

# Do Exporters Pay Higher Wages? Plant-level Evidence from an Export Refund Policy in Chile

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The impact of increased export activity on plant wages is estimated in a developing country context. To avoid potential endogenous selection problems, the empirical analysis benefits from exogenous variation in exports induced by a policy experiment—an export subsidy system implemented in Chile in 1986. Analyses using data from a panel survey of Chilean manufacturing establishments show that while the export subsidy had only a modest positive impact on the industrywide relative high-skilled wage, it significantly increased the plant-level relative high-skilled wage in medium-size establishments, which are most likely to take advantage of the subsidy and enter the export market. JEL codes: O15, F16, J30, F14

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According to the factor proportions theory of international trade (Heckscher-Ohlin), the relative high-skilled wage is expected to fall after trade liberalization in a developing country that is abundant in low-skilled labor. However, most empirical studies focusing on developing countries offer the opposite conclusion—wage inequality (between high- and low-skilled workers) actually rises with entry into the world market (Beyer, Rojas, and Vergara 1999; Hanson and Harrison 1999a, b; Gindling and Robbins 2001; Goldberg and Pavcnik 2007).<sup>1</sup> One potential cause for the growth in inequality following trade liberalization is that the increase in export activity may lead to a rise in exporters' relative high-skilled wage.

The superior performance of exporters over firms producing only for the domestic market is a well-established empirical regularity in international trade (Bernard and Jensen 1997, 1999). Firms selling abroad tend to be larger and

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1. Robertson (2004) is an exception—he finds that the relative skilled wage in Mexico fell after the country joined the North American Free Trade Agreement in 1994.

more productive and to pay higher wages, especially to high-skilled workers. It is also documented that exporters are “better” than nonexporters even before entering foreign markets—Bernard and Jensen (1999, p. 3) find that “... good firms do become exporters...” Hence, any attempt to identify the effect of exports on wages is hindered by endogenous selection. The trouble from an econometric perspective is that the larger, more productive, and higher-wage plants are the ones choosing to enter the export market. This selection process is described theoretically in Melitz (2003). A cross-sectional study relating plants’ export behavior to wages may therefore arrive at the wrong conclusions about the effects of exports on wages.

This article shows empirically that increased export activity in a developing country can lead to a higher relative high-skilled wage.<sup>2</sup> The impact of increased foreign sales on wages is estimated by taking advantage of exogenous variation in exports induced by an export subsidy that the government of Chile introduced in 1986. The Chilean policy allowed producers in industries exporting less than a government-specified threshold level to receive a subsidy of 10 percent of the value of their exports. This interesting policy experiment helps the research design avoid potential endogenous selection issues that could bias the estimate of the effect of exports on wages.

To investigate the impact of the subsidy on the relative white-collar wage, a Chilean plant-level manufacturing panel dataset covering 1979–96 is used. Levinsohn (1999), Pavcnik (2002, 2003), and Lopez and Namini (2006) have used shorter versions of the same panel. No one, however, has yet exploited eligibility for the export subsidy as a source of exogenous variation in export activity. Levinsohn (1999) and Pavcnik (2002) used the data for 1979–86 to investigate plant-level employment and productivity changes, respectively, in response to tariff reductions in the early 1980s. Pavcnik (2003) employed the 1979–86 panel to explore skill upgrading in Chilean plants. Lopez and Namini (2006) used the 1990–99 panel to provide evidence on self-selection into exporting markets.

To circumvent the endogenous selection problem, previous research has used exchange rate shocks as a source of exogenous variation in exports. Verhoogen (2008) takes advantage of the Mexican peso devaluation of 1994 to test his theory of trade, wage inequality, and product quality upgrading. As there is no cross-sectional variation in the export incentive, he evaluates the differential impact of the Mexican peso devaluation on the wages of larger and more productive plants. He finds evidence that following the crisis of 1994, white- and blue-collar wages as well as the skill premium increased more in more productive plants than in less efficient ones.

2. The terms high-skilled and white-collar labor, as well as low-skilled and blue-collar labor, are used interchangeably. The data section contains an outline of the broad subcategories included in the white- and blue-collar labor aggregates used in the Chilean data.

This article uses the cross-sectional nature of the Chilean export subsidy to evaluate its impact by comparing wages in two groups of industries before and after one group was granted access to the subsidy. The main empirical findings suggest that industry exports increased strongly in response to the subsidy. While the average plant-level relative white-collar wage rose only slightly following adoption of the export subsidy, the evidence suggests that the impact of the subsidy was quite uneven across plants. The relative white-collar wage rose significantly more in medium-size establishments—those most likely to enter the export market in response to the subsidy—than in large plants, which are more likely to increase their already existing export sales, or in small plants, which are least likely to take advantage of the export incentive.

More broadly, this article contributes to the empirical literature on international trade and plant-level performance. Most of the studies in this literature find little change in plant behavior in response to increased international trade. Levinsohn (1999) finds little impact of trade liberalization on labor markets; Clerides, Lach, and Tybout (1998) estimate no change in firms' cost structure due to increased export activity; Bernard and Jensen (1999) document no differences in productivity and wage growth for plants when they start exporting. Using an employer–employee matched data set from Morocco, Fafchamps (forthcoming) also fails to find evidence that exporters pay higher relative skilled wage, conditional on plant size. This article, along with Verhoogen (2008), is an exception to many previous investigations, perhaps because the empirical analysis here benefits from exogenous variation in the export incentive. It finds evidence that as exports increase, plant-level relative high-skilled wages rise for establishments that are most likely to enter the export market.

The article is organized as follows. Section I highlights some theoretical reasons why firms might increase (relative) wages as they start serving foreign markets. Section II describes the data and presents summary statistics. Section III outlines the Chilean export subsidy and its use in identifying the impact of increased export activity on wages. Results and discussion follow in section IV. Section V offers some concluding remarks.

## I. WHY MIGHT EXPORTERS PAY HIGHER WAGES?

Theoretically, three mechanisms could potentially explain why rising export activity would lead to higher wages: quasi-rents due to industry- and exporting-specific skills, labor quality upgrading, and efficiency wages. Any of these channels could contribute to higher wages in the export-subsidized manufacturing industries in Chile.

If the export subsidy increases exports without reducing production for domestic consumption, demand for workers in industries granted the subsidy will rise. Because of industry-specific skills and training, workers in the export-subsidized industries would extract quasi-rents as demand for their services increases but their supply is relatively fixed, at least in the short run. Even with

substitution away from domestic market production toward exports, the presence of both industry- and export-specific skills—such as knowledge of a foreign language or ability to operate sophisticated, new, or foreign machinery—would raise demand for workers with such skills, allowing them to extract quasi-rents.

Another reason for higher wages in the subsidized industries could be labor force upgrading as a result of increased demand for the exporting-specific skills. With non-negligible separation costs, upgrading of the quality of labor hired is likely to take place in the medium to long run and to result in higher wages for workers in industries granted the export subsidy.

