

# Quality Upgrading and Establishment Wage Policies: Evidence from Mexican Employer-Employee Data

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PRELIMINARY

## Abstract

This paper draws on employer-employee data from the Mexican social security agency to investigate the relationship between product quality upgrading and wage premia at the plant level. Following Verhoogen (2004), we argue that the peso crisis of late 1994 generated a differential inducement to export and raise product quality within manufacturing industries, with initially larger and more productive establishments increasing exports and producing a greater share of high-quality goods to appeal to U.S. consumers, and initially smaller and less productive establishments remaining focused on the domestic market. We confirm in our data that initially larger manufacturing plants experienced greater wage growth than initially smaller plants within the same industry in 1993-1997, and that this differential change was larger than the differential change in manufacturing in 1997-2001. We then decompose plant-level average wage changes into changes in average workforce skill and changes in plant-year effects, which we take as measures of wage premia. Our baseline estimates suggest that 67%-89% of the difference in differential wage changes between the peso-crisis period and the later period can be attributed to changes in wage premia. In the non-tradable sector, by contrast, there were no significant differential changes in wage premia between initially larger and initially smaller plants in either period, and the modest difference in differential wage changes between the periods is explained largely by differential changes in workforce skill. These results support the hypothesis that the product quality upgrading that accompanied increasing exports led Mexican manufacturers to raise wage premia.

# 1 Introduction\*

Wage heterogeneity among observably similar workers is a pervasive feature of labor markets in both developed and developing countries. Some industries consistently pay more than others (Dickens and Katz, 1987; Krueger and Summers, 1986; Moll, 1993), large firms pay more than small firms (Brown and Medoff, 1989; Velenchik, 1997; Schaffner, 1998), exporting and foreign-owned firms pay more than non-exporting and domestic-owned ones (Bernard and Jensen, 1999; Lipsey and Sjöholm, 2004). A perennial issue is whether this heterogeneity is due to differences in the unobserved skills of workers or to differences in wage premia across establishments (Murphy and Topel, 1987). The growing availability of matched employer-employee data has made it possible to address this question, and a series of recent papers has found (under a particular set of assumptions) that a significant portion of inter-industry and firm-size wage differentials in France and the U.S. can be attributed to differences in firm- and plant-level wage premia (Abowd, Kramarz, and Margolis, 1999; Abowd, Finer, and Kramarz, 1999; Abowd and Kramarz, 1999b,a).

Although the literature using employer-employee data to estimate firm- or plant-level wage premia has been growing quickly, relatively little attention has been paid to the *determinants* of those premia. In this paper, we draw on employer-employee data from Mexico to investigate the role of one possible determinant: product quality. Following Verhoogen (2004), we argue that a trade-related shock — the Mexican peso crisis of late 1994 — generated exogenous variation in the incentive of Mexican producers to increase exports and

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raise product quality, and we relate that variation to variation in wage premia at the plant level. Specifically, we argue that initially larger, more productive establishments within each industry were more likely to increase exports than initially smaller, less productive ones; that exporting plants tended to produce higher-quality goods to export than for the domestic market in order to appeal to richer U.S. consumers; and that product quality upgrading brought about by increasing exports required an upgrading of the workforce as well. The starting point of this paper is the observation that the upgrading of the workforce may plausibly have been of two types: attracting new workers with higher levels of human capital, or improving the productivity of existing workers, for instance by increasing efficiency wages or by investing in training. The two types of workforce upgrading cannot be distinguished in conventional plant-level datasets that only report average wages by occupational category (such as those used by Verhoogen (2004)), but in principle they can be distinguished in employer-employee data, since only the second type has the effect of raising wage premia. A key goal of this paper is to estimate the extent to which wage increases in upgrading plants can be attributed to increasing wage premia.

Our primary data come from the administrative wage records of the Instituto Mexicano del Seguro Social (IMSS), the Mexican social security agency.<sup>1</sup> The data (described in Section 4) contain little information on each worker and almost no information on establishment characteristics. But they have the advantage that we are able to follow individual workers as they move (or not) from establishment to establishment, and can separately estimate the contributions of individual and establishment characteristics in determining their wage. We

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<sup>1</sup> The data have been obtained and cleaned by David Kaplan, and used in several previous papers (Castellanos, Garcia-Verdu, and Kaplan, 2004; Kaplan, Martinez Gonzalez, and Robertson, 2004, 2005).

supplement this information with data from two plant-level surveys conducted by the *Instituto Nacional de Estadísticas, Geografía e Información* (INEGI), the Mexican statistical agency.<sup>2</sup>

Our argument proceeds in several steps. First, we present a simple model to provide intuition for the empirical work (Section 2). A key assumption in the model is that firms (or plants) differ in a single latent variable,  $\lambda$ , which we refer to as entrepreneurial ability. The value of this variable shapes firms' product-quality and wage decisions, as well as their response to the exchange-rate shock. Second, using plant-level data from the manufacturing sector we construct a variety of proxies for the latent entrepreneurial-ability variable and show that the patterns of differential within-industry increases in exports, upgrading of product quality, and increases in wages are robust to the choice of proxy (Section 5). Third, using the employer-employee data for manufacturing and the only proxy available in those data, log employment, we confirm the existence of what might be termed the "quality upgrading" effect of the peso crisis on plant-level average wages: initially larger plants saw relatively greater wage growth than initially smaller plants within the same industry during the peso crisis period (1993-1997), and this differential wage change was greater than in a subsequent period with no devaluation (1997-2001) (Section 6.1).<sup>3</sup> Fourth, we decompose the plant-level average wages into changes in plant-year effects (i.e. plant-year specific wage premia)

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<sup>2</sup> The employer-employee data and the plant-level data are linkable. On our initiative, IMSS and INEGI have entered negotiations about sharing the data, and we are hopeful that we will be able to link the datasets in the near future.

<sup>3</sup> It is important to note that this "quality upgrading" effect is based entirely on a within-industry comparison. A variety of between-industry and macro-level factors are clearly crucial in accounting for the overall pattern of aggregate wage changes in Mexico, but we do not focus on them here. For more on wage changes in Mexico, see Cragg and Epelbaum (1996); Feenstra and Hanson (1997); Revenga (1997); Hanson and Harrison (1999); Meza (1999); Feliciano (2000); Esquivel and Rodriguez-Lopez (2003); Robertson (2000, 2003, 2004).

and changes in average skill of the workforce (Section 6.2). We show that 67%-89% of the difference in differential changes in plant-level average wages is accounted for by changes in plant-level wage premia, depending on the specification we use (Section 6.3). Fifth, we show that the pattern of changes in wage premia we observe in manufacturing is not present in the non-tradable sector, supporting our argument that the changes in wage premia are linked to increased exporting (Section 6.4). Sixth, we show that our results are robust to a number of modifications that address concerns with our basic specification (Section 6.5). Overall, we take these results as fairly strong support for the hypothesis that upgrading plants raised wage premia.

To our knowledge, ours is the first paper to use employer-employee data to estimate wage premia at the plant level in a developing country, and the first to relate those premia to shocks to the trade regime.<sup>4</sup> In addition to the papers cited above, our paper is related to papers on the relationship between trade and skill upgrading in developing countries (Feenstra and Hanson (1996); Davis (1996); Robbins (1996); Kremer and Maskin (2003); Pavcnik (2003); Zhu and Trefler (2005); Antras, Garicano, and Rossi-Hansberg (forthcoming); see Goldberg and Pavcnik (2004) for a review, and Bernard and Jensen (1997) on the U.S.); to models of trade under monopolistic competition with heterogeneous firms (Melitz, 2003; Chaney, 2005); to papers exploring the role of product quality in trade (Shaked and Sutton, 1982; Flam and Helpman, 1987; Stokey, 1991; Copeland and Kotwal, 1996; Murphy and Shleifer, 1997; Brooks, 2003; Schott, 2004; Hallak, forthcoming; Hallak and Schott, 2005); and to papers relating international trade to labor effort (Copeland, 1989; Brecher, 1992; Matusz,

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<sup>4</sup> Haltiwanger and Vodopivec (2003) analyze employer-employee data from Slovenia, but focus on worker and job flows. Muendler and Menezes (2005) use employer-employee data from Brazil to analyze labor reallocation following trade reform.

1996; Leamer, 1999).

## 2 Theory

This section develops a model of monopolistic-competition model of trade with heterogeneous producers, combining ideas from Melitz (2003); McFadden (1978); Anderson, de Palma, and Thisse (1992); Burenstam Linder (1961); Kremer (1993) and others. A number of caveats apply: the model is partial-equilibrium, static, and with special functional forms. The model differs from the model of Verhoogen (2004) in that it distinguishes between (contractible) skill and (non-contractible) effort, but is otherwise similar in structure to that model, to which readers are referred for a less terse exposition.

There are two countries, North ( $n$ ) and South ( $s$ ). Firms are constrained to produce no more than one good for each market. Firms can sell a different quality product and charge a different price in each market. They can also make different optimization decisions on the production line for goods destined for North than the line for goods destined for South. Let  $c = s, n$  index the country in which a plant is located; and  $d = s, n$  index the destination market for a particular production line. In each country, there is a mass of potential entrepreneurs, heterogeneous in a (fixed) productivity parameter  $\lambda$ , which can be thought of as representing entrepreneurial ability, with distribution  $g_c(\lambda)$ . As is standard in monopolistic competition models, it will turn out that each producer differentiates her good and has a monopoly over the production of her own variety; hence  $\lambda$  will also uniquely identify goods within the set of goods produced by firms in a particular location for a particular destination market, and we use  $\lambda$  to index both goods and entrepreneurs.

