the 10% level, but it is much smaller than in the earlier period and is not robust to the instrumenting procedure in column (8).

We now turn to results using the alternative know-how proxy from the principal component procedure. Table 3 reports a correlation matrix of the key variables, deviated from industry means, for 1993. Consistent the theoretical model, log employment, log revenues, log capital-labor ratio, log wage levels for the two types of workers, an indicator for foreign ownership, and exports as a percentage of sales are all significantly correlated with one another in cross-section. The first principal component of these variables captures approximately 34% of their joint variance. Not surprisingly, this proxy is significantly correlated with all of the variables used to construct it. Also, the correlation coefficient between the first principal component and log domestic sales is

[43] The 1994-1997 and 1998-2001 periods, respectively, for export shares.

[44] The change in the export share over 1998-2001 displays a slight gradient in log domestic sales, but the slope is clearly smaller than in the earlier period.

.81, which supports the assertion that the two indexes are capturing a similar underlying pattern in the cross-sectional data. Table 4 reports results similar to those in Table 2 columns (1)-(2) and (5)-(6), but using the first principal component as the productivity proxy.[45] The qualitative pattern of results is identical to that of Table 2.

Table 5 reports results using both proxies for changes in log total revenues and log total employment. Two caveats are important. The first is that the theoretical model did not generate a strong implication for the relationship between changes in these variables and the initial level of the productivity parameter λ. Firms with greater know-how firms see a larger increase in exports but also a larger decrease in domestic sales; the net effect is ambiguous. The second is that the EIA panel is beset by two somewhat unusual forms of measurement error related to the reporting patterns of multi-establishment firms.[46] On one hand, some self-contained production complexes operated by a single firm are divided into several establishments for the purposes of responding to the survey. Moving one stage of the production process from one building to the next will show up in the data as a spurious increase in both employment and sales in one establishment and a spurious decrease in another. On the other hand, the EIA survey allows multi-establishment firms to report information for more than one establishment on a single questionnaire, but INEGI did not keep track of these reporting patterns in a systematic way before 1998. Unobserved changes in the pattern of reporting would have similar consequences in the data to the shifting of production lines between buildings. Both of these peculiarities generate measurement error that is highly correlated between sales and employment, and also correlated over time. This form of measurement error may be part of the explanation for the sharp contrast between results for the changes in log employment and log revenues using the two different productivity proxies. In the 1993-1997 period, we see that the estimates are close to zero and insignificant when using log domestic sales, even instrumenting domestic sales in 1994 with the previous year, and positive and highly significant when using the first principal component. It is possible that the correlated measurement error between domestic sales and employment is biasing the estimates downward, and that the persistence of the measurement over time is such that instrumenting with a previous year does not remove the error. The fact that the results differ to such an extent under the principal-component approach lend credence to this view. It is also possible, however, that changes in revenues and employment were not in fact larger in initially more-productive plants. These results should be interpreted with caution.

[45] Recall that the productivity proxy differs slightly across regressions, since the variable entered in changes on the left-hand side is omitted from the principal component calculation.

[46] Unfortunately, the EIA survey does not report whether or not a plant is part of a multi-establishment firm, and a natural test – to re-estimate the model only with single establishment firms – is unavailable.

We now turn to the EIA 1984-2001 Balanced Panel, in order to compare the 1993-1997 period to earlier periods as well. I re-estimate the basic model for four periods: 1986-1989,[47] 1989-1993, 1993-1997 and 1997-2001. Table 6 reports the results for the same model as in Table 2, columns (1)-(2) for these four periods. The results for 1993-1997 and 1997-2001 are consistent with the results in Table 2. The estimates for the 1989-1993 resemble those for the 1997-2001 period. We see no evidence of differential upgrading in wages or the wage ratio. The only variable for which we have a result that might be interpreted as evidence of upgrading is the capital-labor ratio. In 1986-1989, by contrast, there is evidence of a larger increase in the white-collar wage among initially more productive plants. It is notable that 1986 was also a year in which the peso depreciated markedly (as can be seen in Figure 2a above), losing 46% of its value over the course of the year.[48] The results for 1986-1989 can be seen as providing additional support for the hypothesis of differential quality upgrading induced by exchange rate movements.[49]

The one coefficient estimate in Table 6 that is out of line with the results from tables 2 and 4 is the significant coefficient for the employment ratio in 1993-1997. This finding does not appear to be robust, however. Table 7 reports results analogous to Table 6, but using the first principal component as the productivity proxy. The results are quite similar to those of Table 6, except that the estimate for the employment ratio in 1993-1997 is no longer significant.[50]

The estimates above have been based on balanced panels from which plants that went out of business have been removed completely. Selection bias is a possible concern. In order to test for this possibility, I estimate my basic model using the EIA 1993-2001 Unbalanced Panel and the standard two-step selection correction method (Heckman, 1976).[51] Table 8 reports the results for the 1993-1997 and for the 1997-2001 periods. The first-stage coefficient on initial-year log domestic sales is positive and highly significant, as the theoretical model would lead us to expect. The results for the export percentage, wages, the wage ratio, the capital-labor ratio, and the employment ratio are in line with the results for Table 2, which we have already discussed. We

[47] The choice of 1986 as the initial year is dictated by the lack of data on exports in 1984-1985.

[48] This occurred through ongoing market pressures and adjustments of the exchange rate targets, rather than a single discrete shock.

[49] One reason that the effects may be weaker in this period than the 1993-1997 period is that Mexican firms in the early stages of liberalization had little knowledge or experience with exporting. Although the number of plants with some exports in the EIA 1984-2001 Balanced Panel rose from 22% to 33% over the period, the plants were not successful at selling large volumes of goods. The export share of total output in the panel rose from 5% to just 8% over the same period.

[50] I do not report a separate table results using the IV method of Table 2 columns (3)-(4). Results were similar to those in Table 6.

[51] I do not have instruments that affect only the probability of remaining in the sample and not my other outcome variables. The coefficient on the inverse Mills' ratio term in the second stage is thus identified only by the assumption of bivariate normality of the first and second stage errors. Nevertheless, the selection-corrected estimates may have diagnostic value: if the results were significantly different from the results for balanced panels reported above, we would have cause for concern.

can simply note again that the estimates for the changes in wages and in the wage ratio in the 1993-1997 period, as well as the contrast between results for the 1993-1997 and 1997-2001 period, are quite robust.

4.4 Additional Findings

In this sub-section, I discuss additional findings from the Encuesta Nacional de Empleo, Salarios, Tecnología y Capacitación (ENESTyC) [National Survey of Employment, Wages, Technology and Training], a special survey carried out by INEGI in 1992, 1995 and 1999. Although the ENESTyC has not followed individual plants over time as closely as the EIA and EIM, and consequently appears to contain more measurement error, it has two advantages: first, it collected information on a wider range of plant characteristics; and second, it contains information on the *maquiladora* sector. Details on the sampling design and variable definitions appear in the data appendix. I use the ENESTyC data to address three issues. The first is whether changes in the wider range of plant characteristics we observe are consistent with the hypothesis of differential quality upgrading. The second is whether we can use the wider range of plant characteristics to distinguish between the alternative interpretations of the increases in wages within occupational categories set out in the theoretical part above: whether the wage changes reflect changes in the sorting of heterogeneous workers into plants, changes in training programs, or changes in wage premia offered for efficiency-wage reasons. The third is to what extent Feenstra and Hanson's outsourcing hypothesis is consistent with the information in this newly available dataset.

Table 9 reports means and standard deviations for relevant variables from the ENESTyC in the three waves of the survey, separately for maquiladoras, non-maquila non-exporters, and non-maquila exporters. To begin, I consider only the differences between exporters and non-exporters in the non-maquiladora sector, and return to the maquiladora sector when discussing the Feenstra-Hanson model below. The most relevant variable for this paper is the indicator of ISO 9000 certification, an international production standard certified by independent local certifying agencies.[52] We see that in cross-section, exporters are more likely than non-exporters to have the certification in 1995, and the difference increases in 1998. We also see that exporters have higher average schooling among white-collar workers and blue-collar workers, lower absentee rates, a higher likelihood of providing formal training, and a lower fraction of manual (as opposed

[52] The ISO 9000 group of industry-specific standards mainly concern management and documentation procedures, rather than the quality of output, but the common view among managers in Mexico is that ISO certification also signals high product quality. ISO certification is not cheap talk. The process of obtaining certification is time-consuming and expensive - typically taking between nine months and two years, and costing on average $187,000 (1996 U.S. dollars) (Guler et al, 2002).

to electronic or numerically controlled) equipment.[53] As in the EIA data, exporters on average employ more workers and pay higher wages to both types of workers, but especially to white-collar workers, but employ roughly the same ratio of white- to blue-collar workers in production. Somewhat surprisingly, exporters also have a higher reported accident rate and a higher turnover rate.[54]

We now turn to the question of whether the changes in these variables over time are consistent with the differential quality upgrading hypothesis. I estimate the same econometric model as used with the EIA data above. Because the information from the EIA on domestic sales is more reliable than in the ENESTyC, I focus on the plants that appear in both datasets and use the EIA information. (See the data appendix for details on the linked panels.) Table 10 reports the results, using log domestic sales as the know-how proxy. The strongest result is for the ISO 9000 certification indicator. The magnitude suggests that a doubling of domestic sales in the initial year would lead to an increase in the likelihood of acquiring the certification of 8%. Although I do not have access to data on changes in ISO certification for other periods,[55] and hence am not able to construct an estimate of what would have happened in the absence of the peso crisis, this result appears to provide fairly strong corroborating evidence that plants with initially larger domestic sales did in fact improve their quality levels.

The remaining results for the 1994-1998 period do not provide evidence of changes in plant behavior. The estimates for the changes in whether the plant has a formal training program, the accident rate, the absentee rate, and the turnover rate are statistically indistinguishable from zero and imprecisely estimated. If we believe that higher efficiency wages should be accompanied by less absenteeism and lower turnover, then these results may be interpreted as evidence against the efficiency-wage hypothesis and in favor of the hypothesis that quality upgrading requires hiring workers of higher inherent skill. It is also possible, however, that the non-results for these variables are due to the fact that they are extremely poorly measured, and carry little signal of the true behavior of the plant.

The data from the 1992 and 1999 waves of the ENESTyC (but not the 1995 wave) include data on average schooling by occupational category. These data also bear on the sorting vs. efficiency wage issue. If increasing observed wages only reflected rising efficiency wages, we would not expect

[53] Note that the table reports standard deviations of the data, not standard errors of the means. These differences are all highly statistically significant.

[54] Note that the overall averages for several of these variables, especially the absentee rate, the accident rate and the turnover rate, appear to be quite volatile across waves, which most likely arise from subtle differences in the wording of questions in the different questionnaires. This suggests that the results for these variables should be interpreted with caution.

[55] The ENESTyC was carried out again in 2001, and has recently become available. I hope in the future to gain access to the data, and to be able to compare changes over the 1998-2000 period with earlier changes.

to see evidence of rising average schooling within occupational category. The results, reported in Columns (3) and (4) of Table 10, are mixed. The initial level of log domestic sales is significantly related to the change in average schooling of blue-collar workers. Somewhat surprisingly, the coefficient for the average schooling of white-collar workers, while positive, is markedly smaller than that for blue-collar workers and not significant. Given the finding from the EIA that the relative wage of white-collar workers rose more in firms with greater domestic sales, these findings suggest that efficiency wages may be more important for white-collar than blue-collar workers. It is worth emphasizing again, however, that the sample sizes are small, the data are noisy, and we do not have a counterfactual comparison of what would have happened in the absence of the peso crisis. These results should be seen as suggestive rather than definitive.

Finally, we are in a position to re-evaluate the outsourcing hypothesis mentioned in the introduction in the light of these micro-datasets which were not available to Feenstra and Hanson. This paper has presented evidence that the peso crisis induced differential changes in wages within industries, but not differential changes in the proportion of white- and blue-collar workers in production. The wage changes thus appear not to be explained by a demand shift toward greater numbers of white-collar workers. A defense of the outsourcing hypothesis would be to argue that the results in this paper are focused on the non-maquiladora sector, that outsourcing is likely to be more important in the maquiladora sector, and that the rapid growth of the maquiladora sector following the peso crisis may still have been responsible for the aggregate increase in inequality. But Table 9 shows that maquiladoras employ a greater proportion of *blue-collar* workers – and that the blue-collar workers that they employ have less education on average – than either exporters or non-exporters in the rest of the manufacturing sector.[56] These facts suggest that the first-order consequence of the growth of the maquiladora sector in Mexico was an increase in the relative demand for low-skilled workers.

5 Conclusion

This paper has set out a new model that links expanding trade and increasing inequality through the mechanism of differential quality upgrading within industries. The paper has also presented a new empirical approach, derived directly from the theoretical framework, that uses the exchange rate shock of the late-1994 peso crisis to evaluate the implications of the theoretical model. I find robust evidence of differential quality upgrading in the 1993-1997 period and essentially no

[56] This observation is due originally to Pablo Ibarraran (2003). In unreported results, I find that these patterns hold true even once we control for industry and region effects.

evidence in the 1989-1993 and 1997-2001 periods. These results strongly support the quality-upgrading hypothesis.

A number of questions remain outstanding. One is to what extent the quality-upgrading hypothesis can explain changes in inequality at the aggregate level in Mexico. There are three channels through which differential quality upgrading within industries may contribute to aggregate wage inequality. First, it raises wage dispersion within occupational categories between less-productive and more-productive plants. Second, it may lead to a re-allocation of production toward more-productive plants, which are more unequal than less-productive plants. Third, it generates an increase in wage inequality within the plants that upgrade. Together these channels may account for aggregate within-industry increases in wage inequality that more than offset the between-industry shifts toward industries intensive in unskilled labor illustrated by Figure 1 in the introduction. To what extent they do so is a topic for future work.

A second outstanding question is the generalizability of the quality-upgrading hypothesis. The empirical analysis in this paper is specific to a particular historical event with particular characteristics. Nonetheless, there are reasons to believe that differential quality upgrading may occur in response to other trade-related shocks. We would expect a bilateral reduction in tariffs, transport costs or fixed costs of exporting to have an effect similar to an exchange-rate shock on firms in a country like Mexico: the incentive to export increases and domestic sales decrease (here because of an increase in import competition rather than a contraction of domestic demand) and differential quality upgrading generates an increase in the relative wage of white-collar workers.[57] If this is the case, the quality upgrading hypothesis is likely to be a part of the explanation of the link between trade liberalization and wage inequality more generally.

Final judgment on this issue must await more empirical work, in other countries, responding to other shocks. But if the quality-upgrading hypothesis stands up, then it may help to resolve a second, related puzzle in the political economy of trade policy. The simplest Hecksher-Ohlin model predicts that in developing countries unskilled workers will be winners and skilled workers and owners of capital losers from international integration. From this perspective, it is a mystery why in many developing countries, especially in Latin America, the biggest proponents of integration are educated, urban elites. This paper has also identified winners and losers from trade, but along a different dimension. The winners are the entrepreneurs and employees, especially the white-collar employees, with either the qualifications or the good fortune to be employed in the most modern, technically sophisticated, internationally oriented plants within each industry. The losers are people, especially the unskilled, in less-productive, domestically oriented plants. In this

[57] An appendix that works out the formal logic of this argument is available from the author.

view, both the enthusiasm of the most educated and the pessimism of the poorest and least skilled toward globalization may make economic sense.