Yet another reason for higher wages in exporting establishments might be that exports require greater worker effort. Exporters would pay higher wages to reduce shirking and to increase care in production. This argument, in the spirit of Shapiro and Stiglitz (1984), involves imperfect monitoring and is developed theoretically in Verhoogen (2008), whose general framework is consistent with all three mechanisms discussed here. The model assumes differences in preferences for product quality between developed and developing countries, with consumers in developed countries valuing quality more than consumers in developing countries (see also Murphy and Shleifer 1997). Production of quality in the model is sensitive to workers' effort, especially to high-skilled workers' effort. Therefore, to start exporting, firms in developing countries like Chile need to upgrade output quality and hence to pay higher (efficiency) wages, especially to high-skilled workers, to elicit greater effort. In this environment, an export subsidy raises exports and product quality and leads to higher wages and a higher skill premium.

Finally, there are reasons to expect that the subsidy could provide a greater incentive to enter foreign markets for large and medium-size plants than for small establishments. Empirical research (Roberts and Tybout 1997) has shown that nonexporters incur substantial sunk costs to enter the export market; these costs are incorporated theoretically in Melitz (2003) and in Yeaple (2005). Because larger and more productive establishments are more profitable in the export market, they are more willing to pay the sunk costs to enter foreign markets in response to an incentive such as the Chilean export subsidy.

Further, recent theoretical advances (Verhoogen 2008; Bustos 2007) show that there is a greater effect on wages for plants that enter the export market (extensive margin) than for plants that increase the volume of already positive exports (intensive margin) in response to an export incentive. Because large establishments are more likely to have been exporters before the subsidy is introduced than are small and medium-size plants, wages in large establishments are less likely to be affected than wages in medium-size establishments, which are most likely to enter the export market following the subsidy. Overall, the introduction of an export subsidy would be expected to boost wages and the relative high-skilled wage as a result of increased export activity. However, the effect on wages may be uneven across plants, with the greatest impact in establishments that enter the export market.

## II. DATA

The empirical analysis is based primarily on the Chilean Annual National Industrial Survey (ENIA), a panel survey carried out by the National Statistical Institute of Chile, for the 18 years from 1979 to 1996. The survey covers the universe of all Chilean manufacturing establishments with at least 10 employees. Surveying about 4,800 plants a year, it includes information on approximately 11,000 plants, for a total of more than 80,000 plant-year observations over the time span of the panel. The National Statistical Institute updates the survey annually by incorporating establishments founded during the year and excluding plants that stopped operating for any reason.<sup>3</sup>

An establishment is not necessarily a single-plant firm. However, it is reasonable to assume that most plants in the survey are single-plant firms. The data show that only about 5 percent of establishments purchase materials from other establishments within the firm, and Pavcnik (2002) suggests that about 90 percent of the Chilean manufacturing firms have only one plant.

For each establishment, the survey collects data on production, value added, sales, employment and wages, exports, investment, energy consumption, balance sheets, and other plant characteristics. The white-collar labor aggregate, unlike its wage counterpart, is divided into three subcategories: white-collar executives, white-collar administrative workers, and white-collar production workers. The data on blue-collar workers, again unlike its wage counterpart, is subdivided into blue-collar production workers and blue-collar nonproduction workers.<sup>4</sup> Data on plant-level exports were collected only after 1990, and balance sheet information, which includes the value of capital stock, is available only for 1980 (or 1981).<sup>5</sup> A plant's capital stock for years after 1980 is calculated by the perpetual inventory method based on initial book value in 1980 (or 1981) and only for plants present in the survey in 1980 (or 1981).<sup>6</sup> Plant location was collected only for the first three years of the survey, but more than half of manufacturing output in Chile is produced in the Santiago metropolitan region. All plants are classified into 29 three-digit international standard industrial classification (ISIC, revision 2) industries and then further subdivided into 89 four-digit ISIC industries.

As expected, there are big differences between exporters and nonexporters (table 1). Consistent with previous research (Aw and Batra 1999), Chilean

3. If an establishment contracts in size to fewer than 10 employees, it is no longer included in the Chilean survey and would be considered a "false" exit. Fajnzylber and Maloney (2005) report that for a similar dataset from Colombia, which does not censor for plants with fewer than 10 employees, about 36 percent of plants that fall below the 10-employee cut-off continue to operate and would be considered false exits.

4. Nearly 90 percent of blue-collar workers are production workers, and more than 75 percent of white-collar workers are executives or administrative workers (table 1).

5. Establishment age is not recorded. For all plants present in the first year of the survey (1979), age is imputed by assigning an age of one in 1979. Age is therefore measured with error.

6. See Liu (1993) for detailed description of the construction of the capital measure in this dataset.

TABLE 1. Summary Statistics on Plant Characteristics

Plant characteristic	All plants (1979–96)	Nonexporters (1990–96)	Exporters (1990–96)	No subsidy (1985)	Subsidy (1985)
Log white-collar wage	6.20	6.23	6.92	6.14	5.90
Log blue-collar wage	5.51	5.62	5.97	5.29	5.21
Log relative white-collar wage	0.69	0.61	0.95	0.85	0.69
Log total employment	3.67	3.56	4.75	3.67	3.54
White-collar labor (fraction of total)	0.21	0.21	0.26	0.20	0.19
Log total sales	9.95	9.81	11.74	9.80	9.66
Age (years)	7.19	10.24	10.72	5.98	6.23
White-collar employees	18.82	13.72	53.67	17.15	14.22
White-collar executives	0.18	0.21	0.16	0.19	0.17
White-collar administrative	0.58	0.58	0.55	0.55	0.61
White-collar production workers	0.24	0.21	0.28	0.26	0.22
White-collar female workers	0.32	0.38	0.28	0.23	0.34
Blue-collar employees	53.44	40.27	139.11	55.00	43.07
Blue-collar production workers	0.88	0.89	0.88	0.90	0.86
Blue-collar nonproduction workers	0.12	0.11	0.12	0.10	0.14
Blue-collar female workers	0.19	0.19	0.21	0.10	0.22
Fraction exporters <sup>a</sup> (1990–96)	0.22				
Exports (fraction of sales) (1990–96)			0.28		
FrExpSub <sub>it</sub>	0.96 (0.10)			0.87 (0.15)	1.00 (0.00)
FrExpSubW <sub>it</sub>	0.83 (0.31)			0.43 (0.33)	1.00 (0.00)
Number of observations	78,321	25,218	28,382	1,139	2,860

*Note:* Numbers in parentheses are standard deviations. All numbers are averages. Values for subcategories are fractions of the main category.

<sup>a</sup>Exporters are defined as plants with positive sales abroad.

*Source:* Authors' analysis based on data from the National Statistical Institute of Chile's Annual National Industrial Survey (ENIA); see text for details.

manufacturing exporters are larger than nonexporters, as measured by sales or total employment, and they pay higher wages, especially to white-collar workers.<sup>7</sup> Although comparisons of average wages between exporters and non-exporters are suggestive of the estimate of the effect of exports on wages, such comparisons may be contaminated by endogenous selection if initially higher-wage plants choose to enter the export market. The next section presents the econometric strategy that deals with such selection issues.

### III. ECONOMETRIC STRATEGY

To identify the effect of increased export activity on wages, this article takes advantage of the exogenous variation in exports induced by the Chilean export subsidy implemented in 1986 and still in effect. The industry- and plant-level econometric models are based on the comparison of wages between two groups of plants before and after one group is granted access to an export subsidy. In essence, this is a difference-in-differences framework, with the group of plants given access to the export subsidy in 1986 and thereafter serving as the treatment group and the others as the control group.

