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Exports and Within-Plant Wage Distributions: Evidence from Mexico†

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In many developing countries, increasing international integration has been accompanied by rising wage inequality, and traditional Heckscher-Ohlin models, which rely on between-sector reallocations to link trade and labor-market outcomes, are difficult to reconcile with this pattern (Goldberg and Pavcnik 2007). Recently, researchers have proposed a number of potential within-sector explanations based on the behavior of heterogeneous firms, involving technology choice, quality upgrading, search and bargaining, or fair wages, among other mechanisms.[1] There is evidence at the plant level to support a within-sector link between trade and inequality. For instance, Verhoogen (2008) finds that initially larger, higher-productivity Mexican plants had higher export propensity and wages in cross-section in 1993 and that they were more likely to increase exports and wages in response to the late-1994 devaluation of the peso. The shock to exporting thus arguably increased dispersion in wages between plants within sectors.

At the plant level, however, many of the proposed within-sector mechanisms carry similar observable implications. Distinguishing among the various mechanisms will require moving to a lower level of disaggregation, and exploiting information at the level of individual workers within plants. In this short article and the longer article to which it is a companion (Frías, Kaplan, and Verhoogen 2011), we use employer-employee data from Mexico and an identification strategy from Verhoogen (2008) to examine the effects of exporting on wage outcomes that are not available in standard plant-level datasets. In Frías, Kaplan, and Verhoogen (2011), we estimate the effect of exporting on wage premia, defined as wages above what individual workers would expect to earn elsewhere in the labor market. Wage premia are estimated as plant effects, controlling flexibly for individual heterogeneity (and allowing the return to worker ability to vary over time), implicitly assuming that the plant effect is the same for all employed workers.

In this short article, by contrast, we do not attempt to control for worker heterogeneity, but instead focus on the effect of exporting on the shape of within-plant wage distributions. As we show in more detail below, we find that exporting has little effect on wages at the low end of the wage spectrum within plants, and that it raises within-plant wage dispersion, but not uniformly between all quantiles. The results are consistent with, but add important qualifications to, the finding of Verhoogen (2008) in plant-level data that exporting raised the ratio of white-collar to blue-collar average wages.

This article is related to an active theory literature on trade, matching, and organizations which has proposed a variety of mechanisms

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[1] See, on technology choice, Yeaple (2005) and Bustos (2011); on quality upgrading, Verhoogen (2008); on search and bargaining, Davidson, Matusz, and Shevchenko (2008); Helpman, Itskhoki, and Redding (2010); Felbermayr, Prat, and Schmerer (2011); and Coşar, Guner, and Tybout (2010); on fair wages, Egger and Kreickemeier (2009) and Amiti and Davis (2012).

linking trade and wage distributions within firms.[2] Recent papers using employer-employee data to investigate the consequences of trade for labor-market outcomes (without focusing on the overall within-plant distributions) include Krishna, Poole, and Senses (2011); Hummels et al. (2011); and Davidson et al. (2011); see Frías, Kaplan, and Verhoogen (2011) for a fuller literature review.

I. Data and Econometric Strategy

The dataset we employ is a balanced panel of plants that can be linked between a standard plant panel, the *Encuesta Industrial Anual* (*EIA*), and employer-employee data from the administrative records of the Mexican social security agency, the *Instituto Mexicano del Seguro Social* (*IMSS*). For details, refer to Frías, Kaplan, and Verhoogen (2011).[3] In this article, we focus on 2,531 plants with complete data over the 1993–2001 period. Online Appendix Table A1 reports summary statistics by export status for 1993. It is worth emphasizing that plants in the linked EIA-IMSS sample are larger and have higher wages than the typical Mexican establishment.[4]

To fix ideas, consider the following model:

$$(1)\qquad y_{jt} = \theta e_{jt} + \gamma \hat{\lambda}_{jt} \times t + \mu_j + \xi_{kt} + \psi_{rt} + u_{jt},$$

where j and t index plants and years; y is a plant-level wage outcome; e is a measure of export status, either export share of sales or an indicator for exports being greater than zero; μ_j is a plant fixed effect; ξ_{kt} and ψ_{rt} are industry-year and region (state)-year effects; and u_{jt} is a mean-zero disturbance. $\hat{\lambda}$ is a proxy for a plant's underlying productivity (i.e., a Melitz-type draw), which here will be plant size, and t is year; the $\hat{\lambda}_{jt} \times t$ term captures differential trends by plant size, for reasons that will be made clear below. The primary outcomes, y_{jt}, of interest are wage levels at various quantiles of within-plant distributions; following common practice, we focus on the tenth, twenty-fifth, fiftieth, seventy-fifth, and ninetieth percentiles.[5] Note that the individual at a given percentile within a plant will not be the same over time; our results should be interpreted as characterizing changes in the shape of within-plant distributions, not in wages of particular individuals.

