# Relative Wages, Labor Supplies and Trade in Mexican Manufacturing: Evidence from Two Samples.

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**Abstract:** How do relative wages (between skilled and unskilled workers) respond to technical progress and to relative supply shifts? An empirical model of the wage premium for Mexican manufacturing is employed on two monthly data samples: one, from 1987 to 1995, displays the well documented rising trend in wages right after Mexico joined GATT; the other, from 1994 to 2007, suggests slightly decreasing wages. The model provides support for skill-biased technical change (SBTC) and yields plausible elasticity of substitution for the first sample ($\sigma = 1.03$) and higher elasticity for the second ($\sigma = 1.71$). Allowing export intensity and the real exchange rate to modify the factor augmenting technology ratio, negative relationships are found for the earlier sample: the higher the export intensity or real exchange rate the lower relative wages. The error correction methodology and the bounds approach confirm these results. Combining trade and SBTC, this study supports the view that trade considerations have an impact on wage premiums at the very beginning of trade liberalization. In contrast, the benchmark model seems a more adequate representation when NAFTA and a market-oriented peso help consolidate Mexico in its path towards sustainable growth.

**Keywords:** Export intensity, Mexico, Real exchange rate, Relative wages, SBTC, Wage premium.

**JEL Classification Number:** F16.

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## 1. Introduction

Studies of changing wage inequality across countries or industries rely on several explanations, ranging from labor supply shifts in Katz and Murphy (1992) to trade in Freeman and Katz (1991) and to technology in Berman et al. (1994, 1998). These papers are empirical but a large body of work has provided theoretical support through mechanisms that cause relative wages, or the skill premium, to grow over time. Leamer (1996) combines trade and technology within the Hecksher-Ohlin framework, while Manasse and Turrini (2001, p. 100) show, that: "As barriers to trade fall, more firms will benefit from exporting and will have access to a larger market. Competition for skills boosts skill premia in exporting firms. Hence, trade integration unambiguously leads to a redistribution of income from the workers employed in non-exporting firms to those employed in the export sector."

Studies exploring the trade hypothesis for the U.S. include Revenga (1992), who finds sizable effects of import competition on employment and wages, and Pizer (2000), who finds a negative (positive) association between import penetration (export intensity) and the wage premium between oligopolistic and competitive industries. Empirical evidence of the trade hypothesis has grown lately for less developed countries (LDCs) as well. Mexico, in particular, opened to foreign trade and slashed tariffs and quotas across all sectors from the mid to late 1980s. Perhaps this is the reason why Mexico has become a laboratory for several empirical studies on the effects of trade openness. Tybout and Westbrook (1995), for example, pursue the effects on microeconomic efficiency due to greater openness between 1984 and 1990. Feenstra and Hanson (1997) explore foreign direct investment (FDI) inflows into Mexico and the role of outsourcing by Northern

multinationals. Studies on individual characteristics include Cragg and Epelbaum (1996), Hanson (2003) on regional aspects, and Feliciano (2001) for mixed results depending on whether import license coverage or tariff reductions is used.

Following a long period of reduction from 1965 to 1980, relative wages between skilled workers and unskilled workers in Mexico started to grow after 1985. However, the widening in the wage gap came together with little change in the relative employment of skilled labor. One explanation for this, consistent with the Stolper-Samuelson theorem, is that trade increased the relative price of skill-intensive products, a hypothesis not supported by Hanson and Harrison (1999) over the 1984-1990 period in Mexico. Along these lines, Robertson (2004) regresses changes in prices on relative labor supplies and finds a positive relationship for the first period of liberalization (until 1993), but the pattern of price change reverses with NAFTA. Alternatively, Esquivel and Rodríguez-López (2003) find that trade liberalization would have led to a reduction in the wage gap in Mexico from 1988 to 1994, but that this was offset by the large negative impact of technological progress on the real wage of unskilled workers.