#### *Data Preparation*

On December 19, 1985, the Chilean government established a system of refunds for exporters of nontraditional exports.<sup>8</sup> This policy was adopted partly in response to the severe recession of 1982–83 in Chile—the government sought to diversify exports and stimulate growth through export orientation. Intended to refund duties paid by exporters on imported materials, the program instead simply refunds a percentage of the value of exports—10 percent at the time the system was established. Only products whose export value averaged less than US\$2.5 million for the two years before 1985 qualified for the rebate at its inception in 1986.<sup>9</sup> The Ministry of Economic Affairs, Development, and Reconstruction was responsible for establishing an annual list of products excluded from the benefit for the following year—referred to here as the “exclusion list.” This list remained unchanged for the first three years and changed little thereafter. In 1989, the government expanded eligibility, offering a 5 percent rebate to a small number of previously excluded products. In 1992, a 3 percent export refund was implemented for a few product codes previously excluded from the 10 percent and 5 percent rebates. Because these subsidies were much smaller and applied to only a few product codes, they are ignored. If the export subsidy had an effect on wages, this exclusion

7. Plant-level average wage is a measure of daily wages in constant 1979 Chilean pesos per worker, as information is available only on days of operation, but not on hours.

8. Congress of Chile (Congreso Nacional de Chile) Law 18480 ([www.congreso.cl](http://www.congreso.cl)).

9. This threshold (later applied for years after 1986) increased progressively from US\$2.5 million to \$5 million to \$7.5 million toward the end of the sample period.

will bias the estimates slightly toward zero, since a few plants that received access to a small export subsidy are classified as ineligible.

A concordance was employed to translate products on the exclusion list, which identifies products by the six-digit harmonized system (HS), into industries at the four-digit ISIC classification used in the Chilean Annual National Industrial Survey (Hoekman, Mattoo, and English 2002). Twenty-one four-digit ISIC industries contain at least one product that is ineligible for the export subsidy, and the average for this set of 21 industries is seven excluded products. All products in the remaining 68 industries are allowed the 10 percent refund for their exports.

This information is used to construct an industry measure of export subsidy eligibility,  $FrExpSub$ , which is equal to the fraction of products in each industry eligible for the subsidy. For the 68 industries that produce only eligible products,  $FrExpSub$  is equal to one. For the 21 industries that produce at least one ineligible product,  $FrExpSub$  is some fraction between zero and one, with an average of 0.87 and a standard deviation of 0.15.

Although only a small fraction ( $1 - 0.87 = 0.13$ ) of product codes is ineligible for the export subsidy on average, a much larger fraction of output is likely excluded because not all manufacturing products in the six-digit HS classification are produced. Also, the eligibility cut-off for the export subsidy is based on the volume of exports, not on the share of output exported. This means that the subsidy rule excludes the product lines with some of the largest volumes. Thus, if one could calculate the fraction of output ineligible for the export subsidy in the 21 industries with at least one product line excluded from the export rebate, it would be larger than 0.13 (the average fraction of products excluded from the subsidy). Unfortunately, no output data are available at the six-digit HS level; otherwise, calculating the fraction of output in a four-digit ISIC industry truly excluded from the export subsidy would be trivial. However, based on the dollar amount limit in the subsidy rule and the fact that exports constitute about 0.28 of total sales for the average exporting establishment (table 1), an approximate statistic of 0.50 can be computed for the average output excluded from the export subsidy in the set of 21 industries with at least one ineligible product line.<sup>10</sup>

A benefit of performing the analysis at the more aggregated four-digit ISIC level is that plants may change their (six-digit HS) product selection to benefit from the export subsidy, but in the data, plants (almost) never change their four-digit ISIC

10. The method used to calculate the fraction of output in industries classified as ineligible that comes from product lines truly excluded from the export rebate can be illustrated with the following example. If the three product lines a, b, and c all map into industry A, classified as ineligible for the export subsidy, and both a and b, but not c, are excluded from the refund, there must be at least US\$2.5 million worth of exports of each product a and b. Since establishments servicing foreign markets export about 0.28 of their output (see table 1), total production of a and b must be at least US\$9 million each, and therefore there must be at least US\$18 million worth of output in industry A that is truly excluded from the export subsidy.

industry of operation. Thus, even if a plant in an industry classified as ineligible ( $\text{FrExpSub} \in [0,1)$ ) changes product selection to potentially benefit from the export subsidy, the estimated impact would understate the true effect, since a positive effect of increased export activity on (relative white-collar) wages is expected.

While there are no output data at the six-digit HS product level, export data for eligible and ineligible products can proxy for their output. Thus, for each four-digit ISIC industry, the fraction of exports of eligible products in total exports can be computed,  $\text{FrExpSubW}$ , and used as a proxy for the fraction of output in each industry eligible for the export subsidy. This measure can potentially provide more precise industry eligibility information than does the measure based on the fraction of eligible products,  $\text{FrExpSub}$ .

A complication is that export data at the six-digit HS product level are available starting only in 1990, four years after the subsidy was implemented.<sup>11</sup> Because exports respond to the incentive offered by the subsidy, using exports from 1990 as weights to construct the subsidy eligibility from 1986 onwards is problematic.<sup>12</sup> For this reason, the industry subsidy eligibility variable using the 1990 product-level exports as weights is constructed, but it is used as a robustness check and not in the baseline specification. The summary statistics for the new subsidy eligibility variable,  $\text{FrExpSubW}$ , show that about 43 percent of exports (output) are eligible for the subsidy in the set of 21 industries with at least one ineligible product line (see table 1). This fraction (0.43) is very similar to the 0.50 computed using  $\text{FrExpSub}$ , as outlined above.

Pre-existing differences (as of 1985, the year before the subsidy was implemented) between establishments in the industries eligible for the subsidy and those not eligible are presented in table 1. Establishments in the subsidized industries tend to start out with slightly lower white-collar, blue-collar, and relative white-collar wages. They are also smaller, as measured by total employment or sales. The pre- and post-subsidy trends in exports and total sales for both eligible and ineligible industries suggest that the growth rates of exports and total sales were similar before 1986 when the export subsidy was implemented (see table 2). The growth rate of total sales for the eligible industries was somewhat lower than that of ineligible industries, but the difference is quite small ( $-0.02$  and  $0.01$ ). After 1986, however, exports grew much faster for the eligible industries than for the industries with excluded products. But total sales grew at very similar rates, suggesting substitution of sales away from domestic into foreign markets for the subsidy-eligible industries.

11. The six-digit HS product-level export data are from the United Nations Comtrade database (<http://comtrade.un.org>).

12. This procedure is problematic only for the 21 industries that contain at least one ineligible product line. For the remaining 68 industries, the treatment measure would be equal to one with or without weighting by the export fraction of eligible products.

TABLE 2: Average Annual Growth Rates of Industry Exports and Total Sales before and after the Export Subsidy

Subsidy status	Pre-subsidy (1980–85)			Post-subsidy (1986–93)		
	Exports	Total sales	Exports/sales	Exports	Total sales	Exports/sales
No subsidy	0.10	0.01	0.10	0.17	0.16	0.02
Subsidy	0.10	−0.02	0.13	0.43	0.15	0.28

*Note:* All numbers are averages.

*Source:* Authors' analysis based on data from the National Statistical Institute of Chile's Annual National Industrial Survey (ENIA) and export data from the World Bank and United Nations; see text for details.