Export status in (1) is likely to be endogenous. Suppose, for instance, that plants are subject to unobserved time-varying productivity shocks; a positive shock would be expected both to raise wages and induce the plant to increase exports, generating positive bias in estimates of θ. On the other hand, there may be reverse causation: if firms are subject to labor cost shocks, an increase in wages may induce firms to reduce exports, generating a negative bias in estimates of θ.[6] To address endogeneity, we follow the strategy of Verhoogen (2008) and use the interaction of the late-1994 peso devaluation with initial plant size as a source of exogenous variation in the incentive to export. Online Appendix Figure A1 plots the Mexican real exchange rate over the 1989–2004 period; online Appendix Figure A2 plots export share of sales and the fraction exporter over time in our balanced panel. The magnitudes of the devaluation and the subsequent export response are striking.[7]

[2] This literature is too large to do justice to here; see Antràs and Rossi-Hansberg (2009) for a review.

[3] One detail is particularly relevant: the IMSS top-code varied over the period, and in particular was raised in 1994 and 1995. In results available from the authors, we redo our estimates using plants for which the top- and bottom-codes are never binding at ninetieth and tenth percentiles, respectively, and find that our qualitative results are robust.

[4] Daily wages in the IMSS data are lower on average than eight times hourly wages in the EIA; this may reflect part-time work. Underreporting is also a possible explanation, although Kumler, Verhoogen, and Frías (2012) investigate the extent of underreporting, comparing the administrative records to household data by region and demographic group, and find little evidence of underreporting among large manufacturing plants that respond on a regular basis to the EIA.

[5] In principle, one could write a model like (1) at the individual level and estimate a set of conditional quantile regressions. But to our knowledge the econometric literature does not yet provide an estimator that can deal simultaneously with an endogenous covariate (export status) and the incidental fixed-effect parameters. Here we focus on simple least-squares regressions and leave the development of such a method for future work.

[6] Although the EIA dataset is supposed to exclude *maquiladoras* (assembly-for-export plants), it is possible that some plants shifted to an assembly-for-export model over the study period; this shift may well have generated a negative correlation between the change in export share and the change in wages within plants that should arguably not be attributed to an effect of exporting per se.

[7] Note that wages overall fell in real terms during the peso crisis; the key fact for our purposes is that they fell relatively less in initially larger, higher–export propensity plants.

Theoretical considerations would suggest that larger, more productive plants would be more likely to increase exports in response to the shock; see Verhoogen (2008). This, in turn, suggests that initial size interacted with the shock could serve as an instrument for exporting. A potential concern with this strategy is that there may be differential trends between larger and smaller plants, independent of the effect of the devaluation on the incentive to export. For this reason, we allow for a constant differential trend between larger and smaller plants and compare two periods, the period of the peso crisis and its immediate aftermath (1993–1997) and a later period without a devaluation (1997–2001).

Formally, we first write (1) in first differences, stacking the differences for the 1993–1997 and 1997–2001 periods:[8]

$$\Delta y_{jt} = \theta \Delta e_{jt} + \dot{\xi}_{kt} + \dot{\psi}_{rt} + \gamma \hat{\lambda}_{jt} + \dot{u}_{jt}, \tag{2}$$

where $\Delta y_{jt} = y_{jt+4} - y_{jt}$; $\Delta e_{jt} = e_{jt+4} - e_{jt}$; $\dot{\xi}_{kt}$ and $\dot{\psi}_{rt}$ and $\dot{u}_{jt}$ are again an industry-year effect, a region-year effect, and a mean-zero disturbance. In the context of (2), our strategy is to instrument Δe_{jt} with $\hat{\lambda}_{jt}$ interacted with an indicator for the peso-crisis period—that is, $\hat{\lambda}_{jt} \times T_{93-97}$ where T_{93-97} is an indicator for the 1993-1997 period. The instrumental-variables (IV) estimate of θ will thus be based on changes in export status induced by the interaction of initial size and the devaluation, separate from underlying differential trends between larger and smaller plants. Here we use log employment from the IMSS data as our proxy for a plant's underlying Melitz draw, $\hat{\lambda}_{jt}$.[9]

A thorny issue of interpretation is that exports and wages in a given period are simultaneous choices, presumably outcomes of the same firm optimization problem, and it may be unclear how to think about an "effect" of one on the other. One way to sidestep this difficulty is to focus on the reduced form corresponding to the IV model and think of the devaluation as generating exogenous variation in the incentive to export. Here we present both IV and reduced-form results.[10]

II. Results

Table 1 presents simple cross-sectional regressions of wage outcomes on export status (either a 0/1 exporter indicator or export share of sales) or plant size (log employment), for the year 1993.[11] As is common in other datasets, exporting and plant size are associated with higher wages on average. Overall, there is little evidence of systematic differences in the cross-sectional wage-exporting or wage-size relationships across quantiles; the differences in estimates across columns 3–7 are generally not significant.