Appendix A: Theory Appendix

A.1: Micro-Foundations of Aggregate Demands

Following Andersen et al (1992)'s extension of McFadden (1978, 1981), we assume the random match-specific term ε_{ij} in (1) has c.d.f. $F(\varepsilon) = \exp\left(-\exp\left(-\frac{\varepsilon}{\mu}+\gamma\right)\right)$, where $\gamma = .5772$ (Euler's constant) is included to ensure that the expectation of ε_{ij} is zero, and μ is a positive constant. The probability that consumer i chooses variety j, call it $\overline{P}_{ij}$ can be derived as follows:

$$\begin{aligned}
\overline{P}_{ij} &= \Pr(\varepsilon_{i1}, \varepsilon_{i2}, ...\varepsilon_{iJ} | V_{ij} > V_{it}, \forall t \neq j) \\
&= \Pr(\varepsilon_{it} < \theta q_j - p_j - (\theta q_t - p_t) + \varepsilon_{ij}, \forall t \neq j)
\end{aligned}$$

Because the $\varepsilon' s$ are i.i.d., the probability that variety j is chosen conditional on a particular value of ε_{ij}, is the product of the c.d.f.s:

$$\overline{P}_{ij} | \varepsilon_{ij} = \prod_{t \neq j} F(\theta q_j - p_j - (\theta q_t - p_t) + \varepsilon_{ij})$$

The marginal probability is the expected value of the conditional probability over all ε_{ij}:

$$\begin{aligned}
\overline{P}_{ij} &= \int_{-\infty}^{\infty} \left\{ \prod_{t \neq j} F(\theta q_j - p_j - (\theta q_t - p_t) + \varepsilon) \right\} f(\varepsilon) d\varepsilon \\
&= \int_{-\infty}^{\infty} \left\{ \prod_{t \neq j} \exp\left[-\exp\left(-\frac{\theta q_j - p_j - (\theta q_t - p_t) + \varepsilon}{\mu} + \gamma\right)\right] \right\} f(\varepsilon) d\varepsilon
\end{aligned}$$

The p.d.f. of the distribution of ε_{ij} is given by:

$$f(\varepsilon) = \frac{1}{\mu} \exp\left(-\frac{\varepsilon}{\mu} + \gamma\right) \exp\left[-\exp\left(-\frac{\varepsilon}{\mu} + \gamma\right)\right]$$

Let $\phi \equiv \exp\left(-\frac{\varepsilon}{\mu} + \gamma\right)$. Then $d\phi = -\frac{1}{\mu} \exp\left(-\frac{\varepsilon}{\mu} + \gamma\right) d\varepsilon$.
The marginal probability can then be written:

$$\begin{aligned}
\overline{P}_{ij} &= -\int_{\infty}^{0} \exp(-\phi) \prod_{t \neq j} \exp\left[-\phi \exp\left(-\left[\theta q_j - p_j - (\theta q_t - p_t)\right]/\mu\right)\right] d\phi \\
&= -\int_{\infty}^{0} \exp(-\phi) \exp\left\{-\phi \exp\left[-(\theta q_j - p_j)/\mu\right] \sum_{t \neq j} \exp\left[(\theta q_t - p_t)/\mu\right] \right\} d\phi \\
&= -\int_{\infty}^{0} \exp\left\{-\phi \left[1 + \exp\left[-(\theta q_j - p_j)/\mu\right] \sum_{t \neq j} \exp\left[(\theta q_t - p_t)/\mu\right]\right]\right\} d\phi
\end{aligned}$$

$$= -\int_{\infty}^{0} \exp\left\{-\phi\left[\frac{\sum_{t=1}^{J}\exp\left[\left(\theta q_t - p_t\right)/\mu\right]}{\exp\left[\left(\theta q_j - p_j\right)/\mu\right]}\right]\right\} d\phi$$

$$= \frac{\exp\left[\left(\theta q_j - p_j\right)/\mu\right]}{\sum_{t=1}^{J}\exp\left[\left(\theta q_t - p_t\right)/\mu\right]}$$

The expected aggregate demand for variety j, for all N consumers, is then:

$$E\left(x_j\right) = N \cdot \overline{P}_{ij} = \frac{N\exp\left[\left(\theta q_j - p_j\right)/\mu\right]}{\sum_{t=1}^{J}\exp\left[\left(\theta q_t - p_t\right)/\mu\right]}$$

A.2: Wage Ratios

From (9), we have: $w^{h*}(\lambda) = \underline{w}^h + \alpha^h\theta q^*(\lambda)$ and $w^{l*}(\lambda) = \underline{w}^l + \alpha^l\theta q^*(\lambda)$, where $q^*(\lambda) = (\eta\lambda\theta^{\alpha})^{1/1-\alpha}$. Making the dependence of the wage and quality terms on λ implicit to reduce clutter:

$$\begin{aligned}
\frac{\partial\left(w^{h*}/w^{l*}\right)}{\partial\lambda} &= \frac{\left(\partial w^{h*}/\partial\lambda\right)\left\{\underline{w}^l + \alpha^l\theta q^*\right\} - \left(\partial w^{l*}/\partial\lambda\right)\left\{\underline{w}^h + \alpha^h\theta q^*\right\}}{\left\{w^{l*}\right\}^2} \\
&= \frac{\frac{\theta q^*}{(1-\alpha)\lambda}\left\{\alpha^h\underline{w}^l - \alpha^l\underline{w}^h\right\}}{\left\{w^{l*}\right\}^2}
\end{aligned}$$

The numerator will be positive, and $\frac{\partial\left(w^{h*}/w^{l*}\right)}{\partial\lambda} > 0$ if and only if $\alpha^h/\alpha^l > \underline{w}^h/\underline{w}^l$.
Similarly:

$$\begin{aligned}
\frac{\partial\left(w^{h*}/w^{l*}\right)}{\partial\theta} &= \frac{\left(\partial w^{h*}/\partial\theta\right)\left\{\underline{w}^l + \alpha^l\theta q^*\right\} - \left(\partial w^{l*}/\partial\theta\right)\left\{\underline{w}^h + \alpha^h\theta q^*\right\}}{\left\{w^{l*}\right\}^2} \\
&= \frac{\frac{q^*}{(1-\alpha)}\left\{\alpha^h\underline{w}^l - \alpha^l\underline{w}^h\right\}}{\left\{w^{l*}\right\}^2}
\end{aligned}$$

Again, the numerator will be positive and $\frac{\partial\left(w^{h*}/w^{l*}\right)}{\partial\theta} > 0$ if and only if $\alpha^h/\alpha^l > \overline{w}_h^k/\overline{w}_l^k$.

A.3: Export Share of Output vs. Know-how Parameter

From (18):

$$\begin{aligned}
\frac{\partial x_{cd}^*(\lambda)}{\partial\lambda} &= \frac{\partial I_{cd}^*}{\partial\lambda}x_{cd} \\
&= \left(\eta\delta_{cd}^{\alpha}\lambda^{\alpha}\theta_d\right)^{1/1-\alpha} x_{cd}^*(\lambda)
\end{aligned} \tag{26}$$

conditional on the firm entering consumer market d. From the definition of $\chi_s^*(\lambda)$, we have:

$$\begin{aligned}
\frac{\partial \chi_s^*(\lambda)}{\partial \lambda} &= \frac{\left(x_{ss}^*(\lambda)+x_{sn}^*(\lambda)\right)\frac{\partial x_{sn}^*(\lambda)}{\partial \lambda}-x_{sn}^*(\lambda)\left(\frac{\partial x_{ss}^*(\lambda)}{\partial \lambda}+\frac{\partial x_{sn}^*(\lambda)}{\partial \lambda}\right)}{\left(x_{ss}^*(\lambda)+x_{sn}^*(\lambda)\right)^2} \\
&= \frac{x_{ss}^*(\lambda)\,x_{sn}^*(\lambda)\left(\frac{\partial I_{sn}^*(\lambda)}{\partial \lambda}-\frac{\partial I_{ss}^*(\lambda)}{\partial \lambda}\right)}{\left(x_{ss}^*(\lambda)+x_{sn}^*(\lambda)\right)^2} \\
&= \chi_s^*(\lambda)\left(1-\chi_s^*(\lambda)\right)\left(\eta\delta_{cd}^{\alpha}\lambda^{\alpha}\right)^{1/1-\alpha}\left\{\theta_n^{1/1-\alpha}-\theta_s^{1/1-\alpha}\right\}>0
\end{aligned}$$

for all $\lambda>0$, since $\theta_n>\theta_s$ by assumption.

A.4: Export Share of Sales vs. Know-how Parameter

Noting that $S_{ss}^*(\lambda)=p_{ss}^*(\lambda)\,x_{ss}^*(\lambda)$ and $S_{sn}^*(\lambda)=p_{sn}^*(\lambda)\,x_{sn}^*(\lambda)$, take the partial derivative of both sides of the definition of $\sigma_s(\lambda)$ with respect to λ. (For notational convenience, I drop the explicit dependence of p_{ss}, p_{sn}, x_{ss} and x_{sn} on λ.)

$$\frac{\partial \sigma_s}{\partial \lambda}=\sigma_s(\lambda)\left(1-\sigma_s(\lambda)\right)\left\{\left[\frac{\frac{\partial x_{sn}}{\partial \lambda}}{x_{sn}}-\frac{\frac{\partial x_{ss}}{\partial \lambda}}{x_{ss}}\right]+\left[\frac{\frac{\partial p_{sn}}{\partial \lambda}}{p_{sn}}-\frac{\frac{\partial p_{ss}}{\partial \lambda}}{p_{ss}}\right]\right\} \tag{27}$$

From (26) in Appendix A.3 we have that

$$\frac{\partial x_{sd}}{\partial \lambda}=\left[\eta\theta_d\left(\lambda\delta_{sd}\right)^{\alpha}\right]^{1/1-\alpha}x_{sd}$$

Simplifying the first term in square brackets in (27), we have:

$$\left[\frac{\frac{\partial x_{sn}}{\partial \lambda}}{x_{sn}}-\frac{\frac{\partial x_{ss}}{\partial \lambda}}{x_{ss}}\right]=\eta\lambda^{\alpha/1-\alpha}\left[\left(\theta_n\delta_{sn}^{\alpha}\right)^{1/1-\alpha}-\left(\theta_s\right)^{1/1-\alpha}\right]>0 \tag{28}$$

Since $\theta_n>\theta_s$ and $\delta_{sn}\geq 1$.
From (18), we have that

$$\frac{\partial p_{sd}}{\partial \lambda}=\alpha\delta_{sd}\theta_d\frac{\partial q_{sd}}{\partial \lambda}=\frac{\alpha\delta_{sd}\theta_d q_{sd}}{(1-\alpha)\,\lambda}$$

Simplifying the second term in square brackets in (27), we have:

$$\left[\frac{\frac{\partial p_{sn}}{\partial \lambda}}{p_{sn}}-\frac{\frac{\partial p_{ss}}{\partial \lambda}}{p_{ss}}\right]=\frac{\alpha}{(1-\alpha)\,\lambda}\left[\frac{\delta_{sn}\theta_n q_{sn}}{p_{sn}}-\frac{\delta_{ss}\theta_s q_{ss}}{p_{ss}}\right]$$

Substituting for p_{sn} and p_{ss} from (18) and defining $\phi\equiv\mu+w_s^h+w_s^l$, the term in brackets on the right hand side can be re-written:

$$\begin{aligned}
\frac{\delta_{sn}\theta_n q_{sn}}{p_{sn}}-\frac{\delta_{ss}\theta_s q_{ss}}{p_{ss}} &= \frac{1}{p_{sn}p_{ss}}\left\{\delta_{sn}\theta_n q_{sn}\left[\mu+w_s^h+w_s^l+\alpha\theta_s q_{ss}\right]\right\}- \\
&\quad \frac{1}{p_{sn}p_{ss}}\left\{\theta_s q_{ss}\left[\mu+w_s^h+w_s^l+\alpha\delta_{sn}\theta_n q_{sn}\right]\right\}
\end{aligned} \tag{29}$$

$$= \quad \frac{1}{p_{sn}p_{ss}} \left(\mu + w_s^h + w_s^l\right) \left(\delta_{sn}\theta_n q_{sn} - \theta_s q_{ss}\right) > 0$$

From (27), (28) and (29), we can conclude that $\partial \sigma_s / \partial \lambda > 0$.

A.5: Comparative-Static Results for Entry

The zero-profit conditions determining entry (20) can be rewritten:

$$D_s^*\left(\lambda_{ss}^{\min}, \lambda_{ns}^{\min}\right) = \frac{\mu N_s}{f_{ss}} \exp\left\{I_{ss}^*\left(\lambda_{ss}^{\min}\right)\right\} \tag{30}$$

$$D_s^*\left(\lambda_{ss}^{\min}, \lambda_{ns}^{\min}\right) = \frac{\mu N_s}{f_{ns}} \exp\left\{I_{ns}^*\left(\lambda_{ns}^{\min}\right)\right\} \tag{31}$$

$$D_n^*\left(\lambda_{sn}^{\min}, \lambda_{nn}^{\min}\right) = \frac{\mu N_n}{f_{sn}} \exp\left\{I_{sn}^*\left(\lambda_{sn}^{\min}\right)\right\} \tag{32}$$

$$D_n^*\left(\lambda_{sn}^{\min}, \lambda_{nn}^{\min}\right) = \frac{\mu N_n}{f_{nn}} \exp\left\{I_{nn}^*\left(\lambda_{nn}^{\min}\right)\right\} \tag{33}$$

Consider first the effect of the devaluation of the peso. Taking the derivative of both sides of (30) and (31) with respect to δ_{sn} and solving for $\partial \lambda_{ss}^{\min} / \partial \delta_{sn}$, we have:

$$\frac{\partial \lambda_{ss}^{\min}}{\partial \delta_{sn}} = \frac{1}{\Omega_s} \left\{ \frac{\partial I_{ns}^*\left(\lambda_{ns}^{\min}\right)}{\partial \lambda_{ns}^{\min}} \left[\frac{\partial D_s^*}{\partial \delta_{sn}}\right] + \frac{\partial D_s^*}{\partial \lambda_{ns}^{\min}} \left[\frac{\partial I_{ss}^*\left(\lambda_{ss}^{\min}\right)}{\partial \delta_{sn}} - \frac{\partial I_{ns}^*\left(\lambda_{ns}^{\min}\right)}{\partial \delta_{sn}} \right] \right\} \tag{34}$$

where

$$\Omega_s \equiv D_s^* \frac{\partial I_{ss}^*\left(\lambda_{ss}^{\min}\right)}{\partial \lambda_{ss}^{\min}} \frac{\partial I_{ns}^*\left(\lambda_{ns}^{\min}\right)}{\partial \lambda_{ns}^{\min}} - \frac{\partial D_s^*}{\partial \lambda_{ss}^{\min}} \frac{\partial I_{ns}^*\left(\lambda_{ns}^{\min}\right)}{\partial \lambda_{ns}^{\min}} - \frac{\partial D_s^*}{\partial \lambda_{ns}^{\min}} \frac{\partial I_{sd}^*\left(\lambda_{ss}^{\min}\right)}{\partial \lambda_{ss}^{\min}} > 0 \tag{35}$$

It follows from the definition of $I_{cd}^*\left(\lambda_{cd}^{\min}\right)$ that $\partial I_{ss}^*\left(\lambda_{ss}^{\min}\right) / \partial \delta_{sn} = 0$ and $\partial I_{ns}^*\left(\lambda_{ns}^{\min}\right) / \partial \lambda_{ns}^{\min} > 0$. Since $\delta_{ns} \equiv 1/\delta_{sn}$, we also have that $\partial I_{ns}^*\left(\lambda_{ns}^{\min}\right) / \partial \delta_{sn} < 0$. From (16) and (17), we have that $\partial D_s^* / \partial \lambda_{ns}^{\min} < 0$. Note further that:

$$\frac{\partial D_s^*}{\partial \delta_{sn}} = \int_{\lambda_{sd}^{\min}}^{\lambda_s^{\max}} \frac{\partial I_{ns}^*(\lambda)}{\partial \delta_{sn}} \exp\left\{I_{ns}^*(\lambda)\right\} M_s g_s(\lambda)\, d\lambda < 0$$

Given these relationships, we can conclude from (34) that $\partial \lambda_{ss}^{\min} / \partial \delta_{sn} < 0$.

To sign $\partial \lambda_{sn}^{\min} / \partial \delta_{sn}$, we simply take the derivative of both sides of (32). The assumption that aggregate supply to the Northern market is approximately unaffected by the peso crisis is equivalent to assuming that the aggregate quantity D_n^* is unchanged. Thus we have:

$$\frac{\mu N_n}{f_{sn}} \exp\left\{I_{sn}^*\left(\lambda_{sn}^{\min}\right)\right\} \left\{ \frac{\partial I_{sn}^*\left(\lambda_{sn}^{\min}\right)}{\partial \lambda_{sn}^{\min}} \frac{\partial \lambda_{sn}^{\min}}{\partial \delta_{sn}} + \frac{\partial I_{sn}^*\left(\lambda_{sn}^{\min}\right)}{\partial \delta_{sn}} \right\} = 0$$

which yields:

$$\frac{\partial\lambda_{sn}^{\min}}{\partial\delta_{sn}} = -\frac{\frac{\partial I_{sn}^{*}\left(\lambda_{sn}^{\min}\right)}{\partial\delta_{sn}}}{\frac{\partial I_{sn}^{*}\left(\lambda_{sn}^{\min}\right)}{\partial\lambda_{sn}^{\min}}}$$

It follows again from the definition of $I_{cd}^{*}(\cdot)$ that $\partial I_{sn}^{*}\left(\lambda_{sn}^{\min}\right)/\partial\delta_{sn} > 0$ and $\partial I_{sn}^{*}\left(\lambda_{sn}^{\min}\right)/\partial\lambda_{sn}^{\min} > 0$. Thus $\partial\lambda_{sn}^{\min}/\partial\delta_{sn} < 0$.