Assume that aggregate demand for differentiated good produced by entrepreneur with

ability  $\lambda$  in country  $c$  and sold in domestic market  $d$  is given by:

$$x_{cd}(\lambda) = \frac{N_d}{D_d} \exp \left\{ \theta_d q_{cd}(\lambda) - \frac{p_{cd}(\lambda)}{\delta_{cd}} \right\} \quad (1)$$

where

$$D_d \equiv \sum_{c=n,s} \int_{\lambda_{cd} \in \Lambda_{cd}} \exp \left\{ \theta_d q_{cd}(\lambda) - \frac{p_{cd}(\lambda)}{\delta_{cd}} \right\} g_c(\lambda) d\lambda \quad (2)$$

$N_d$  is the number of consumers in destination market  $d$ , taken to be exogenous.  $\theta_d$  is the willingness of consumers in country  $d$  to pay for quality, assumed to be greater in North than in South:  $\theta_n > \theta_s$ . This difference can be derived from differences in income, and is the only difference between Northern and Southern consumers.  $q_{cd}(\lambda)$  is the quality of good produced by entrepreneur with ability  $\lambda$  in country  $c$  for market  $d$ .  $p_{cd}(\lambda)$  is the price of the good, *denominated in terms of goods in country  $c$* .  $\delta_{cd}$  is the ratio of goods prices in country  $d$  to goods prices in country  $c$ , otherwise known as the real exchange rate. The term  $\frac{p_{cd}(\lambda)}{\delta_{cd}}$  thus represents the price charged by producer  $\lambda$  in location  $c$  for market  $d$ , denominated in terms of goods in country  $d$ , which is the appropriate measure in the consumer demand equation in country  $d$ .  $\Lambda_{cd}$  represents the set of all producers from country  $c$  that enter destination market  $d$ .  $D_d$  thus represents a sum over all producers (from either country) that enter market  $d$ . This expression for aggregate demand follows from a standard multinomial-logit specification of discrete choices by individuals, where each individual purchases one unit of a good in the industry and the utility of the good includes a random term with an extreme-value distribution.<sup>5</sup>

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Production of one unit of output is assumed to require one worker. Production of quality

<sup>5</sup> See Verhoogen (2004) for details.

is assumed to depend on three different inputs: entrepreneurial ability,  $\lambda$ , human capital,  $h$ , and effort,  $e$ , in Cobb-Douglas fashion:

$$q_{cd} = A\lambda(h_{cd})^{\beta_h}(e_{cd})^{\beta_e} \quad (3)$$

where  $\beta \equiv \beta_h + \beta_e$  and we assume  $\beta < 1$  to ensure an interior solution in the choice of quality.  $A$  is a constant term that captures the general level of technology. This specification borrows from Kremer (1993) the insight that when producing high-quality goods, a large number of low-quality workers cannot substitute for one high-quality one.<sup>6</sup> Note that this specification nests two special models:  $\beta_e = 0$  yields a model in which quality depends only on the inherent skill of workers, as in Kremer (1993);  $\beta_h = 0$  yields a model in which quality depends only on effort, as in Dalmazzo (2002).

We assume that there exists an exogenously given wage-skill schedule in the outside labor market, linear and passing through the origin:

$$h_{cd} = z_h \tilde{w}_{cd} \quad (4)$$

where  $z_h$  is a positive constant. The wage  $\tilde{w}$  represents the minimum required to induce a worker of human capital  $h$  to show up for work; we refer to it as the *market wage* for a worker with human capital  $h_{cd}$ . We assume that the wage-skill schedule has the same slope  $z_h$  in both countries, although this assumption could easily be relaxed. We assume that effort is a function of the gap between what a worker receives in the firm and her market wage. Let  $v$  represent the gap, or quasi-rent, where  $v \equiv w_{cd} - \tilde{w}_{cd}$ . For simplicity, we assume that effort

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<sup>6</sup> A situation in which high-quality workers are assigned to high-quality jobs may also arise in an assignment model of the type reviewed by Sattinger (1993).



is a linear function of the quasi-rent passing through the origin:

$$e_{cd} = z_e v_{cd} = z_e (w_{cd} - \tilde{w}_{cd}) \quad (5)$$

where  $z_e$  is a positive constant.<sup>7</sup>

The key distinction between human capital and effort is that the former is contractible and transferable across firms such that it earns the same return in a particular firm as on the outside market, while the latter is imperfectly transferable or subject to contracting problems such that workers earn a premium above what they would earn in the outside market. If workers have firm-specific skills that are not valued but that generate benefits to the firm, some of which are in turn shared with workers, those firm-specific skills are best interpreted as part of “effort” in our model. Similarly, if effort is observable and verifiable, as for instance in Leamer (1999), such that firms and workers can write contracts over the amount of effort that workers supply, then it is best interpreted as part of “human capital.”

Firms choose the price, ( $p$ ), the level of human capital, ( $h$ ), and the wage ( $w$ ), which in turn determine worker effort. Marginal cost per unit of output is constant, and equal to the wage of the single worker producing a single unit of output. The fixed cost to produce in the domestic market is  $f$ ; the fixed cost to produce for the export market is  $f_x$ , where the domestic fixed cost has to be paid prior to exporting. It is convenient to write these fixed costs as  $f_{cd}$  where  $f_{cd} = f + f_x$  for exporters and  $f$  for domestic producers.

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. As a

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<sup>7</sup> The assumptions that the wage-skill and wage-effort schedules are linear and passing through the origin are clearly simplistic, but more flexible specifications would seem to impose a greater algebraic burden with few compensating insights.

consequence, all firms differentiate and have a monopoly in the market for their particular variety. Profit is given by:

$$\pi_{cd}(\lambda) = (p_{cd} - w_{cd}) x_{cd} - f_{cd} \quad (6)$$

Each firm is assumed to treat the aggregate quantity in the denominator of the expression for demand,  $D_d$  in equation 1, as unaffected by its own choices. Optimizing over the choice of  $p$ ,  $h$ , and  $w$ , and solving for  $q$ , we have:

$$q_{cd}(\lambda) = \eta(\lambda)^{\frac{1}{1-\beta}} (\delta_{cd}\theta_d)^{\frac{\beta}{1-\beta}} \quad (7)$$

where  $\eta \equiv (A)^{\frac{1}{1-\beta}} (z_h\beta_h)^{\frac{\beta_h}{1-\beta}} (z_e\beta_e)^{\frac{\beta_e}{1-\beta}}$ . The optimal quality of good produced is increasing in both  $\lambda$  and  $\theta$ . An entrepreneur with higher ability will produce higher quality for a given market. Conditional on a given level of ability, an entrepreneur will produce a higher quality good for the Northern than for the Southern market.

The optimal choices of human capital,  $h$ , and wage,  $w$ , and the corresponding level of effort,  $e$ , can be summarized in terms of the optimal quality level:

$$h_{cd}(\lambda) = \theta_d \delta_{cd} z_h \beta_h q_{cd}(\lambda) \quad (8)$$

$$w_{cd}(\lambda) = \theta_d \delta_{cd} \beta q_{cd}(\lambda)$$

$$e_{cd}(\lambda) = \theta_d \delta_{cd} z_e \beta_e q_{cd}(\lambda)$$

These variables follow the same pattern as quality: all are increasing in both  $\lambda$  and  $\theta$ . Note further that the Cobb-Douglas form of the production function for quality implies that the shares of the wage made up of the “market wage”,  $\tilde{w}$ , and the quasi-rent,  $v$ , are constant:

$$\tilde{w}_{cd}(\lambda) = \frac{\beta_h}{\beta} w_{cd}(\lambda) \quad (9)$$

$$v_{cd}(\lambda) = \frac{\beta_e}{\beta} w_{cd}(\lambda)$$

Thus the “market wage” and quasi-rent components of wages also follow the same cross-sectional pattern as product quality.

Although we have modeled plants as making separate wage decisions on each production line, we do not observe wages by production line. The wages (and hence market wages and quasi-rents) we observe correspond to weighted averages across production lines, where the weights are given by the export and domestic-market shares of production.

The analysis of plants’ entry decisions (and hence the shares of export and domestic production that enter the wage averaged across production lines) is somewhat tedious and is the same as that presented in Verhoogen (2004), so we do not reproduce it here. Instead, we sketch the intuition informally, focusing on Southern plants. Figure 1 summarizes the relationship between product quality and  $\lambda$  in the cross-section of Southern plants, prior to an exchange-rate shock. The  $q_{ss}(\lambda)$  curve represents the quality that a Southern plant with ability  $\lambda$  produces for the Southern market; the  $q_{sn}(\lambda)$  curve represents the corresponding quality for the Northern market. As mentioned above, both are increasing in  $\lambda$ , and for a given  $\lambda$ ,  $q_{sn}(\lambda) > q_{ss}(\lambda)$ . The  $\bar{q}_s(\lambda)$  (dotted) curve represents the counterfactual average quality that would obtain if all plants entered both markets. Not all plants enter both markets: entrepreneurs with  $\lambda < \lambda_{ss}^{\min}$  enter neither market; those with  $\lambda_{ss}^{\min} \leq \lambda < \lambda_{sn}^{\min}$  enter only the domestic market; those with  $\lambda_{sn}^{\min} \leq \lambda$  enter both markets. What we actually expect to observe, both in product quality (to the extent that it is observable) and in wages, is given by the curve in bold.

We model an exchange-rate shock as having two exogenous effects: first, the real exchange

rate,  $\delta_{sn}$ , rises; second, total consumer demand in South, represented by the number of Southern consumers,  $N_s$ , falls. Figure 2 summarizes the effect of the exchange rate shock on the level average quality within each plant. The  $q_{sn}(\lambda)$  curve shifts up in response to the shock, since quality is cheaper to produce relative to what consumers are willing to pay for it. The dotted counterfactual  $\bar{q}_s$  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 reproduces the solid curve in Figure 1. The thicker, gray solid line represents the actual average quality curve post-crisis. Figure 3 depicts the *change* in average quality as a function of  $\lambda$ , that is, the difference between the thick gray and the thin black curves in Figure 2.

It can be shown that the change in the plant-level average wage,  $d\bar{w}_s(\lambda)$ , follows the same pattern as the change in average quality illustrated in Figure 3. From equation 9, it follows directly that the changes in market wages and quasi-rents are proportional to the overall wage change:

$$\begin{aligned} d\bar{w}_s(\lambda) &= \frac{\beta_h}{\beta} d\bar{w}_s(\lambda) \\ d\bar{v}_s(\lambda) &= \frac{\beta_e}{\beta} d\bar{w}_s(\lambda) \end{aligned} \tag{10}$$

A central goal of our empirical work will be to estimate the parameters  $\beta_h$  and  $\beta_e$ , and thereby decompose the differential plant-level wage changes (observed in Verhoogen (2004) and confirmed in our data) into market wage and quasi-rent components. Note that if we only had data on average wages at the plant level, these parameters would not be separately identified, and we would not be able to distinguish sorting by unobserved skill from changes in wage premia at the plant level. An important caveat to our empirical implementation is that

imprecision in our proxies for  $\lambda$  prevent us from estimate the exact functional form illustrated in Figure 3. We must content ourselves with estimating linear relationships between initial  $\lambda$  and subsequent changes in product quality and wage measures.

### 3 Econometric Approach

The model discussed above suggests an empirical approach based on comparisons of differential changes between initially higher- $\lambda$  and initially lower- $\lambda$  plants across time periods and sectors: if the quality-upgrading story is correct, we expect to see greater differential product quality and wage changes in the period of the peso crisis than in a period without a devaluation, and greater differential changes in tradables (manufacturing) than in non-tradables (construction, retail, services, transportation). These comparisons form the basis of our identification strategy.

Our basic econometric model has the following form:

$$\Delta y_j = \mu + \hat{\lambda}_{0j}\gamma + D_j\pi + \epsilon_j \tag{11}$$

where  $j$  indexes establishments,  $\Delta y_j$  is a change in an outcome variable of interest,  $\hat{\lambda}_{0j}$  is a proxy for the latent entrepreneurial ability variable taken from an initial year (discussed in more detail in Section 5 below), and  $D_j$  is a vector of industry and state dummies. The outcome variables we focus on are the export share, an indicator of product quality, and wage measures. In Section 6.2 below, we describe how we estimate measures of the “market wage” and quasi-rent components of wages, and include them as outcome variables in this equation.