### Model and Analysis

Formally, the identification strategy at the industry level can be written as:

$$(1) \quad \log w_{jt} = \alpha_1 \text{FrExpSub}_j * I_{1986} + \mu_j + \tau_t + \varepsilon_{jt},$$

where  $\log w_{jt}$  is the natural logarithm of the white-, blue-, or the relative white-collar average wage in industry  $j$  (four-digit ISIC) in year  $t$ ,  $t = 1979, 1980, \dots, 1996$ , and  $I_{1986}$  is a dummy variable equal to one for 1986 and thereafter and equal to zero otherwise. Industry,  $\mu_j$ , and year,  $\tau_t$ , fixed effects, are included to absorb interindustry wage differentials and aggregate shocks. As already defined,  $\text{FrExpSub}_j$  is the fraction of products in each industry eligible for the subsidy. The coefficient of interest is  $\alpha_1$ , which represents the impact of increasing the share of subsidy-eligible products in a given industry on the average industrywide wage.

Further, to take advantage of the detailed plant-level panel data, the following econometric model is specified:

$$(2) \quad \log w_{ijt} = \beta_1 \text{FrExpSub}_j * I_{1986} + \mathbf{X}_{ijt} \beta_2 + \lambda_i + \tau_t + \xi_{ijt},$$

where  $\log w_{ijt}$  is the natural logarithm of the white-, blue-, or the relative white-collar average daily wage for plant  $i$ , in industry  $j$  (four-digit ISIC) in year  $t$ ,  $t = 1979, 1980, \dots, 1996$ . Relevant plant-level controls, such as plant age and the fraction of female employees, are included in the vector  $\mathbf{X}_{ijt}$ . Plant-specific time-invariant heterogeneity is captured in the plant fixed effect,  $\lambda_i$ . The coefficient of interest here is  $\beta_1$ , which represents the impact of increasing the share of subsidy-eligible products in a given industry on the average plant-level wage.

With an export incentive such as the subsidy, the least productive (smallest) plants, which had likely served only the domestic market, would not enter the export market because it would not be profitable enough to do so (Melitz 2003). Hence, wages in small establishments would not necessarily be affected by the export policy. At the other end of the spectrum, the most productive

(largest) plants would find exporting much more rewarding. Many of them were probably already serving international markets before the subsidy was introduced. The new export incentive would likely induce the largest establishments to increase their already positive exports (intensive margin). Recent theoretical developments (Verhoogen 2008; Bustos 2007), however, suggest that wages (and productivity) would be affected mainly in establishments that enter the export market (extensive margin) as a result of the export incentive. In that case, the large plants that had previously served foreign markets might not experience large wage effects following the subsidy.

There is an additional reason why wages in very large establishments might not respond to the incentive. Because the eligibility cut-off is in volume of exports at the product level, a plant would not be eligible for the subsidy if it exports more than the cut-off volume. This is most likely to happen with large establishments, which tend to be the large exporters. If small, medium size, and large plants manufacture the same products, all plants would then be ineligible. If, however, small and medium-size establishments manufacture a different set of products than do large plants within the same four-digit ISIC industry, only the large plants would be ineligible for the export rebate and their wages would not change following the subsidy.

To assess the within-industry heterogeneity that might arise in response to the export subsidy, industry-level eligibility is interacted with plant size, proxied by the logarithm of the plant's initial employment.<sup>13</sup> In particular, the following specification is estimated:

$$(3) \quad \log w_{ijt} = \gamma_1 \text{FrExpSub}_j^* I_{1986} + \gamma_2 \text{FrExpSub}_j^* \log(\text{Employment})_{ij} \\ + \gamma_3 \text{FrExpSub}_j^* I_{1986}^* \log(\text{Employment})_{ij} + \lambda_i + \tau_t + \zeta_{ijt}.$$

As discussed, the impact of the export subsidy might not be linear in plant size. In particular, the export subsidy might not affect small and large plants as much as medium-size plants. To this end, two additional specifications are estimated. One is similar to equation (3) but includes a quadratic term in plant size. Another explicitly splits the sample into three groups of plants based on initial employment—small (10–16 employees), medium (17–32 employees), and large (more than 32 employees):

$$(4) \quad \log w_{ijt} = \delta_1 \text{Small}_i^* \text{FrExpSub}_j^* I_{1986} + \delta_2 \text{Medium}_i^* \text{FrExpSub}_j^* I_{1986} \\ + \delta_3 \text{Large}_i^* \text{FrExpSub}_j^* I_{1986} + \mathbf{X}_{ijt} \delta_4 + \mu_j^S * \text{Small}_i + \mu_j^M * \text{Medium}_i \\ + \mu_j^L * \text{Large}_i + \tau_t^S * \text{Small}_i + \tau_t^M * \text{Medium}_i + \tau_t^L * \text{Large}_i + \lambda_i + \tau_t + \psi_{ijt},$$

13. Initial rather than contemporaneous employment is used because it is exogenous with respect to the export subsidy. The two measures are highly positively correlated, with a correlation coefficient of about 0.80. The results are similar if initial sales are used instead as a proxy for plant size.

where *Small*, *Medium*, and *Large* are the three plant size group indicators.<sup>14</sup> To check how robust the results are to the size category cut-offs used, another version of specification 4 that uses the (within-industry) deciles of the size distribution as cut-offs is also estimated.

Problems with heteroskedasticity and serial correlation arise in a plant-level panel data setup such as this one, but the consequences from potential serial correlation can be even more severe if the identification involves difference-in-differences with multiple time periods, where the main variable of interest varies by industry rather than by plant. To solve this problem, this study follows Bertrand, Duflo, and Mullainathan (2004), who recommend calculating robust standard errors clustered by industry, not by industry-year cell or by plant.<sup>15</sup> For the industry-level regressions specified in equation (1), standard errors are also clustered by industry.

#### IV. RESULTS

This section examines the effect of the export subsidy on exports and wages.

##### *Effects of the Subsidy on Exports*

The survey did not record plant-level exports before 1990, which makes it impossible to estimate the impact of the subsidy on plant-level exports since the subsidy was introduced in 1986. However, industry-level (four-digit ISIC) export data for the period before and after the introduction of the subsidy are available from the World Bank (Nicita and Olarreaga 2001).<sup>16</sup>

The following equation is estimated for industry-level exports:

$$(5) \quad \log \text{Exports}_{jt} = \phi_1 \text{FrExpSub}_j^* I_{1986} + \mu_j + \tau_t + \mu_j^* \text{time}_t + \eta_{jt},$$

which includes an industry-specific time trend,  $\mu_j^* \text{time}_t$ , to further capture any differences in export trends across industries. As expected, when industries gain access to the export subsidy, their exports rise (table 3, regression 3.1). To control for industry-specific business cycle fluctuations in exports, a full set of interactions between the economywide GDP growth rate and the industry

14. The plant size cut-offs divide the initial size distribution of all plants into three equal parts. The average establishment-level exports (in 1990) for large plants are more than US\$7 million—higher than the product line cut-off for the export subsidy throughout the sample period. It is therefore unlikely that large plants that export predominantly in one product category would qualify. The average establishment-level exports for medium-size and small plants are about 10 times smaller than those of large plants, well below the cut-off. If medium-size and small plants manufacture products that are different from those produced and exported by large establishments, the introduction of the export subsidy will strengthen the export incentive for them.

15. This estimator of the variance–covariance matrix is consistent in the presence of any correlation pattern within industries over time.