Now consider the effect of the reduction in the number of consumers in South. Note that D_{s}^{*} and $I_{cs}^{*}\left(\lambda_{cs}^{\min}\right)$ do not depend directly on N_s; they are affected only through $\lambda_{ss}^{\min}$ and $\lambda_{ns}^{\min}$. Hence taking the derivative of both sides of (30) and (31) with respect to N_s and solving for $\partial\lambda_{ss}^{\min}/\partial N_s$, we have:

$$\frac{\partial\lambda_{ss}^{\min}}{\partial N_s} = -\frac{\partial I_{ns}^{*}\left(\lambda_{ns}^{\min}\right)}{\partial\lambda_{ns}^{\min}}\left(\frac{D_s^{*}}{\Omega_s N_s}\right) \tag{36}$$

where Ω_s is defined as in (35) above. Since $\partial I_{ns}^{*}\left(\lambda_{ns}^{\min}\right)/\partial\lambda_{ns}^{\min} > 0$ by the definition of $I_{cd}^{*}(\cdot)$ in (19), it follows that $\partial\lambda_{ss}^{\min}/\partial N_s < 0$. Finally, since no other terms in (32) depend on N_s it follows immediately that $\lambda_{ss}^{\min}$ does not depend on N_s, that is, $\partial\lambda_{sn}^{\min}/\partial N_s = 0$.

Finally, we state precisely the condition required to ensure that the exchange-rate shock induces exit of Southern firms from the Southern market. The net effect of the shock on the cut-off for domestic firms to enter the Southern market is:

$$d\lambda_{ss}^{\min} = \frac{\partial\lambda_{ss}^{\min}}{\partial\delta_{sn}}d\delta_{sn} + \frac{\partial\lambda_{ss}^{\min}}{\partial N_s}dN_s$$

Given our results for $\partial\lambda_{ss}^{\min}/\partial\delta_{sn}$ and $\partial\lambda_{ss}^{\min}/\partial N_s$ in (34) and (36) above, $d\lambda_{ss}^{\min} > 0$ (and the exchange-rate shock induces exit of Southern firms from the Southern market) if and only if:

$$\frac{-dN_s}{d\delta_{sn}} > \left\{-\frac{\partial D_s^{*}}{\partial\delta_{sn}} + \frac{\frac{\partial D_s^{*}}{\partial\lambda_{ns}^{\min}}\cdot\frac{\partial I_{sn}^{*}\left(\lambda_{ns}^{\min}\right)}{\partial\delta_{sn}}}{\frac{\partial I_{sn}^{*}\left(\lambda_{ns}^{\min}\right)}{\partial\lambda_{ns}^{\min}}}\right\}\frac{N_s}{D_s^{*}} \tag{37}$$

We will maintain this assumption in what follows.

A.6: Changes in Firm-level Observables

Note that the total effect of the exchange rate shock on output on each production line can be expressed as:

$$dx_{sd}(\lambda) = \frac{\partial x_{sd}(\lambda)}{\partial\delta_{sn}}d\delta_{sn} + \frac{\partial x_{sd}(\lambda)}{\partial N_s}dN_s \tag{38}$$

for $d = s, n$. Recall from (18) that

$$x_{sd}^{*}(\lambda) = \frac{N_d}{D_d^{*}}\exp\left\{I_{sd}^{*}(\lambda)\right\} \tag{39}$$

Consider first the effect on output on the domestic production line, x_{ss}. Taking the partial derivative of both sides of (39) with respect to δ_{sn}, and noting that $I_{ss}^{*}(\lambda)$ does not depend on δ_{sn}, we have:

$$\frac{\partial x_{ss}(\lambda)}{\partial \delta_{sn}} = -\frac{N_s}{(D_s^*)^2} \exp\{I_{ss}^*(\lambda)\} \frac{\partial D_s^*}{\partial \delta_{sn}} > 0$$

since $\partial D_s^* / \partial \delta_{sn} < 0$. Taking the partial derivative of both sides of (39) with respect to N_s, we have:

$$\frac{\partial x_{ss}(\lambda)}{\partial N_s} = \frac{1}{D_s^*} \exp\{I_{ss}^*(\lambda)\} > 0$$

Substituting these into (38), we have:

$$\begin{aligned} dx_{ss}(\lambda) &= \frac{\exp\{I_{ss}^*(\lambda)\}}{D_s^*} \left\{ -\frac{N_s}{D_s^*} \frac{\partial D_s^*}{\partial \delta_{sn}} d\delta_{sn} + dN_s \right\} \\ &= x_{ss}(\lambda) \left\{ -\frac{1}{D_s^*} \frac{\partial D_s^*}{\partial \delta_{sn}} d\delta_{sn} + \frac{1}{N_s} dN_s \right\} \end{aligned} \tag{40}$$

If the condition required to ensure that the exchange-rate shock induced exit of Southern firms from the domestic market (equation (37) in Appendix A.5) is satisfied, then the term in curly braces will be negative. Thus in the empirically relevant case, $dx_{ss}(\lambda) < 0$.

Now consider the effect on the output of the export production line. Taking the partial of both sides of (39) with respect to δ_{sn}, recalling the assumption that the aggregate D_n^* is approximately unaffected by the changes in exports of Southern firms, we have:

$$\frac{\partial x_{sn}(\lambda)}{\partial \delta_{sn}} = \frac{N_n}{D_n^*} \exp\{I_{sn}^*(\lambda)\} \frac{\partial I_{sn}^*(\lambda)}{\partial \delta_{sn}} > 0$$

since $\partial I_{sn}^*(\lambda) / \partial \delta_{sn} > 0$ by (19). Because of the segmentation of consumer markets, the change in the number of Southern consumers does not affect output of Southern firms for the Northern market. The only change in output for export is the change induced by the change in the exchange rate. Hence:

$$\begin{aligned} dx_{sn}(\lambda) &= \frac{N_n}{D_n^*} \exp\{I_{sn}^*(\lambda)\} \frac{\partial I_{sn}^*(\lambda)}{\partial \delta_{sn}} d\delta_{sn} \\ &= x_{sn}(\lambda) \frac{\partial I_{sn}^*(\lambda)}{\partial \delta_{sn}} d\delta_{sn} \end{aligned} \tag{41}$$

Hence we know $dx_{sn}(\lambda) > 0$.

Now consider the change in the export share of output for Southern firms in case (5) above, the always exporters. Recall that $\chi_s(\lambda) \equiv x_{sn}(\lambda) / (x_{ss}(\lambda) + x_{sn}(\lambda))$. Thus:

$$\begin{aligned} d\chi_s &= \frac{(x_{ss}(\lambda) + x_{sn}(\lambda)) dx_{sn}(\lambda) - x_{sn}(\lambda)(dx_{ss}(\lambda) + dx_{sn}(\lambda))}{(x_{ss}(\lambda) + x_{sn}(\lambda))^2} \\ &= \chi_s(\lambda)(1 - \chi_s(\lambda)) \left\{ \frac{dx_{sn}(\lambda)}{x_{sn}(\lambda)} - \frac{dx_{ss}(\lambda)}{x_{ss}(\lambda)} \right\} \\ &= \chi_s(\lambda)(1 - \chi_s(\lambda)) \left\{ \frac{\partial I_{sn}^*(\lambda)}{\partial \delta_{sn}} d\delta_{sn} + \frac{1}{D_s^*} \frac{\partial D_s^*}{\partial \delta_{sn}} d\delta_{sn} - \frac{1}{N_s} dN_s \right\} \end{aligned}$$

where $\chi_s(\lambda)$ is the export share of output prior to the shock.

Let $\Sigma_s \equiv \left\{ \frac{\partial I^*_{sn}(\lambda)}{\partial \delta_{sn}} d\delta_{sn} + \frac{1}{D^*_s} \frac{\partial D^*_s}{\partial \delta_{sn}} d\delta_{sn} - \frac{1}{N_s} dN_s \right\}$. Given the results (40) and (41) above, and the facts that $d\delta_{sn} > 0$ and $dN_s < 0$, we know that $\Sigma_s > 0$. From the definition of $I^*_{sn}(\lambda)$ in (19), it follows that $\partial I^*_{sn}(\lambda) / \partial \delta_{sn} = \left[\eta \theta_n (\delta_{sn} \lambda)^{\alpha} \right]^{1/(1-\alpha)}$. Hence Σ_s is increasing in λ.

A.7 Effects of Exchange Rate Shock on the Wage Ratio

Note that:

$$d\overline{\omega}^*_s(\lambda) = d\left(\frac{\overline{w}^{h*}_s(\lambda)}{\overline{w}^{l*}_s(\lambda)} \right) = \frac{\overline{w}^{l*}_s(\lambda)\, d\overline{w}^{h*}_s(\lambda) - \overline{w}^{h*}_s(\lambda)\, d\overline{w}^{l*}_s(\lambda)}{\left(\overline{w}^{l*}_s(\lambda) \right)^2} \tag{42}$$

where $\overline{w}^{h*}_s(\lambda)$ and $\overline{w}^{l*}_s(\lambda)$ were defined in (21). Substituting the expressions for $\overline{w}^{h*}_s(\lambda)$ and $\overline{w}^{l*}_s(\lambda)$ from (21) and for $d\overline{w}^{h*}_s(\lambda)$ and $d\overline{w}^{l*}_s(\lambda)$ from (24) and simplifying, we have:

$$d\overline{\omega}^*_s(\lambda) = \left[\frac{\alpha^h}{\alpha^l} - \frac{\underline{w}^h_s}{\underline{w}^l_s} \right] \frac{d\overline{w}^{l*}_s(\lambda)}{\overline{w}^{l*}_s(\lambda)}$$

If $\alpha^h / \alpha^l > \underline{w}^h_s / \underline{w}^l_s$ (which was the condition for the wage ratio to be increasing in λ in cross-section) the wage ratio will increase whenever the average blue-collar wage increases, and the pattern of qualitative changes will be similar to that of the other observables, average wages for each type of worker and capital intensity.

A.8 Effects of Exchange Rate Shock on the Export Share of Sales

Consider the now-familiar counterfactual where all firms enter both markets. First, note that $\sigma_s(\lambda)$ can be rewritten (dropping the explicit dependence of p_{sd} and x_{sd} on λ) as

$$\sigma_s(\lambda) = \frac{1}{1 + \frac{p_{ss} x_{ss}}{p_{sn} x_{sn}}}$$

Thus $\sigma_s(\lambda)$ increases in response to the shock if and only if $\frac{p_{ss} x_{ss}}{p_{sn} x_{sn}}$ decreases. We can write:

$$d\left(\frac{p_{ss} x_{ss}}{p_{sn} x_{sn}} \right) = \left(\frac{p_{ss}}{p_{sn}} \right) d\left(\frac{x_{ss}}{x_{sn}} \right) + \left(\frac{x_{ss}}{x_{sn}} \right) d\left(\frac{p_{ss}}{p_{sn}} \right)$$

Note that $x_{ss}/x_{sn} = -1 + 1/\chi_s(\lambda)$. Hence:

$$d\left(\frac{x_{ss}}{x_{sn}} \right) = -\frac{1}{\left(\chi_s(\lambda) \right)^2} d\chi_s(\lambda)$$

In the case that $\chi_s(\lambda)$, which ensures that $d\chi_s(\lambda) > 0$, we have that $d\left(\frac{x_{ss}}{x_{sn}} \right) < 0$.

We can also write:

$$d\left(\frac{p_{ss}}{p_{sn}} \right) = \frac{p_{sn} dp_{ss} - p_{ss} dp_{sn}}{\left(p_{sn} \right)^2}$$

Note from (18) that neither p_{ss} not p_{sn} depend on N_s, and p_{ss} does not depend on δ_{sn}. This expression can thus be rewritten:

$$d\left(\frac{p_{ss}}{p_{sn}}\right) = \frac{-p_{ss}}{(p_{sn})^2} \cdot \frac{\partial p_{sn}}{\partial \delta_{sn}} < 0$$

since, from (18), $\partial p_{sn}/\partial \delta_{sn} > 0$. This fact, and the fact that $d\left(\frac{x_{ss}}{x_{sn}}\right) < 0$, together imply that $d\left(\frac{p_{ss}x_{ss}}{p_{sn}x_{sn}}\right) < 0$. Hence we can conclude that $d\sigma_s(\lambda) > 0$. The export share of sales increases in response to the exchange-rate shock, conditional on firms entering both markets. The pattern of entry will give us an actual pattern of changes in $\sigma_s(\lambda)$ similar to that for the export share of output.

A.9: Effects of Exchange Rate Shock on Total Employment, Sales

For firms that enter both markets, we have from (40) and (41) in Appendix A.6 that:

$$\begin{aligned} dx_{ss}(\lambda) &= x_{ss}(\lambda)\left\{-\frac{1}{D_s^*}\frac{\partial D_s^*}{\partial \delta_{sn}}d\delta_{sn} + \frac{1}{N_s}dN_s\right\} < 0 \\ dx_{sn}(\lambda) &= x_{sn}(\lambda)\frac{\partial I_{sn}^*(\lambda)}{\partial \delta_{sn}}d\delta_{sn} > 0 \end{aligned}$$

For firms that enter only the domestic market, total output is given by $x_{ss}(\lambda)$. These firms will see total output fall in response to the exchange rate shock. Firms that enter both markets have total output $x_{ss}(\lambda) + x_{sn}(\lambda)$. For these firms, there are two effects. On one hand, $x_{ss}(\lambda)$ is increasing in λ (and the term in braces in the top equation is independent of λ); thus the decline in domestic output is larger for higher-λ firms. On the other hand, we have seen above that $x_{sn}(\lambda)$ and $\partial I_{sn}^*(\lambda)/\partial \delta_{sn}$ are increasing in λ; thus the increase in output for the export market is larger for higher-λ firms. It is not clear which effect will predominate. We can make an analogous argument for total sales.

A.10: Relationship Between log(domestic sales) and Underlying Productivity

From the definition of sales and (18), we have:

$$\begin{aligned} S_{ss} &= p_{ss}x_{ss} \\ &= \left[\mu + \delta_s \underline{w}_s^h + \delta_s \underline{w}_s^l + \alpha\theta_s q_{ss}\right]\left[\frac{N_s}{D_s}\exp\{I_{ss}\}\right] \end{aligned}$$

Taking the log of both sides and substituting for I_{ss} from (19), we have:

$$\begin{aligned} \ln(S_{ss}) &= \ln\left[\mu + \delta_s \underline{w}_s^h + \delta_s \underline{w}_s^l + \alpha(\eta\lambda\theta_s)^{\frac{1}{1-\alpha}}\right] + \ln N_s - \ln D_s \\ &\quad + \left[(1-\alpha)(\eta\theta_s\lambda)^{\frac{1}{1-\alpha}} - \mu - \delta \underline{w}_s^h - \delta \underline{w}_s^l\right]/\mu \end{aligned}$$

Taking a first-order Taylor expansion of this expression around the mean value of λ, call it $\overline{\lambda}$, yields:

$$\ln(S_{ss}) = A + B\left(\lambda - \overline{\lambda}\right)$$

where

$$A = \ln\left[\mu + \delta_s \underline{w}_s^h + \delta_s \underline{w}_s^l + \alpha\left(\eta\overline{\lambda}\theta_s\right)^{\frac{1}{1-\alpha}}\right] + \left[(1-\alpha)\left(\eta\theta_s\overline{\lambda}\right)^{\frac{1}{1-\alpha}} - \mu - \delta\underline{w}_s^h - \delta\underline{w}_s^l\right]/\mu$$

$$B = (\eta\theta_s)^{\frac{1}{1-\alpha}}\left(\overline{\lambda}\right)^{\frac{\alpha}{1-\alpha}}\left\{1 + \frac{\alpha}{(1-\alpha)\left[\mu + \delta_s \underline{w}_s^h + \delta_s \underline{w}_s^l + \alpha\left(\eta\overline{\lambda}\theta_s\right)^{\frac{1}{1-\alpha}}\right]}\right\}$$

Note that $B > 0$. Taking the mean of $\ln(S_{ss})$ over all plants, we have:

$$\overline{\ln(S_{ss})} = A + B\left(\overline{\lambda} - \overline{\lambda}\right) = A$$

Thus we can write:

$$\lambda - \overline{\lambda} = \frac{1}{B}\left(\ln(S_{ss}) - \overline{\ln(S_{ss})}\right)$$

where $1/B > 0$. Thus the underlying productivity parameter is proportional to the log of domestic sales, where both variables are deviated from industry means.

