

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

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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 significantly larger than the differential changes either in manufacturing in 1997-2001 or in non-tradable sectors in the peso-crisis period. The matched employer-employee nature of the data allows us to distinguish wage changes due to changes in the composition of the workforce from changes due to changes in plant-level wage premia. We estimate plant-year effects in a person-level wage regression, controlling for person fixed effects, and then relate within-plant changes in those plant-year effects to initial size in different periods. Our baseline estimates suggest that approximately two-thirds of the difference in differential wage changes between the peso-crisis period and the later period can be attributed to changes in plant-year effects. These results support the argument that product quality upgrading led Mexican manufacturers to raise wage premia.

1 Introduction*

Labor economists have long noted that some firms pay more to observably similar workers than others: profitable industries pay more than highly competitive industries (Dickens and Katz, 1987; Krueger and Summers, 1986), large firms pay more than small ones (Brown and Medoff, 1989). A central issue in this literature is whether such differentials are due to differences in workers' unobserved skill 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 have found (under a particular set of assumptions) that approximately half of inter-industry wage differentials and 70 percent of firm size-wage differentials in France and the U.S. are attributable to differences in firm-level wage policies (Abowd, Kramarz, and Margolis, 1999; Abowd, Finer, and Kramarz, 1999; Abowd and Kramarz, 1999a). But the employer-employee literature has so far focused more on estimating the size of wage premia than on identifying their determinants.

In this paper, we investigate the role of one possible determinant of wage premia: product quality. Our contribution is to combine the analysis of matched employer-employee data with a research design that identifies arguably exogenous variation in the incentive of Mexican plants to undertake quality upgrading. Following Verhoogen (2004), we argue that the peso crisis of late 1994 generated a differential inducement to export and upgrade within 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

* We would like to thank David Card, participants of the 4th annual UT/ITAM conference, and a number of seminar participants for helpful comments. All errors are ours.

smaller and less productive establishments remaining focused on the domestic market. For the reason hypothesized by Kremer (1993) — that high-quality products are particularly sensitive to mistakes, and therefore require careful workers — this upgrading of product quality requires an upgrading of the workforce as well. The upgrading of the workforce may plausibly take two different forms: 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 mechanisms cannot be distinguished in conventional plant-level datasets which typically report only average wages by occupational category. But they are distinguishable in employer-employee data, since only the second mechanism would have the effect of raising wage premia.

We draw on evidence from the administrative wage records of the Instituto Mexicano del Seguro Social (IMSS), the Mexican social security agency, to address this issue.¹ The data (described in Section 3) contain precious little information on each worker and almost no information on establishment characteristics.² But they have the great 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.

Our econometric approach proceeds in three steps. First, we verify in our data the patterns of changes in plant-level average wages observed in Mexican plant-level datasets

¹ The data have been obtained and cleaned by David Kaplan, and used in several previous papers (Kaplan, Martinez Gonzalez, and Robertson, 2003; Castellanos, Garcia-Verdu, and Kaplan, 2004; Kaplan, Martinez Gonzalez, and Robertson, 2004, 2005).

² The data are potentially linkable to the detailed establishment-level surveys conducted by the Instituto Nacional de Estadísticas, Geografía e Información (INEGI), the Mexican government statistical agency. We have applied for permission to link these datasets, and are hopeful that we will be able to do so in the near future.

(Section 4). Second, we estimate person fixed effects and plant-year fixed effects in a person-level wage regression. Third, we relate within-plant *changes* in plant-year effects and average person effects to initial plant employment *levels* in different periods and in different sectors (Section 5). As a further test, we relate changes in the average wages of workers continuously employed at their plants (whom we call stayers) to initial employment; in the absence of non-random selection out of plants, these changes should resemble changes in plant-year effects.

Our main results are as follows. We confirm the finding of Verhoogen (2004) that initially larger manufacturing plants experienced greater wage growth than initially smaller plants within the same industry in 1993-1997 and that this change was significantly larger than the differential change in manufacturing in the 1997-2001 period. We find that the differential changes in both plant-year effects and average person effects were larger in 1993-1997 than the corresponding changes in 1997-2001, but that the effect is twice as large, in percentage terms, for the plant-year effects. That is, approximately two-thirds of the difference in differential wage changes between the peso-crisis period and the later period can be attributed to changes in establishment wage policies. The pattern of changes for stayers are quite similar to those for plant-year effects. The social-security data also include non-tradable sectors, and we show that outside of manufacturing the differential wage changes are almost entirely accounted for by changes in average person effects. These results provide strong support for the hypothesis that upgrading plants raised wage premia, above and beyond the compensation for fixed characteristics of individual workers.

Our paper is related to two distinct literatures. It is related to the literature using employer-employee data in developed countries, especially France and the U.S., mentioned

above and surveyed in Abowd and Kramarz (1999b). It is also to a small theoretical literature on labor effort, wages and international trade (Copeland, 1989; Brecher, 1992; Matusz, 1996; Leamer, 1999). To our knowledge, ours is the first paper to link effort explicitly to product quality — and to shocks to the incentive to produce quality — in a trade model, and the first to test the implications of such a model empirically.

2 Theory

This section develops a model of trade and quality upgrading to organize and motivate our empirical work.³ A number of caveats apply: the model is partial-equilibrium, static, with special functional forms. Our goal is to spell out the logic of the hypotheses we will test empirically, not to present a fully specific general-equilibrium model.

There are two countries, North (n) and South (s). We 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. We 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. We thus potentially have two sets of optimization equations for each firm in each country, one for each production line. 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)

³ The model is closely related to the model presented in Verhoogen (2004), and readers are referred to the earlier paper for a somewhat less terse exposition.

productivity parameter λ , which can be thought of as representing entrepreneurial ability, with distribution $g_c(\lambda)$. Each producer is constrained to produce just one variety for each market. 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; as a result, λ will uniquely identify goods within the set of goods produced by firms in a particular location for a particular destination market, as well as entrepreneurs. To keep the notation simple, we use λ to index goods from the outset.

Assume that aggregate demand for differentiated good produced by entrepreneur with ability λ 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. θ_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 λ in country c for market d . $p_{cd}(\lambda)$ is the price of the good, *denominated in terms of goods in country c* . δ_{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 λ 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 . Λ_{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.⁴ The key points of this specification are that the demand for a given good is decreasing in price and increasing in quality, and that the extent to which it is increasing in quality depends on the willingness of consumers to pay for quality. The number of consumers and the real exchange rate will become important below in the modeling of the exchange-rate devaluation.

Production on each production line is assumed to be governed by a modified O-ring-type production function (Kremer, 1993). We assume that the quantity and quality of output are governed by separate production functions. Production of one unit of output is assumed to require one worker. Production of quality is assumed to depend on three different inputs: entrepreneurial ability, λ , human capital, h , and effort, e , in Cobb-Douglas fashion:⁵

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

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. Note that this specification nests two special models: $\beta_e = 0$ yields a model in which quality depends only

⁴ See Verhoogen (2004) for details.

⁵ Our model thus differs from the model of Verhoogen (2004) which had two types of workers and physical capital, but did not distinguish between human capital and effort.

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.⁶ 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 in both countries, although this assumption could easily be relaxed.

We assume that effort is a function of the gap between what the workers receives in the firm and her market wage, either because workers perceive the rent to be a gift and choose to reciprocate as in Akerlof (1982) or because the rent raises the cost of job loss and induces workers to shirk less to avoid being fired as in Bowles (1985) or Shapiro and Stiglitz (1984).⁷ Let v represent the gap, or quasi-rent, where $v \equiv w_{cd} - \tilde{w}_{cd}$. For simplicity we assume that effort 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.

We assume that the firm chooses the price of its goods (p), the level of human capital

⁶ Here we assume that skill is general. An alternative possibility is that skill is partly job-specific. We return to this question below.

⁷ The key distinction between human capital and effort in our model is that the former is perfectly observable and hence can be contracted upon, while the latter is not. If effort is observable and contractible as in Leamer (1999), then it is best interpreted as part of h_{cd} in our model.

of each worker employed in the firm (h), and the wage (w), which in turn determines the worker's choice of effort. The marginal cost of producing one unit of output is simply the wage, w , and this marginal cost does not depend on the total amount of output. There are two (potential) fixed costs: one to set up a plant and enter the domestic market, f , and another to enter the export market, f_x . Firms have to pay to set up their plant in their home country before they can export. Thus by assumption there are no firms that enter only the export market. It will be convenient to write these fixed costs as f_{cd} where $f_{cd} = f + I_{c=d}f_x$, where $I_{c=d}$ takes the value 1 if $c = d$ and 0 if not.

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 profitability of a plant with entrepreneurial ability λ producing in country c for market d is given by:

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

As is standard in monopolistic competition models, we 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 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 $\eta \equiv (A)^{\frac{1}{1-\beta}} (z_h\beta_h)^{\frac{\beta_h}{1-\beta}} (z_e\beta_e)^{\frac{\beta_e}{1-\beta}}$. Note that the optimal quality of good produced is increasing in both λ and θ . 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 levels of effort, e , and quasi-rents, v , 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 λ and θ . 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:

$$\begin{aligned} \tilde{w}_{cd}(\lambda) &= \frac{\beta_h}{\beta} w_{cd}(\lambda) \\ v_{cd}(\lambda) &= \frac{\beta_e}{\beta} w_{cd}(\lambda) \end{aligned} \quad (9)$$

In terms of the optimal quality choice, output is given by:

$$x_{cd}(\lambda) = \frac{N_d}{D_d} \exp \left\{ \frac{\theta_d (1 - \beta) q_{cd}(\lambda)}{\mu} - 1 \right\} \quad (10)$$

Output and hence employment are increasing in λ for each location-destination pair, since $q_{cd}(\lambda)$ is increasing in λ . They increase more steeply for the Northern consumer market, since θ_d influences both the slope of $q_{cd}(\lambda)$ with respect to λ and the slope of $x_{cd}(\lambda)$ with respect to $q_{cd}(\lambda)$. As is standard in logit models, the mark-up is constant: $p_{cd}(\lambda) - w_{cd}(\lambda) = \mu \delta_{cd}$.

Profitability at the optimum is given by:

$$\pi_{cd}(\lambda) = \mu\delta_{cd}x_{cd}(\lambda) - f_{cd} \quad (11)$$

Profitability is thus increasing in λ as well.