16. The export data from the World Bank cover 1980–96 (Nicita and Olarreaga 2001). Mirrored exports (Chilean exports as recorded by export partners) are used because they tend to be more accurate.

TABLE 3. Impact of the Export Subsidy on Industry-level Exports and Total Sales

Variable	Log exports			Log sales			Log exports/sales		
	3.1	3.2	3.3	3.4	3.5	3.6	3.7	3.8	3.9
FrExpSub <sub>it</sub> *I <sub>1986</sub>	1.79** (0.91)	1.63* (0.90)		-0.10 (0.21)	-0.13 (0.22)		1.89** (0.95)	1.75** (0.90)	
FrExpSubW <sub>it</sub> *I <sub>1986</sub>			0.55** (0.25)			0.01 (0.13)			0.56* (0.32)
Industry-specific GDP growth-rate interaction		Yes	Yes		Yes	Yes		Yes	Yes
R <sup>2</sup>	0.91	0.93	0.93	0.97	0.97	0.97	0.86	0.88	0.88
Number of observations	1,080	1,080	1,080	1,080	1,080	1,080	1,080	1,080	1,080

\*Significant at the 10 percent level; \*\*significant at the 5 percent level; and \*\*\*significant at the 1 percent level.

*Note:* Numbers in parentheses are robust standard errors clustered by industry. All regressions contain a full set of year and (four-digit ISIC) industry dummy variables, as well as (four-digit ISIC) industry-specific time trends.

*Source:* Authors' analysis based on data from the National Statistical Institute of Chile's Annual National Industrial Survey and export data from the World Bank and United Nations; see text for details.

dummies is included (regression 3.2). The estimates from regressions 3.1 and 3.2 are quite similar. The estimated coefficient of 1.79 (regression 3.1) is positive and statistically significant, implying that if an industry's fraction of eligible products rose 0.10, or 10 percentage points (the sample standard deviation of  $\text{FrExpSub}_{it}$ ; see table 1), exports in that industry would rise 17.9 percent.<sup>17</sup> While this estimate is statistically significantly different from zero, it is not significantly different from 10 percent, implying that the hypothesis that the elasticity of exports with respect to  $\text{FrExpSub}_{it}$  is unity cannot be rejected. The results with the alternative subsidy eligibility measure,  $\text{FrExpSubW}_{it}$ , (regression 3.3) imply that if an industry's fraction of eligible output rose 0.31, or 31 percentage points (the sample standard deviation of  $\text{FrExpSubW}_{it}$ ; see table 1), exports in that industry would rise 17.1 percent. This estimate is quite similar to that obtained using  $\text{FrExpSub}_{it}$ .

It is informative to see how the expansion of exports affected overall industry sales. As the export subsidy was introduced, the producer price for exports increased relative to the producer price in the domestic market. This could, for example, induce some substitution away from production for the domestic market and toward production for the international market. Another possibility is that there would be no substitution but only an increase in exports. This can happen if the production decisions for the domestic market and for exports are decoupled.<sup>18</sup>

The estimates of the effect of the export subsidy on total industry sales (see table 3, regressions 3.4–3.6) are close to zero, indicating that the export subsidy had little impact on industry output. Hence, domestic industry sales must have declined almost as much as industry exports rose, leading to a nearly one-for-one production substitution. Supporting this conclusion are the last three specifications in table 3 (regressions 3.7–3.9), which report the estimated impact of the export subsidy on the ratio of industry exports to sales. The coefficients are very similar to the effect of the subsidy on exports for regressions 3.1–3.3.

### *Effects of the Subsidy on Wages*

While imprecisely estimated, the effect of the export subsidy on the industry white-collar wage is close to zero (table 4). The estimated coefficients are  $-0.02$ ,  $0.15$ , and  $0.10$  (regressions 4.1–4.3). There is a small negative effect of the export subsidy on the industry blue-collar wage (regressions 4.4–4.6). Consequently, the results indicate a moderate positive effect on the relative white-collar wage (regressions 4.7–4.9). The estimates imply that if the fraction of industry products eligible for the export subsidy increases 0.10

17. As pointed out earlier, excluding 13 percent of products in an industry from the export subsidy ( $\text{FrExpSub}_{it} = 0.13$ ) is equivalent to excluding about 50 percent of industry output. Therefore, the impact of increasing the industry fraction of products eligible for the subsidy by 0.10 is roughly equivalent to increasing the industry output eligible for the subsidy by 0.50, or 50 percentage points.

18. That is the theoretical set-up in Verhoogen (2008).

TABLE 4: Impact of the Export Subsidy on Industry-level Wages

Variable	Log white-collar wage			Log blue-collar wage			Log relative white-collar wage		
	4.1	4.2	4.3	4.4	4.5	4.6	4.7	4.8	4.9
FrExpSub <sub>it</sub> *I <sub>1986</sub>	−0.02 (0.54)	0.15 (0.45)		−0.20 (0.60)	−0.49 (0.30)		0.18 (0.16)	0.63** (0.30)	
FrExpSubW <sub>it</sub> *I <sub>1986</sub>			0.10 (0.17)			0.00 (0.14)			0.10 (0.12)
Industry-specific GDP growth-rate interaction		Yes	Yes		Yes	Yes		Yes	Yes
R <sup>2</sup>	0.95	0.97	0.97	0.96	0.98	0.98	0.52	0.65	0.65
Number of observations	1,210	1,210	1,210	1,210	1,210	1,210	1,210	1,210	1,210

\*Significant at the 10 percent level; \*\*significant at the 5 percent level; and \*\*\*significant at the 1 percent level.

*Note:* Numbers in parentheses are robust standard errors clustered by industry. All regressions contain a full set of year and (four-digit ISIC) industry dummy variables, as well as (four-digit ISIC) industry-specific time trends.

*Source:* Authors' analysis based on data from the National Statistical Institute of Chile's Annual National Industrial Survey and export data from the World Bank and United Nations; see text for details.

( $\text{FrExpSub}_{jt} = 0.10$ ), the relative industry white-collar wage would rise 1.8–6.3 percent. Alternatively, if the fraction of industry output eligible for the subsidy increases 0.31 ( $\text{FrExpSubW}_{jt} = 0.31$ ), the relative white-collar wage would rise 3.1 percent.

The next set of results, from equation (2), estimate the effect of the export subsidy on the average plant-level wage using the detailed plant-level data from the Chilean Annual National Industrial Survey. Similar to the estimates with aggregate industry data, the results indicate that the export subsidy has almost no impact on the average plant-level white-collar wage (table 5, regression 5.1). The next specification (regression 5.2) includes a number of covariates—plant age and age squared (see Brown and Medoff 2003) and the fraction of female white-collar employees. The estimated nonlinear effect of plant age shows that wages increase with age but at a decreasing rate. Additionally, having a higher fraction of female employees is associated with a lower plant-level wage. The coefficient on the export subsidy, however, does not change much from that in regression 5.1.

To assess the impact of labor composition on the average plant-level white-collar wage, the next specification explicitly controls for the fraction of white-collar administrative workers and executives (regression 5.3).<sup>19</sup> Plant-level white-collar labor composition may change in response to the export subsidy. For example, a firm may choose to substitute one type of white-collar labor for another, substituting a relatively less expensive type for one that has become relatively more expensive as a result of the subsidy. As expected, plants with a larger fraction of white-collar executives also have higher white-collar wages. The estimate of the effect of the export subsidy, however, while slightly higher, is very similar to the previous estimates (regressions 5.1 and 5.2). Using the alternative measure of subsidy eligibility,  $\text{FrExpSubW}_{jt}$ , produces similar results (regression 5.4).