Appendix B: Data Appendix

The results in this paper are primarily based on two surveys conducted by the Instituto Nacional de Estadísticas, Geografía, e Información (INEGI), the Mexican government statistical agency: the Encuesta Industrial Anual (EIA) [Annual Industrial Survey] and the Encuesta Nacional de Empleo, Salarios, Tecnología y Capacitación (ENESTyC) [National Survey of Employment, Wages, Technology and Training]. This appendix describes the sampling design of each survey and the process by which I cleaned and linked the datasets. A note on classifications: the industrial classification on which the surveys are mainly based is the Clasificación Mexicana de Actividades y Productos 1994 (CMAP 94) (Mexican Classification of Activities and Products). It is organized in 6-digit industries called *clases* (classes), 4 digit industries called *ramas* (branches), and 2 digit industries called *divisiones* (divisions). In the manufacturing sector, there are 309 *clases*, 50 *ramas*, and 9 *divisiones*. Earlier versions of the CMAP were in use prior to 1994 (for the EIA 1984-1994 panel and the ENESTyC 1992, described below), and a new version, the CMAP 99, was created for the 1999 Industrial Census, in order to harmonize the Mexican system with the North American Industrial Classification System (NAICS), now in use in all three NAFTA countries (and used for the 1999 ENESTyC).

B.1: Encuesta Industrial Anual (EIA) and Encuesta Industrial Mensual (EIM)

The Encuesta Industrial Anual (EIA) is carried out yearly, in the spring, with data referring to the previous calendar year. It contains information on employment, hours, wages, expenditures (including whether inputs are domestic or foreign), sales, other revenues, (including revenue from subcontracting), inventories, and capital assets and investment. A companion survey, the Encuesta Industrial Mensual (EIM) [Monthly Industrial Survey], is carried out monthly using a less extensive survey at the same set of plants. I have linked two separate panel datasets, one covering the 1993-2001 period and an earlier on covering the 1984-1994 period.

B.1.1 Sampling Design

The sample for the 1993-2001 EIA and EIM panels was drawn by the following procedure. In 1993, 205 of the 309 6-digit industries (*clases*) in the CMAP-94 were chosen to be included in the EIA. From a list of the universe of manufacturing plants in Mexico (generated in preparation for the upcoming 1994 industrial census), plants within each *clase* were ranked in decreasing order of the value of production (*valor de produccion*), the value of the output of the plant priced at the "factory" price (*venta de fabrica*). Plants were then selected in decreasing order of value of output until the set of selected plants made up 85% of the total value of output (not including maquiladoras) of the *clase*. This rule was subject to the following qualifications:

1. If a plant employed 100 or more workers, it was added to the sample, regardless of whether the 85% level had been reached.

2. If more than 100 plants were required to cover 85% of output within the *clase*, the number of plants was limited to 100. (There were no cases in which this rule conflicted with (a).)

3. If four or fewer plants made up 85% or more of total output of the *clase*, then to preserve the confidentiality of those plants, all plants in the *clase* were included, superseding rule (b).

These criteria generated a sample of 7,042 plants in 1993. These plants were subsequently followed over time. Specific analysts in the INEGI offices in Aguascalientes, Mexico, are assigned to follow particular plants over time, and to double-check inconsistencies or sudden changes in the plant, in many cases by calling the establishment on the phone. The same is true for the monthly EIM survey the same plants. As a result, the quality of the data in the EIA and EIM is better than in surveys with less regular contact between the INEGI analysts and the survey respondents.

A small number of plants were added to the survey after 1993, but they were not added in a systematic way, and I ignore the new establishments. The questionnaire used in the EIA changed over time from 1993 to 1997. In 1997, a consistent format for the questionnaire was settled upon, and has since remained in effect. Variables collected prior to 1997 that are no longer included in the survey have been discarded from the dataset. An exception is foreign ownership, which was collected in 1994 but not thereafter, which has been preserved and can still be linked to the ongoing panel.

The data on employment and wages by occupational category comes from the EIM, not the EIA. The occupational categories are white-collar workers (*empleados*) and blue-collar workers (*obreros*), which correspond to the more familiar categories of non-production and production workers.

The convention in the EIA and EIM for plants producing under subcontract is to report earnings from subcontracting as income from subcontracting services (*ingresos por servicios de maquila*), not under value of production or sales. For plants subcontracting out, payments to subcontractors are classified under costs, and sales of the subcontracted goods are reported under total value of production and sales. For this reason, total revenues (defined as total sales plus income from subcontracting services provided to other plants minus subcontracting services purchased from other plants) may be a more reliable indicator of the volume of output than reported total sales.

Participants in the governments maquiladora program typically report zero value of production, and hence are excluded by the EIA sampling procedure. It is possible, however, that a few exporting maquiladoras (*maquiladoras de exportacíon*) with non-zero total values of production were mistakenly included in the sample.

An important advantage of the EIA and the EIM is that the analysts tracking each establishment keep track of why some establishments fail to respond or respond in an irregular way. I classified plants that left the sample into two categories. I classified plants that went out of business are classified as deaths. I classified plants that switched to industries not covered by the survey, that switched from manufacturing to wholesale or retail sales, that merged with other establishments, or that failed to provide data are classified as other exiters. I assume that these other exiters are missing "at random", and can be ignored in my estimation.

The design of the 1984-1994 panel sample was similar to that of the 1993-2001 sample. In 1984, under an earlier industrial classification system, 129 *clases* were selected for the panel. Establishments within each *clase* were chosen following the same criteria described for the 1993-2001 panel above. The principal difference was that establishments were included within each industry until 80% of the total value of production in the industry was covered. The original sample consisted of approximately 3,200 plants. The EIA surveys prior to 1992 did not elicit information on exports. Information on exports is available from supplementary surveys funded by the World Bank for the same sample of plants, but only for the years 1986-1990. Data on exports are not available for 1984-1985 or 1991.

In the 1984-1994 panel there is no information on why plants exited. Rather than construct separate balanced and unbalanced panels for the 1984-1994 period, I focus on the plants that can be linked to the 1993-2001 panel (and hence can be followed during the peso crisis.) Many plants can not be linked from the 1984-1994 to 1993-2001 samples because the set of industries surveyed changed. Because the EIA focuses on large plants, however, it is still possible to link a significant number of establishments. After following the cleaning procedure outlined above, there are 706 establishments that appear with complete data over the entire 1984-2001 period.

B.1.2 Cleaning Procedure

I cleaned 1993-2001 EIA panel as follows:

1. I removed establishments that report data for more than one establishment or that have their

data reported by another establishment. In some multi-establishment firms, survey respondents are unwilling or unable to report information separately for each establishment. In these cases, respondents report joint data for the establishments on a single survey. Only in 1998 did INEGI begin keeping track systematically of the reporting patterns. Establishments which have their information reported elsewhere are easy to identify; their records appears in zeros. Establishments that report information for more other establishments in the same firm are more difficult to identify. My approach was to discard any establishment that had its information reported elsewhere in any year, and to discard any establishment that reported information from another establishment in 1998 or later. Although reporting patterns have not changed much over time, it is possible that this procedure fails to catch all cases of consolidation of information. This is one possible reason for the correlated measurement error between sales/revenues and total employment discussed in section 4.3 above.

2. I removed establishments owned in whole or in part by government entities.

3. I removed establishments that appear in any year to be maquiladoras, in the sense that exports make up 100% of their sales. Since 1989 maquilas have been allowed to sell some of their output on the domestic market. It is thus possible that some maquilas with non-zero domestic sales are still included in the sample after cleaning.

4. I removed establishments that were missing data in any year.

5. I removed establishments that in any year had a reporting irregularity, i.e. a strike or suspension of operations during part of the year.

6. I removed establishments that had a change of more than 500% from one year to the next in the following variables: employment and wage of both categories of workers; total revenues; or hours worked per worker.

Table B1 summarizes the cleaning procedure. In the end I am left with a balanced panel of 3,003 plants that have complete data over the entire 1993-2001 period, and an unbalanced panel that includes 602 additional plants that went out of business during the period. I discard the establishments that exited at random.

Following a recommendation to reduce measurement error of Angrist and Krueger (1999), I censored the key variables at the tails, replacing values in the lower or upper 1% tails with values at the 1st and 99th percentiles. I carried out this procedure for revenues, domestic sales, employment, wages for each category of worker, the capital-labor ratio, the wage ratio, and the employment ratio.

B.1.3: Variable Definitions

The variables total employment and employment of each type of worker (white-collar and blue-collar) are drawn directly from the EIA datasets. The remaining variables were constructed as follows.

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

Blue-collar real 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 wage/blue-collar wage.

Employment ratio = white-collar employment/blue-collar employment.

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

Revenues = total sales + income from subcontracting for other plants - expenditures on subcontracting by other plants, in thousands of 1994 pesos, deflated by producer price index.

Export percentage of sales = 100*export sales/total sales.

Capital-labor ratio = book value of capital assets at beginning of year/total employment, in thousands of 1994 pesos, deflated by producer price index.

Foreign ownership indicator = 1 if plant has any foreign ownership, 0 if not. (Based on 1994 data for 1993-2001 panel; year not specified for 1984-1994 panel.)

B.2: Encuesta Nacional de Empleo, Salarios, Tecnología y Capacitación (ENESTyC)

The ENESTyC is a special supplementary survey that includes detailed quantitative and qualitative questions regarding training, turnover, technology use and a variety of workplace practices. The was carried out in 1992, 1995, 1999, with many of the questions referring to the previous calendar year.[58] The survey covered 5,039 establishments in 1992, 5,240 in 1995 and 6,876 in 1999. Unlike the EIA, the ENESTyC in each year is based on a representative sample of plants. Also unlike the EIA, it includes maquiladoras. The sampling design in each year was stratified by total employment, with plants with 100 or more employees being sampled with certainty, and a sample of plants with fewer than 100 employees drawn at random. In 1995, two samples were drawn. One was a probabilistic sample similar to the samples in 1992 and 1999, with the difference that maquiladoras were excluded. The second sample was a follow-up sample for the 1992 sample. All respondents to the 1992 survey that could be located were included. A few maquiladoras were included in the 1995 survey through this follow-up sample.

Sampling weights are reported based on the probability of inclusion in the sample. The surveys were designed as separate cross-sections, not as a panel, but because of the fact that large establishments are sampled with certainty, it is possible to link a fair number of plants over time. The different waves of the survey did not employ the same identification codes, and many of the establishments had to be linked across years by establishment name and street address. In the end, it was possible to link 2,248 plants between the 1992 and 1995 surveys, 1,056 plants between the 1992 and 1999 surveys, 1,552 plants between the 1995 and 1999 surveys., and 891 plants across all three waves. Although many questions changed in the questionnaires between waves, several key variables are comparable across waves.

My considered judgment after working with these data is that for the variables the EIA and ENESTyC have in common, the EIA is the more reliable source. For this reason, in the estimation reported in Table 10 in the text, I limited the ENESTyC sample to plants that could be linked between the two datasets. Of the 1,552 plants that appear in both the 1995 and 1999 ENESTyCs, 886 also appear in the 1993-2001 EIA Balanced Panel. Because the ENESTyC mainly reports data from the previous year, I linked the ENESTyC observations to the EIA data for 1994 and 1998. I call the resulting dataset the 1994-1998 EIA-ENESTyC panel. Because focusing only on the plants that appear in both these waves and the 1984-2001EIA panel would leave us with a very small sample (approximately 100 plants), I link the 1992 ENESTyC to the 1993 EIA. Of the 1,056 plants that appear in both the 1992 and 1999 surveys, 282 can be linked to the EIA 1993-2001 Balanced Panel. I call the resulting dataset the 1991-1998 EIA-ENESTyC Panel. I also re-did the analysis using the ENESTyC panels not linking to the EIA, and results were qualitatively similar.

Variable definitions:

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

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.

[58] The survey was also carried out in 2001, but I do not yet have access to the 2001 data.

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

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 = same as blue-collar average schooling, but for white-collar workers.

Manual equipment % = Percentage of total value of machines and equipment in plant composed of manual machines and equipment.

The ENESTyC asked a number of different questions about employee training. It appears from the pattern of responses, however, that respondents misunderstood many of the specific questions, or used different rules of thumb to guide their answers. The most reliable measure of training appears to be simply whether or not a formal training program exists at the plant, rather than how many workers received training. The ENESTyC reports wage and employment data by 4 occupational categories: unskilled blue-collar workers, skilled blue-collar workers, administrative and technical workers, and managers. The definitions of the categories in the official documentation, however, are imprecise, especially on the distinction between unskilled and skilled blue-collar workers, and this seems to have led to a significant amount of noise in the data, with some plants reporting all their blue-collar workers under the skilled blue-collar category, and others under the unskilled blue-collar category. The coarser distinction between blue-collar and white-collar workers is more familiar to survey respondents and appears to be more reliable.

I censored outliers for the turnover rate, the accident rate, and the absentee rate following the same procedure as for the EIA, assigning all variables in the bottom or top 1% to the values at the 1st and 99th percentiles respectively.

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Table 1
Summary Statistics by Export Status, EIA Balanced Panel, 1993, 1997, 2001

		1993			1997			2001		
		Non-exporters	Exporters	All	Non-exporters	Exporters	All	Non-exporters	Exporters	All
Employment	Mean	194.0	345.1	240.8	187.7	339.3	259.4	201.2	376.1	280.2
	S.D.	246.2	377.7	301.4	252.9	385.8	331.4	278.9	430.0	365.6
Revenues	Mean	45.4	96.3	61.2	46.8	114.4	78.8	49.7	122.0	82.4
	S.D.	91.9	155.7	117.8	92.2	207.5	161.2	110.3	229.1	177.9
Domestic sales	Mean	43.8	74.3	53.3	45.3	81.5	62.4	48.1	86.5	65.5
	S.D.	83.7	118.2	96.7	83.2	138.9	114.4	99.2	156.5	129.7
K/L ratio	Mean	41.5	58.0	46.6	32.0	46.8	39.0	43.7	67.5	54.5
	S.D.	65.4	82.6	71.5	47.8	68.3	58.8	68.9	91.0	80.5
White-collar hourly wage	Mean	21.0	28.8	23.4	14.4	21.2	17.6	16.7	26.1	20.9
	S.D.	12.8	15.4	14.1	10.4	13.7	12.5	12.4	16.7	15.3
Blue-collar hourly wage	Mean	8.10	9.71	8.60	5.78	7.12	6.41	7.03	9.18	8.00
	S.D.	3.79	4.74	4.18	3.01	3.92	3.54	3.90	5.34	4.73
White-collar/blue-collar wage	Mean	2.80	3.18	2.92	2.59	3.11	2.83	2.46	3.02	2.72
	S.D.	1.78	1.61	1.74	1.61	1.67	1.66	1.59	1.74	1.68
white-collar % of employment	Mean	30.9	33.6	31.7	31.5	30.2	30.9	31.2	31.7	31.5
	S.D.	17.2	16.5	17.0	19.0	17.0	18.1	18.7	18.2	18.5
export % of sales	Mean		17.43			20.65			21.64	
	S.D.		23.01			23.71			24.88	
% with foreign ownership	Mean	8.7	31.9	16.3						
	S.D.	28.2	46.6	36.9						
N		2074	929	3003	1583	1420	3003	1647	1356	3003
% of sample		69.1	30.9		52.7	47.3		54.8	45.2	

Notes: Establishment classified as an exporter if it had any export sales for the year. Establishment classified as having foreign ownership if foreigners owned any positive share. Data on foreign ownership from 1994; all other data from year indicated. Revenues and domestic sales measured in millions of 1994 pesos, K/L ratio in thousands of 1994 pesos, wages in 1994 pesos per hour. Average 1994 exchange rate: 3.38 pesos/US$1. Further variable definitions in data appendix.