The fact that profitability is increasing in λ implies that for each location-destination pair there is a single cut-off value of productivity, call it λ_{cd}^{\min} , above which all firms enter and earn positive profits, and below which no firms enter. The cut-off is defined implicitly by the requirement that the marginal firm have zero profits. There are four such cutoffs, one for each location-destination pair. At a given point in time, the set of all potential entrepreneurs is divided into three subsets: (1) *non-entrants*, potential plants do not enter either market; (2) *non-exporters*, plants that produce only for the domestic market; and (3) *exporters*, plants that produce for both markets.

Consider the export share of output for these groups. Hereafter to simplify the exposition we focus on Southern producers, but the analysis for Northern producers is analogous. Define the export share to be: $\chi_s(\lambda) \equiv \frac{x_{sn}(\lambda)}{x_{sn}(\lambda) + x_{ss}(\lambda)}$. For non-entrants, output is undefined. For non-exporters, $\chi_s(\lambda) = 0$. For exporters, we saw in equation (10) that output on each production line is increasing in λ and that output on the export line increases more steeply in λ . Consequently, the export share of output is also increasing in λ . There is a discontinuity in $\chi_s(\lambda)$ at the cut-off for entry into the export market.

Equations (7) and (8) give the values of quality, wages, human capital, effort and quasi-rents on each production line, but typically such variables are observed only at the plant level. The quantities corresponding to observable variables in standard datasets correspond to weighted average of the production-line values, and we construct such averages here. The

plant-level average product quality is:

$$\bar{q}_s(\lambda) = \chi_s(\lambda)q_{sn}(\lambda) + (1 - \chi_s(\lambda))q_{ss}(\lambda) \quad (12)$$

Figure 1 summarizes the relationship between $\bar{q}_s(\lambda)$ 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 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 . Let $\bar{h}_s(\lambda)$, $\bar{w}_s(\lambda)$, $\bar{e}_s(\lambda)$, and $\bar{v}_s(\lambda)$ be the plant-level averages for human capital, wages, effort and quasi-rents, defined as weighted averages on the two production lines like $\bar{q}_s(\lambda)$. Each of these averages follows a pattern similar to that of average product quality.

We model an exchange-rate shock as having two effects, both of which we take as exogenous. First, the real exchange rate, δ_{sn} , defined as the Northern price level over the Southern price level, rises. Second, the shock reduces total consumer demand in South, represented by a reduction in the number of Southern consumers, N_s . To reduce algebra, we 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. We continue to focus on Southern plants.

Appendix A shows that as a consequence of these two effects, more Southern firms enter the export market (λ_{sn}^{\min} falls). The net effect on entry of Southern firms into the Southern market is ambiguous, but in actual fact the number of bankruptcies of manufacturing firms in Mexico rose sharply during the peso crisis; it appears that the empirically relevant case is the one in which λ_{ss}^{\min} rises, and we focus on this case hereafter. Appendix B gives a precise statement of the condition required to obtain it.

The continuum of potential entrepreneurs in South can now be divided into five categories based on the pre-crisis and post-crisis cutoffs: (1) *never entrants* ($\lambda < \lambda_{ss,pre}^{\min}$) do not enter either market in either period; (2) *exiters* ($\lambda_{ss,pre}^{\min} \leq \lambda < \lambda_{ss,post}^{\min}$) 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.

Consider the change in the export share. For the always non-exporters: $d\chi_s(\lambda) = 0$. For the switchers into exporting, 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, Appendix C shows the change in export share of output is given by: $d\chi_s = \Sigma_s \chi_{s,pre} [1 - \chi_{s,pre}]$ 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 $\chi_s(\lambda) < 1/2$. Verhoogen (2004) reports that among the minority of Mexican plants that export, the average fraction of sales derived from exports is 20-25%.⁸ It thus appears reasonable to focus on the case where $\chi_s < 1/2$, and we do so hereafter.

Now consider the change in plant-level average quality, under the counterfactual that all

⁸ Exports made up more than half of total sales in only 3-7% of plants over the 1993-2001 period. 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.

plants enter both markets:

$$d\bar{q}_s(\lambda) = \chi_{s,pre}(\lambda)dq_{sn}(\lambda) + (q_{sn}(\lambda) - q_{ss}(\lambda))d\chi_s(\lambda) \quad (13)$$

From equation (7), we have that $dq_{sn}(\lambda) = \frac{\beta q_{sn}(\lambda)}{(1-\beta)\delta_{sn}}$ which, like $q_{sn}(\lambda)$, is increasing in λ . We also know from equation (7) that $q_{sn} - q_{ss}$ for all λ , conditional on entry into both markets. We have seen above that both χ_s and $d\chi_s$ are positive and increasing in λ ; hence $d\bar{q}_s(\lambda)$ is positive and increasing in λ as well. Note, however, that not all plants enter both markets. For the always non-exporters, $d\bar{q}_s(\lambda) = 0$. For the switchers into exporting, $d\bar{q}_s(\lambda) = (q_{sn}(\lambda) - q_{ss}(\lambda))\chi_{s,post}(\lambda)$.

Figure 2 summarizes the effect of the exchange rate shock on the level average quality. The $q_{sn}(\lambda)$ curve shifts up in response to the shock. 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 and represents the actual average quality curve pre-crisis. 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 λ . The change in 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 discontinuities in the export share of output.

The expressions for changes in the plant-level average values of the average wage, the

average market wage, and the average quasi-rent can be written:

$$d\overline{w}_s(\lambda) = \theta_n \delta_{sn} \beta d\overline{q}_s(\lambda) + (\theta_n \delta_{sn} - \theta_s \delta_{ss}) \beta q_{ss} d\chi_s \quad (14)$$

$$d\widetilde{w}_s(\lambda) = \frac{\beta_h}{\beta} d\overline{w}_s(\lambda) \quad (15)$$

$$d\overline{v}_s(\lambda) = \frac{\beta_e}{\beta} d\overline{w}_s(\lambda) \quad (16)$$

There are two key implications of these equations. First, since the change in the plant-average wage, $d\overline{w}_s(\lambda)$, is a linear combination of the changes in average quality and the export share, we know that the change in the average wage — and hence the changes in the “market wage,” the quasi-rent, human capital and effort — follow the same qualitative pattern as the changes in plant-average quality illustrated in Figure 3. As a consequence, we expect to see greater increases in the average wage, the market wage, and the quasi-rent in high- λ than in small- λ plants in response to an exchange-rate devaluation. This is the prediction that we take to data in the remainder of the paper.

Second, if we only had data on average wages at the plant level, the Cobb-Douglas coefficients, β_h and β_e , would not be separately identified. This is a formalization of the argument in the introduction that plant-level data on average wages left open the question of whether the differential wage changes in response to the peso crisis identified by Verhoogen (2004) are due to sorting or changing-wage-premia. To address this issue, we need to be able to estimate the market wage and the quasi-rent separately, for which we need employer-employee data. With the employer-employee data, we can recover the parameters governing the contribution of human capital and effort to the quality production function by comparing the shares of the differential changes in overall average wages explained by changes in the

market wages and quasi-rents.

3 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. Our data can be considered a census of private, formal-sector establishments and their workforces. The number of workers in the dataset 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 Sept. 30 for each year 1985-2001. At the establishment level, the data contain information only on location and industry (using the IMSS's own 271 4-digit industrial categories).

From these data, we construct three balanced panels of establishments, with their employees. We have been guided in part by the desire to maximize comparability between our panels and the plant-level panels used in Verhoogen (2004), and in part by the desire to use the same panel of establishments in our various specifications. In the first panel, which we refer to as the IMSS short manufacturing panel, we include manufacturing plants that have at least 50 workers in all years 1992-2001, 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. In our second panel, which we refer to as the IMSS short non-tradables panel, we use the same criteria for establishments in non-tradable sectors: construction, retail trade, transportation and services. In our third panel, which we refer to as the IMSS long manufacturing panel, we include plants with at least 50 workers in every year from 1985-2001, and meet the other criteria. Within each panel, we further required 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 reasons for this requirement are discussed below. The three panels contain 3632, 3659, and 2132 establishments respectively.

Two aspects of our cleaning procedure merit explanation. First, 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. Second, 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.

Summary statistics on these three 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.

In Section 4 below, we also report results using the Encuesta Industrial Anual (EIA), a panel dataset based on an establishment survey conducted by the Instituto Nacional de Estadísticas, Geografía e Información (INEGI), the Mexican government statistical agency. The EIA contains data on total employment, total hours worked, the total wage bill, expenditures, sales, inventories, and capital assets at the establishment level. A companion

survey, the Encuesta Industrial Mensual (EIM), contains information on employment, wages and hours worked for two separate occupational categories, *empleados* and *obreros*, corresponding to white-collar and blue-collar workers. Neither survey covers the *maquiladoras*, the assembly-for-export plants located mainly along the U.S. border, which are covered by a separate survey. In this paper, we focus on a panel dataset for 1993-2001. For summary statistics and a fuller description of these data, we refer the reader to Verhoogen (2004).

4 Comparison of Results from IMSS and EIA Data

This section compares the results for average wages at the plant level from the IMSS employer-employee data to the EIA plant-level data. Before we begin, let us briefly review the key facts about the peso crisis.

At the end of December, 1994, the Mexican peso lost approximately 50% of its value in a matter of days. A major recession ensued; Mexican GDP dropped by 6.7% from 1994 to 1995. The shock to the terms of trade dwarfed the effect of tariff changes under the North American Free Trade Agreement (NAFTA), in effect since the previous January, which in most cases was phasing out tariffs at a rate of a few percentage points per year.

The price level and labor costs dropped sharply in Mexico relative to the U.S. Figure 4 plots the real exchange rate over the period for which we have data, and illustrates the sharp real depreciation in 1995.⁹ Exports rose sharply, with approximately 85% of them destined for the U.S. market. Figures 5a-c illustrate the key facts: domestic sales dipped, export sales rose (interestingly, total sales remained roughly on trend for the EIA sample

⁹ Note that there was also a real depreciation of the peso, less sudden but nearly as large, in 1986. We return to this below.

of large plants), 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. The model presented above, and case study evidence reported in Verhoogen (2004), suggest that this shift in production towards the export market was accompanied by an increase in the average quality of good produced in exporting plants.

We now to the evidence on quality upgrading and differential wage changes within each industry, first using the EIA plant-level panel. The first issue to be confronted is the choice of a proxy variable for λ , the underlying productivity parameter in the model. Verhoogen (2004) presents evidence that a number of different proxies for productivity — log domestic sales, the first principal component of a number of variables hypothesized to be correlated with entrepreneurial ability, and log total factor productivity, among others — are highly correlated with one another and yield similar results. That paper suggests that log domestic sales should be the preferred proxy, in large part because it is well-measured. The basic theoretical justification for this proxy is that plants with greater managerial talent will sell more goods and hire more workers; we can then infer underlying productivity from size. In the IMSS employer-employee data, we have just one proxy available: log employment. Empirically, log employment is also highly correlated with other productivity proxies in the plant-level data (correlation coefficient .78 with log domestic sales and .75 with the first principal component measure); we therefore have confidence that log employment is a reasonable proxy.