Equation (2) is next re-estimated using the blue-collar plant-level wage as a dependent variable. The results reveal that the export subsidy had a small negative impact on blue-collar wages (see table 5, regressions 5.5–5.8). The effect is similar to that obtained with industry-level data. The results further indicate that the export subsidy had a small positive impact on the relative white-collar wage (see table 5, regressions 5.9–5.11). The estimates imply that if the fraction of industry products eligible for the export subsidy rose 0.10 ( $\text{FrExpSub}_{jt} = 0.10$ ), the average plant-level relative white-collar wage would rise 1.51–1.57 percent. Alternatively, the results imply that if the fraction of industry output eligible for the subsidy increases 0.31 ( $\text{FrExpSubW}_{jt} = 0.31$ ), the relative white-collar wage would rise 0.90 percent.

Overall, the industry-level and the average plant-level results show that adoption of the export subsidy was followed by a mild increase in the relative white-collar wage, driven mostly by the small decline in the blue-collar wage.

19. The omitted category is white-collar production workers.

TABLE 5. Impact of the Exports Subsidy on Plant-level Wages

Variable	Log white-collar wage				Log blue-collar wage				Log relative white-collar wage		
	5.1	5.2	5.3	5.4	5.5	5.6	5.7	5.8	5.9	5.10	5.11
FrExpSub <sub>it</sub> *I <sub>1986</sub>	−0.008 (0.073)	0.006 (0.072)	0.011 (0.071)		−0.159 (0.106)	−0.146 (0.103)	−0.147 (0.103)		0.151 (0.099)	0.157 (0.094)	
FrExpSubW <sub>it</sub> *I <sub>1986</sub>				−0.013 (0.039)				−0.041 (0.040)			0.029 (0.057)
Plant age		0.080*** (0.003)	0.082*** (0.003)	0.048*** (0.004)		0.080*** (0.003)	0.080*** (0.003)	0.044*** (0.002)		0.001 (0.003)	0.003 (0.004)
Plant age squared		−0.001*** (0.000)	−0.001*** (0.000)	−0.001*** (0.000)		−0.001*** (0.000)	−0.001*** (0.000)	−0.001*** (0.000)		−0.001* (0.000)	−0.001* (0.000)
White/blue-collar female workers		−0.024 (0.037)	−0.020 (0.034)	−0.023 (0.034)		−0.199*** (0.045)	−0.191*** (0.037)	−0.196*** (0.037)		0.022 (0.038)	0.027 (0.038)
White-collar executives			0.145*** (0.015)	0.146*** (0.016)							
White-collar administrative workers			0.015 (0.012)	0.012 (0.013)							
Blue-collar non-production workers							−0.094*** (0.031)	−0.095*** (0.032)			
R <sup>2</sup>	0.98	0.99	0.99	0.99	0.98	0.99	0.99	0.99	0.67	0.68	0.68
Number of observations	78,321	78,321	78,321	78,321	78,321	78,321	78,321	78,321	78,321	78,321	78,321

\*Significant at the 10 percent level; \*\*significant at the 5 percent level; and \*\*\*significant at the 1 percent level.

*Note:* Numbers in parentheses are robust standard errors clustered by industry. All specifications contain a full set of year dummy variables and plant fixed effects. The omitted category is white-collar production workers.

*Source:* Authors' analysis based on data from the National Statistical Institute of Chile's Annual National Industrial Survey and export data from the World Bank and United Nations; see text for details.

It is somewhat surprising that an export incentive that substantially increased industry exports did not raise the relative white-collar wage as much. One possible explanation is that increased export activity does not boost wages—meaning that the positive association between wages and exporting status usually observed in the data is driven by selection. Alternatively, the lack of industrywide effects might be due to the fact that the export subsidy does not affect wages in large plants, which drive the industry results.

As discussed, if exports boost wages, there are two reasons why the subsidy might fail to do so in large plants. First, large plants may have already been serving foreign markets before the introduction of the subsidy and so would simply be increasing the volume of their exports following the subsidy.<sup>20</sup> Theory posits that the impact of the export incentive on wages will be greater for firms entering the export market (extensive margin) than for firms increasing their already positive volume of exports (Verhoogen 2008; Bustos 2007). Large plants are therefore unlikely to experience wage increases as high as those of small or medium-size plants that begin to export in response to the incentive. Second, because large plants may export more than the government specified subsidy cut-off for the products they manufacture, they may not be eligible to receive the export subsidy. As exports in eligible industries rose strongly following adoption of the subsidy, the more likely explanation is that large establishments did experience a rise in already existing exports that led to only a small change in wages, as predicted in Verhoogen (2008) and Bustos (2007).

The next set of results is from equation (3), which assesses heterogeneity in the impact of the export subsidy (table 6). The first three specifications include a linear interaction term of the export subsidy measure and plant size as represented by (the logarithm of) initial employment (regressions 6.1, 6.3, and 6.5). The slight positive effect of the subsidy on the white-collar wage and the negative impact on the blue-collar wage decline with plant size (employment). The estimated coefficients for the relative white-collar wage show that the positive impact also declines with plant size. The magnitudes imply that the positive effect of the subsidy on the relative white-collar wage declines to zero for the average size plant.

As discussed, the impact of the export subsidy on (relative) plant wages may be nonlinear in plant size. To test this hypothesis, the following three specifications also include quadratic (in employment) interaction terms in addition to the linear (in employment) interaction terms (regressions 6.2, 6.4, and 6.6). The results confirm the presence of nonlinear effects. Almost all quadratic terms are economically and statistically significant, tracing the declining portion of a parabola that opens upward. These estimates indicate a positive

20. Based on the size cut-offs for small, medium-size, and large plants defined in this study, about 6 percent of small plants, 12 percent of medium-size plants, and 40 percent of large plants export to foreign countries.

TABLE 6. Heterogeneity of the Impact of the Exports Subsidy on Plant-level Wages

Variable	Log white-collar wage		Log blue-collar wage		Log relative white-collar wage		Log white-collar labor	Log blue-collar labor	Log relative white-collar labor	Subsidies received	Log total sales
	6.1	6.2	6.3	6.4	6.5	6.6	6.7	6.8	6.9	6.10	6.11
FrExpSub <sub>it</sub> *I <sub>1986</sub>	0.063 (0.070)	0.382*** (0.117)	−0.244** (0.120)	−0.259 (0.188)	0.307*** (0.108)	0.641*** (0.138)	0.270** (0.111)	0.237** (0.093)	0.032 (0.144)	26,627.65 (55,025.27)	−0.259 (0.298)
FrExpSub <sub>it</sub> * Log employment	0.015 (0.017)	0.009 (0.097)	0.109 (0.090)	0.142** (0.067)	−0.095 (0.108)	−0.133 (0.104)	−0.240*** (0.084)	−0.753*** (0.120)	0.513*** (0.158)	−122,218.30** (55,504.28)	−0.506** (0.201)
FrExpSub <sub>it</sub> * Log employment <sup>2</sup>		0.002 (0.001)		−0.008*** (0.001)		0.010*** (0.002)	0.104*** (0.002)	0.128*** (0.003)	−0.024*** (0.004)	1,459.34** (642.03)	0.010*** (0.004)
FrExpSub <sub>it</sub> *I <sub>1986</sub> * Log employment	−0.024*** (0.009)	−0.193*** (0.046)	0.030** (0.013)	0.045 (0.049)	−0.055*** (0.013)	−0.237*** (0.042)	−0.183*** (0.045)	−0.262*** (0.037)	0.080 (0.065)	−95,961.44*** (35,026.34)	−0.152** (0.068)
FrExpSub <sub>it</sub> *I <sub>1986</sub> * Log employment <sup>2</sup>		0.020*** (0.005)		−0.002 (0.005)		0.022*** (0.005)	0.014** (0.005)	0.024*** (0.004)	−0.011 (0.008)	15,193.69*** (4,941.65)	0.020*** (0.008)
R <sup>2</sup>	0.99	0.99	0.99	0.99	0.68	0.68	0.86	0.89	0.66	0.65	0.96
Number of observations	78,321	78,321	78,321	78,321	78,321	78,321	78,321	78,321	78,321	78,321	78,321