Table 2
Differential Effects of Peso Crisis by Initial Productivity, EIA 1993-2001 Balanced Panel
Proxy for productivity: log domestic sales

	Period							
	1993-1997				1997-2001			
	OLS		IV		OLS		IV	
	Coeff. Est.	R-sqr.	Coeff. Est.	R-sqr.	Coeff. Est.	R-sqr.	Coeff. Est.	R-sqr.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Independent variables: initial-year log(domestic sales), industry and state effects								
Dependent variable:								
Δ (export % of sales)	1.960***	0.171	1.295***	0.156	0.869***	0.142	0.470***	0.124
	[0.238]		[0.251]		[0.188]		[0.181]	
Δ log(white-collar wage)	0.060***	0.15	0.050***	0.147	0.007	0.093	0.000	0.085
	[0.007]		[0.007]		[0.006]		[0.006]	
Δ log(blue-collar wage)	0.029***	0.137	0.029***	0.123	0.001	0.104	0.001	0.106
	[0.006]		[0.005]		[0.005]		[0.004]	
Δ log(wage ratio)	0.031***	0.087	0.021***	0.091	0.006	0.085	-0.001	0.085
	[0.008]		[0.008]		[0.007]		[0.007]	
Δ log(K/L ratio)	0.064***	0.115	0.043***	0.105	0.020*	0.108	0.01	0.100
	[0.012]		[0.011]		[0.011]		[0.009]	
Δ log(employment ratio)	0.005	0.115	0.008	0.099	0.001	0.112	-0.002	0.107
	[0.009]		[0.008]		[0.007]		[0.006]	

Notes: OLS estimates from unweighted regressions of variable in left-hand column on log domestic sales in initial year, industry and state effects. IV estimates from similar regressions treating initial years as 1994 and 1998 and instrumenting log domestic sales with previous year. Coefficient estimates for log domestic sales in odd-numbered columns, R-squared for each regression in even-numbered columns. Robust standard errors in brackets. N=3003 for all regressions. Further variable definitions in data appendix. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

Table 3
Bivariate Correlations, 1993, EIA Balanced Panel
All variables deviated from industry means

	log (employment)	log (revenues)	log (K/L ratio)	white-collar hourly wage	blue-collar hourly wage	employment ratio	foreign ownership	export % of sales	1st princ. comp. proxy	log (domestic sales)
log (employment)	1									
log (revenues)	0.8297**	1								
log (K/L ratio)	0.1704**	0.3880**	1							
white-collar hourly wage	0.2980**	0.4296**	0.2178**	1						
blue-collar hourly wage	0.2162**	0.3753**	0.2071**	0.4560**	1					
employment ratio	-0.0083	0.0822**	0.1045**	-0.0012	0.1513**	1				
foreign ownership indicator	0.1920**	0.2748**	0.1948**	0.2381**	0.2092**	0.1806**	1			
export % of sales	0.1768**	0.1682**	0.1246**	0.0901**	0.0161	-0.0406*	0.1237**	1		
1st princ. comp. proxy	0.7464**	0.8821**	0.5127**	0.6546**	0.5979**	0.1662**	0.4830**	0.2711**	1	
log (domestic sales)	0.7810**	0.9526**	0.3512**	0.4038**	0.3735**	0.0929**	0.2416**	-0.0850**	0.8093**	1

Notes: Table reports bivariate correlation coefficients for indicated variables, deviated from industry means, in 1993. The productivity proxy is the first principal component of the variables above it (deviated from industry means). Variable definitions in data appendix. ** indicates significance at 1% level, * at 5% level.

Table 4
Differential Effects of Peso Crisis by Initial Productivity, EIA 1993-2001 Balanced Panel
Proxy for productivity: first principal component of plant characteristics

	Period			
	1993-1997		1997-2001	
	Coeff. Estimate	R-sqr.	Coeff. Estimate	R-sqr.
	(1)	(2)	(3)	(4)
Independent variables: productivity proxy (first principal component of initial-year characteristics), industry and state effects				
Dependent variable:				
Δ (export % of sales)	1.698***	0.162	0.529**	0.136
	[0.234]		[0.223]	
Δ log(white-collar wage)	0.067***	0.151	0.002	0.092
	[0.008]		[0.007]	
Δ log(blue-collar wage)	0.034***	0.138	0.000	0.104
	[0.006]		[0.005]	
Δ log(wage ratio)	0.036***	0.087	0.001	0.085
	[0.009]		[0.008]	
Δ log(K/L ratio)	0.067***	0.114	0.035***	0.11
	[0.012]		[0.012]	
Δ log(employment ratio)	-0.003	0.115	-0.005	0.112
	[0.009]		[0.008]	

Notes: All estimates by unweighted OLS. Coefficients estimates for productivity proxy in odd-numbered columns, R-squared for each regression in even-numbered columns. Productivity proxy calculated by taking first principal component of the following variables in the initial year, deviated from industry means: export % of sales, log(employment), log(K/L ratio), log(revenues), log(white-collar employment/blue-collar employment), log(white-collar wage), log(blue-collar wage), foreign ownership indicator. Level of a given variable is omitted from principal component calculation when its change is the dependent variable. In calculating the change in the wage ratio, the overall average wage is used to calculate principal component, rather than white-collar or blue-collar wage. Robust standard errors in brackets. N=3003 for all regressions. Further variable definitions in data appendix. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

Table 5
Differential Effects of Peso Crisis on Employment and Revenues by Initial Productivity, EIA 1993-2001 Balanced Panel

	Period							
	1993-1997				1997-2001			
	OLS		IV		OLS		IV	
	Coeff. Est.	R-sqr.	Coeff. Est.	R-sqr.	Coeff. Est.	R-sqr.	Coeff. Est.	R-sqr.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Independent variables: initial-year log(domestic sales), industry and state effects								
Dependent variable:								
Δ log(employment)	0.000	0.125	0.005	0.129	0.018***	0.147	0.015***	0.135
	[0.007]		[0.006]		[0.006]		[0.005]	
Δ log(revenues)	-0.004	0.176	0.014	0.175	0.008	0.206	0.014**	0.187
	[0.010]		[0.009]		[0.007]		[0.007]	
Independent variables: productivity proxy (first principal component of initial-year characteristics), industry and state effects								
Δ log(employment)	0.036***	0.134			0.025***	0.148		
	[0.008]				[0.007]			
Δ log(revenues)	0.055***	0.186			0.029***	0.209		
	[0.009]				[0.008]			

Notes: OLS estimates from unweighted regressions of variable in left-hand column on proxy (indicated above), industry and state effects. IV estimates from similar regressions treating initial years as 1994 and 1998 and instrumenting log domestic sales with previous year. First principal component extracted from the following variables in the initial year, deviated from industry means: export % of sales, log(employment), log(K/L ratio), log(revenues), log(white-collar employment/blue-collar employment), log(white-collar wage), log(blue-collar wage), foreign ownership indicator. Level of a given variable is omitted from principal component calculation when its change is the dependent variable. In calculating the change in the wage ratio, the overall average wage is used to calculate principal component, rather than white-collar or blue-collar wage. Robust standard errors in brackets. N=3003 for all regressions. Further variable definitions in data appendix. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

Table 6
Differential Effects of Peso Crisis by Initial Productivity, EIA 1984-2001 Balanced Panel
Proxy for productivity: log domestic sales

	Period							
	1986-1989		1989-1993		1993-1997		1997-2001	
	Coeff. Est.	R-sqr.	Coeff. Est.	R-sqr.	Coeff. Est.	R-sqr.	Coeff. Est.	R-sqr.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Independent variables: initial-year log(domestic sales), industry and state effects								
Dependent variable:								
Δ (export % of sales)	0.743**	0.277	0.504*	0.269	1.170**	0.271	0.770**	0.257
	[0.298]		[0.285]		[0.459]		[0.377]	
Δ log(white-collar wage)	0.036***	0.25	0.000	0.234	0.062***	0.311	-0.008	0.27
	[0.012]		[0.015]		[0.016]		[0.013]	
Δ log(blue-collar wage)	0.01	0.298	0.004	0.288	0.029***	0.31	0.005	0.242
	[0.010]		[0.012]		[0.011]		[0.012]	
Δ log(wage ratio)	0.027*	0.263	-0.003	0.221	0.033**	0.269	-0.013	0.244
	[0.014]		[0.017]		[0.017]		[0.016]	
Δ log(K/L ratio)	0.059*	0.274	0.071**	0.251	0.061**	0.26	0.006	0.24
	[0.033]		[0.036]		[0.024]		[0.028]	
Δ log(employment ratio)	0.005	0.31	-0.02	0.275	0.049***	0.316	0.009	0.27
	[0.014]		[0.015]		[0.018]		[0.013]	
Δ log(employment)	-0.003	0.343	-0.039***	0.294	0.001	0.324	0.012	0.334
	[0.009]		[0.015]		[0.013]		[0.012]	
Δ log(revenues)	-0.037**	0.333	-0.091***	0.403	0.001	0.311	-0.007	0.406
	[0.015]		[0.019]		[0.022]		[0.017]	

Notes: All estimates by unweighted OLS. Coefficients estimates for log domestic sales in odd-numbered columns, R-squared for each regression in even-numbered columns. Data on exports vs. domestic sales not available prior to 1986, whence the choice of 1986-1989 as periodization in columns (1)-(2). Robust standard errors in brackets. N=706 for all regressions. Further variable definitions in data appendix. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

Table 7
Differential Effects of Peso Crisis by Initial Productivity, EIA 1984-2001 Balanced Panel
Proxy for productivity: first principal component of plant characteristics

	Period							
	1986-1989		1989-1993		1993-1997		1997-2001	
	Coeff. Est.	R-sqr.	Coeff. Est.	R-sqr.	Coeff. Est.	R-sqr.	Coeff. Est.	R-sqr.
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Independent variables: productivity proxy (first principal component of initial-year characteristics), industry and state effects								
Dependent variable:								
Δ (export % of sales)	0.463	0.273	-0.279	0.267	1.183**	0.271	0.464	0.252
	[0.328]		[0.353]		[0.471]		[0.426]	
Δ log(white-collar wage)	0.034***	0.249	-0.005	0.235	0.066***	0.314	-0.001	0.269
	[0.013]		[0.014]		[0.016]		[0.014]	
Δ log(blue-collar wage)	0.011	0.298	0.001	0.288	0.029**	0.31	0.004	0.242
	[0.010]		[0.011]		[0.012]		[0.012]	
Δ log(wage ratio)	0.029**	0.263	-0.007	0.221	0.045***	0.274	-0.007	0.243
	[0.013]		[0.016]		[0.017]		[0.017]	
Δ log(K/L ratio)	0.066**	0.275	0.065*	0.25	0.064***	0.261	-0.007	0.24
	[0.032]		[0.035]		[0.023]		[0.030]	
Δ log(employment ratio)	-0.002	0.31	-0.034**	0.279	0.025	0.307	-0.006	0.27
	[0.014]		[0.015]		[0.018]		[0.015]	
Δ log(employment)	0.015*	0.347	-0.029**	0.287	0.01	0.325	0.026**	0.34
	[0.008]		[0.014]		[0.012]		[0.012]	
Δ log(revenues)	-0.014	0.324	-0.039**	0.375	0.051***	0.321	0.017	0.407
	[0.014]		[0.019]		[0.018]		[0.018]	

Notes: All estimates by unweighted OLS. Coefficients estimates for productivity proxy in odd-numbered columns, R-squared for each regression in even-numbered columns. Productivity proxy calculated by taking first principal component of the following variables in the initial year, deviated from industry means: export % of sales, log(employment), log(K/L ratio), log(revenues), log(white-collar employment/blue-collar employment), log(white-collar wage), log(blue-collar wage), foreign ownership indicator. Level of a given variable is omitted from principal component calculation when its change is the dependent variable. In calculating the change in the wage ratio, the overall average wage is used to calculate principal component, rather than white-collar or blue-collar wage. Robust standard errors in brackets. N=706 for all regressions. Further variable definitions in data appendix. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

Table 8
Differential Effects of Peso Crisis by Initial Productivity, Selection-Correction Model, EIA 1993-2001 Unbalanced Panel
Proxy for productivity: log domestic sales

		Period			
		1993-1997		1997-2001	
		Coeff. Est.	chi-sqr.	Coeff. Est.	chi-sqr.
		(1)	(2)	(3)	(4)
Independent variables: initial-year log(domestic sales), industry and state effects					
Dependent variable:					
First stage:	Remains in sample	0.239***	241.2	0.354***	334.2
		[0.032]		[0.032]	
Second stage:	Δ (export % of sales)	1.917***	874.8	0.473*	767.2
		[0.305]		[0.250]	
	Δ log(white-collar wage)	0.059***	694.1	0.012	607.8
		[0.011]		[0.009]	
	Δ log(blue-collar wage)	0.027***	699.7	-0.003	644.3
		[0.008]		[0.007]	
	Δ log(wage ratio)	0.032***	500.1	0.016	577.6
		[0.012]		[0.010]	
	Δ log(K/L ratio)	0.063***	622.1	-0.002	655.5
		[0.017]		[0.015]	
	Δ log(employment ratio)	0.015	577.8	0.013	666.1
		[0.012]		[0.010]	
	Δ log(employment)	0.004	684.2	0.013*	786.6
		[0.009]		[0.008]	
	Δ log(revenues)	0.011	848.8	0.014	1067.0
		[0.013]		[0.010]	
	N (remain in sample)	3297		3003	
	N (exit)	308		294	
	N (total)	3605		3297	

Notes: Estimates follow the two-step selection-correction procedure of Heckman (1976), with a first-stage probit and second-stage OLS with an inverse Mills ratio term. Estimates are unweighted. Coefficients estimates for log(domestic sales) in odd-numbered columns, chi-squared statistic for test that all coefficients (except the intercept) are zero in even-numbered columns. Standard errors in brackets. Monetary values in 1994 pesos. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

Table 9
Summary Statistics by Maquiladora and Export Status, ENESTyC, 1992, 1995, 1999

		1992				1995				1999			
			Non-maquilas				Non-maquilas				Non-maquilas		
		Maquilas	Non-exp	Exporters	All	Maquilas	Non-exp	Exporters	All	Maquilas	Non-exp	Exporters	All
ISO 9000 certification	Mean					0.2	0.0	0.1	0.0	0.5	0.1	0.3	0.2
	S.D.					0.4	0.2	0.3	0.2	0.5	0.2	0.4	0.4
White-collar avg. schooling	Mean	12.0	11.7	12.6	12.0					12.6	11.3	12.9	12.1
	S.D.	1.8	2.6	1.6	2.3					2.4	3.0	2.0	2.7
Blue-collar avg. schooling	Mean	6.7	6.9	7.4	7.0					7.2	7.6	8.2	7.7
	S.D.	1.2	1.6	1.5	1.6					1.4	1.9	1.6	1.7
Absentee rate	Mean	3.1	3.6	2.7	3.2	1.1	1.4	1.2	1.3	1.3	1.2	1.0	1.1
	S.D.	2.4	2.9	2.5	2.7	0.9	1.0	1.0	1.0	1.1	1.0	0.9	1.0
Accident rate	Mean					4.5	2.8	3.4	2.9	2.2	2.1	2.8	2.3
	S.D.					3.4	4.5	4.1	4.4	2.6	3.3	3.2	3.2
Formal training indicator	Mean	0.65	0.29	0.66	0.39	0.49	0.26	0.68	0.34	0.82	0.32	0.73	0.52
	S.D.	0.5	0.5	0.5	0.5	0.5	0.4	0.5	0.5	0.4	0.5	0.4	0.5
Turnover rate	Mean	130.2	72.9	82.4	78.3	121.4	48.7	60.0	51.5	111.4	52.6	73.3	68.0
	S.D.	112.6	117.0	92.9	113.4	64.8	66.5	56.7	65.5	67.0	63.7	59.6	66.7
% manual equipment	Mean	26.8	25.2	21.5	24.5					26.4	38.7	23.3	32.3
	S.D.	33.9	35.3	29.7	34.2					30.3	38.1	30.4	35.6
Employment	Mean	439.7	137.4	415.1	214.8	261.6	81.4	278.4	117.7	505.2	116.4	364.8	252.1
	S.D.	237.4	207.9	249.4	253.0	218.1	141.9	202.2	172.8	224.4	191.7	256.3	267.9
White-collar % of emp.	Mean	18.7	32.0	31.8	31.0	15.9	34.7	31.5	34.0	17.7	36.5	28.3	31.0
	S.D.	10.6	28.8	17.7	26.3	12.6	32.8	20.5	30.9	15.9	31.5	19.9	27.4
White-collar hourly wage	Mean	3.02	2.98	4.07	3.32	4.62	2.82	4.21	3.30	3.89	2.53	4.03	3.35
	S.D.	1.3	1.7	1.8	1.8	2.6	2.0	2.2	2.2	2.5	1.9	2.5	2.4
Blue-collar hourly wage	Mean	1.03	0.99	1.28	1.05	1.09	1.04	1.32	1.10	1.15	0.91	1.16	1.03
	S.D.	0.4	0.4	0.4	0.4	0.5	0.5	0.5	0.5	0.4	0.4	0.5	0.5
N		394	3049	1213	4656	78	3591	1560	5229	589	4115	2168	6872

Notes: Table reports means and standard deviations (not standard errors) weighted by ENESTyC sampling weights times employment. Variable definitions in data appendix.