The coefficient of interest in these regressions is  $\gamma$ , which captures a differential change

in the outcome by the initial value of the entrepreneurial ability proxy. We estimate the model separately for the peso crisis period (1993-1997) and a later period without a devaluation (1997-2001), and separately for tradables and non-tradables.<sup>8</sup> If there were no other differential influences on large and small plants besides the exchange rate, we would expect  $\gamma > 0$  in 1993-1997 in tradables and  $\gamma = 0$  in 1997-2001 or in tradables. However, a more plausible specification would allow for some differential evolution between high- $\lambda$  and low- $\lambda$  plants, as would be predicted by an industrial evolution model such as that of Jovanovic (1982). For our purposes, the crucial prediction of the model is that  $\gamma$  is larger in 1993-1997 and in tradables than in the other period or sector.

The key assumption underlying our approach is that differences between the estimates of  $\gamma$  in tradables in 1993-1997 and the estimates for non-tradables or the non-tradables sector can be attributed to the peso crisis and the quality-upgrading mechanism outlined above. Our approach is analogous to a difference-in-difference-in-differences (D-in-D-in-D) strategy: we compare differences in changes in outcomes between high- $\lambda$  and low- $\lambda$  plants in a “treated” period, 1993-1997, and a “control” period, 1997-2001. When we add the comparison with the non-tradables sector, our approach is analogous to a four-difference design. The analogy is not fully accurate in the sense that  $\lambda$  is a continuous rather than a binary variable, but it points out that our approach shares the vulnerability of nearly all difference-in-difference-type designs that there may be unrelated differential trends confounding our results. We address what we see as the main counter-arguments as we proceed below.

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<sup>8</sup> Our choice of the periods 1993-1997 and 1997-2001 is determined largely by availability of the EIA plant-level data. The main EIA panel begins in 1993, and we have access until 2001. We have experimented with alternative periodizations, and have found results similar to those reported here.

## 4 The Data

The employer-employee data used in this paper are drawn from the administrative records of the Instituto Mexicano del Seguro Social (IMSS), the Mexican social security agency. All private, formal-sector Mexican employers are required to report wages for their employees, and pay social-security taxes on the basis of their reports. The raw data can be considered a census of private, formal-sector establishments and their workforces for 1985-2001. The number of workers in the raw data at a given point in time ranges from approximately 5 million in 1985 to approximately 11 million in 2001. Kaplan, Martinez Gonzalez, and Robertson (2004, 2005) compared IMSS employment figures for manufacturing to the 1993 Industrial Census, and found that approximately 90% of manufacturing workers in private firms were included in the IMSS data.

The IMSS data contain information on the age, sex and daily wage of individuals, in addition to the state and year of the individual's first registration with IMSS and an individual identifier. The wage figures are based on a measure of total compensation, called the *salario base de cotizacion*, which includes both wages and benefits, including payments made in cash, bonuses, commissions, room and board, overtime payments, and in-kind benefits. The raw data include a start date and an end date for each wage earned for each individual in each establishment. When an individual's wage changes, the record for the old wage is closed and a record for the new wage is opened. We extracted data for September 30 for each year. At the establishment level, the data contain information only on location and industry (using the IMSS's own 271 4-digit industrial categories). Further details are in the data appendix.

We supplement the employer-employee data with information from two plant-level sur-

veys of the manufacturing sector conducted by INEGI, the Mexican statistical agency. The Encuesta Industrial Anual (EIA) [Annual Industrial Survey], a yearly plant panel, contains basic balance sheet information as well as data on employment, hours worked, and total wage bill for two separate occupational categories, *empleados* and *obreros*, corresponding to white-collar and blue-collar workers.<sup>9</sup> The EIA does not include information on the *maquiladoras*, the assembly-for-export plants located mainly along the U.S. border. The sampling design of the EIA is somewhat unorthodox: the largest plants in 205 6-digit industries were selected deterministically in 1993, and those plants were followed over time, with minimal and haphazard refreshing of the sample. After cleaning, we are left with a balanced panel of 3,255 plants over the period 1993-2001.<sup>10</sup> For details, refer to the data appendix of Verhoogen (2004).

The Encuesta Nacional de Empleo, Salarios, Tecnologia y Capacitacion (ENESTyC) [National Survey of Employment, Wages, Technology and Training], contains information on a variety of qualitative plant characteristics. The ENESTyC includes information on whether a plant has ISO 9000 certification, an international production standard commonly associated with product quality. We draw on information from four waves of the ENESTyC, in 1992, 1995, 1999 and 2001, each of which mainly contains information from the previous calendar year (1991, 1994, 1998 and 2000). A subset of plants from the EIA 1993-2001 panel can be linked to the ENESTyC: 1791 in 1994; 850 in all three years 1994, 1998 and 2000. The ENESTyC includes some maquiladoras, but we exclude them since they have no matching

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<sup>9</sup> The white-collar/blue-collar breakdown comes from the Encuesta Industrial Mensual (EIM), a less extensive but higher-frequency survey conducted on the same sample as the EIA.

<sup>10</sup> An earlier EIA panel is also available for the 1984-1994 period, but we do not use it in this version of this paper.



record in the EIA.

In anticipation of being able to link the IMSS employer-employee data to the INEGI plant-level data, we construct a balanced panel of establishments in the employer-employee data that is comparable to the EIA plant-level panel. We have also been guided by the goal of maintaining the same sample size in our various specifications. In what we refer to as the IMSS manufacturing panel, we include manufacturing plants that have at least 50 workers in all years 1993-2001, that have at least one leaver (worker who switches out) and one entrant (worker who switches in) in each year, and that have at least one worker who has stayed and one who will stay for four years or the maximum possible given our time period, whichever is smaller. We also construct an IMSS non-tradables panel, using the same criteria for establishments in non-tradable sectors: construction, retail trade, transportation and services. For the purposes of identifying plant effects, we further require that each establishment be part of the largest group of “connected” establishments, where connected means sharing at least one worker at some point during the relevant period with another connected firm. The manufacturing and non-tradables panels contain 3628 and 3659 establishments respectively.

Summary statistics on these panels appear in Table 1. Within each sector, larger establishments tend to pay higher wages. Worker separation rates in our data are somewhat larger than the rates in the U.S., but of the same order of magnitude. (For further details, see Kaplan, Martinez Gonzalez, and Robertson (2004).) Importantly, the cumulative turnover rates are large enough to be consistent with significant changes in the composition of the workforce in each establishment. Summary statistics for the EIA panel for 1993, broken down by export status, appear in Table 2. The patterns are similar to those documented

for the U.S. by Bernard and Jensen (1999), and subsequently for many other countries: exporters are larger, more capital-intensive, higher-wage, more likely to be foreign-owned than non-exporters, and are a minority of plants in each industry. In addition, the second panel of the table presents information from the ENESTyC and shows that exporters are more likely to have ISO 9000 certification and more likely to have a formal employee training program than non-exporters.

## 5 Results from EIA Plant-Level Data

A crucial first step in our econometric approach is to choose a proxy for the latent entrepreneurial-ability variable,  $\lambda$ . The model suggests that a number of observable variables at the plant level — the export share, indicators of product quality, sales in each market (export and domestic), employment, and wages — will be correlated with the underlying latent variable in cross-section. In a model with capital, we would also expect capital-intensity to be correlated with  $\lambda$ . If receiving foreign direct investment (FDI) affords a plant access to the technical knowledge of the parent firm, we might expect FDI to be correlated with  $\lambda$  as well. In principle, there are a variety of ways in which we might combine these variables to construct a proxy for  $\lambda$ .

In this version of the paper, we have not yet been able to link the IMSS employer-employee data to the EIA plant-level data, and hence we just have one proxy available in the employer-employee data: log employment.<sup>11</sup> In this section, we use information from the EIA plant-level data to show that log employment is highly correlated with a number of other plausible proxies, and that the basic results on plant-level differential changes in

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<sup>11</sup> We avoid using wage levels as a proxy for  $\lambda$ , since changes in wages are our primary outcome variable.

exports, product quality and wages are robust to the choice of proxy.

In the EIA panel, we construct six different proxies. In all cases, we use information from an initial year (or years), i.e. 1993 or 1997 (1993-94 or 1997-98 in the case of TFP).

- Log domestic sales: This proxy has the practical advantage that it is relatively well-measured and the theoretical advantage that it bears a smooth, continuously differentiable relationship to  $\lambda$ , since, unlike other variables, sales are measured separately by production line.
- Export share index: We run a tobit of the export share of sales on log hours worked (separately by occupation), log domestic sales, log capital-labor ratio, and an indicator for whether the plant has  $\geq 10\%$  foreign ownership; we then recover the predicted values ( $x'_j \hat{\beta}$  where  $x_j$  is a vector of the independent variables and  $\hat{\beta}$  are the estimated coefficients), deviate from industry means, and standardize the variable.
- ISO 9000 index: Using the subset of observations that can be linked between the EIA and the ENESTyC, we run a probit of whether or not a plant has ISO 9000 certification on log hours worked (separately by occupation), log domestic sales, log capital-labor ratio, the FDI indicator, and export share of sales.<sup>12</sup> We recover the estimated coefficients, and calculate the predicted  $x'_j \hat{\beta}$  for all plants in our EIA panel, deviate from industry means and standardize.
- First principal component: We take the first principal component (the vector capturing the maximum proportion of joint variance) of log total hours worked (separately by

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<sup>12</sup> We regress the ISO 9000 certification information from the 1995 ENESTyC (pertaining to 1994) on information from the 1993 EIA, and ISO 9000 from the 1999 ENESTyC (pertaining to 1998) on information from the 1997 EIA.

occupation), log domestic sales, log capital-labor ratio, the FDI indicator, and export share of sales. We deviate from industry means and standardize.

- Total factor productivity (TFP): We regress log value-added (sales minus costs) on log hours worked (separately by occupation), log capital stock,<sup>13</sup> year dummies and plant dummies for the two initial years of a period (i.e. 1993 and 1994, for the 1993-1997 period), separately for 9 two-digit industries. We then recover the coefficient estimates on the plant dummies, deviate them from industry means, and standardize.<sup>14</sup> This approach has the advantage that year-specific measurement error in value-added is not treated as part of TFP.
- Log employment: This proxy has the advantage of being well-measured, and, as mentioned above, the only proxy available in the IMSS employer-employee data.

Table 3 presents bivariate correlations between these proxies for 1993. Five of these six proxies are highly correlated. The exception is the TFP measure, but even for TFP we can reject the hypotheses of no bivariate correlations with ease.

We now turn to the evidence on the differential effect of the peso crisis within industries, using our various proxies for the entrepreneurial-ability variable,  $\lambda$ . It is important to note, first of all, that the crisis represented an enormous shock. The Mexican peso lost approximately 50% of its nominal value, and nearly as much of its real value, in a matter of days at the end of December, 1994. Figure 4 plots the real exchange rate over the 1984-2001 period. GDP fell by 6.7% from 1994 to 1995. Exports rose sharply, with approximately 85%

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<sup>13</sup> Calculated following the permanent-inventory method described in Olley and Pakes (1996).