\*Significant at the 10 percent level; \*\*significant at the 5 percent level; and \*\*\*significant at the 1 percent level.

*Note:* Numbers in parentheses are robust standard errors clustered by industry. All specifications contain a full set of year dummy variables and plant fixed effects. The omitted category is white-collar production workers.

*Source:* Authors' analysis based on data from the National Statistical Institute of Chile's Annual National Industrial Survey and export data from the World Bank and United Nations; see text for details.

impact of the export subsidy on white-collar and relative white-collar wages for smaller plants, an effect that declines quickly with plant size. The implied impact on the median plant is zero. The estimates with industry-specific time trend and industry-specific GDP growth rate interaction are similar to those in regressions 6.1–6.6 but are not reported here because of space constraints.

The results from specification 3 using white-, blue-, and relative white-collar labor as dependent variables are also reported (regressions 6.7–6.9). The estimates for white- and blue-collar labor mimic those for the relative white-collar wage, implying that the export subsidy led to increased employment of both white- and blue-collar workers in smaller plants. As a result, relative employment for white-collar labor did not change much in these establishments, in contrast to the relative white-collar wage.

While there are no data on plant-level exports before 1990, data are available on the overall amount of subsidies received by each plant, which can include other government subsidies in addition to the export subsidy. To investigate the changes in overall subsidies following the introduction of the export subsidy, equation (3) is estimated with overall subsidies received as a dependent variable. The results reveal the same pattern as that for the relative white-collar wage—a positive impact for smaller plants that quickly declines with size (see table 6, regression 6.10).

The estimates of specification 3 with total sales as a dependent variable show a pattern that is broadly similar to that for the relative white-collar wage (see table 6, regression 6.11). In this case, it implies a small negative impact on total sales for smaller plants and a greater negative effect for larger ones, perhaps as a result of the greater competition in the export market stemming from the increased export activity of small and medium-size producers. All of the patterns found for the white-, blue-, and relative white-collar wages, as well as for total subsidies and sales, are robust to inclusion of industry-specific time trends and industry-specific GDP interaction to account for trends and cyclical changes (results not reported). Estimates similar to those presented in table 6 are obtained using the alternative measure of the export subsidy eligibility,  $FrExpSubW_{jt}$  (not reported).

Overall, the estimates indicate that the relative white-collar wage rose in smaller plants, which are perhaps most likely to be affected by the export subsidy and to enter foreign markets. This could be because the skills of white-collar workers, but not of blue-collar workers, are more important in exporting—perhaps because white-collar workers know a foreign language, are able to operate the machinery that produces the higher quality output needed to compete abroad, or are involved in monitoring production workers to ensure higher quality output (Verhoogen 2008).

Further, the evidence suggests that the relative employment of white-collar labor did not change much in either smaller or larger plants due to the export rebate. Therefore, at least in the short run, the relative supply of white-collar labor must be fairly inelastic. Perhaps, as already discussed, this is due to

industry-specific and exporting-specific skills that allow incumbent white-collar employees to extract quasi-rents, at least in the short run.

Another possibility may be that employers upgrade the quality of skilled workers by replacing white-collar workers with more qualified employees so as to improve competitiveness in the new export market they enter as a result of the export subsidy. Recent empirical evidence from German employer–employee matched data in Schank, Schnabel, and Wagner (2007) suggests that after controlling for plant size and worker characteristics, the exporter’s wage premium declines (but does not vanish), which implies that exporters employ higher quality labor.

This evidence would support the idea that labor quality upgrading may be partially responsible for the increase in skilled wages as a result of the export subsidy. While this proposition cannot be tested directly with the Chilean data, which are not employer–employee matched data and do not contain information on employee education, this scenario is fairly unlikely in the short run because of the high separation costs in Chile at the time of the subsidy implementation—one month of severance pay for each year of service up to five years (and then up to 11 years after 1990; Cortazar 1997). In the long run, the upgrading of high-skilled labor quality could explain the increase in white-collar wages in smaller plants as employers strive to enter and successfully compete in foreign markets in response to the export subsidy incentive.

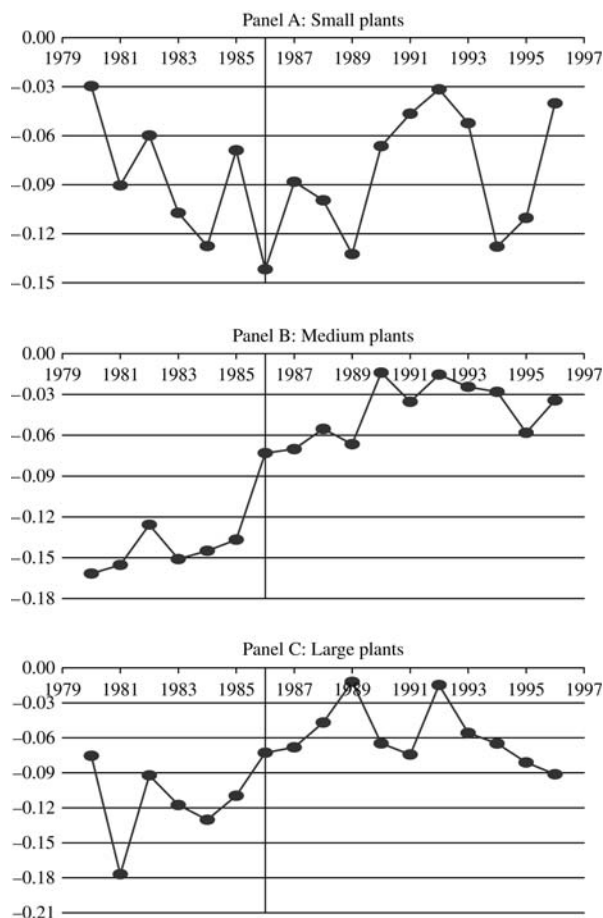
An alternative explanation that could account for the increase in relative high-skilled wages in smaller establishments without changing the relative quantity of high-skilled labor hired even in the short run is an efficiency wage model with imperfect monitoring. Employers may pay higher white-collar wages to induce higher on-the-job effort directed at better monitoring of blue-collar workers to prevent shirking.