Table 10
Differential Effects on Variables from Auxiliary Dataset, 1994-1998 and 1991-1998 EIA-ENESTyC Panels
Proxy for productivity: log domestic sales

	1994-1998		1991-1998	
	(1)	(2)	(3)	(4)
Independent variables: initial-year log(domestic sales), industry and state effects				
Dependent variable:				
Δ indicator for ISO 9000 certification	0.079***	R2=.35		
	[0.018]	N=767		
Δ indicator for formal training program	-0.009	R2=.21		
	[0.022]	N=885		
Δ accident rate	0.05	R2=.19		
	[0.178]	N=876		
Δ absentee rate	-0.038	R2=.32		
	[0.080]	N=468		
Δ turnover rate	-2.563	R2=.26		
	[2.829]	N=728		
Δ avg. schooling, white-collar			0.218	R2=.59
			[0.225]	N=281
Δ avg. schooling, blue-collar			0.390**	R2=.53
			[0.177]	N=282

Notes: Estimates from unweighted regressions of variable in left-hand column on log domestic sales in base year, industry and state effects. For changes over 1994-1998 period, base year is 1994. For changes over 1991-1998 period, base year is 1993. Coefficients estimates for log(domestic sales) in columns (1), (3); coefficients on industry, state dummies omitted. Robust standard errors in brackets. Further variable definitions in data appendix. *** indicates significance at 1% level, ** at 5% level, * at 10% level.

Fig. 1a: Shift Toward Less Skill-Intensive Industries

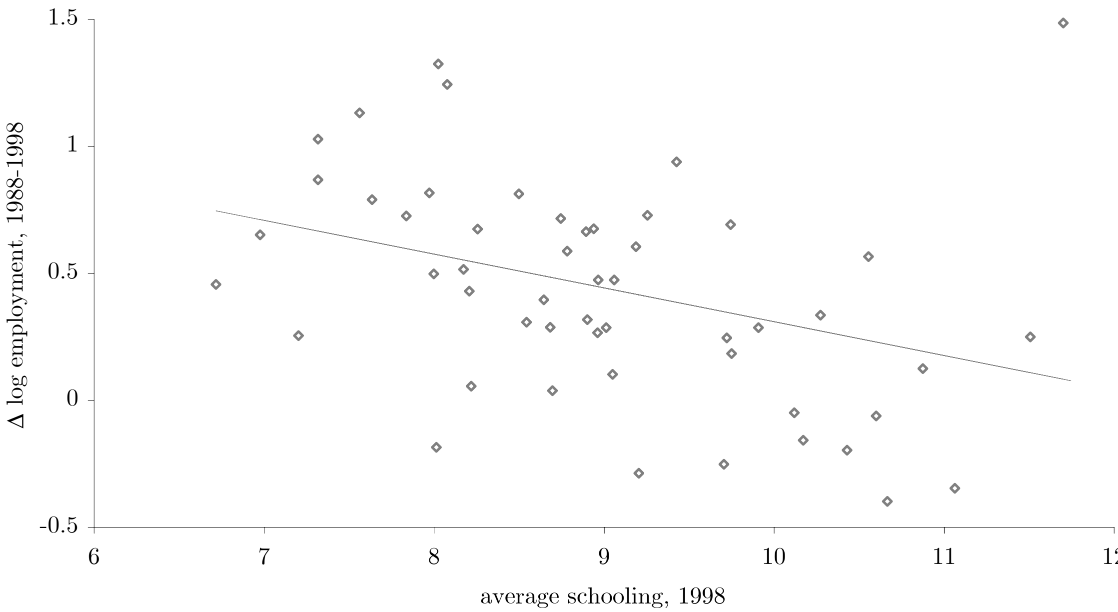
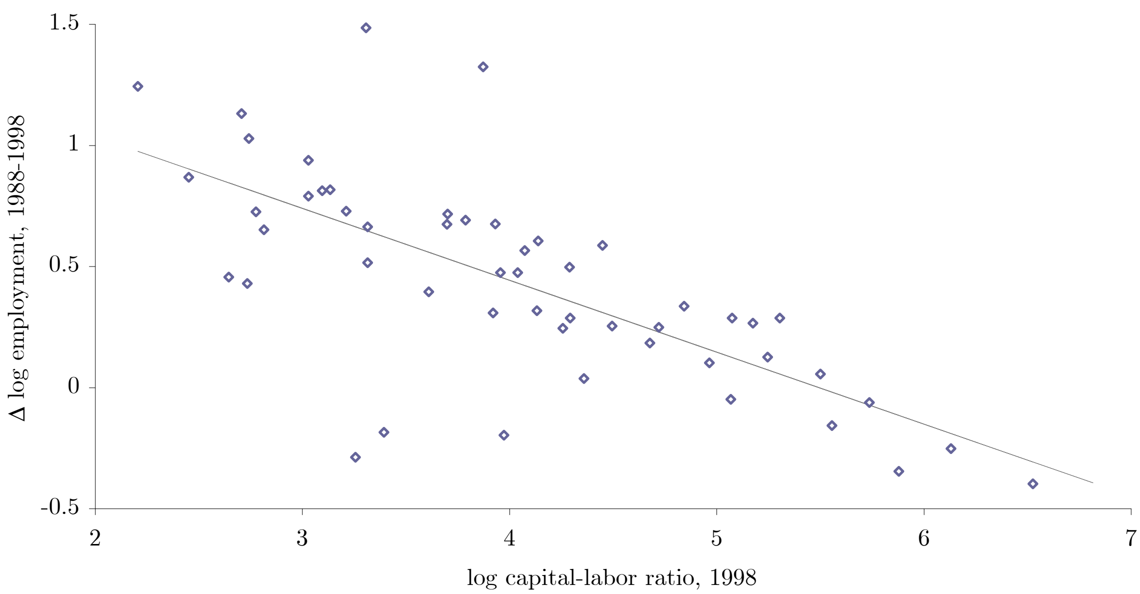


Fig. 1b: Shift Toward Less Capital-Intensive Industries



Notes: Each point represents one 4-digit industry (*rama*); see data appendix for description of industry categories. Data on employment in 1988 and 1998 and capital in 1998 from the Censos Industriales (Industrial Censuses). Data on schooling is from Encuesta Nacional de Empleo Urbano (ENEU), a household survey similar to the CPS for full-time employees (male and female), ages 12-64, in 16 cities in original ENEU sample.

Fig. 2a: Nominal and Real Exchange Rates

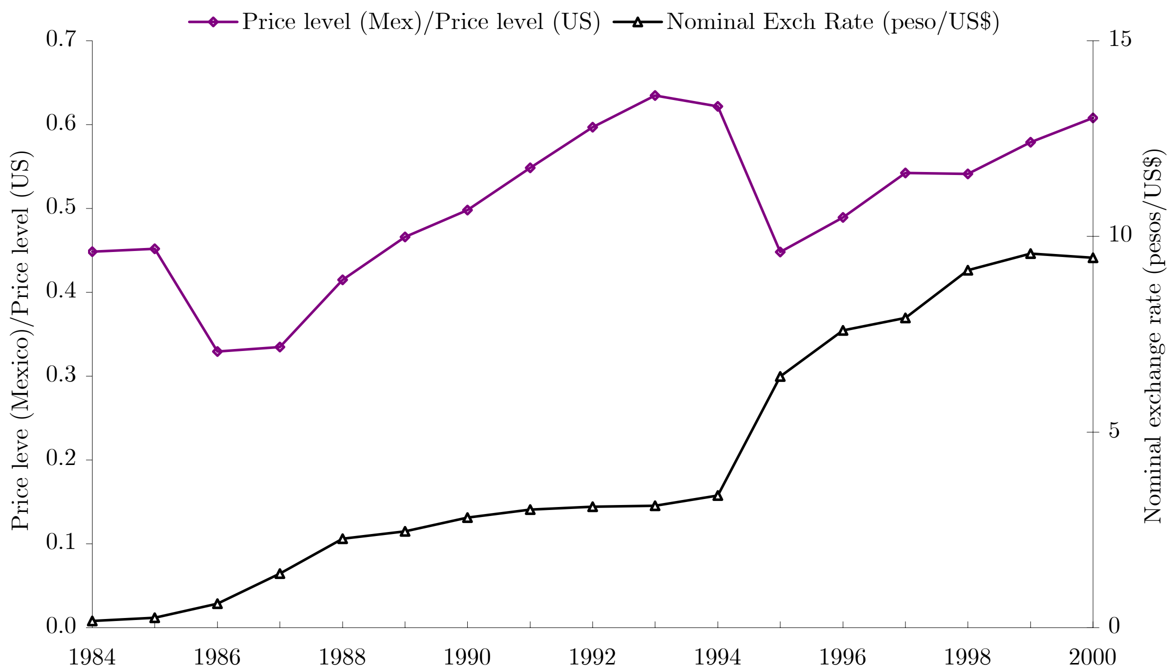
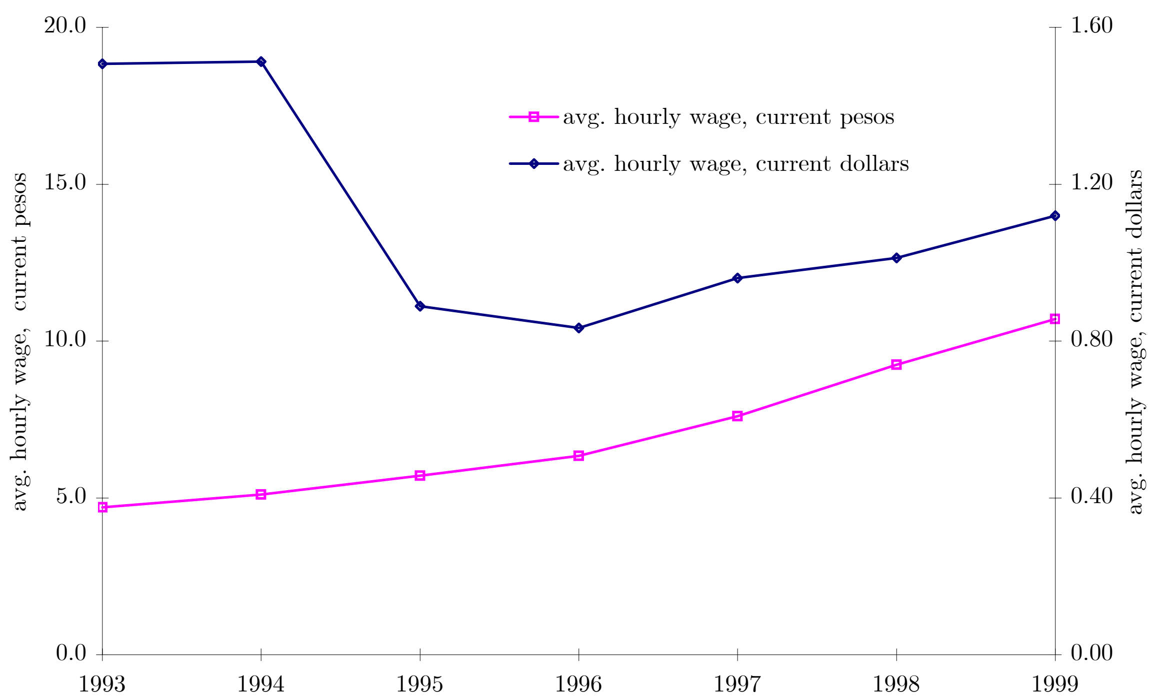


Figure 2b: Wage Changes, 1993-1999



Notes: Data on price levels and exchange rate from Penn World Table 6.1. Wage data for full-time male workers with 9 years schooling from Encuesta Nacional de Empleo Urbano (ENEU), converted to dollars at current exchange rates.

Figure 3a: VW production in Mexico, by destination, 1988-2002

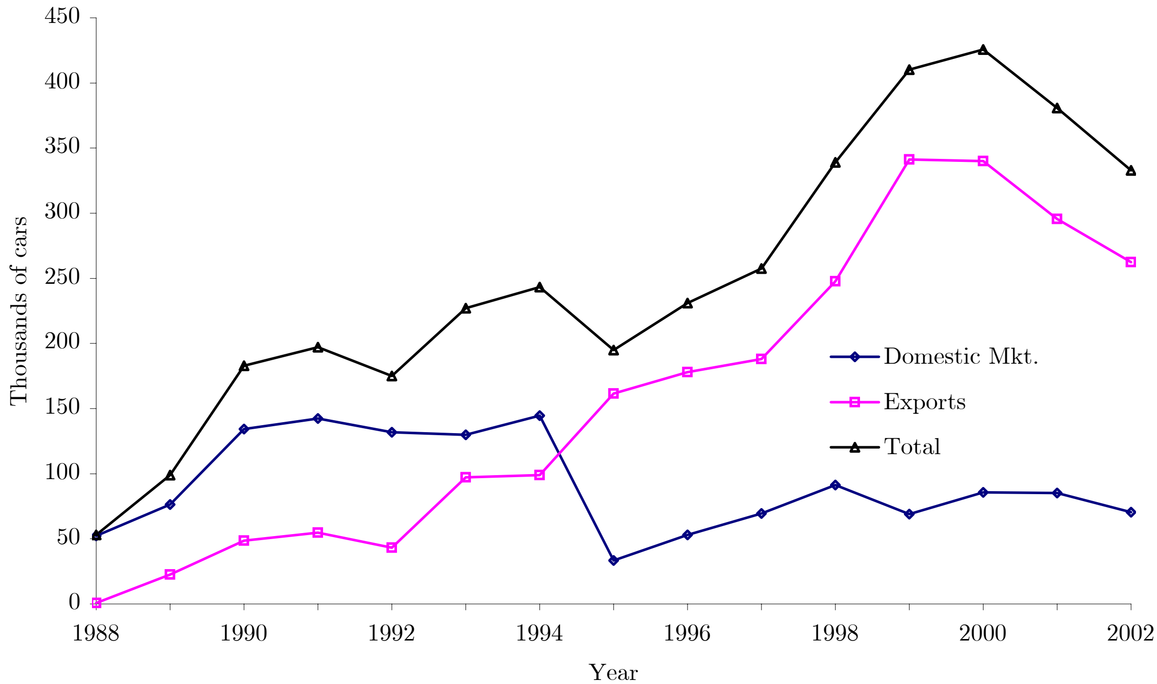
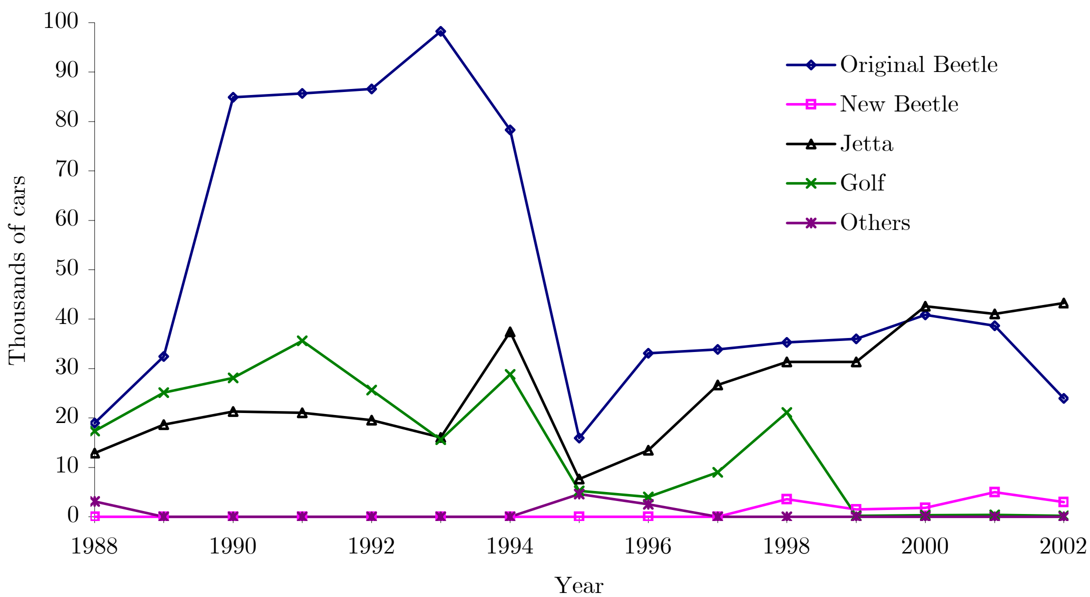


Figure 3b: VW production in Mexico for domestic market, by model, 1988-2002



Source: Bulletins of the Asociacion Mexicana de la Industria Automotriz (Mexican Automobile Industry Association).