<sup>14</sup> We thank Matthias Schuendeln for suggesting this approach.

destined for the U.S. market. Figures 5a-c, which are based on the EIA 1993-2001 balanced panel, illustrate that domestic sales dipped and export sales rose, with the result that the export share of sales rose sharply. The number of plant with positive exports rose from approximately 30% to 45% of the sample, but the new entry was slower than the increase in the export share of sales.<sup>15</sup>

Consider the differential export response to the shock. Columns 1 and 2 of Table 4 present results of 12 regressions of the form of equation 11 (six proxies, two periods each), where the change in the export share is the dependent variable and the coefficients on the industry and region dummies have been omitted. (Column 3 presents the differences between the 1993-1997 and 1997-2001 coefficients.) We see that in the 1993-1997 period, higher- $\lambda$  plants within each industry saw greater export growth than lower- $\lambda$  plants, and this differential change was significantly greater than in the 1997-2001 period. This basic pattern is robust to the choice of proxy. To gain a sense of the magnitudes of the effects, note that if we compare two plants, one of which was a factor  $e$  larger than the other in domestic sales terms, the larger plant saw export growth approximate 1.9% greater than the smaller plant over the 1993-1997 period, and .6% greater over the 1997-2001 period.

One concern with the specification in Columns 1-2 is that domestic sales, which appear in the denominator of the export share, also enter the initial level of the various proxies on the right-hand side; a spuriously large coefficient on the entrepreneurial-ability proxy may

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<sup>15</sup> It is worth emphasizing that the peso crisis was a much larger shock than the North American Free Trade Agreement (NAFTA), which took effect in January 1994. Mexico's main trade liberalization came with its entrance into the General Agreement on Tariffs and Trade in the mid-1980s, and by 1994 the vast majority of Mexican imports were covered by tariffs of 20% or less. Average U.S. tariffs on goods from Mexico were on the order of 3-5%. In the majority of cases, NAFTA phased out existing tariffs slowly over time. Relative to the exchange-rate devaluation, the year-by-year tariff changes were quite small.

result. To address this concern, we also estimate an instrumental variables model, where we use the value of the proxy from 1993 as an instrument for the 1994 value, and use the change in export share from 1994-1997 as the dependent variable. Columns 4-6 present the results. The qualitative message is unchanged.

Product quality is difficult to capture in large datasets. The best measure we have of product quality for a large set of plants is ISO 9000 certification. The ISO 9000 certification process is largely focused on how systematized and rationalized management processes are, but the certification is commonly believed to reflect product quality and reliability as well. We have information on 850 plants that appear both in the EIA and in the three waves of the ENESTyC in which ISO 9000 information is available. Table 5 reports regressions of the form of equation 11, with the change in an ISO 9000 indicator variable as the outcome. Columns 1-3 report the results from simple linear probability models. While we observe significant differences between the coefficients on the  $\lambda$  proxies for 1993-1997 and 1997-2001, the coefficients for 1993-1997 are in several cases not significantly different from zero, and the coefficients for in 1997-2001 are significantly negative, which seems counter-intuitive in the context of our model. Both of these puzzling results may be due to censoring: plants that had already acquired ISO 9000 initially — which tended to be high- $\lambda$  — could not have experienced a positive change in ISO 9000 over the subsequent period, and plants that did not have the certification initially could not have experience a negative change. An arguably more appropriate specification is to estimate a censored normal regression model, where plants with ISO 9000 initially are treated as censored above, and plants without ISO 9000 initially are treated as censored below. Results from this specification are reported in

Columns 4-6. The results are consistent with the quality upgrading story: high- $\lambda$  plants are significantly more likely to acquire ISO 9000 over the 1993-1997 period than low- $\lambda$  plants, and this difference is larger in 1993-1997 than in 1997-2001. Again, the results are robust to the choice of proxy.<sup>16</sup>

Finally, consider wage changes. Table 6 is organized like Table 4, with the change in the log average wage at the plant as the outcome variable, using the full EIA panel. The results are again consistent with the quality-upgrading story: the coefficients for 1993-1997 are significantly greater than for 1997-2001, and the results are not sensitive to the choice of proxy.<sup>17</sup> The motivation for including the IV results is again measurement error. Average wages in the EIA are calculated by dividing the wage bill by employment; measurement error in the initial level of employment would induce a spurious positive correlation between initial employment and the change in wages. A related concern is that establishments were heterogeneous and inconsistent in how they reported low-wage temporary blue-collar workers. If an establishment reported temporary workers as employees in an initial year but not in a later year, we would expect a similar bias. The IV results address both of these concerns, and are similar to the OLS results.

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<sup>16</sup> These results are consistent with case study evidence suggesting that the shift in production towards the export market was accompanied by an increase in the average quality of good produced in exporting plants. For instance, Verhoogen (2004) presents evidence that the Volkswagen plant in Puebla, Mexico, shifted production from Old Beetles for the Mexican market to Jettas and Golfs for the U.S. market, with a corresponding increase in technical sophistication and skill requirements.

<sup>17</sup> Note that this is a purely within-industry comparison. Real wages overall fell sharply in the 1993-1997 period; the results indicate that they fell less in higher- $\lambda$  plants within each industry.

## **6 Results from Employer-Employee Data**

### **6.1 Replicating Plant-Level Results**

We now turn to the IMSS employer-employee data, using log employment as our proxy for entrepreneurial ability. We first attempt to replicate the results from the plant-level data in the employer-employee dataset. Part A of Table 7 reproduces (with some reorganization) the results from the bottom row of Table 6. Part B of the table presents analogous results using the IMSS employer-employee tradables panel. The results are quite similar across the two datasets. There is again a statistically significant difference between the results for 1993-1997 and 1997-2001 periods. The hypothesis that the estimates of the differences between the 1993-1997 and 1997-2001 coefficients are equal across datasets cannot be rejected at conventional levels of confidence. The similarity of the results across datasets is especially reassuring for two reasons. First, establishments have an incentive to under-report employees' wages to the IMSS. The fact that the results are similar to results using reported wages in the EIA, which have no bearing on social security taxes or other government payments, suggests that the bias from under-reporting is small. Second, the IMSS data report only a daily wage; we do not know how many hours of work the daily wage reflects. In the EIA data, by contrast, we know the number of hours of work; the average wage variable is an average hourly wage. The fact that the results using the different datasets are so similar again suggests that the bias from using the daily rather than the hourly wage is small.

### **6.2 Decomposing Plant-Level Average Wages**

The results for plant-level average wages leave open the question of whether the differential wage changes are due to sorting of workers of different skill into plants or to changing wage



premia. We now turn to the task of decomposing the plant-level average wage changes into these two components. The basic econometric model we use to conduct this decomposition is the following:

$$w_{it} = \alpha_i + \psi_{j(i,t)t} + x'_{i,t}\phi + u_{it} \quad (12)$$

where  $w_{it}$  is the log wage of individual  $i$  in year  $t$ ,  $\alpha_i$  is a person fixed effect,  $j(i, t)$  indicates the plant at which individual  $i$  is employed in period  $t$ ,  $\psi_{j(i,t)t}$  is a plant-year effect,  $x'_{i,t}$  is a vector of observable time-varying characteristics of individual  $i$ , and  $u_{it}$  is a mean-zero disturbance. We use the years 1993, 1997 and 2001.

Three notes about this model are important. First, the model allows plant effects to vary across years, but constrains individual effects to be fixed; we relax this assumption below. Second, the person effects  $\alpha_i$  and the plant-year effects  $\psi_{jt}$  are separately identified in this regression only by individuals switching between establishments. (This is the reason for the requirement that all establishments in our panels be part of a “connected” group of establishments; for details on identification of plant effects, see Abowd, Creecy, and Kramarz (2002).) Third, estimating this model by OLS requires the assumption that the error term is uncorrelated with the co-variates. In particular, we assume that individual innovations in productivity (which enter the error term) are uncorrelated with the plant-year effects. This assumption is often referred to as the assumption of conditional random mobility, or simply exogenous mobility. Note that it does not require that mobility be completely random; it requires that mobility be random conditional on time-invariant individual characteristics and the other co-variates. In imposing this assumption, we are following Abowd, Kramarz, and Margolis (1999) and the subsequent employer-employee literature. Nevertheless, the

assumption is clearly a strong one, and we plan to investigate its validity in future work.

We estimate this model on individual-level data for the years 1993, 1997 and 2001. We have few individual-level observables to include in the  $x'_{i,t}$  vector: tenure and tenure squared, age squared, and experience squared. (Age and experience in levels are not identified separately from year dummies; age squared and experienced squared are separately identified from each other because we observe the date an individual first registered with the IMSS, and do not have to calculate experience as a function of age.) We implement the regression as a simple “within” regression, where the outcome and covariates are deviated from individual-specific means over the period. The computational problems discussed in Abowd, Kramarz, and Margolis (1999) and Abowd, Creecy, and Kramarz (2002) do not arise here because we estimate approximately 11,000 (approximately 3,600 plants \* 3 years) plant-year effects, and matrices of this size can be inverted by standard software packages.

Once we have estimated this model, we recover the estimated plant-year effects,  $\hat{\psi}_{jt}$ , and treat them as measures of wage premia. We also construct measures of average skill of the workforce at the plant level. Under our assumptions, the estimated person effects,  $\hat{\alpha}_i$ , are estimates of the value of all time-invariant individual characteristics, including what is commonly thought of as unobserved skill as well as observed skill.<sup>18</sup> The estimates  $x'_{i,t}\hat{\phi}$  are estimates of the value of the available time-varying observable characteristics. Thus  $\hat{\alpha}_i + x'_{i,t}\hat{\phi}$  is an estimate of the overall value of the skills of individual  $i$  at time  $t$ . We take the average

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<sup>18</sup> Note that this measure of skill should be interpreted as a measure of *general* skill — the skill that earns equal rewards within a given plant or in the outside market. Workers may develop plant-specific skills, and may reap some of the benefits of those skills. In our framework, the returns to plant-specific skills would count as part of the wage premium, where the wage premium should be interpreted as the difference between what a given worker earns in the plant and what he or she would earn on the outside market.

of this measure at the plant level in each year:

$$\bar{\alpha}_{jt} = \sum_{i|j(i,t)=j} (\hat{\alpha}_i + x'_{i,t}\hat{\phi}) \quad (13)$$

We take this measure as an estimate of the average skill of the workforce at the plant-year level.

To illustrate the role of the time-varying observables and to check the robustness of our results, we also estimate equation 12 imposing the restriction that  $\phi = 0$ , i.e. excluding the time-varying observables in equation 12 and assuming that the individual fixed effects capture all relevant dimensions of employee skill.

### 6.3 Estimating Changes in Wage Premia and Average Workforce Skill

With our plant-year-specific measures of wage premia and average skill, we estimate models of the form of equation 11, with the average log wage, the estimated plant-year effect,  $\hat{\psi}_{jt}$ , and the average skill of the workforce,  $\bar{\alpha}_i$ , as dependent variables.