Some studies on Latin American economies have shown that trade explains only part of the variation in relative wages.[1] This article extends the trade hypothesis to one in which *trade and technology interact to produce changes in relative wages*. Lacking reliable data on computer usage or R&D in Mexico, such as Autor et al. (1998) on U.S. manufacturing, this paper attempts to add to the literature in three ways.

First, we use data from Mexico's INEGI from two different samples. The earlier, from 1987 to 1995, comprehends a time interval when a substantial amount of change in the degree of openness occurred. There is also an increasing wage inequality in Mexican

manufacturing as measured by wages of skilled workers relative to unskilled. On the other hand, relative labor supply of skilled labor rises from 1991 onwards. The second sample, from 1994 to 2007, shows in contrast a slightly declining wage premium. As in the previous case, relative labor supply of skilled labor rises about halfway, from 1999 onwards. The very contrasting patterns in the two samples on Mexican manufacturing could shed some light on the relationship between trade and the wage premium across different stages of the trade liberalization process.

Second, we estimate a model, borrowed from Acemoglu (2002) and Katz and Murphy (1992), in which relative wages are explained as a function of a technology trend and the relative supply of skilled (H) against unskilled workers (L). As H/L increases, the skill premium (w) should fall. But this tendency of falling w could be compensated by changes in technology. The U.S. recent experience in Berman et al. (1994, 1998) and Autor et al. (1998) has shown that a rapid increase in the relative supply of skills (H/L) occurred without a corresponding fall in w. This suggests the demand for skills must have increased to prevent the relative wages of skilled workers from declining: the so-called skill-bias technical change (SBTC) hypothesis.[2]

We provide in this paper an alternative approach to those studies that explain relative wages using only trade related variables. Recent work by Beaudry and Green (2005) finds that the Katz and Murphy (1992) model offers a poor explanation of the change in the U.S. college premium over 1976-2000. The approach in this paper is to allow trade considerations to modify the factor augmenting technology ratio, through examination of the impact of export intensity or the real exchange rate on the wage premium. In the words of Acemoglu (2002, p. 56): “SBTC accelerates, neither

exogenously nor in response to the changes in the supply of skilled workers, but due to trade opening". Autor et al. (1998, p. 1202) find that U.S. "industries with expanding export shares appear to have faster growth in nonproduction worker payroll share", a result already present in Bernard and Jensen (1997).

Trade considerations should certainly be important for Mexico, a small open economy with higher share of trade in its GDP than the U.S. Our basic insight is to explore how export intensity or the real exchange rate affects the traditional model of SBTC. Following studies on worker's characteristics [e.g., Li and Xu (2003) captures the role of export intensity in explaining the skilled labor share in China] or on real exchange rate effects [e.g., Robertson (2003) and Verhoogen (2007)], we modify the basic model to allow for export intensity or the real exchange rate.

Two applications of the relationship between relative labor supplies and wage premiums to LDCs have been attempted. The first is on a set of Latin American countries by Sánchez-Páramo and Schady (2003), who argue that in Argentina, Brazil, Chile, Colombia, and Mexico relative wages and labor supply are rising (or falling) almost monotonically. Their estimates of the elasticity of substitution ($\sigma$) are very imprecise, however, leading them to simulation exercises adopting differing values of $\sigma$ from 1 to 3. The second focuses on the export-oriented sector of Mexico by Mollick (2007), who examines capital expenditures with respect to total expenditures by Mexican maquiladoras and relates it to the growing wage premium of 16% over the years from 1990 to 2006.

Third, in addition to the specification problems associated with the omission of trade variables, there is the possibility of unit roots in the series associated with the dramatic variation in relative wages during the transition to an open economy. This leads

to cointegration and error correction modeling (ECM) techniques to take into account non-stationary data. In particular, after discussing benchmark estimations we employ two methodologies: Banerjee et al. (1998) to test for cointegration in a single-equation framework and Pesaran et al. (2001) through their bounds test approach to either I (0) or I (1) series.