Figure 3c: VW production in Mexico for export, by model, 1988-2002

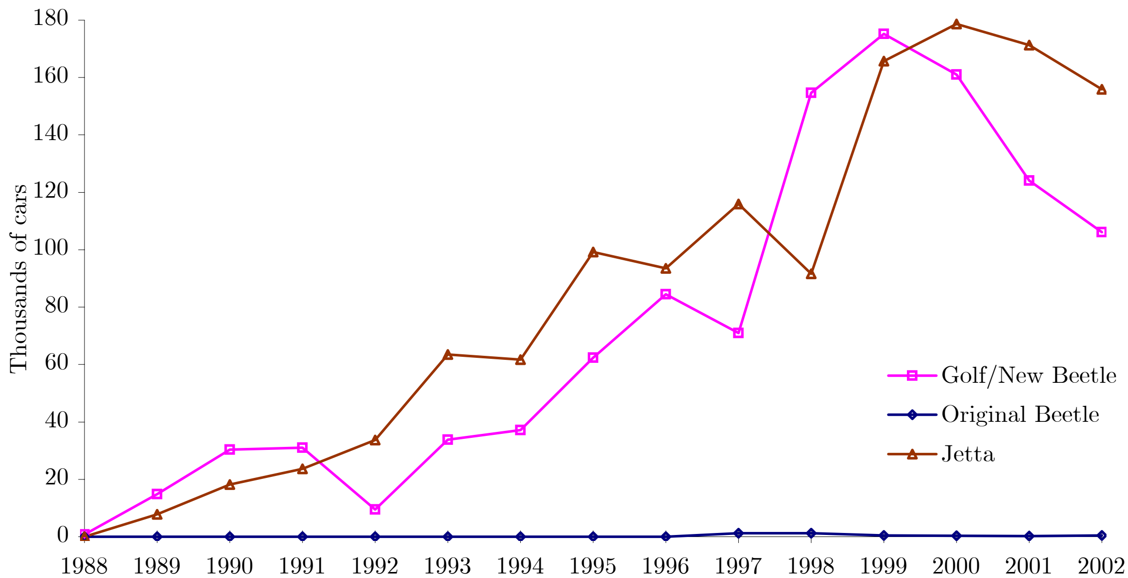
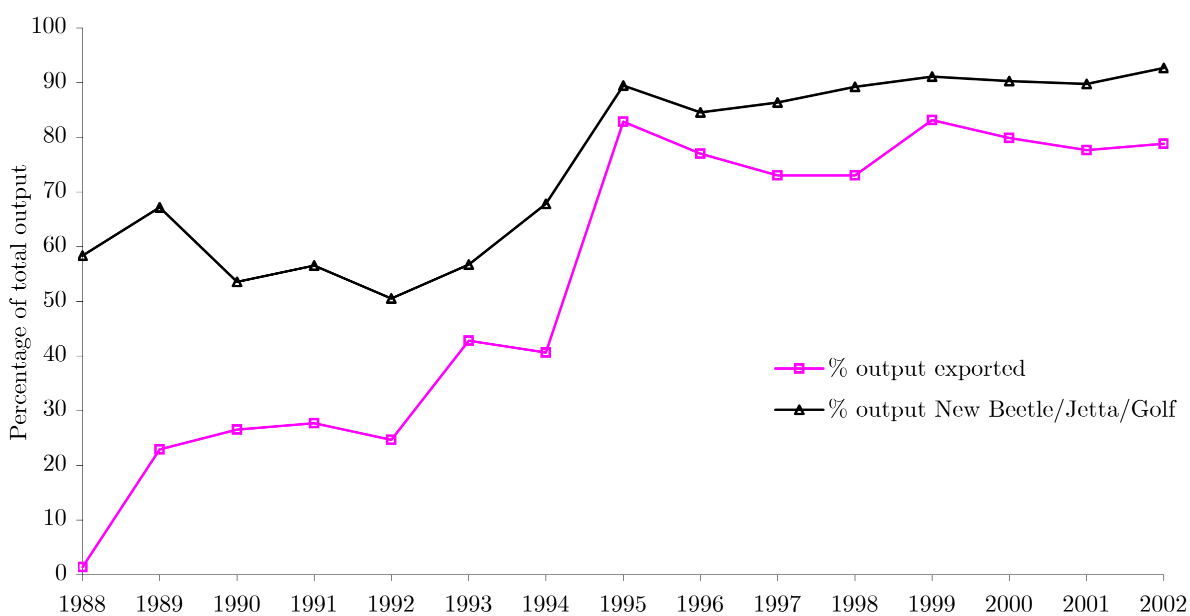


Figure 3d: Exports, High-quality Models as Percentage of Output, 1988-2002



Source: Bulletins of the Asociación Mexicana de la Industria Automotriz (Mexican Automobile Industry Association).

Fig. 4a: Sales, Domestic and Export, 1993-2001

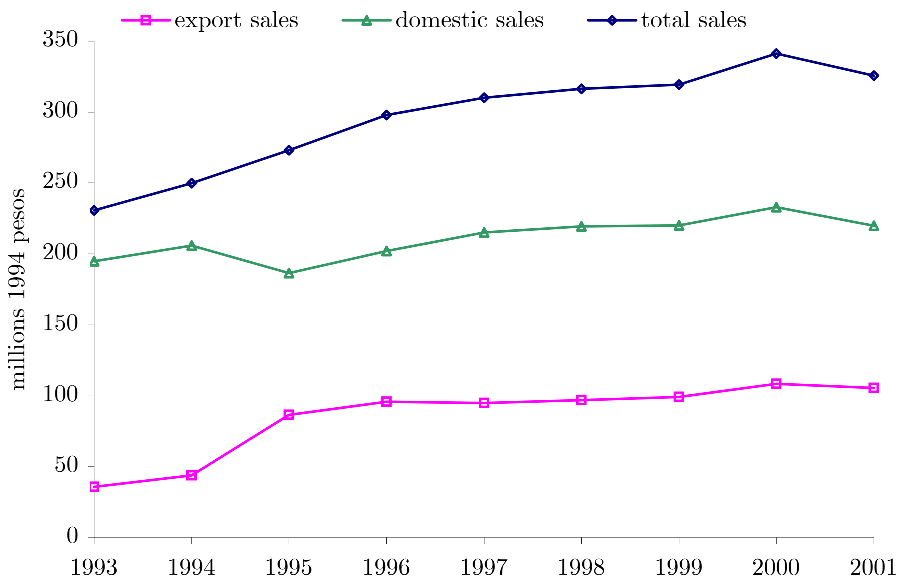


Fig. 4b: Export percentage of total sales

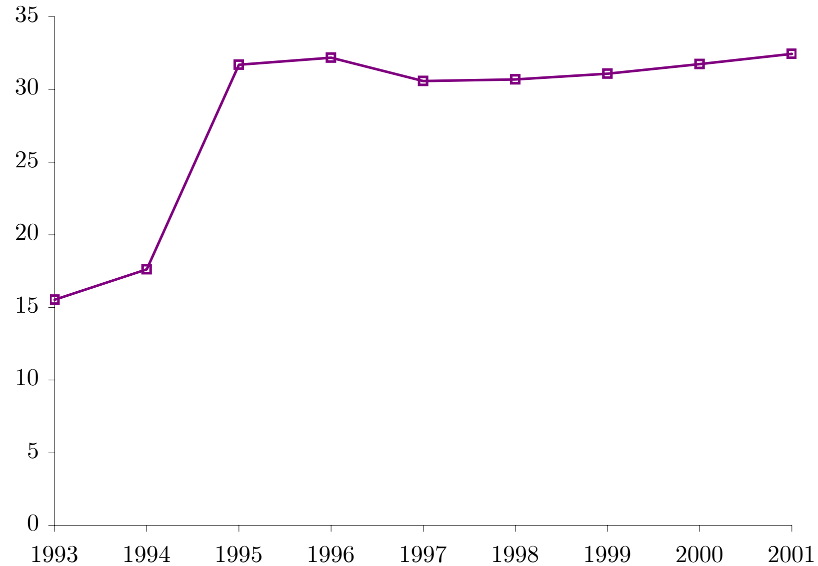
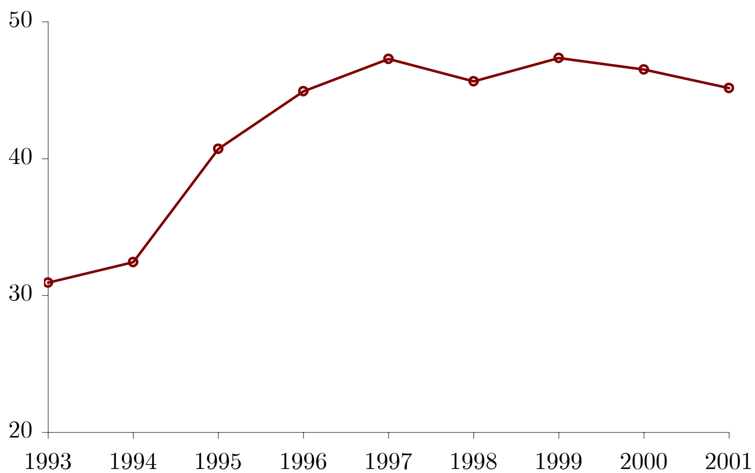


Fig. 4c: Percentage of Plants Exporting



Notes: Data from EIA 1993-2001 Balanced Panel. Export percentage of total sales calculated as total exports for all plants/total sales for all plants. Plants with exports greater than zero classified as exporting.

Fig. 5a: Average Quality as a Function of Know-how Parameter

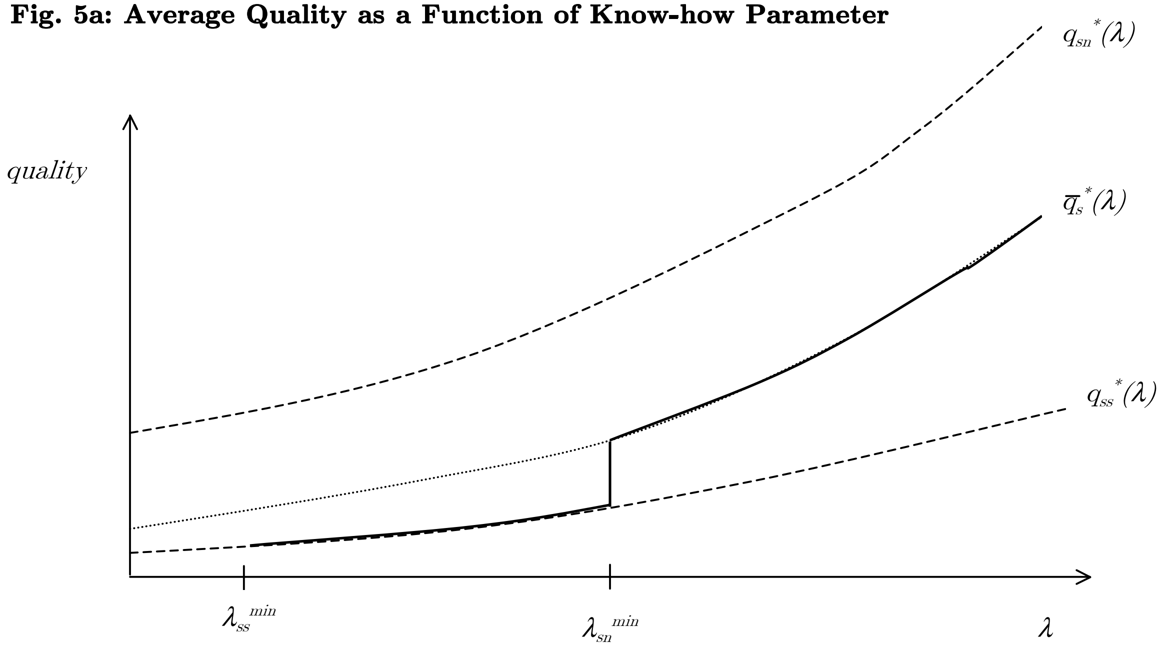


Fig. 5b: Avg. Quality as a Function of Know-how Parameter, pre- and post-shock

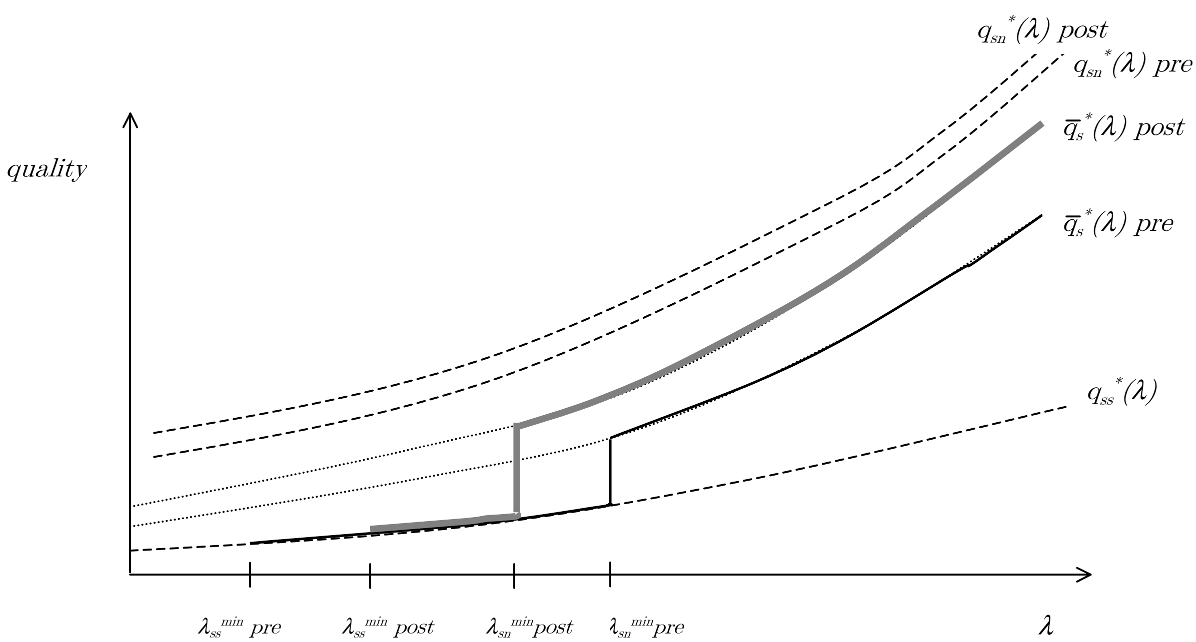


Fig. 5c: Change in Average Quality as a Function of Know-how Parameter

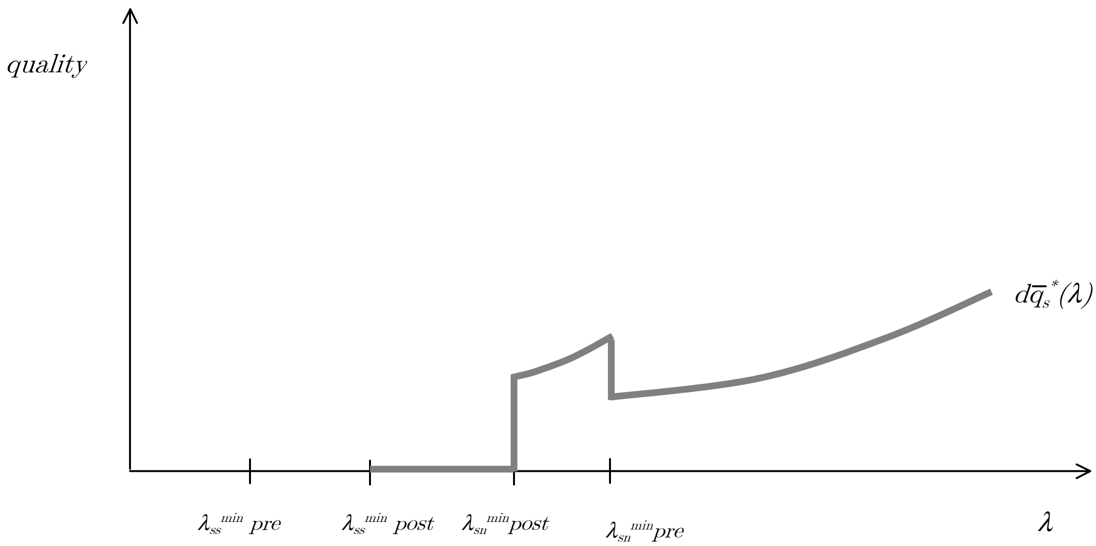


Fig. 6: Differential Effect of Peso Crisis by Degree of Engagement in Export Market