Part A of Table 8 reports the results under the assumption that  $\phi = 0$  in equation 12. The advantage of beginning with this more restrictive specification is that the decomposition of plant-level average wages into wage premia and average skill is particularly transparent. The results for the average log wage in Column 1, which do not reflect the decomposition described above, are (unsurprisingly) similar to the results for the log average wage in Part B of Table 7. Part A, Columns 2-3 indicate that initially higher-employment plants have a greater change in both wage premia and average workforce skill than initially lower-employment plants in both periods, but that the differential change was significantly larger in 1993-1997 than in 1997-2001. The key message of Part A, Columns 1-3, however, is in the magnitudes: for

plant-level average wages, the difference in the 1993-1997 and 1997-2001 coefficients on log employment is 0.036; approximately two-thirds (0.024/0.036) of this difference is explained by the changing wage premia. As a further test, we calculate the average log wage at the establishment-year level for stayers, employees continuously employed in an establishment.<sup>19</sup>

Note that the stayers results are subject to a potential bias due to selection on observables (i.e. individual fixed effects or time-varying observed characteristics) in addition to the potential bias due to selection on unobservables (i.e. period-specific shocks to individual productivity) mentioned above. Nonetheless, the stayers results are a simple, transparent check on the decomposition of average wage changes into wage premia changes and average workforce skill changes. In the absence of extensive selection on observables, we would expect the stayers results to resemble the results for the changes in wage premia, and they indeed are reassuringly similar to the wage premia results of Column 2.

Part B of Table 8 reports the results where  $\phi$  in equation 12 is unrestricted. Part B, Column 1 reproduces the results in Part A, Column 1, which are not sensitive to the restriction on  $\phi$ . The key point to notice in Part B, Column 2 is that the coefficient representing the differential change in wage premia over 1997-2001 goes essentially to zero. In other words, the significant differential change in wage premia in 1997-2001 in Part A, Column 2 can be attributed to almost entirely to changes in observable workforce characteristics. The coefficient for average skill for 1997-2001 in Part B, Column 3 reflects this fact: it is roughly twice as large as when time-varying observable characteristics were excluded. The results in Part B suggest that an even larger proportion of the difference in differential changes in average

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<sup>19</sup> We keep track of the set of stayers from 1993-1997 and from 1997-2001 separately.

wages — approximately 89% — can be attributed to changes in wage premia.

Once we include time-varying observables in equation 12, the results for changes in wage premia (Part B, Column 2) are no longer directly comparable to the results for raw changes in average wages for stayers (Part A, Column 4). To construct comparable figures for stayers, we regress the change in the average wage of stayers on initial log employment and changes in the average values of the time-varying observables. The coefficients on the log employment term in each period are reported in Part B, Column 4. The results for stayers are again not statistically different from the wage premia results.

#### 6.4 Results for Non-Tradable Sectors

A possible objection to our approach is that the difference in differential changes in our dependent variables was indeed due to the peso crisis, but that the effect operated through a channel other than the quality-upgrading channel we have emphasized here. For example, the peso devaluation was accompanied by a major crisis in the Mexican banking sector, which resulted in a contraction of the availability of credit. If this credit crunch adversely affected small firms, then we might again observe a greater coefficient on initial employment in 1993-1997, even in the absence of quality upgrading. To address this objection, we compare the manufacturing sector to a collection of non-tradable sectors: construction, retail trade, transportation and services. If the differential wage changes were a consequence of the credit crunch or another aspect of the macro shock unrelated to trade and product quality, then we would expect to see similar results in non-tradables as in manufacturing.<sup>20</sup>

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<sup>20</sup> Additional evidence that the differential wage changes were driven by quality upgrading and not some other aspect of the macroeconomic crisis is provided by Kandilov (2005), who explicitly tests the predictions of Verhoogen (2004) in the context of an export subsidy in Chile. He finds that industries offered export subsidies displayed greater wage growth for white-collar workers than non-subsidized industries, and that

The results for the non-tradable sectors appear in Table 9. As in Table 8, Panel A reports results constraining  $\phi = 0$  in 12, and Panel B reports results where  $\phi$  is unrestricted. The results in Column 1 indicate that there was a significant difference in differential average wage changes between 1993-1997 and 1997-2001. The key point, however, is that this difference in differential changes is *not* attributable to differences in differential changes in wage premia. The results for changes in wage premia, in both Part A and Part B, indicate that we cannot reject the hypothesis of no differential change in wage premia in the non-tradables sector, in either period. The majority of the difference in differential changes in average wages can be attributed to changes in average skill of the workforce. Interestingly, the results for the differential changes in average workforce skill (Column 3) are similar to the results for manufacturing. One possible interpretation of this result is that the peso crisis threw many relatively high-skill workers into unemployment, and that large establishments generally were better-positioned than small establishments to attract these workers, independent of export status. The stayers results, calculated in the same way as in Table 8, reinforce the results for wage premia.

## 6.5 Alternative Specifications

Another possible objection to our approach is that the differential wage changes were indeed due to the increase in exporting brought about by the peso crisis, but that the wage changes reflect rent-sharing from increasingly profitable export operations, rather than a consequence of quality upgrading. It is difficult to investigate this hypothesis directly, since the measures of profitability that can be constructed from plant-level panels are accounting profits, which

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within subsidized industries initially larger and more productive plants saw greater wage increases for both white-collar and blue-collar workers than initially smaller or less productive plants.

are typically poor measures of economic profits, and since the employer-employee data and the plant-level data have not yet been linked. Given these constraints, one straightforward way to investigate the rent-sharing hypothesis is to control for the employment growth of establishments. Establishments presented with new profit-making opportunities typically expand; as a result, employment growth can be interpreted as a proxy, albeit imperfect, for the growth of profitability. Table 10 reports results for our basic model including the change in employment as an additional co-variate. Two patterns are notable. First, the change-in-scale term is associated with a significant *decrease* in the average person effect. This is consistent with the idea that establishments face a limited supply of skill in their local labor markets, and must hire lower-skill workers as they expand. Second, the change-in-scale term is associated with a significant increase in the plant effect. This is consistent with the rent-sharing hypothesis. It is also consistent with the hypothesis that plants are face internal equity constraints or institutional constraints that require them to pay low-productivity workers a wage above their marginal product; in this case, newly hired low-productivity workers may receive higher wage premia than existing workers. The first association is stronger than the second, hence the overall association of the change in scale with average log wages is negative. But the key point for our argument is that the results for plant effects are largely unaffected — indeed, they are strengthened — by the inclusion of the scale term. Comparing Tables 8 and 10, we see that the difference in differential changes in wage premia is larger when we control for scale effects than when we do not. We conclude that rent-sharing is unlikely to be the entire explanation for the greater differential change in wages and plant effect during 1993-1997.

The final objection that we consider is that the return to skill varies over time. Equation 12 imposes the assumption that the individual effects,  $\alpha_i$  are fixed, and that the coefficients on the time-varying observable variables are constant over time. But if high- $\lambda$  plants employ higher-skill workers than low- $\lambda$  plants and the return to skill rises over time, equation 12 will interpret the rising wages of high-skilled workers as an increase in wage premia in high- $\lambda$  plants.

To address this concern, we estimate a model of the following form, a modified version of equation 12:

$$w_{it} = (\alpha_i + x'_{i,t}\phi) \eta_t + \psi_{j(i,t)t} + u_{it} \quad (14)$$

where  $\eta_t$  can be interpreted as the general-equilibrium return to skill. There are at least two approaches one could take to estimating a model of this form. One approach would be to estimate the model with non-linear least squares, estimating  $\eta_t$  simultaneously with the other parameters. The large number of individual and plant-year effects makes this a computationally difficult task.<sup>21</sup> A second approach would be first to construct estimates of  $\eta_t$ , and then to use them to estimate the other parameters. In this paper, we follow the second approach.

Our procedure is the following:

1. We take data from the four years prior to the initial year of our period of interest (i.e. 1989-1992 if we are interested in 1993-1997) and estimate a model of the form of equation 12, to generate preliminary estimates of  $\alpha_i$  and  $x'_{i,t}\phi$ , call them  $\tilde{\alpha}_i$  and  $x'_{i,t}\tilde{\phi}$ .

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<sup>21</sup> We are pursuing a non-linear version of a Chamberlain (1982) correlated random effects model, where we linearly project the individual effects,  $\alpha_i$  on the value of the covariates from all periods before estimating. This approach has the advantage that we avoid estimating directly the million-plus individual effects for each year. We thank Steven Lich-Tyler for suggesting it.



2. We take data for stayers between two periods (call them  $t$  and  $t + 1$ ),<sup>22</sup> and estimate equation 14 in differences:

$$\Delta w_i = \left( \tilde{\alpha}_i + x'_{i,t} \tilde{\phi} \right) \Delta \eta + \psi_{j(i,t)} + \Delta u_i \quad (15)$$

This yields an estimate of  $\hat{\Delta} \eta$ . Normalizing  $\eta_t$  for an initial year (i.e. 1993) to 1, we can recover estimates  $\hat{\eta}_t$  for other years (i.e. 1997, 2001).

3. We substitute  $\hat{\eta}_t$  into equation 14. The model can then be estimated on the full sample by a quasi-differencing procedure as described in Wooldridge (2002, pp. 317-322). Stacking observations for a given individual and multiplying both sides of equation 14 by a quasi-differencing operator, we can write:

$$[I_T - P_\eta] w_i = [I_T - P_\eta] \left( \alpha_i + x'_{i,t} \phi \right) \eta_t + [I_T - P_\eta] \psi_{j(i,t)t} + [I_T - P_\eta] u_i \quad (16)$$

where  $P_\eta = \eta (\eta' \eta)^{-1} \eta'$  and  $\eta \equiv (\eta_1 \ \eta_2 \ \dots \ \eta_T)'$ . The individual-effect term,  $\alpha_i$  is differenced out, and the equation can be estimated by OLS, yielding estimates  $\hat{\alpha}_i$ ,  $\hat{\phi}$ , and  $\hat{\psi}_{jt}$ . We then construct an average workforce skill measure and estimate a model of the form of equation 11, as above.

The values we calculate for  $\eta_t$  in step 2 for 1993, 1997 and 2001 are 1, 0.94 and 0.87 respectively. Table 11 presents results for differential changes in plant-year effects and average skill from this procedure. The results are nearly identical to the results in Part B of Table 8, where the general-equilibrium return to skill is held constant. This approach must be accompanied by a number of caveats. The estimates  $\tilde{\alpha}_i$  and  $\tilde{\phi}$  in step 1 assume that  $\eta_t$  is

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<sup>22</sup> The advantage of focusing on stayers here is that we can capture the change in wages due to changing plant-year effects with a plant dummy.

held constant, and serial correlation in the measurement error in wages from the pre-period data used in step 1 to wages in the initial year of the period used in step 2 may lead to spurious underestimates of  $\eta_t$ . Nevertheless, the fact that the results of this procedure so closely resemble the results above reassures us that the difference in differential changes in wage premia is not being driven entirely by changes in the general-equilibrium return to skill.