Table 2 examines changes in the wage outcomes over time. Panels A and B present OLS regressions of (2). Perhaps surprisingly, there is little robust evidence that within-plant increases in exports are associated with increases in wages, when controlling for differential trends by plant size. These results should be treated with caution, however, as the possible endogeneity of export status has not been addressed.

Columns 1–7 of panel C present the reduced-form estimates corresponding to our IV model. In columns 1 and 2, the significant coefficients for initial log employment interacted with the indicator for the 1993–1997 period indicate that the differential trend between larger and smaller plants was significantly greater during the peso

[8] The results are robust to changes in the beginning and end dates of the periods.

[9] Verhoogen (2008) focused on domestic sales as the preferred proxy for λ but showed that the basic patterns were robust to the choice of proxy. Here we use employment from the IMSS data to avoid difficulties arising from the fact that domestic sales enter in the denominator of export share.

[10] The reduced form can be written:

$$\Delta y_{jt} = \pi_1 \hat{\lambda}_{jt} \times T_{93-97} + \pi_2 \hat{\lambda}_{jt} + \ddot{\xi}_{kt} + \ddot{\psi}_{rt} + \ddot{u}_{jt}. \tag{3}$$

Note that estimating (3) and testing whether $\pi_1 = 0$ is equivalent to running a regression of Δy_{jt} on $\hat{\lambda}_{jt}$ and industry and region effects separately by period and testing whether the coefficients on $\hat{\lambda}_{jt}$ are equal across periods, which is the approach of Verhoogen (2008).

[11] Note that the outcome variable in column 1, log mean hourly wage from the EIA, is not directly comparable to the outcome in column 2, the mean of log daily wage from the IMSS data, both because of the hourly versus daily issue discussed above, and because of the order of taking logarithms. The results in the two columns (and also below) are qualitatively similar, however.

Table 1—Wage Outcomes, Cross-Sectional Patterns, 1993

	Log mean hourly wage (EIA) (1)	Mean log daily wage (IMSS) (2)	Quantiles of within-firm log wage distribution 10th (3)	25th (4)	50th (5)	75th (6)	90th (7)
Exporter	0.184***	0.091***	0.084***	0.096***	0.092***	0.105***	0.101***
	(0.022)	(0.015)	(0.015)	(0.017)	(0.018)	(0.019)	(0.020)
Export share	0.302***	0.117**	0.140***	0.138**	0.078	0.111*	0.192***
	(0.085)	(0.055)	(0.054)	(0.060)	(0.062)	(0.067)	(0.070)
Log employment (IMSS)	0.148***	0.084***	0.087***	0.086***	0.088***	0.081***	0.083***
	(0.010)	(0.008)	(0.007)	(0.009)	(0.009)	(0.009)	(0.010)

Notes: Table reports 21 separate regressions, of dependent variable at top against covariate at left, industry and region (state) effects. All regressions have $N = 2{,}531$ and include six-digit industry-year effects and region (state)-year effects. Export share is fraction of total sales derived from exports. Exporter indicator takes the value 1 if export share is greater than zero, and 0 otherwise. Changes are for periods 1993–1997 or 1997–2001; initial log employment refers to employment as reported in IMSS data in first year of period (1993 or 1997). Robust standard errors in parentheses.

***Significant at the 1 percent level.

**Significant at the 5 percent level.

*Significant at the 10 percent level.

crisis period than during the period without a devaluation, suggesting that the differential shock to the incentive to export raised average plant-level wages. We see that this effect differed significantly by quantile, however. Column 3 indicates that there was no such difference in differential trends at the tenth percentile. The difference in differential trends was significantly greater at the twenty-fifth, fiftieth and seventy-fifth percentiles (columns 4–6). Interestingly, the coefficients at the seventy-fifth and ninetieth percentiles are nearly identical, suggesting little increase in dispersion in the top quartile in response to the shock.

We now turn to the IV results. Column 8 of panel C, Table 2 presents the first stage. Our instrument has explanatory power: the coefficient on initial log employment interacted with the 1993–1997 indicator is positive with a t-statistic of approximately 3. Panel D presents the IV estimates. The message is similar to that of the reduced-form results: the shock to exporting had no effect at the tenth percentile, but a significant positive effect at higher quantiles, with the magnitude of the effect increasing up to the seventy-fifth percentile. (Appendix Table A2 reports reduced-form and IV estimates for various log wage ratios, which correspond to differences in estimates from Table 2. Except for the log 90–75 and 90–50 ratios, the increase in log wage ratios between quantiles is significant).