Figures 6a-d illustrate the main results from the EIA panel. Figure 6a plots non-

parametric local bivariate regressions of the export share of sales in 1994 and 1997 against log domestic sales 1993, separately in 1994 and 1997, where all variables have been deviated from industry-year means. Because of the deviation from industry-year means, the information in this graph is contained in the slope. The fact that the slope is greater in 1997 than in 1994 indicates that the change in the export share from 1994-1997 was greater for plants with greater initial domestic sales within each industry.¹⁰ Figure 6b presents a similar graph for log average wages, and the difference in slopes again suggests that the change in log average wages from 1993-1997 was greater in plants with greater initial domestic sales. Figures 6c-d present non-parametric regressions of changes in the export share and log average wages against initial log domestic sales, over two periods, 1993-1997 and 1997-2001. The graphs indicate that the differential changes in the export share and log average wages are larger in 1993-1997 than in 1997-2001.

We now turn to a linear regression specification to test these observations. Our basic model is the following:

$$\Delta \overline{w}_j = \beta_0 + \beta_1 S_{j0} + \beta_2 D_j + \beta_3 R_j + \epsilon_j \quad (17)$$

where j indexes establishments, \overline{w}_{jt} is the log average wage, S_{j0} is initial size of establishment, D_j is a vector of industry indicator variables, R_j is a vector of region indicator variables, and ϵ_{jt} is a mean-zero, serially uncorrelated disturbance.

The coefficient of interest in these regressions is β_1 , which captures a differential change in the log average wage by initial size, controlling for industry and state effects. We estimate

¹⁰ We focus on 1994 rather than 1993 when focusing on the export share of sales since domestic sales appears in the denominator of the export share, and we want to minimize the effect of possible measurement error in domestic sales on our results.

this model separately by period. If there were no other differential influences on large and small plants besides the exchange rate, the strict prediction of the theoretical model would be that $\beta_1 > 0$ in 1993-1997 and β than in 1997-2001. However, a more plausible specification would allow for some differential evolution of the export share in large and small plants. The crucial prediction of the model is that β_1 will be significantly larger in 1993-1997 than in 1997-2001. Part A of Table 2 presents the results for this model from the EIA 1993-2001 panel. Consistent with Figures 6c-d and with the predictions of the model, we find that the coefficients on the initial log domestic sales term are significantly larger in the earlier period. Part B of Table 2 reports a similar regression using a smaller subset of plants for which data on ISO 9000 certification are available. ISO 9000 is an international production standard measure the rationalization of management procedures that is commonly associated with high product quality. The data are available on for two years, 1994 and 1998, but the positive significant coefficient on the initial size term provides corroborative evidence that the initially larger plants differentially raised product quality during the peso crisis years.

Table 3 presents the most closely comparable results using the EIA and IMSS panels. Because information on domestic sales is not available in the IMSS data, we rely on log employment as the productivity proxy. Comparing the results for the EIA using the log employment proxy in Column (1) of Part A to the results from Column 3 of Table 2, we note that the coefficients on initial size are larger in both periods. But the coefficient in 1993-1997 remains significantly larger than the coefficient in the later period. 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. To address this concern, Column 2 reports similar regressions using the change in the dependent variables from 1994-1997 and 1998-2001, and the level of initial employment in 1993 and 1997, respectively. Our coefficient estimates are smaller than in Column 1 as expected, but follow the same pattern. Compare the results for the EIA to the results for the IMSS short manufacturing panel in Part B. The results tell a very consistent story. The similarity of the results for the two datasets is 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.

5 Estimation of Plant and Person Effects

The results of the previous section leave open the question of whether the differential wage changes are due to sorting or changing wage premia. In this section, we take advantage of the matched employer-employee data to estimate the separate contributions of individual effects and plant effects to wage changes at the plant level.

We begin by estimating a simple person-level model of the following form:

$$w_{ijt} = \alpha_i + \psi_{jt} + \epsilon_{ijt} \quad (18)$$

where w_{ijt} is the log wage of individual i at plant j in year t , α_i is a person fixed effect, ψ_{jt} is a plant-year effect, and ϵ_{ijt} is a mean-zero disturbance. For the short panels, we use the years 1993, 1997 and 2001. For the long panel, we use the years 1985, 1989, 1993, 1997, and 2001.

Two notes about this model are important. First, the person effects α_i and the plant-year effects ψ_{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; otherwise we would not be able to identify the person and establishment effects for those establishments. Second, estimating this model by OLS requires the assumption that the person effects of the switchers are uncorrelated with the error term. That is, we impose the assumption of exogenous mobility. In doing so, we are following Abowd, Kramarz, and Margolis (1999) and much of the employer-employee literature. Investigating the validity of the assumption of exogenous mobility is a subject for future work.

Once we have estimated this model, we recover the estimated person effects $\hat{\alpha}_i$ and plant-year effects $\hat{\psi}_{jt}$.¹¹ Using the estimated person effects, we construct a measure of the average person effect at the establishment level, $\overline{\hat{\alpha}_i}$. We also calculate the average log wage at the

¹¹ In implementing this regression, we first deviate all variables from the means for individual workers, and then recover the estimates of the coefficients on the individual fixed effects from the estimated coefficients for the other covariates. The computational problems discussed in Abowd, Kramarz, and Margolis (1999) do not arise here because we estimate approximately 3,600 (or fewer for the long panel) establishment effects, and matrices of this size can be inverted by standard software packages.

establishment level. We then estimate models of the form of equation (17), with the average log wage, the estimated plant-year effect, $\hat{\psi}_{jt}$, and the average person effect, $\hat{\alpha}_i$, as dependent variables. As a further test, we also calculate the average log wage at the establishment-year level for stayers, employees continuously employed in an establishment. If there is no non-random selection of stayers — that is, if the assumption of exogenous mobility holds — then we would expect the results for stayers to be similar to the results for the plant-year effects, since the person effects, α_i drop out of the wage-change calculation.

Figures 7a-d present non-parametric local bivariate regressions, analogous to those in Figures 6a-b, to illustrate the main results for the IMSS short manufacturing panel. Again, all variables have been deviated from industry-year means. The graph for avg log wages (Figure 7a) shows that in 1993 avg log wages are higher in establishments with greater 1993 employment. This is the well-known firm size-wage effect noted by Brown and Medoff (1989) for the U.S. and by Velenchik (1997) and Schaffner (1998) for developing countries. The graph also shows that wage growth from 1993 to 1997 was greater in large establishments. Figures 7b-c present analogous graphs for the plant-year effects and average person effects. We see that plant effects are greater in larger establishments within each year, and again that the change in plant effects was greater in larger plants. The graphs for the average person effects are remarkably flat. This is surprising given the prediction of the model that human capital will be an increasing function of λ in cross section. But the lack of correlation between size and person effects is consistent with results for developed countries; Abowd, Kramarz, Lenger mann, and Pérez-Duarte (2004) find no significant correlation between firm and person effects in the U.S. and a negative correlation in France. The graph for stayers

(Figure 7) is similar to the plant effects graph, as expected. Figures 8 compare changes for the periods 1993-1997 and 1997-2001, analogous to Figures 6c-d. Although the graphs are not flat for the 1997-2001 period for any of the dependent variables, the slopes of the 1993-1997 graphs are clearly greater. It is also evidence that the difference in slopes is greater for the average log wage, plant effect, and average log stayers' wage than for the average person effect.

Table 4 presents linear regression results for a model of the form of equation (17) for these same four dependent variables in the IMSS short manufacturing panel. The results confirm the observations on the figures made in the previous paragraph. We simply note that the differences between the coefficients on initial log employment are strongly statistically significant in all cases. We also note that the difference in coefficients on initial log employment for the plant effect, 0.024, is twice as large as the difference in coefficients on the average person effect, 0.012.

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 product quality, then we would expect to see similar results in non-tradables as in manufacturing. The results for the non-tradable sectors appear in Table 5. The results for the differential changes in average person effects (Column (3)) are similar to the results for manufacturing, but the results for the differential changes in plant effects stand in sharp contrast. There was essentially no differential within-industry change in plant effects by initial employment in either 1993-1997 or 1997-2001 in non-tradables. This is what we would have expected if the plant effects depend on product quality. The stayers results again reinforce the results for establishment effects.

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. One simple way to investigate this possibility is to control for the the growth of establishments in a model of the form of (17). Establishments presented with new profit-making opportunities typically expand; as a result, employment growth is correlated, albeit imperfectly, with growth of profitability. Table 6 reports results for our basic model using the change in employment as a proxy for the change in profitability. Two patterns are notable. First, the change-in-scale term is associated with a significant increase in the plant effect. This is consistent with the rent-sharing hypothesis. Second, the term is associated with a significant *decrease* in the average person effect. This is consistent with the idea that establishments face an upward-sloping supply curve for skill: the more an establishment expands, the further down the supply curve it must move in order to hire new

workers. The second association is stronger than the first, hence the overall association of the change in scale with average log wages is negative. But the key point is that the results for plant effects are largely unaffected — indeed, strengthened — by the inclusion of the scale term. 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.

Finally, we estimate our baseline model on the IMSS long manufacturing model, using data from 1985-2001. As we saw in Figure 4, there was a real depreciation of the peso in 1986, followed by a long appreciation until the late-1994 peso crisis. Verhoogen (2004) found in a panel of 700 plants from the EIA that results for differential wage changes in the 1985-1989 devaluation period were weak but similar to the 1993-1997 crisis period, and that the results for the periods of appreciation, 1989-1993 and 1997-2001, were similar as well. In the IMSS data, by contrast, the results from the pre-crisis years are less clear-cut. The differential changes in log average wage and average log wage are significantly greater in 1993-1997 than in any of the other periods, as we would have expected. But for the plant effects and stayers regressions, the differential changes in 1989-1993 are not significantly lower than in 1993-1997. The difference between these results and those of Verhoogen (2004) may be attributable to the inclusion of a large number of additional plants in the IMSS long panel.

6 Conclusion

This paper has presented evidence that plant effects rose more in initially larger plants than in initially smaller plants within each industry in response to the peso crisis. Our results strongly support the hypothesis that quality upgrading in response to the exchange-rate shock led plants to increase wage premia. It remains an open question precisely why higher product

quality requires higher wage premia. Above we hypothesized that high-quality production requires particularly motivated workers, and that higher wages increase motivation, along the lines of the efficiency-wage models of Bowles (1985); Shapiro and Stiglitz (1984) or Akerlof (1982). Another possibility is that high-quality production requires providing training, some of which is general, and that subsequently establishments must pay higher wages to prevent their trained employees from leaving for other establishments. We leave the resolution of this question to future work.