## 7 Conclusion

In this paper, we have confirmed in a new dataset the result of Verhoogen (2004) that initially larger manufacturing plants increased wages relative to initially smaller plants within each industry during the peso crisis period, and that this differential change was greater than in a subsequent period without a devaluation. We have shown further that, under the assumption of conditional random mobility, 67%-89% of the difference in differential wage changes can be attributed to changes in wage premia, depending on the specification we use. This pattern holds in manufacturing but not in a set of non-tradable sectors, which supports our argument that the changes in average wages and wage premia were driven by quality upgrading due to increasing exports. Our results are robust to a variety of modifications of our econometric model, addressing possible concerns with our approach. We interpret the evidence presented in this paper as providing reasonably strong support for the hypothesis that quality upgrading in response to the late-1994 peso crisis led Mexican manufacturing plants to increase wage premia.

It remains an open question precisely why higher product quality requires higher wage premia. One possibility is that high-quality production requires particularly motivated workers, and that plants pay higher wages to increase motivation, along the lines of the efficiency-

wage models of Bowles (1985); Shapiro and Stiglitz (1984) or Akerlof (1982). A second possibility is that quality upgrading leads to investments by plants or by workers in plant-specific skills, and that the gains to such investments are subsequently shared with workers (Hashimoto, 1981). A third possibility is that quality upgrading leads plants or workers to invest in *general* training, and then plants raise wages to prevent workers from leaving for other establishments. This third possibility may be distinguished from the first two by examining what happens to workers who have been employed by upgrading plants after they leave the plants. The first two possibilities may be distinguishable from one another using information on training programs in the ENESTyC survey. We leave the resolution of this question to future work.

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## A Data Appendix

Several aspects of our data cleaning procedure for the IMSS employer-employee data merit explanation.

### A.1 Minimum Wages

Mexico has three regional minimum wages, with a higher minimum in Mexico City than in outlying areas. Prior to 1990, it was common for establishments to report wages below the corresponding regional minimum wage. In 1990, IMSS initiated a campaign to require establishments to report at least the minimum wage (whether or not they were in fact paying the minimum wage), and many establishments appear to have complied. A small fraction (on the order of 1/10 of 1%) continued report wages lower than the applicable minimum wage. In all years, we replace values below the minimum wage with the corresponding minimum wage.

### A.2 Top Codes and Outliers

The wage data were top-coded over the period we study, at values ranging from 10 to 25 times the Mexico City minimum wage. Nonetheless, a number of individuals have reported wages above the top-code. In an effort to minimize the effects of the changing top-code and the outliers, we “winsorized” our data, setting wage values above the 95th percentile for a given year equal to the value at the 95th percentile, and values below the 5th percentile to the value at the 5th percentile. The 95th percentile was below the top-code in all years. In several years, more than 5% of individuals in our data were receiving the lowest minimum wage.

### A.3 Geographical Selection

We dropped establishments in municipalities along the Northern border with the U.S. The vast majority of such plants are maquiladoras and since maquiladoras are not included in the EIA panels, for reasons of comparability we also dropped them from the IMSS panels.

**Table 1: Summary Statistics, IMSS Panels**

	Manufacturing panel			Non-tradables panel		
	emp. <200 (1)	emp. >=200 (2)	all (3)	emp. <200 (4)	emp. >=200 (5)	all (6)
average daily wage (1994 pesos)	43.59 (0.045)	54.91 (0.025)	52.78 (0.022)	53.26 (0.060)	61.75 (0.039)	59.74 (0.033)
average daily wage (current US \$)	13.70 (0.014)	17.43 (0.008)	16.72 (0.007)	16.77 (0.019)	19.57 (0.013)	18.91 (0.011)
Fraction male	0.71 (0.001)	0.72 (0.000)	0.72 (0.000)	0.72 (0.001)	0.63 (0.000)	0.65 (0.000)
Age	32.0 (0.013)	31.4 (0.006)	31.5 (0.005)	32.7 (0.012)	34.6 (0.007)	34.1 (0.006)
% stayers (previous 4 years)	0.43 (0.0007)	0.44 (0.0003)	0.44 (0.0003)	0.39 (0.0007)	0.41 (0.0004)	0.41 (0.0003)
N establishments	2425	1203	3628	2800	879	3679

Notes: Panels include years 1993, 1997, 2001. Standard errors of means in parentheses. Stayers calculated as (# employees continuously in establishment over previous 4 years)/(current employment); only available in 1997, 2001.

**Table 2: Summary Statistics by Export Status, EIA Panel, 1993**

	Non-exporters	Exporters	All
A. EIA Panel			
Employment	183.3 (5.0)	335.0 (12.2)	227.5 (5.1)
Revenues	43.1 (1.8)	91.1 (4.8)	57.1 (1.9)
Domestic sales	41.2 (1.7)	71.6 (3.7)	50.1 (1.6)
K/L ratio	42.8 (1.4)	55.6 (2.6)	46.6 (1.3)
Avg. hourly wage	12.1 (0.1)	16.5 (0.3)	13.4 (0.1)
white-collar % of employment	30.4 (0.4)	33.5 (0.5)	31.3 (0.3)
export % of sales		15.4 (0.6)	
% with foreign ownership (1994)	8.8 (0.6)	29.7 (1.5)	14.9 (0.6)
TFP	-0.09 (0.02)	0.21 (0.03)	0.00 (0.02)
N	2306	949	3255
B. EIA-ENESTyC Panel (1994)			
Has ISO 9000	8.6 (0.8)	14.1 (1.4)	10.5 (0.7)
Has formal training program	65.6 (1.4)	75.8 (1.7)	69.2 (1.1)
N	1166	625	1791

Notes: Standard errors (of means) in parentheses. Exporter defined as having exports greater than zero. Having foreign ownership means at least 10% FDI. Data on foreign ownership, ISO 9000, training from 1994; all other data from 1993. 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. TFP is estimate of firm fixed effect in regression of log value-added on log inputs for 1993-94, separately for 9 two-digit industries; estimates then deviated from 6-digit industry means.

**Table 3: Proxies for  $\lambda$ , 1993, EIA Panel**

	log (emp- loyment)	log (dom- estic sales)	exp. share index	ISO 9000 index	1st princ. comp.	log tfp (fixed effect)
log (employment)	1					
log (domestic sales)	0.8014*	1				
export share index	0.8719*	0.8322*	1			
ISO 9000 index	0.6987*	0.6090*	0.8049*	1		
1st princ. component	0.9020*	0.8789*	0.9660*	0.8398*	1	
TFP (fixed effect)	0.3326*	0.4754*	0.3875*	0.2720*	0.4123*	1

Notes: All variables deviated from industry means. Asterisk indicates significance at the 99% level. Export share index is Xb from tobit of export share on log employment, log revenues, log K/L ratio, foreign ownership indicator. ISO 9000 index is Xb from probit of ISO 9000 on log employment, log revenues, log K/L ratio, foreign ownership indicator, and export share of sales. First principal component is of the following variables: log employment, log revenues, log K/L ratio, foreign ownership indicator, and export share of sales. TFP is estimate of firm fixed effect in regression of log value-added on log inputs for 1993-94, separately for 9 two-digit industries; estimates then deviated from 6-digit industry means. \* = 1% level.

**Table 4: Results on Export Shares, EIA Panel**

	Dependent variable: $\Delta$ export share of sales					
	OLS		IV			
	1993-1997 (1)	1997-2001 (2)	Difference (3)	1994-1997 (3)	1998-2001 (4)	Difference (5)
Proxy for $\lambda$ :						
log dom. sales	0.019** [0.002]	0.006** [0.002]	0.013** [0.003]	0.014** [0.002]	0.004** [0.001]	0.010** [0.002]
export share index	0.017** [0.002]	0.003 [0.002]	0.014** [0.003]	0.013** [0.002]	0.003 [0.002]	0.010** [0.003]
ISO 9000 index	0.008** [0.002]	-0.004 [0.002]	0.012** [0.003]	0.007** [0.002]	-0.002 [0.002]	0.009** [0.003]
1st prin. component	0.013** [0.002]	-0.002 [0.002]	0.015** [0.003]	0.011** [0.002]	-0.001 [0.002]	0.012** [0.003]
TFP (fixed effect)	0.009** [0.002]	0.002 [0.002]	0.007* [0.003]			
log employment	0.016** [0.003]	0.004 [0.002]	0.012** [0.004]	0.013** [0.003]	0.003 [0.002]	0.010** [0.004]

Notes: Each coefficient reflects a separate regression. Robust standard errors in brackets. R-squareds omitted (all are between .10 and .17). All regressions include industry and state fixed effects (coefficients omitted). N = 3255. IV uses initial year value (1993 or 1997) as instrument for subsequent value (1994 or 1998) of proxy, and takes change over three years (1994-1997, 1998-2001) as outcome variable. See text or notes to Table 3 for details on proxies. \*\* = 1% level, \* = 5% level.

**Table 5: Results on ISO 9000 Certification, EIA-ENESTyC panel**

Proxy for $\lambda$ :	Dependent variable: $\Delta$ ISO 9000					
	OLS			Censored Normal Regression		
	1994-1998 (1)	1998-2000 (2)	Difference (3)	1994-1998 (4)	1998-2000 (5)	Difference (6)
log domestic sales	0.064** [ 0.021]	-0.078** [ 0.017]	0.142** [0.027]	0.216** [ 0.038]	0.038** [ 0.013]	0.178** [0.041]
export share index	0.032 [0.023]	-0.087** [0.020]	0.119** [0.030]	0.211** [ 0.039]	0.052** [ 0.014]	0.159** [0.041]
ISO 9000 index	0.012 [ 0.023]	-0.107** [ 0.018]	0.119** [0.029]	0.178** [ 0.037]	0.030* [ 0.013]	0.208** [0.039]
1st prin. component	0.048* [0.024]	-0.095** [0.020]	0.143** [0.031]	0.253** [ 0.040]	0.044** [ 0.014]	0.209** [0.042]
TFP (fixed effect)	0.026 [ 0.026]	-0.078** [ 0.021]	0.104** [0.033]	0.070* [ 0.034]	0.016 [ 0.013]	0.056 [0.036]
log employment	0.044 [ 0.031]	-0.089** [ 0.024]	0.143** [0.039]	0.248** [ 0.054]	0.032 [ 0.018]	0.216** [0.057]

Notes: Each coefficient reflects a separate regression. Each regression includes industry and state fixed effects (coefficients omitted). N = 850. Robust standard errors in brackets. R-squareds omitted. \*\* = 1% level, \* = 5% level.