The difference between the OLS and the IV results is notable. In our view, the most likely explanation is the reverse causality mentioned above: idiosyncratic shocks to wages may adversely affect plants' competitiveness on export markets and thus generate a spurious negative bias in the OLS estimates.

The magnitudes of the coefficients in panel D of Table 2 appear large, but are plausible. Consider two plants that differ in initial log employment by one standard deviation, approximately 1.0. Column 8 of panel C, Table 2 suggests that this gives rise to a 1.2 percent greater increase in export share in the larger plant over the 1993–1997 period than the 1997–2001 period. This (arguably exogenous) relative increase in export share is associated with approximately a 3.2 percent relative increase in mean hourly wage or a 6.4 percent relative increase in mean daily wage.

III. Conclusion

This article has had the modest goal of establishing several facts about the effect of exporting on within-plant wage distributions, focusing on wages at the tenth, twenty-fifth, fiftieth, seventy-fifth, and ninetieth percentiles. There are three key findings: (i) there is no evidence of an effect of exporting on wages at the tenth percentile; (ii) the wage effects of exporting are larger at higher percentiles, up to the

Table 2—Changes in Wage Outcomes

	Δlog mean hourly wage (EIA) (1)	Δmean log daily wage (IMSS) (2)	Δ quantiles of within-firm log wage distribution 10th (3)	25th (4)	50th (5)	75th (6)	90th (7)	Δexport share (8)
Panel A. OLS								
Δ exporter	0.025**	0.009	0.001	0.003	0.003	0.016	0.023*	
	(0.011)	(0.007)	(0.008)	(0.008)	(0.008)	(0.010)	(0.012)	
initial log emp.	0.030***	0.044***	0.037***	0.043***	0.044***	0.045***	0.054***	
	(0.005)	(0.004)	(0.004)	(0.004)	(0.005)	(0.005)	(0.006)	
Panel B. OLS								
Δ export share	0.011	−0.019	−0.014	−0.056*	−0.035	−0.029	−0.004	
	(0.043)	(0.027)	(0.030)	(0.030)	(0.032)	(0.036)	(0.044)	
initial log emp.	0.031***	0.044***	0.037***	0.044***	0.044***	0.045***	0.054***	
	(0.005)	(0.004)	(0.004)	(0.004)	(0.005)	(0.005)	(0.006)	
Panel C. Reduced form and first stage								
init. log emp. ×	0.032***	0.048***	−0.001	0.030***	0.048***	0.065***	0.065***	0.012***
T_{93-97}	(0.010)	(0.007)	(0.008)	(0.009)	(0.009)	(0.010)	(0.012)	(0.004)
initial log emp.	0.016	0.021***	0.037***	0.029***	0.022**	0.015	0.024*	0.004
	(0.006)	(0.005)	(0.005)	(0.006)	(0.006)	(0.006)	(0.008)	(0.002)
Panel D. IV								
Δ export share	2.647**	3.928***	-0.058	2.455**	3.965***	5.296***	5.333***	
	(1.227)	(1.443)	(0.639)	(1.113)	(1.532)	(1.945)	(2.026)	
initial log emp.	0.006	0.007	0.037***	0.021*	0.007	−0.004	0.004	
	(0.012)	(0.014)	(0.007)	(0.011)	(0.015)	(0.019)	(0.020)	

Notes: All regressions have $N = 5{,}062$ and include six-digit industry-year effects and region (state)-year effects. Export share is fraction of total sales derived from exports. Exporter indicator takes the value 1 if export share is greater than zero, and 0 otherwise. Changes are for periods 1993–1997 or 1997–2001; initial log employment refers to employment as reported in IMSS data in first year of period (1993 or 1997). T_{93-97} is indicator variable that takes the value 1 for 1993–1997 period, 0 for 1997–2001. Robust standard errors in parentheses.

***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

seventy-fifth; and (iii) there is no evidence of an increase in dispersion between the seventy-fifth and ninetieth percentiles. The third fact suggests that the increase in within-plant wage dispersion is not due solely to wage increases for top managers.

An interesting issue that remains largely unexplored is the dynamic adjustment of the wage distribution within plants. It may be that in the response to the export shock plants must initially raise the wages of engineers, technicians, and other skilled workers, but that wages at the low end catch up over the medium run. The results here are consistent with this interpretation, but data constraints limit our ability to pursue it further. Investigating such dynamics is a topic for future work.

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