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A Changes in Entry Patterns in Response to Shock

The zero-profit conditions determining entry (19) can be written:

$$\pi_{cd}(\lambda_{cd}^{\min}) = \frac{\mu \delta_{cd} N_d}{D_d} \exp \left\{ \frac{\theta_d q_{cd}^*(\lambda_{cd}^{\min}) - \frac{p_{cd}^*(\lambda_{cd}^{\min})}{\delta_{cd}}}{\mu} \right\} - f_{cd} = 0 \quad (\text{A.19})$$

for $c, d = n, s$. These can be rewritten:

$$D_s(\lambda_{ss}^{\min}, \lambda_{ns}^{\min}) = \frac{\mu \delta_{ss} N_s}{f_{ss}} \exp \{ I_{ss}(\lambda_{ss}^{\min}) \} \quad (\text{A.20})$$

$$D_s(\lambda_{ss}^{\min}, \lambda_{ns}^{\min}) = \frac{\mu \delta_{ns} N_s}{f_{ns}} \exp \{ I_{ns}(\lambda_{ns}^{\min}) \} \quad (\text{A.21})$$

$$D_n(\lambda_{sn}^{\min}, \lambda_{nn}^{\min}) = \frac{\mu \delta_{sn} N_n}{f_{sn}} \exp \{ I_{sn}(\lambda_{sn}^{\min}) \} \quad (\text{A.22})$$

$$D_n(\lambda_{sn}^{\min}, \lambda_{nn}^{\min}) = \frac{\mu \delta_{nn} N_n}{f_{nn}} \exp \{ I_{nn}(\lambda_{nn}^{\min}) \} \quad (\text{A.23})$$

where $I_{cd}(\cdot)$ is defined as:

$$I_{cd}(\lambda) \equiv \frac{1}{\mu} \left[\theta_d q_{cd}(\lambda) - \frac{p_{cd}(\lambda)}{\delta_{cd}} \right] \quad (\text{A.24})$$

To sign $\partial \lambda_{ss}^{\min} / \partial \delta_{sn}$, differentiate (20) and (21) with respect to δ_{sn} and solve for $\partial \lambda_{ss}^{\min} / \partial \delta_{sn}$:

$$\frac{\partial \lambda_{ss}^{\min}}{\partial \delta_{sn}} = \frac{1}{\Omega_s} \left\{ \frac{\partial I_{ns}(\lambda_{ns}^{\min})}{\partial \lambda_{ns}^{\min}} \frac{\partial D_s}{\partial \delta_{sn}} + \frac{\partial D_s}{\partial \lambda_{ns}^{\min}} \left[\frac{1}{\delta_{sn}} - \frac{\partial I_{ns}^*(\lambda_{ns}^{\min})}{\partial \delta_{sn}} \right] \right\} \quad (\text{A.25})$$

where

$$\Omega_s \equiv \frac{\partial I_{ss}(\lambda_{ss}^{\min})}{\partial \lambda_{ss}^{\min}} \frac{\partial I_{ns}(\lambda_{ns}^{\min})}{\partial \lambda_{ns}^{\min}} - \frac{\partial D_s}{\partial \lambda_{ss}^{\min}} \frac{\partial I_{ns}(\lambda_{ns}^{\min})}{\partial \lambda_{ns}^{\min}} - \frac{\partial D_s}{\partial \lambda_{ns}^{\min}} \frac{\partial I_{ss}(\lambda_{ss}^{\min})}{\partial \lambda_{ss}^{\min}} \quad (\text{A.26})$$

It follows from the definition of $I_{cd}(\cdot)$ that $\partial I_{ss}(\lambda_{ss}^{\min}) / \partial \delta_{sn} = 0$, $\partial I_{ns}(\lambda_{ns}^{\min}) / \partial \lambda_{ns}^{\min} > 0$, and $\partial I_{ns}(\lambda_{ns}^{\min}) / \partial \delta_{sn} < 0$. From (2), we have that $\partial D_s / \partial \lambda_{ns}^{\min} < 0$. Under standard regularity conditions,¹² we can write:

$$\begin{aligned} \frac{\partial D_s}{\partial \delta_{sn}} &= \int_{\lambda_{sd}^{\min}}^{\infty} \frac{\partial \exp \{ I_{ns}(\lambda) \}}{\partial \delta_{sn}} g_s(\lambda) d\lambda \\ &= \int_{\lambda_{sd}^{\min}}^{\infty} \frac{\partial I_{ns}(\lambda)}{\partial \delta_{sn}} \exp \{ I_{ns}(\lambda) \} g_s(\lambda) d\lambda < 0 \end{aligned} \quad (\text{A.27})$$

since $\partial I_{ns}(\lambda) / \partial \delta_{sn} < 0$. Given these relationships, we can conclude that $\Omega_s > 0$ and $\partial \lambda_{ss}^{\min} / \partial \delta_{sn} < 0$.

¹² See Amemiya (1985, Theorem 1.3.2, p. 17) or Newey and McFadden (1994, Lemma 3.6, p.2152).

To sign $\partial\lambda_{sn}^{\min}/\partial\delta_{sn}$, differentiate (22), recalling the assumption that profitability (and hence entry, captured in D_n) in the Northern market is approximately unaffected by the devaluation, and solve for $\partial\lambda_{sn}^{\min}/\partial\delta_{sn}$:

$$\frac{\partial\lambda_{sn}^{\min}}{\partial\delta_{sn}} = -\frac{\frac{\partial I_{sn}(\lambda_{sn}^{\min})}{\partial\delta_{sn}} + \frac{1}{\delta_{sn}}}{\frac{\partial I_{sn}(\lambda_{sn}^{\min})}{\partial\lambda_{sn}^{\min}}} \quad (\text{A.28})$$

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

To sign $\partial\lambda_{ss}^{\min}/\partial N_s$, note that D_s , $I_{ss}(\cdot)$ and $I_{sn}(\cdot)$ depend directly on N_s only through λ_{ss}^{\min} and λ_{ns}^{\min} , differentiate (20) and (21) with respect to δ_{sn} , and solve:

$$\frac{\partial\lambda_{ss}^{\min}}{\partial N_s} = -\frac{D_s \frac{\partial I_{ns}(\lambda_{ns}^{\min})}{\partial\lambda_{ns}^{\min}} - \frac{\partial D_s(\lambda_{ns}^{\min})}{\partial\lambda_{ns}^{\min}}}{\Omega_s N_s} \quad (\text{A.29})$$

where Ω_s is defined as in (26) above. Since $\partial I_{ns}(\lambda_{ns}^{\min})/\partial\lambda_{ns}^{\min} > 0$ by the definition of $I_{cd}(\lambda)$ and $\partial D_s/\partial\lambda_{ns}^{\min} < 0$ as noted above, we have $\partial\lambda_{ss}^{\min}/\partial N_s < 0$.

Since the markets are segmented and no terms in (22) depend on N_s , it follows immediately that λ_{sn}^{\min} does not depend on N_s ; that is, $\partial\lambda_{sn}^{\min}/\partial N_s = 0$.

B Sufficient Condition for Exchange-Rate Shock to Induce Exit of Southern Plants from Domestic Market

The following condition is sufficient to ensure required to ensure that the exchange-rate shock induces exit of Southern plants from the Southern market:

$$\frac{-dN_s}{N_s} + \left\{ \frac{\frac{\partial D_s}{\partial\delta_{sn}} \frac{\partial I_{ns}(\lambda_{ns}^{\min})}{\partial\lambda_{ns}^{\min}} - \frac{\partial D_s}{\partial\lambda_{ns}^{\min}} \left[\frac{\partial I_{ns}(\lambda_{ns}^{\min})}{\partial\delta_{sn}} - \frac{1}{\delta_{sn}} \right]}{D_s \frac{\partial I_{ns}(\lambda_{ns}^{\min})}{\partial\lambda_{ns}^{\min}}} \right\} d\delta_{sn} > 0 \quad (\text{B.30})$$

Using (28), (29) and this condition, we have:

$$d\lambda_{ss}^{\min} = \frac{\partial\lambda_{ss}^{\min}}{\partial\delta_{sn}} d\delta_{sn} + \frac{\partial\lambda_{ss}^{\min}}{\partial N_s} dN_s > 0 \quad (\text{B.31})$$

Intuitively, the condition (30) ensures that decline in the number of consumers in South is sufficiently large in magnitude to outweigh the increased competitiveness of goods produced in South. We maintain this assumption in what follows. Note that the condition is stronger than necessary to ensure exit; the condition as written also ensures that production of Southern plants for the Southern market declines, as we will see in Appendix C.

C Change in Export Share of Output

Differentiating the definition of $\chi_s(\lambda)$ and simplifying:

$$d\chi_s = \chi_s(\lambda)(1 - \chi_s(\lambda)) \left\{ \frac{dx_{sn}(\lambda)}{x_{sn}(\lambda)} - \frac{dx_{ss}(\lambda)}{x_{ss}(\lambda)} \right\} \quad (\text{C.32})$$

Using the expression for $x_{ss}(\lambda)$ in (10) and noting that $\frac{\partial I_{ss}(\lambda)}{\partial \delta_{sn}} = \frac{\partial I_{ss}(\lambda)}{\partial N_s} = 0$, we have:

$$\begin{aligned} dx_{ss}(\lambda) &= \frac{\partial x_{ss}(\lambda)}{\partial \delta_{sn}} d\delta_{sn} + \frac{\partial x_{ss}(\lambda)}{\partial N_s} dN_s \\ &= x_{ss}(\lambda) \left\{ -\frac{\frac{\partial D_s}{\partial \delta_{sn}}}{D_s} d\delta_{sn} + \frac{1}{N_s} dN_s \right\} \end{aligned} \quad (\text{C.33})$$

If the condition required to ensure that the exchange-rate shock induced exit of Southern plants from the domestic market (equation (30) in Appendix B) is satisfied, then the term in curly braces will be negative. Thus in the empirically relevant case, $dx_{ss}(\lambda) < 0$.