Two more flexible specifications—equation (4), with 3 and 10 size categories of plants—provide more detailed evidence on the heterogeneity of the export subsidy impact on wages. First, however, the following equation is estimated separately for each of the three plant size categories to show the evolution of the relative white-collar wages in industries that are eligible for the export subsidy and those that are not:

$$(6) \quad \log \frac{w_{ijt}^{\text{white-collar}}}{w_{ijt}^{\text{blue-collar}}} = \lambda_i + \sum_t \sigma_t \text{Year}_t + \sum_{j,t} \kappa_{jt} \text{ISIC}_j^* \text{Year}_t + o_{ijt},$$

where  $\text{Year}_t$  is a year indicator;  $\text{ISIC}_j$  is a four-digit ISIC industry indicator; and  $\lambda_i$  is a plant fixed effect, as before. The estimates of the  $\kappa_{jt}$  terms are then split into the two groups of industries—those eligible for the export subsidy and those ineligible—and the difference between each group average for each year is plotted (figure 1). Note that the (magnitude of the) difference in the

FIGURE 1. Differences in the Relative Skilled Wage between Establishments in Eligible and Ineligible Industries (Eligible minus ineligible)



Source: Authors' analysis based on data from the National Statistical Institute of Chile's Annual National Industrial Survey; see text for details.

relative white-collar wage between the eligible and the ineligible industries declines over time, but the (magnitude of the) decrease is especially pronounced for the medium size plants, with a sharp decline in 1986, the year the export subsidy went into effect.

To formally document the changes in wages due to the export subsidy in the three plant size categories, equation (4) is estimated. With the white-collar wage as a dependent variable, and using  $FrExpSub_{it}$  as the industry measure of subsidy eligibility, the estimates are  $-0.210$  (with a standard error of  $0.203$ ) for  $\delta_1$ ,  $0.324$  ( $0.194$ ) for  $\delta_2$ , and  $-0.097$  ( $0.081$ ) for  $\delta_3$ . The estimates for the relative white-collar wage are quite similar. These results indicate that medium-size establishments experienced an increase in the white-collar wage as a result

of the export subsidy. The magnitudes imply that if the average industry's fraction of products eligible for the subsidy rose 0.10 (the sample standard deviation), high-skilled wages in medium-size plants would increase 0.032, or 3.2 percent. White-collar wages are estimated to decrease 0.97 percent in the large plants and by 2.10 percent in the small plants.

To check how robust the results are to the three size category cut-offs, another version of specification 4 is estimated that includes 10 size groups using the (within-industry) deciles of the size distribution as cut-offs.<sup>21</sup> The results support the findings for the three size categories (table 7). The positive impact of the export subsidy on both white-collar and relative white-collar wages is particularly strong for the fourth, fifth, and sixth deciles, which include the medium-size establishments.

While data limitations make it impossible to test whether exports of plants in those three deciles rose in response to the subsidy, the behavior of total sales can be examined. Consistent with the previous findings (table 6), regression 7.4 in table 7 documents that total sales for plants in the fourth, fifth, and sixth deciles rose on average more than sales in the bottom three or the top four deciles in response to the export subsidy. While the industrywide estimates implied that the increase in industry exports was not accompanied by a rise in industry sales (table 3), the plant-level results suggest that establishments most likely to enter foreign markets and increase exports in response to the subsidy did experience an increase in total sales (table 7).

## V. CONCLUSION

Using the exogenous variation in export activity brought about by an interesting policy experiment—an export subsidy—this study identifies the effect of increased exports on plant-level wages in the Chilean manufacturing industry. Avoiding potential endogenous selection, the research design uses plant-level panel data to show that while there was only a mild increase in the relative white-collar wage for the average plant as a result of the subsidy, the relative white-collar wage in smaller (medium-size) establishments—those most likely to take advantage of the subsidy and enter foreign markets—rose more substantially following implementation of the export subsidy. This evidence highlights the importance of skill in the export process, and it conforms to previous work documenting a rise in the relative high-skilled wage following trade liberalization in developing countries.

While the relative white-collar wage rose in response to the export subsidy in smaller (medium-size) plants, the relative employment of white-collar workers in these establishments did not change much. In the short run, in part due to large separation costs, incumbent white-collar workers may have gained

21. Size groups 1–10 are also based on initial employment level. Size group 1 contains the smallest establishments, and size group 10 the largest.

TABLE 7. Impact of the Exports Subsidy on Plant-level Wages by Plant-size Group

Variable	Log white-collar wage 7.1	Log blue-collar wage 7.2	Log relative white-collar wage 7.3	Log total sales 7.4
Size group 1*FrExpSub <sub>it</sub> *I <sub>1986</sub>	−0.179 (0.195)	−0.082 (0.184)	−0.097 (0.143)	0.336 (0.346)
Size group 2*FrExpSub <sub>it</sub> *I <sub>1986</sub>	0.013 (0.313)	0.436 (0.299)	−0.423 (0.120)	0.223 (0.331)
Size group 3*FrExpSub <sub>it</sub> *I <sub>1986</sub>	−0.155 (0.231)	−0.238 (0.188)	0.083 (0.226)	−0.250 (0.395)
Size group 4*FrExpSub <sub>it</sub> *I <sub>1986</sub>	0.238 (0.290)	0.089 (0.194)	0.148 (0.383)	0.119 (0.393)
Size group 5*FrExpSub <sub>it</sub> *I <sub>1986</sub>	0.463* (0.280)	−0.125 (0.126)	0.587** (0.300)	0.179 (0.319)
Size group 6*FrExpSub <sub>it</sub> *I <sub>1986</sub>	0.320 (0.433)	0.258 (0.264)	0.062 (0.268)	0.562 (0.419)
Size group 7*FrExpSub <sub>it</sub> *I <sub>1986</sub>	−0.331* (0.183)	−0.023 (0.135)	−0.308* (0.159)	−0.542 (0.528)
Size group 8*FrExpSub <sub>it</sub> *I <sub>1986</sub>	−0.400 (0.189)	−0.467** (0.193)	0.066 (0.217)	−0.933 (0.606)
Size group 9*FrExpSub <sub>it</sub> *I <sub>1986</sub>	−0.194 (0.222)	−0.572*** (0.183)	0.378 (0.282)	−0.932* (0.494)
Size group 10*FrExpSub <sub>it</sub> *I <sub>1986</sub>	−0.092 (0.195)	−0.313*** (0.102)	0.220 (0.237)	0.582 (0.476)
R <sup>2</sup>	0.99	0.99	0.68	0.96
Number of observations	78,321	78,321	78,321	78,321

\*Significant at the 10 percent level; \*\*significant at the 5 percent level; and \*\*\*significant at the 1 percent level.

*Note:* Numbers in parentheses are robust standard errors clustered by industry. All specifications contain a full set of year dummy variables and plant fixed effects, as well as full sets of year and industry dummy variables interacted with the plant size categories. Size groups 1–10 are based on initial employment level, with size group 1 containing the smallest establishments, and size group 10 the largest.

*Source:* Authors' analysis based on data from the National Statistical Institute of Chile's Annual National Industrial Survey and export data from the World Bank and United Nations; see text for details.

higher wages because of either industry- and exporting-specific skills they possess or an effort-elicitation mechanism put in place to ensure output quality suitable for foreign markets. In the long run, plant-level relative white-collar wages in smaller (medium-size) establishments may have increased because of labor quality upgrading as well.

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