**Table 6: Results on Plant-Level Average Wages, EIA Panel**

Proxy for $\lambda$ :	Dependent variable: $\Delta$ avg. log hourly wage					
	OLS			IV		
	1993-1997 (1)	1997-2001 (2)	Difference (3)	1994-1997 (4)	1998-2001 (5)	Difference (6)
log dom. sales	0.044** [0.006]	0.012** [0.004]	0.032** [0.007]	0.041** [0.005]	0.009* [0.004]	0.032** [0.006]
export share index	0.051** [0.006]	0.023** [0.005]	0.029** [0.008]	0.049** [0.005]	0.019** [0.005]	0.030** [0.007]
ISO 9000 index	0.035** [0.006]	0.006 [0.006]	0.029** [0.008]	0.034** [0.006]	0.001 [0.005]	0.033** [0.008]
1st prin. component	0.048** [0.006]	0.017** [0.005]	0.031** [0.008]	0.046** [0.005]	0.013** [0.005]	0.033** [0.007]
TFP (fixed effect)	0.016** [0.006]	0.002 [0.005]	0.014 [0.008]			
log employment	0.070** [0.007]	0.024** [0.006]	0.046** [0.009]	0.058** [0.007]	0.020** [0.005]	0.038** [0.009]

Notes: Each coefficient reflects a separate regression. Robust standard errors in brackets. R-squareds omitted (all are between .10 and .17). All regressions include industry and state fixed effects (coefficients omitted). N = 3255. IV uses initial year value (1993 or 1997) as instrument for subsequent value (1994 or 1998) of proxy, and takes change over three years (1994-1997, 1998-2001) as outcome variable. See text or notes to Table 3 for details on proxies. \*\* = 1% level, \* = 5% level.

**Table 7: Comparing EIA and IMSS Manufacturing Panels**

		Dependent variable:	
		OLS (1)	IV (2)
A. EIA 1993-2001 panel			
1993-1997	log employment 1993	0.070** [0.007]	0.058** [0.007]
		0.156	0.154
1997-2001	log employment 1997	0.024** [0.006]	0.020** [0.005]
		0.11	0.112
	Difference	0.046** [0.009]	0.038** [0.009]
B. IMSS 1993-2001 panel			
1993-1997	log employment 1993	0.056** [0.004]	0.042** [0.004]
		0.185	0.189
1997-2001	log employment 1997	0.022** [0.004]	0.015** [0.003]
		0.117	0.105
	Difference	0.034** [0.006]	0.027** [0.005]

Notes: All regressions include industry and state fixed effects (coefficients omitted). N = 3255 in part A, N = 3628 in part B. Robust standard errors in brackets. R-squareds below standard errors. IV uses initial year value (1993 or 1997) as instrument for subsequent value (1994 or 1998) of proxy, and takes change over three years (1994-1997, 1998-2001) as outcome variable. \*\* = 1% level, \* = 5% level.



**Table 8: Baseline Results, IMSS Manufacturing Panel**

		Dependent variable:			
		$\Delta$ avg. log(wage)	$\Delta$ plant effect	$\Delta$ avg. skill	$\Delta$ avg. log(wage) of stayers
		(1)	(2)	(3)	(4)
A. Excluding time-varying observables					
1993-1997	log employment 1993	0.055** [0.004]	0.034** [0.004]	0.021** [0.003]	0.031** [0.004]
		0.197	0.179	0.164	0.174
1997-2001	log employment 1997	0.019** [0.004]	0.010** [0.003]	0.009** [0.003]	0.010** [0.003]
		0.124	0.143	0.091	0.147
Difference		0.036** [0.006]	0.024** [0.005]	0.012** [0.004]	0.021** [0.005]
B. Including time-varying observables					
1993-1997	log employment 1993	0.055** [0.004]	0.033** [0.004]	0.022** [0.003]	0.026** [0.004]
		0.197	0.186	0.183	0.2201
1997-2001	log employment 1997	0.019** [0.004]	0.001 [0.003]	0.018** [0.003]	-0.004 [0.003]
		0.124	0.135	0.098	0.2271
Difference		0.036** [0.006]	0.032** [0.005]	0.004 [0.004]	0.030** [0.005]

Notes: All regressions include industry and state fixed effects (coefficients omitted). N = 3628. Robust standard errors in brackets. R-squareds below standard errors. Part B, Column (4) includes plant-level averages of tenure squared, age squared, and experience squared (coefficients omitted). \*\* = 1% level, \* = 5% level.

**Table 9: Baseline Results, IMSS Non-Tradables Panel**

		Dependent variable:			
		$\Delta$ avg. log(wage)	$\Delta$ plant effect	$\Delta$ avg. skill	$\Delta$ avg. log(wage) of stayers
		(1)	(2)	(3)	(4)
A. Excluding time-varying observables					
1993-1997	log employment 1993	0.014** [0.004]	-0.002 [0.004]	0.016** [0.003]	-0.005 [0.004]
		0.156	0.157	0.106	0.162
1997-2001	log employment 1997	-0.002 [0.006]	-0.007 [0.006]	0.005 [0.003]	-0.006 [0.006]
		0.100	0.128	0.118	0.137
Difference		0.017* [0.007]	0.005 [0.007]	0.012** [0.004]	0.001 [0.007]
B. Including time-varying observables					
1993-1997	log employment 1993	0.014** [0.004]	0.000 [0.004]	0.014** [0.003]	0.001 [0.004]
		0.156	0.152	0.082	0.2068
1997-2001	log employment 1997	-0.002 [0.006]	-0.006 [0.006]	0.004 [0.003]	-0.002 [0.004]
		0.100	0.117	0.100	0.2405
Difference		0.017* [0.007]	0.006 [0.007]	0.010 [0.004]	0.003 [0.006]

Notes: Non-tradables include construction, retail trade, transportation, and services. All regressions include industry and state fixed effects (coefficients omitted). N = 3679. Robust standard errors in brackets. R-squareds below standard errors. Part B, Column (4) includes plant-level averages of tenure squared, age squared, and experience squared (coefficients omitted). \*\* = 1% level, \* = 5% level.

**Table 10: Controlling for Change in Scale, IMSS Manufacturing Panel**

		Dependent variable:		
		$\Delta$ avg. log(wage)	$\Delta$ plant effect	$\Delta$ avg. skill
		(1)	(2)	(3)
1993-1997	log employment 1993	0.043** [0.004]	0.043** [0.004]	0.000 [0.002]
	$\Delta$ log employment	-0.086** [0.010]	0.069** [0.008]	-0.154** [0.007]
	R-squared	0.225	0.208	0.388
1997-2001	log employment 1997	0.014** [0.003]	0.006 [0.003]	0.008** [0.002]
	$\Delta$ log employment	-0.072** [0.011]	0.066** [0.007]	-0.138** [0.009]
	R-squared	0.15	0.162	0.292
Difference		0.030** [0.005]	0.037** [0.005]	-0.008** [0.003]

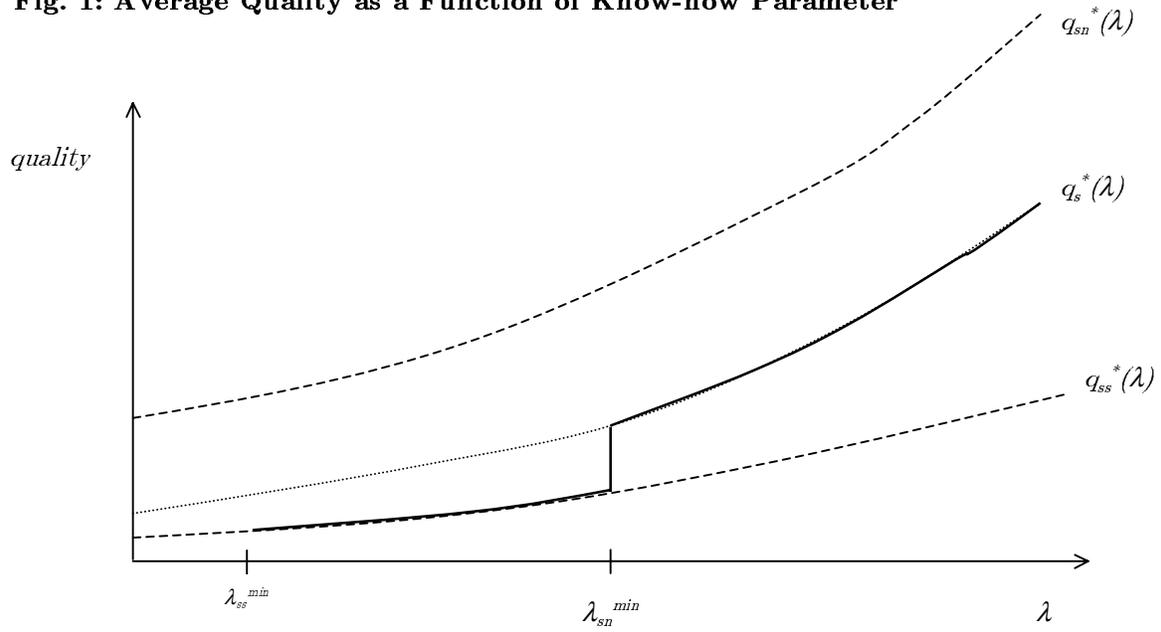
Notes: All regressions include industry and state fixed effects (coefficients omitted). N = 3628. Robust standard errors in brackets. \*\* = 1% level, \* = 5% level.

**Table 11: Allowing for Changing Return to Skill, IMSS Manufacturing Panel**

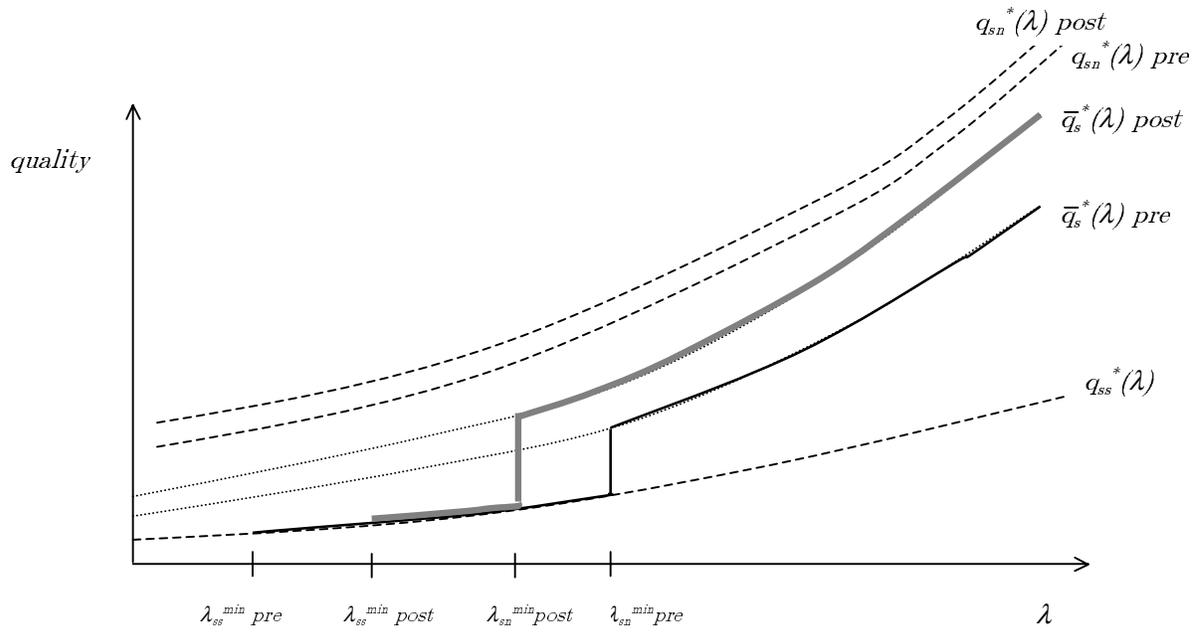
	$\Delta$ avg. log(wage)	$\Delta$ plant effect	$\Delta$ avg. skill
	(1)	(2)	(3)
1993-1997			
log employment 1993	0.055** [0.004]	0.034** [0.004]	0.021** [0.003]
	0.197	0.194	0.176
1997-2001			
log employment 1997	0.019** [0.004]	0.003 [0.003]	0.016** [0.003]
	0.124	0.141	0.121
Difference	0.036** [0.006]	0.031** [0.005]	0.005 [0.004]

Notes: All regressions include industry and state fixed effects (coefficients omitted). N = 3628. Robust standard errors in brackets. R-squareds below standard errors. \*\* significant at 1% level, \* at 5% level.