Again using the expression for $x_{ss}(\lambda)$ in (10), and recalling the assumption that the aggregate D_n is approximately unaffected by the changes in exports of Southern plants, we have:

$$\begin{aligned} dx_{sn}(\lambda) &= \frac{\partial x_{sn}(\lambda)}{\partial \delta_{sn}} d\delta_{sn} + \frac{\partial x_{sn}(\lambda)}{\partial N_s} dN_s \\ &= x_{sn}(\lambda) \frac{\partial I_{sn}(\lambda)}{\partial \delta_{sn}} d\delta_{sn} \end{aligned} \quad (\text{C.34})$$

Since $\frac{\partial I_{sn}(\lambda)}{\partial \delta_{sn}} > 0$, we have $dx_{sn}(\lambda) > 0$.

Substituting (33) and (34) into (32):

$$d\chi_s = \chi_s(\lambda)(1 - \chi_s(\lambda)) \left\{ \frac{\partial I_{sn}(\lambda)}{\partial \delta_{sn}} d\delta_{sn} + \frac{1}{D_s} \frac{\partial D_s}{\partial \delta_{sn}} d\delta_{sn} - \frac{1}{N_s} dN_s \right\} \quad (\text{C.35})$$

Let $\Sigma_s \equiv \left\{ \frac{\partial I_{sn}(\lambda)}{\partial \delta_{sn}} d\delta_{sn} + \frac{1}{D_s} \frac{\partial D_s}{\partial \delta_{sn}} d\delta_{sn} - \frac{1}{N_s} dN_s \right\}$. Condition (30) ensures that $\frac{1}{D_s} \frac{\partial D_s}{\partial \delta_{sn}} d\delta_{sn} - \frac{1}{N_s} dN_s > 0$ and since $\partial I_{sn}(\lambda) / \partial \delta_{sn} > 0$, we have $\Sigma_s > 0$. In addition, $\partial I_{sn}(\lambda) / \partial \delta_{sn} = [\eta \theta_n (\delta_{sn} \lambda)^\alpha]^{1/(1-\alpha)}$, which is increasing in λ . Since only the first term in Σ_s varies with λ , we can conclude that Σ_s is increasing in λ .

Table 1: Summary Statistics: IMSS 1993-2001 Panel

	Manufacturing, short panel			Non-tradables, short panel			Manufacturing, long panel		
	emp. <200	emp. >=200	all	emp. <200	emp. >=200	all	emp. <200	emp. >=200	all
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
average daily wage (1994 pesos)	43.69 (0.045)	55.10 (0.025)	52.99 (0.022)	53.26 (0.060)	61.75 (0.039)	59.74 (0.033)	45.78 (0.045)	58.32 (0.024)	56.37 (0.021)
average daily wage (current US \$)	13.72 (0.015)	17.48 (0.008)	16.79 (0.007)	16.77 (0.019)	19.57 (0.013)	18.91 (0.011)	15.13 (0.022)	19.20 (0.010)	18.65 (0.009)
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)	0.76 (0.001)	0.78 (0.000)	0.78 (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)	32.2 (0.014)	32.0 (0.006)	32.0 (0.005)
% stayers (previous 4 years)	0.43 (0.0008)	0.44 (0.0003)	0.44 (0.0003)	0.39 (0.0007)	0.41 (0.0004)	0.41 (0.0003)	0.44 (0.0007)	0.47 (0.0003)	0.47 (0.0003)
N individuals (000s)	674.4	2966.5	3640.9	785.2	2538.5	3323.7	615.2	3336.9	3952.2
N establishments	2410	1222	3632	2800	879	3679	1461	671	2132
Avg. emp (establishments)	279.9	2427.5	1002.5	280.4	2887.9	903.4	421.1	4973.1	1853.7

Notes: Manufacturing short panel and non-tradables short panels include years 1993, 1997, 2001. Manufacturing long panel includes 1985, 1989, 1993, 1997, 2001. Standard errors of means in brackets. Stayers calculated as (# employees continuously in establishment over previous 4 years)/(current employment).

Table 2: Results from Plant-level Datasets

A. EIA 1993-2001 panel, log domestic sales as size measure			
	Δ export share of sales	Δ export share of sales, sub. to base year	$\Delta \log(\text{avg. wage})$
	(1)	(2)	(3)
1993-1997 log domestic sales, 1993	1.964** [0.238]	1.295** [0.251]	0.043** [0.006]
R-squared	0.171	0.156	0.154
1997-2001 log domestic sales 1997	0.869** [0.188]	0.470** [0.181]	0.009* [0.004]
R-squared	0.142	0.124	0.112
Difference between 1993-1997 and 1997-2001 coefficients	1.091** [0.303]	0.825** [0.309]	0.034** [.007]

Notes: All regressions include industry and state fixed effects (coefficients omitted). N = 3003. Robust standard errors in brackets. In Column 2, dependent variable calculated as change from year after base year to end year (i.e. 1994-1997 and 1998-2001). ** significant at 1% level, * at 5% level.

B. EIA-ENESTyC 1994-1998 panel, log domestic sales as size measure	
	Δ ISO 9000 certification
1994-1998 log domestic sales, 1994	0.079** [0.018]
R-squared	0.350

Notes: Includes industry and state fixed effects (coefficients omitted). N = 767. Robust standard error in brackets. ** significant at 1% level.

Table 3: Comparing EIA 1993-2001 and IMSS 1993-2001 Manufacturing Panels

		$\Delta \log(\text{avg. wage})$	$\Delta \log(\text{avg. wage}),$ sub. to base year
		(1)	(2)
A. EIA 1993-2001 panel, log employment as size measure			
1993-1997	log employment 1993	0.072** [0.008]	0.057** [0.007]
	R-squared	0.163	0.151
1997-2001	log employment 1997	0.020** [0.006]	0.016** [0.005]
	R-squared	0.114	0.118
B. IMSS 1993-2001 panel, log employment as size measure			
1993-1997	log employment 1993	0.056** [0.004]	0.041** [0.004]
	R-squared	0.185	0.18
1997-2001	log employment 1997	0.022** [0.004]	0.015** [0.003]
	R-squared	0.117	0.101

Notes: All regressions include industry and state fixed effects (coefficients omitted). N = 3003 in part A, N = 3632 in part B. Robust standard errors in brackets. In Column 2, dependent variable calculated as change from year after base year to end year (i.e. 1994-1997 and 1998-2001). ** significant at 1% level, * at 5% level.

Table 4: Baseline Results, IMSS 1993-2001 Manufacturing Panel

		Δ avg. log(wage) (1)	Δ plant effect (2)	Δ avg. person effect (3)	Δ avg. log(wage) of stayers (4)
1993-1997	log employment 1993	0.055** [0.004]	0.034** [0.004]	0.021** [0.003]	0.031** [0.004]
	R-squared	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]
	R-squared	0.124	0.143	0.091	0.147
	Difference between 1993-1997 and 1997-2001 coefficients	0.036** [0.006]	0.024** [0.005]	0.012** [0.004]	0.021** [0.005]

Notes: All regressions include industry and state fixed effects (coefficients omitted). N = 3632. Robust standard errors in brackets. ** significant at 1% level, * at 5% level.

Table 5: Baseline Results, IMSS 1993-2001 Non-Tradables Panel

		Δ avg. log(wage) (1)	Δ plant effect (2)	Δ avg. person effect (3)	Δ avg. log(wage) of stayers (4)
1993-1997	log employment 1993	0.014** [0.004]	-0.002 [0.004]	0.016** [0.003]	-0.005 [0.004]
	R-squared	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]
	R-squared	0.100	0.128	0.118	0.137
Difference between 1993-1997 and 1997-2001 coefficients		0.017* [0.007]	0.005 [0.007]	0.012** [0.004]	0.001 [0.007]

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. ** significant at 1% level, * at 5% lev

Table 6: Controlling for Change in Scale, IMSS 1993-2001 Manufacturing Panel

		Δ avg. log(wage) (1)	Δ plant effect (2)	Δ avg. person effect (3)	Δ avg. log(wage) of stayers (4)
1993-1997	log employment 1993	0.043** [0.004]	0.044** [0.004]	-0.001 [0.002]	0.043** [0.004]
	Δ log employment	-0.086** [0.010]	0.071** [0.008]	-0.157** [0.007]	0.091** [0.009]
	R-squared	0.225	0.202	0.394	0.208
1997-2001	log employment 1997	0.014** [0.003]	0.016** [0.003]	-0.002 [0.002]	0.016** [0.003]
	Δ log employment	-0.072** [0.011]	0.077** [0.008]	-0.149** [0.009]	0.080** [0.008]
	R-squared	0.15	0.176	0.316	0.18
Difference between 1993-1997 and 1997-2001 coefficients on log employment		0.030** [0.005]	0.028** [0.005]	0.001 [0.003]	0.028** [0.005]

Notes: All regressions include industry and state fixed effects (coefficients omitted). N = 3632. Robust standard errors in brackets. ** significant at 1% level, * at 5% level.

Table 7: IMSS Long Panel

		$\Delta \log(\text{avg. wage})$ (1)	$\Delta \text{avg.} \log(\text{wage})$ (2)	$\Delta \text{plant effect}$ (3)	$\Delta \text{avg. person effect}$ (4)	$\Delta \text{avg.} \log(\text{wage})$ (5)
1985-1989	log employment 1985	0.017** [0.005]	0.016** [0.005]	0.009* [0.004]	0.007* [0.003]	0.007 [0.005]
	R-squared	0.213	0.223	0.220	0.168	0.203
1989-1993	log employment 1989	0.015* [0.006]	0.030** [0.006]	0.030** [0.008]	0.000 [0.004]	0.028** [0.008]
	R-squared	0.215	0.236	0.233	0.107	0.222
1993-1997	log employment 1993	0.059** [0.006]	0.059** [0.006]	0.043** [0.005]	0.016** [0.004]	0.035** [0.005]
	R-squared	0.238	0.252	0.238	0.204	0.223
1997-2001	log employment 1997	0.020** [0.005]	0.017** [0.005]	0.006 [0.004]	0.010** [0.004]	0.005 [0.004]
	R-squared	0.117	0.125	0.164	0.115	0.171

Notes: All regressions include industry and state fixed effects (coefficients omitted). N = 2132. Robust standard errors in brackets. ** significant at 1% level, * at 5% level.