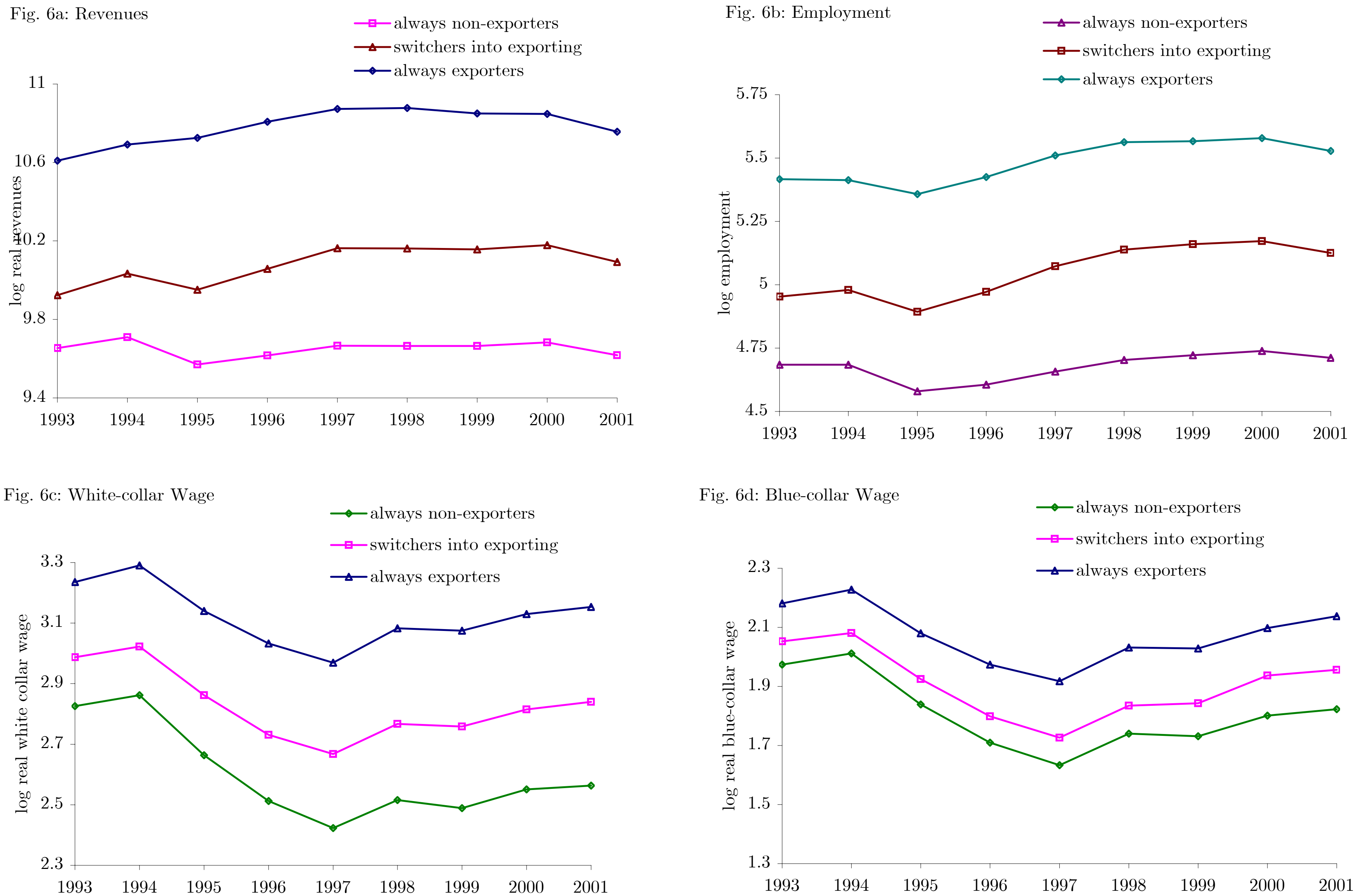


Fig. 7: Non-Parametric Regressions, Levels of Key Variables vs. log Domestic Sales, 1993 and 1997

Fig. 7a: Export % of Sales

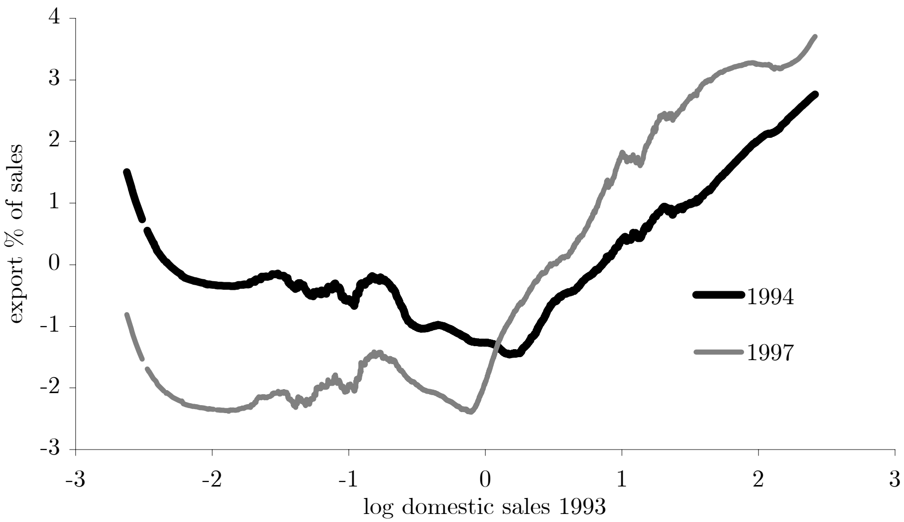


Fig. 7b: White-collar wage

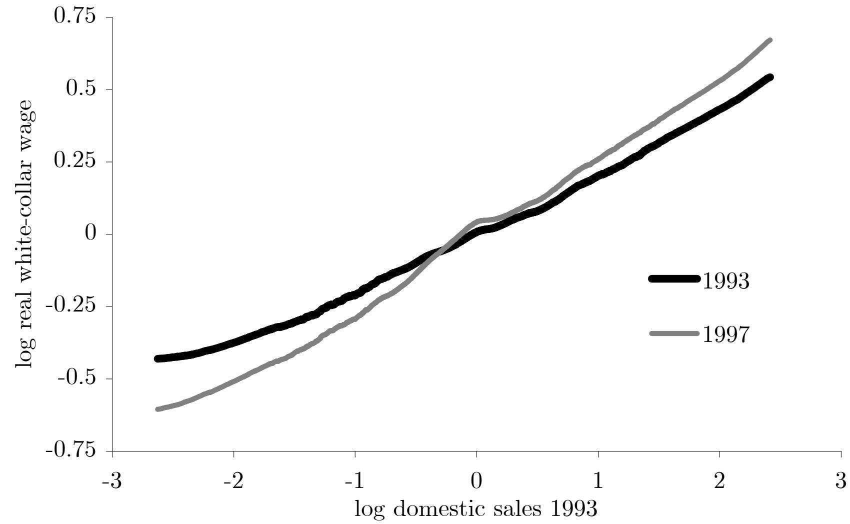


Fig. 7c: Blue-collar wage

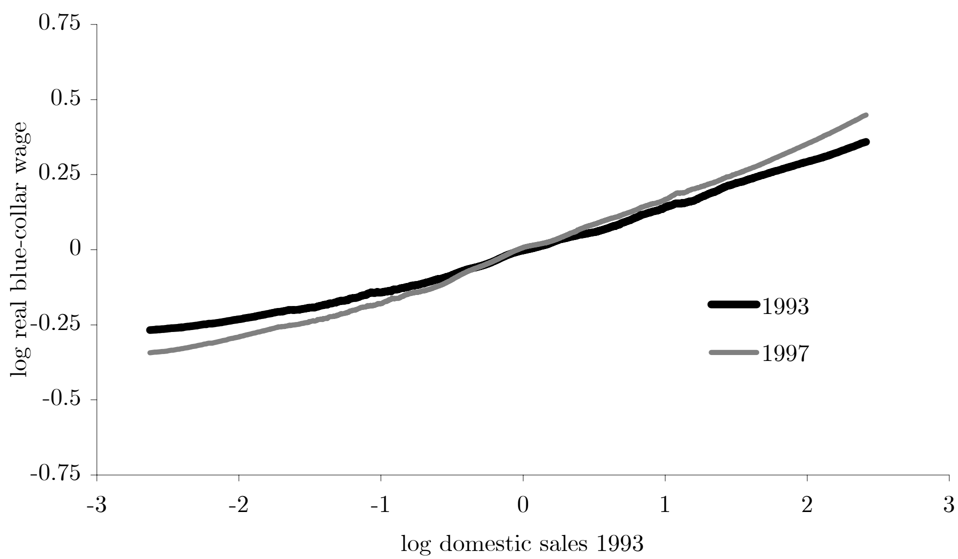


Fig. 7d: Wage ratio

Notes: All variables deviated from industry means. Graphs are locally smoothed non-parametric regressions (bandwidth = .5), of levels of indicated variables in indicated year on log domestic sales in 1993, using EIA 1993-2001 Balanced Panel.

Fig. 8: Non-Parametric Regressions, Changes 1993-1997 and 1997-2001

Fig. 8a: Changes in export percentage of sales

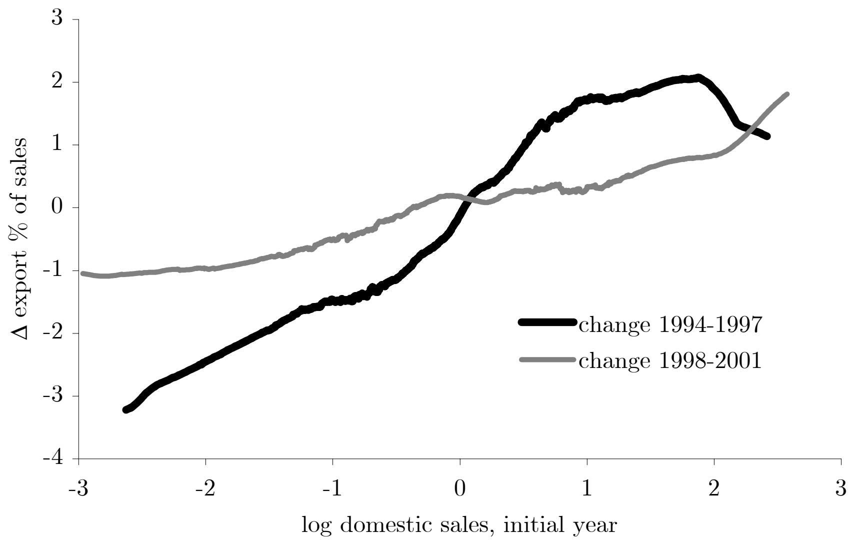


Fig. 8b: Changes in white-collar wage

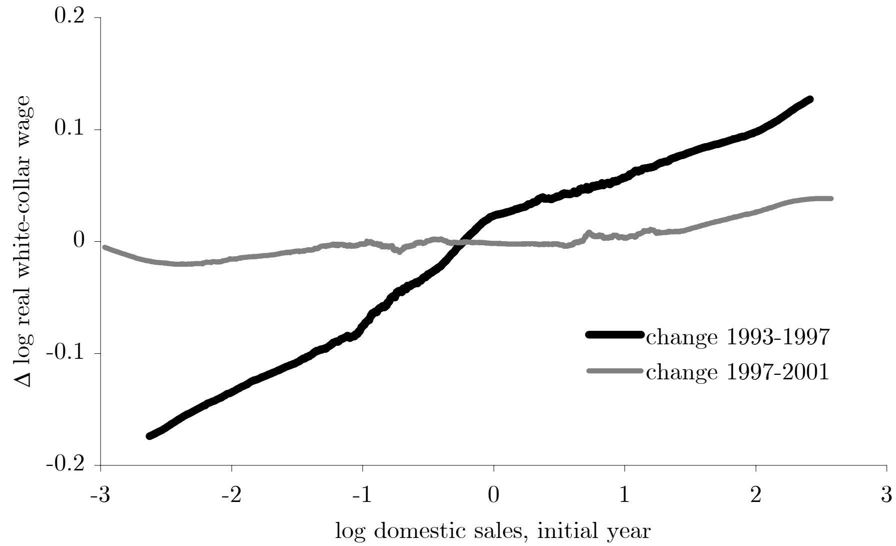


Fig. 8c: Changes in blue-collar wage

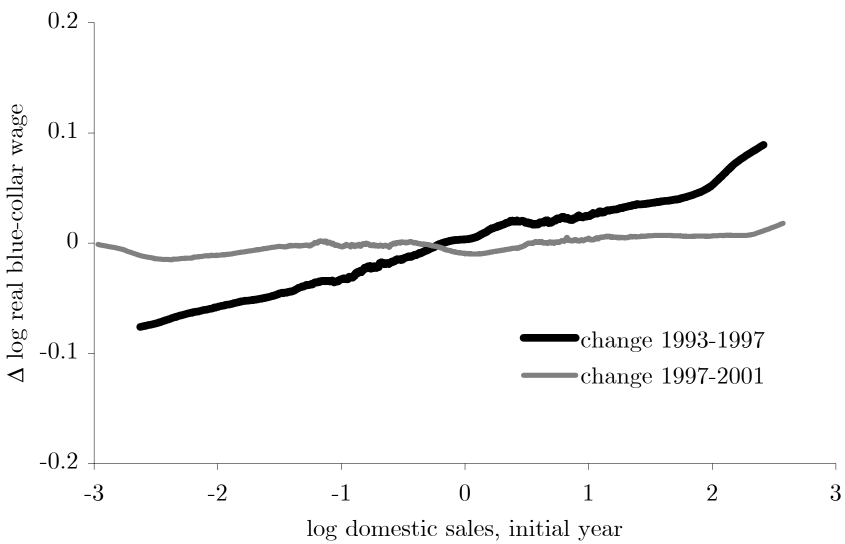
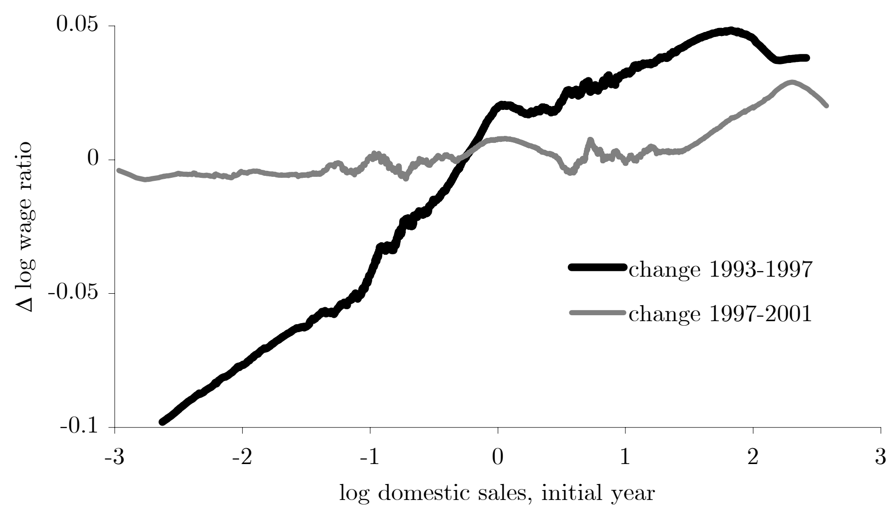


Fig. 8d: Changes in wage ratio



Notes: All variables deviated from industry means. Graphs are locally smoothed non-parametric regressions (bandwidth = .5), of changes of indicated variables over indicated periods on log domestic sales in initial year (1993 or 1997), using EIA 1993-2001 Balanced Panel. For export percentage, changes omit initial year to avoid bias from mean reversion.