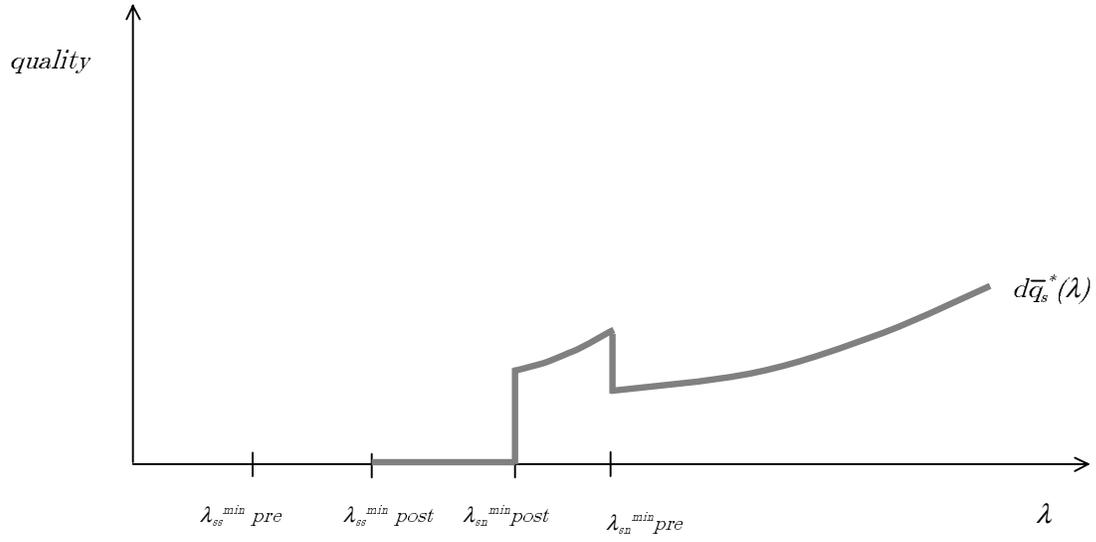
**Fig. 1: Average Quality as a Function of Know-how Parameter**



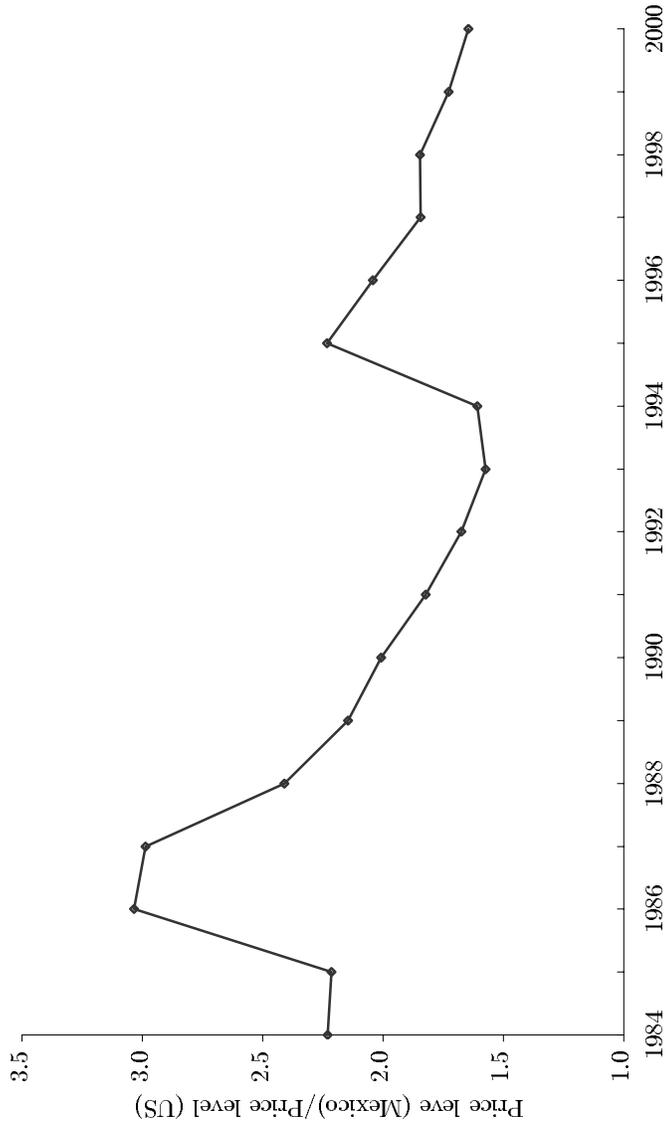
**Fig. 2: Change in Average Quality as a Function of Know-how Parameter**



**Fig. 3: Change in Average Quality as a Function of Know-how Parameter**

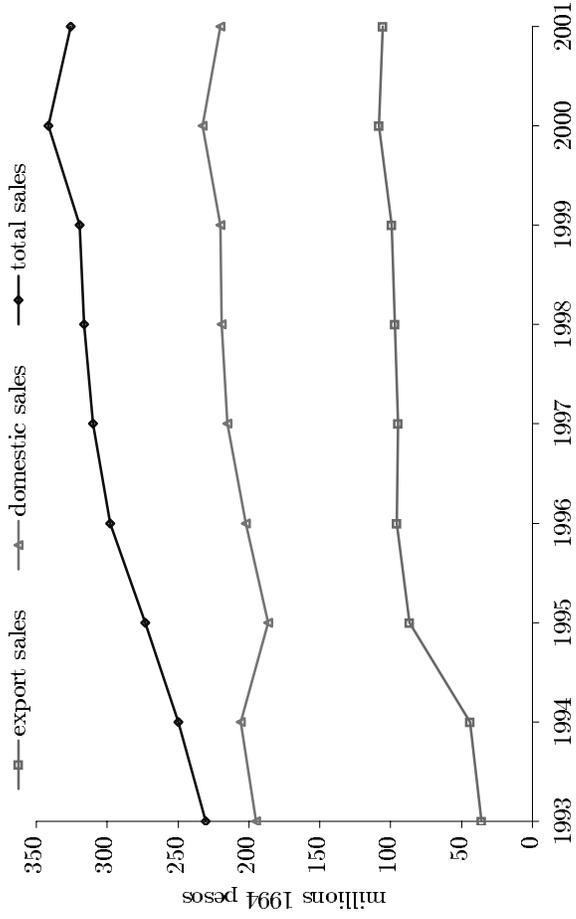


**Fig. 4: Real Exchange Rate**

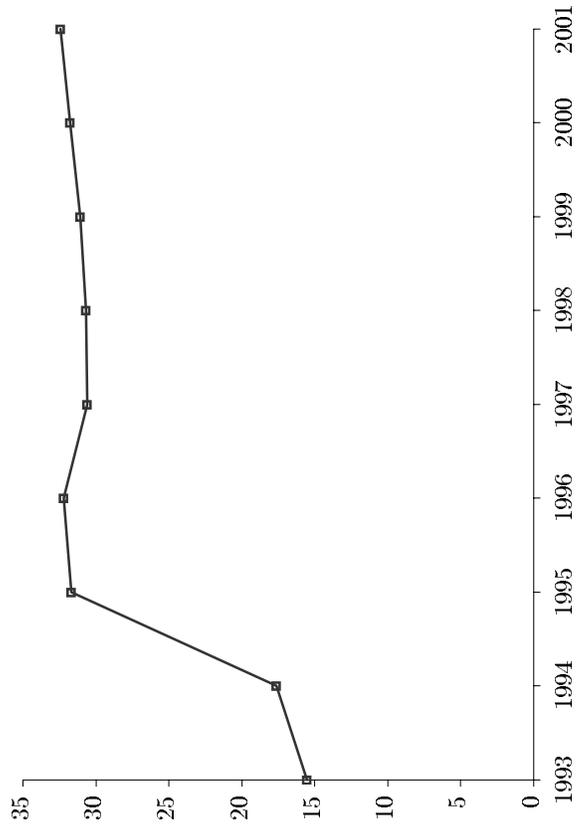


Note: Price levels and exchange rate from Penn World Table 6.1.

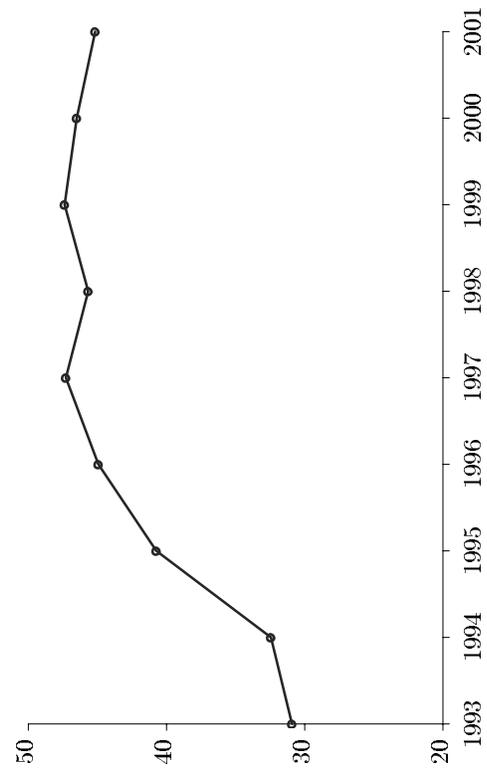
**Fig. 5a: Sales, Domestic and Export, 1993-2001**



**Fig. 5b: Export percentage of total sales**



**Fig. 5c: Percentage of Plants Exporting**



Notes: Data from EIA 1993-2001 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.