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

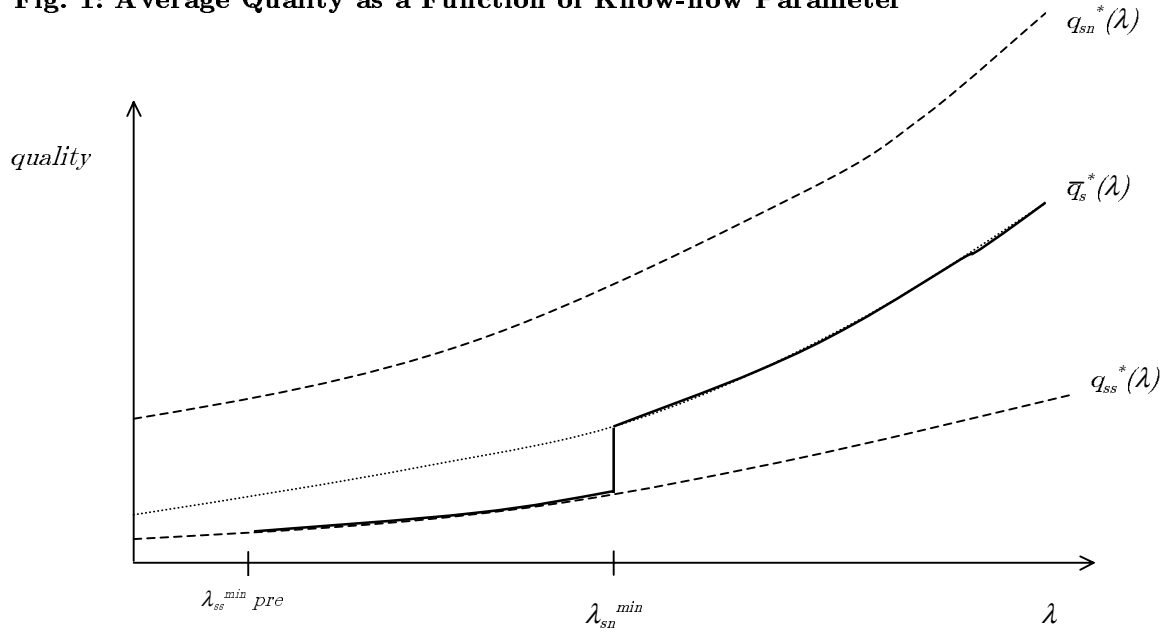


Fig. 2: Change in Average Quality as a Function of Know-how Parameter

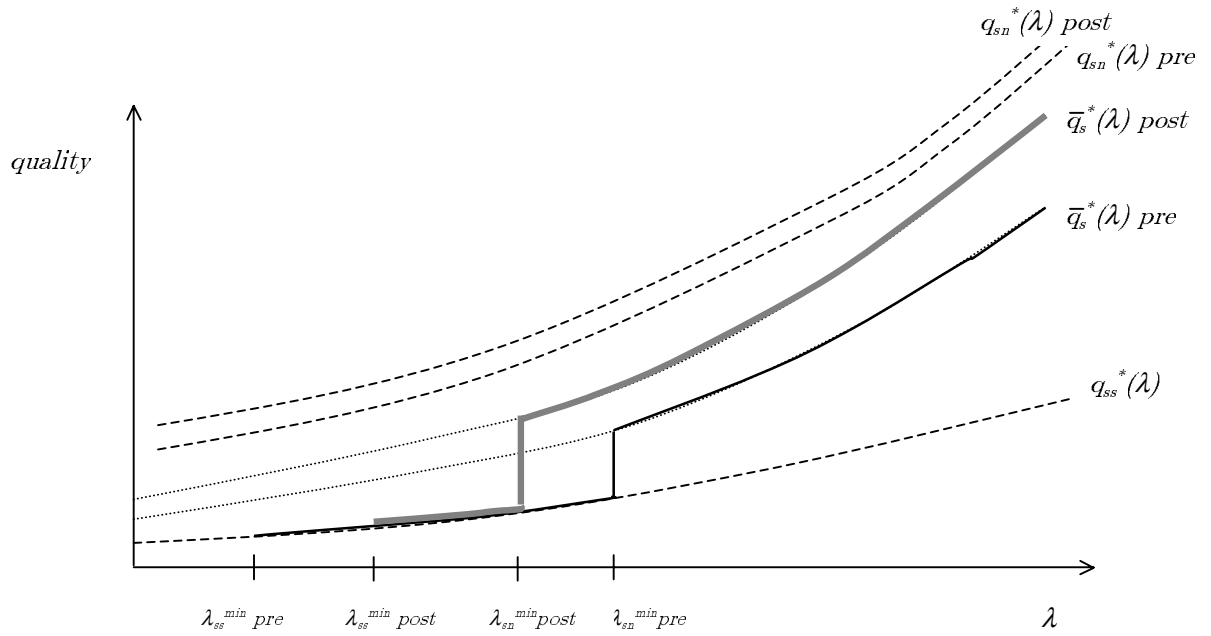


Fig. 3: Change in Average Quality as a Function of Know-how Parameter

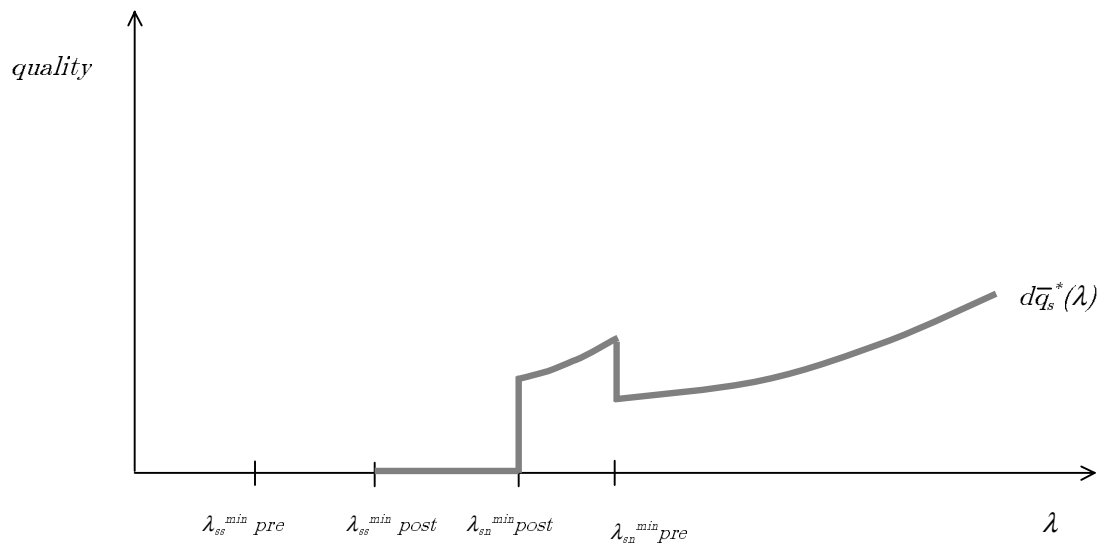
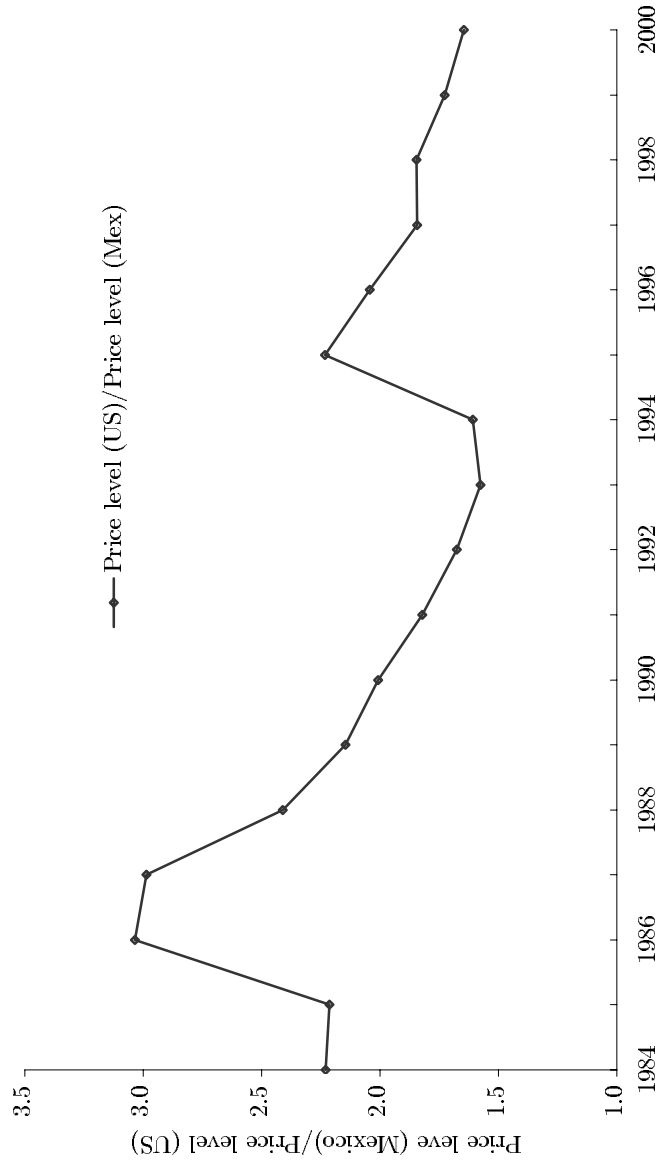


Fig. 4: Real Exchange Rate



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

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

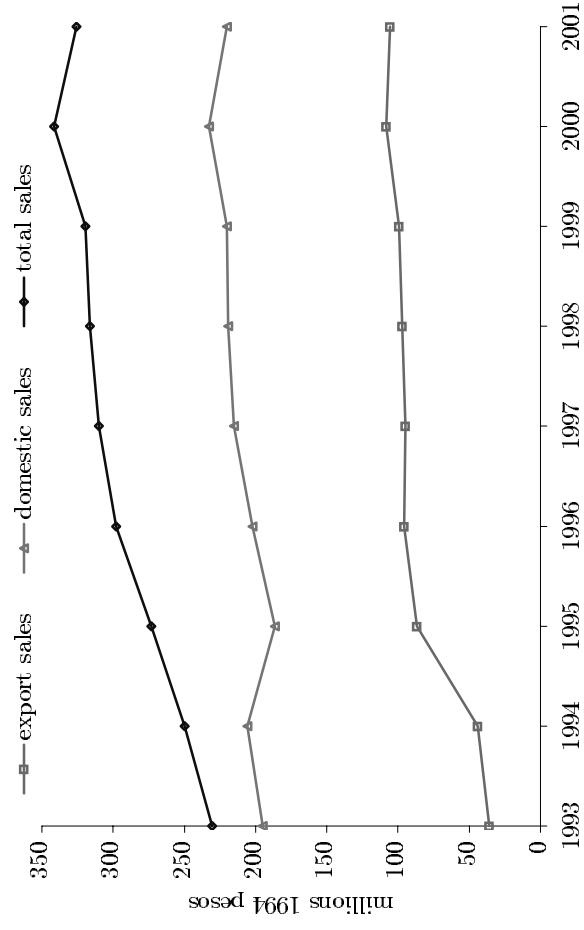


Fig. 5b: Export percentage of total sales

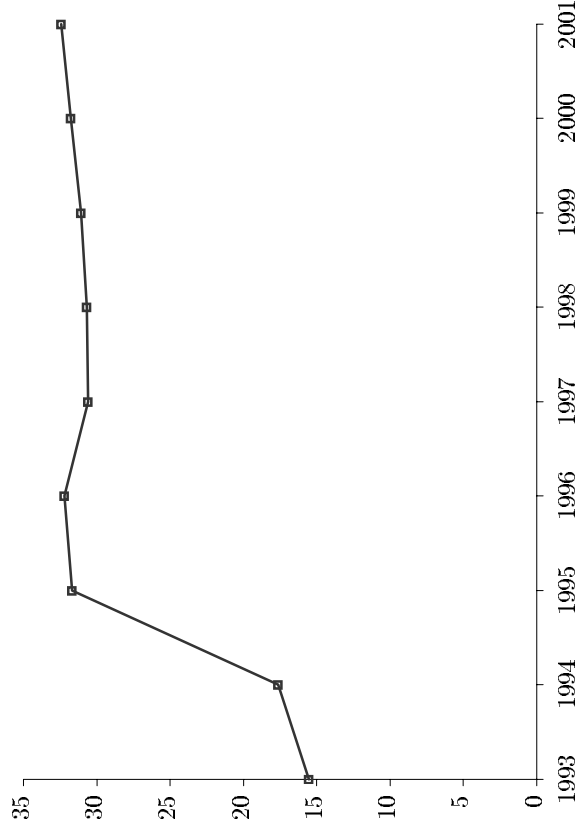
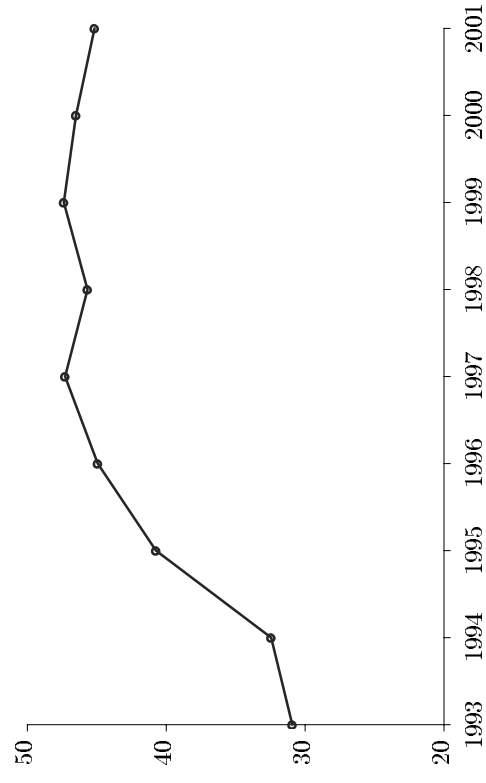


Fig. 5c: Percentage of Plants Exporting



Notes: Data from EIA 1993-2001 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. 6: Non-Parametric Regressions, 1993 and 1997, EIA Short Panel

Fig. 6a: Export % of Sales

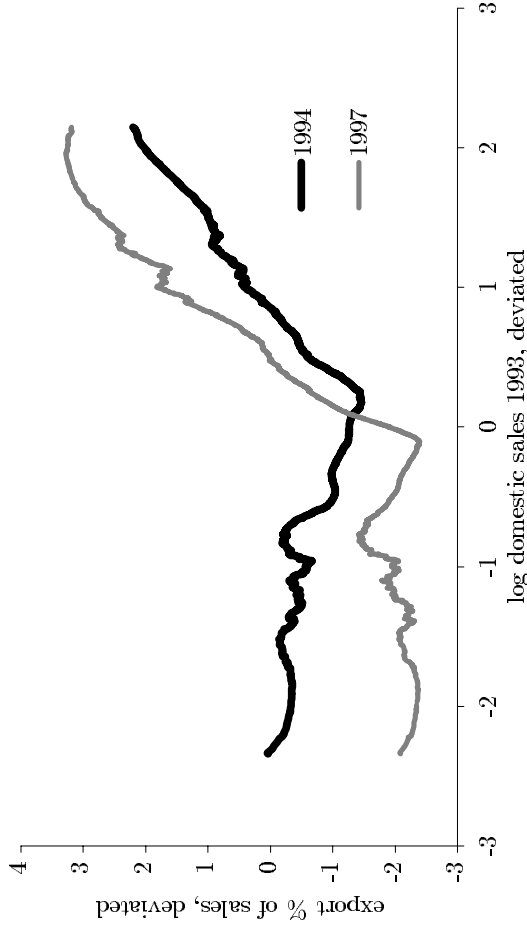


Fig. 6b: Log Avg. Wage

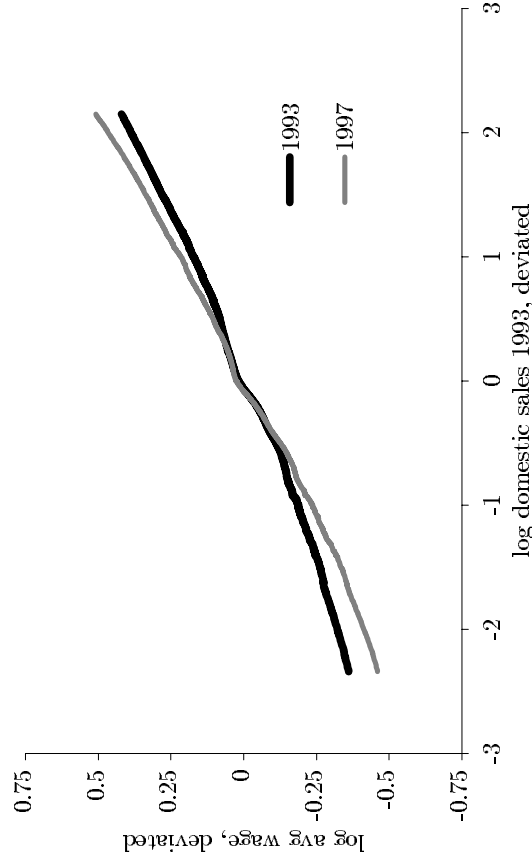


Fig. 6c: Change in Export % of Sales

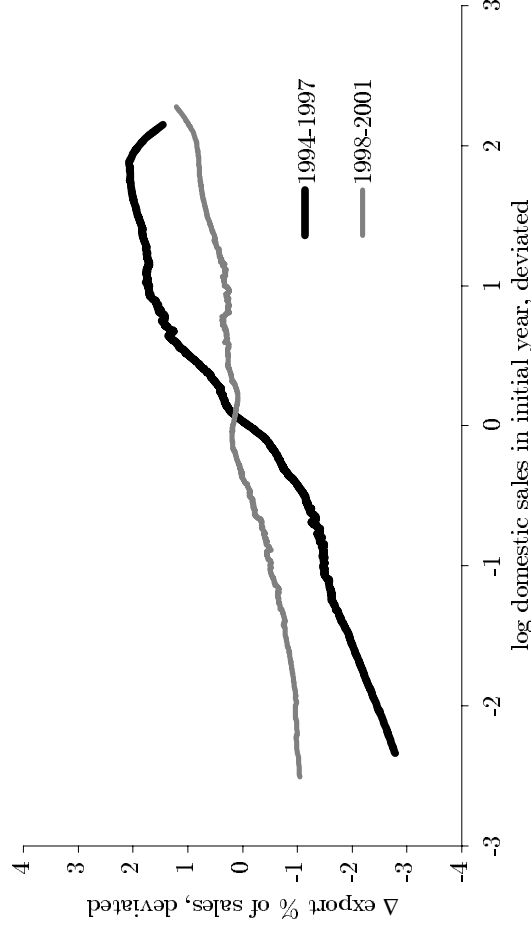
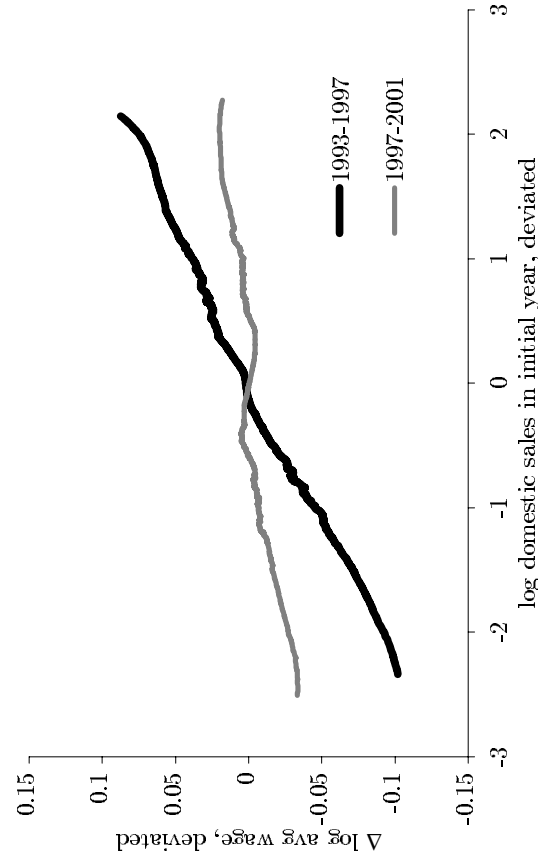
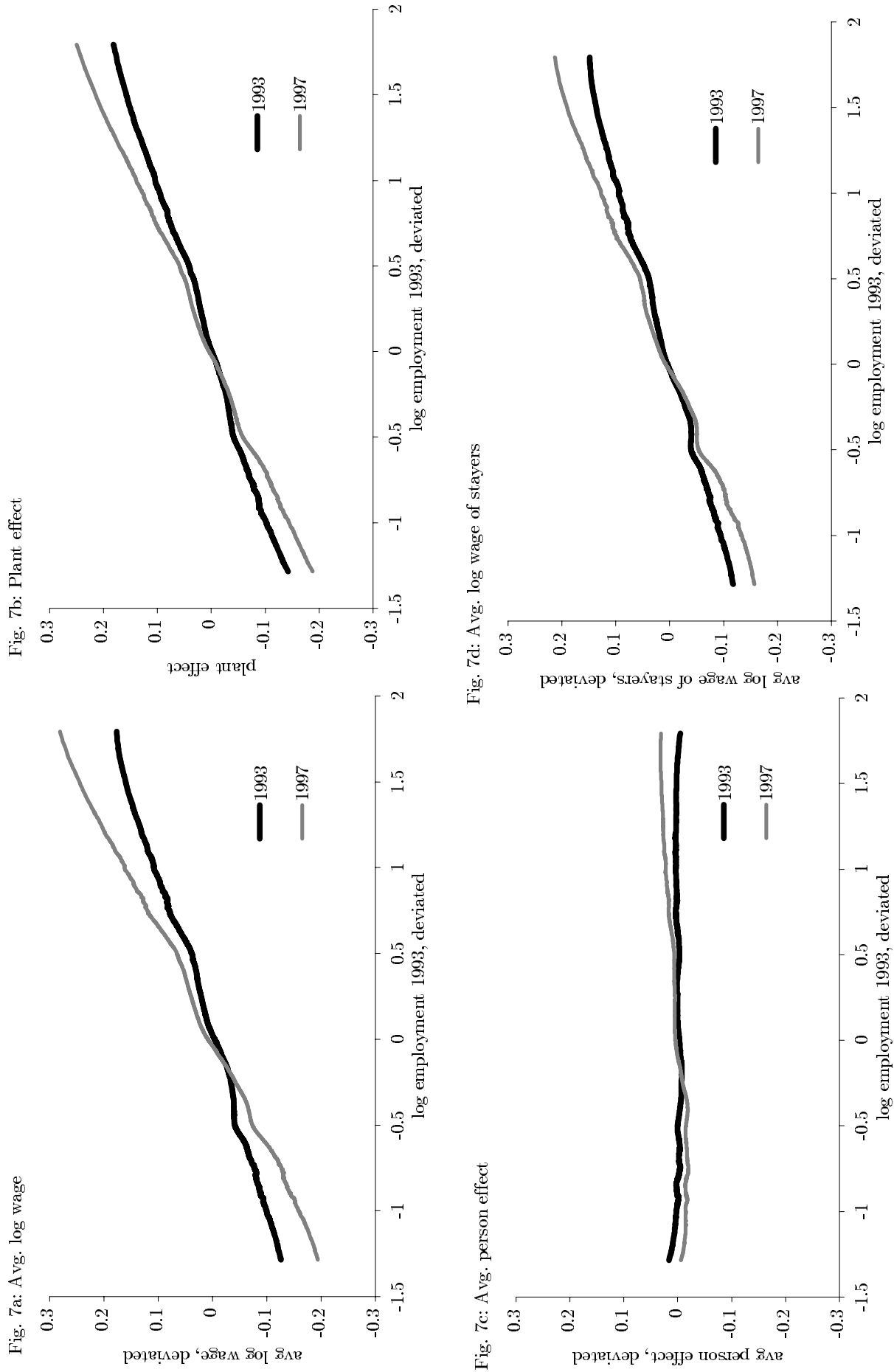


Fig. 6d: Change in Log Avg. Wage



Notes: All variables deviated from industry means. Graphs are locally smoothed non-parametric regressions (bandwidth = .5), of y-axis variable on log employment 1993, using IMSS 1993-2001 panel, trimmed at 2nd and 98th percentiles.

Fig. 7: Non-Parametric Regressions, 1993 and 1997, IMSS Short Panel



Notes: All variables deviated from industry-year means. Graphs are locally smoothed non-parametric bivariate regressions (bandwidth = .3), of y-axis variable on log employment 1993, using IMSS 1993-2001 panel, trimmed at 2nd and 98th percentiles.

Fig. 8: Non-Parametric Regressions, 1993-1997 and 1997-2001, IMSS Short Panel

Fig. 8a: Change in Avg. Log Wage

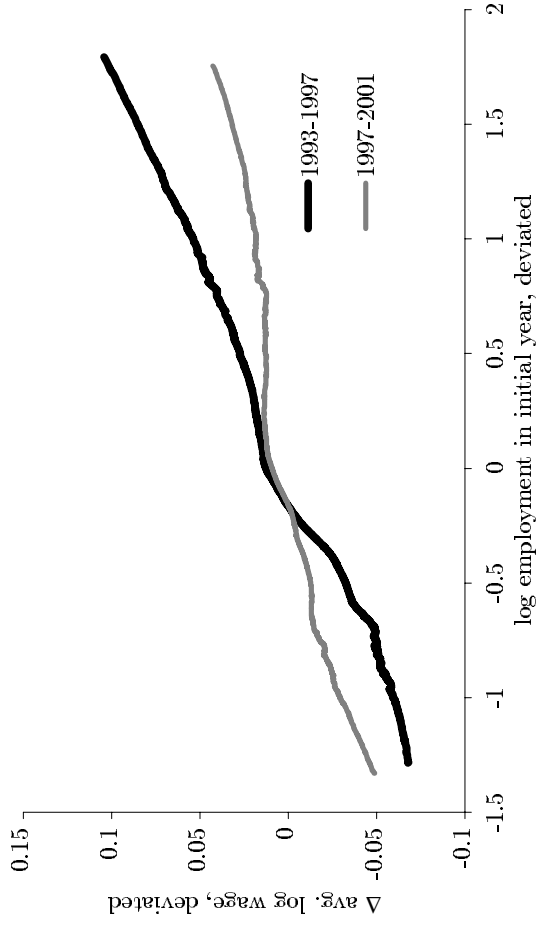


Fig. 8b: Change in Plant Effect

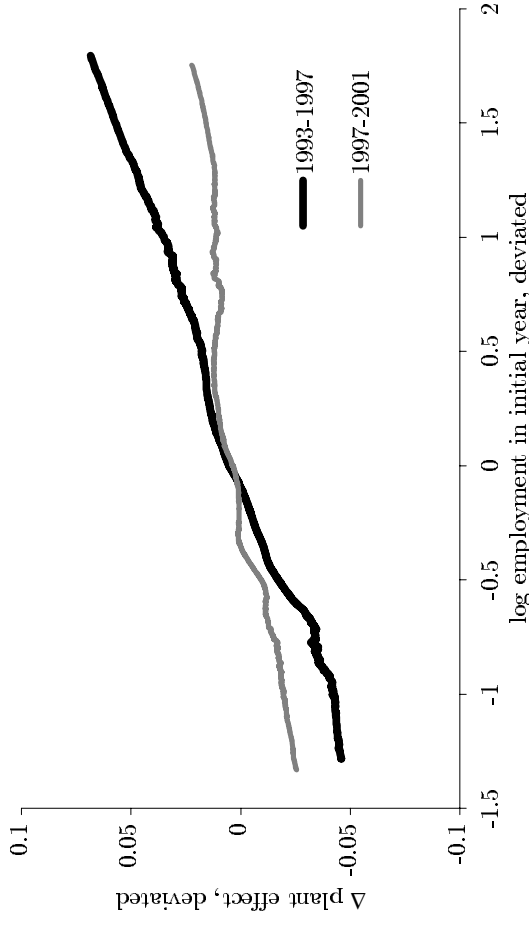


Fig. 8c: Change in Avg. Person Effect

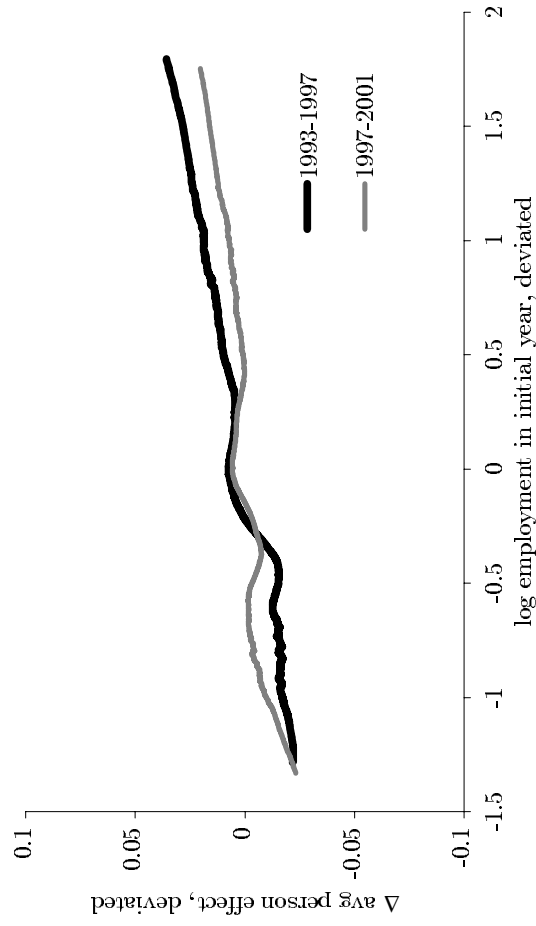
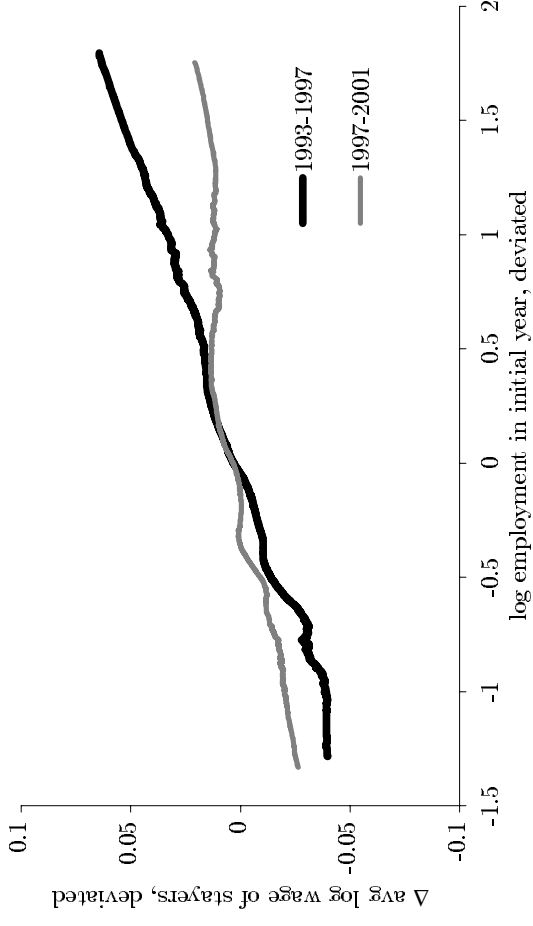


Fig. 8d: Change in Avg. Log Wage of Stayers



Notes: All variables deviated from industry-year means. Graphs are locally smoothed non-parametric bivariate regressions (bandwidth = .3), of y-axis variable on log employment 1993, using IMSS 1993-2001 panel, trimmed at 2nd and 98th percentiles.