Our major results are as follows. The elasticity of substitution ($\sigma$) between skilled and unskilled workers is estimated at 1.03 (earlier sample) or 1.71 (recent sample), in line with the survey by Johnson (1997) suggesting $\sigma$ is likely to be between 1 and 2. Further estimations suggest, for the earlier sample, decreases in the wage gap with higher export intensity and with a weaker peso. This is consistent with goods exports in a given industry sector being on average relatively intensive in unskilled labor. No such forces are observed in the later sample under higher trade patterns as well as a market oriented peso. The error correction methodology and the bounds approach confirm these results. Combining trade and SBTC, this study supports the view that trade considerations have an impact on wage premiums at the very beginning of trade liberalization. In contrast, the benchmark model seems a more adequate representation when NAFTA and a market-oriented peso help consolidate Mexico in its path towards sustainable growth.

This paper is organized as follows. Section 2 introduces the data, section 3 presents the theoretical framework and section 4 reports the econometric results. Section 5 reviews the work and presents extensions.

## 2. Data Description and Main Features

This study examines wages, employment, shipments, exports, and effective real exchange rate monthly data for the Mexican manufacturing sector during 1987-1995 and 1994-2007. The approach studies two different samples: the earlier sample covers 129 classes of activity and the other comprehends 205 classes in the later sample. The 1987-1995 period is such that maximal degree of change in openness occurs (the entrance of Mexico into GATT in 1986 and the NAFTA signature in 1994), while the 1994-2007 is the one that immediately witnesses the currency shock of late 1994 and early 1995, together with the NAFTA signature and subsequent stabilization of the Mexican economy under more competitive economic conditions.

The dataset comes from Mexico's Instituto Nacional de Geografia, Estadística e Informatica (INEGI: http://www.inegi.gob.mx) and is based on a representative survey of Mexican manufacturing collected at monthly frequency. More detailed comparison on the two samples can be found in Robertson (2004). As measures of industry employment, the number of production workers (L, unskilled) and non-production workers (H, skilled) are considered. This two-tier classification has been adopted by many studies, such as Berman et al. (1994, 1998) and Acemoglu and Zilibotti (2001). For the particular Mexican case, Robertson (2004) looks at additional data sources and show that production workers have less education in every industry than non-production workers. He concludes that the two-tier distinction to (imperfectly) classify skill intensity seems valid in the Mexican case.

The average nominal earnings in pesos ("salarios" for blue-collar workers; "sueldos" for white-collar workers) are deflated by the Mexican aggregate consumer price index. We define the (real) relative wage as the real wage of skilled workers divided by

unskilled workers. We next obtain the real wage adjusted for hours worked by type of workers. We finally define the (real) relative wage (**rwhr**) as the real wage of skilled workers adjusted by hours-worked, divided by the real wage of unskilled workers also adjusted by hours-worked.

Figure 1a contains the growing trend of relative wages over the earlier period, a well documented fact by Hanson and Harrison (1999), Robertson (2004) and others. In Figure 1a real relative wages corrected for hours-worked (**rwhr**) move from 2.07 to 3.20 in the period: a 55% rise. The relative participation of non-production workers in the total workforce (**lsratio** or **H/L ratio**) has fluctuated but not by much, at around 0.30. In contrast to industrial economies such as the U.S. and the U.K. that have had an increase in the share of skilled workers in employment [Berman et al. (1998)], Mexican manufacturing has not changed by much. Gera et al. (2001), for example, report no strong evidence that skill intensity increased at the aggregate level in Canada over 1981-1994.

According to Figure 1a, the relative labor supply of H/L decreases, bottoms out and then increases from 1991 onwards. Therefore, something else must be responsible for relative wages to rise when the relative supply of skilled workers rise. Figure 1b shows the flat trend for relative wages over the later period and subsequent slight decline of about 11%: from 2.91 in January of 1994 to 2.59 in January of 2007. Relative labor supplies also rise from 1999, coinciding with falling wage premiums. The latter matches the idea that an increase in the relative supply of skills should place a negative effect on the wage premium of skilled workers.

[Figures 1a and 1b here]

The following definitions are used:

*Employment:* persons employed during a month in a given establishment (or outside it but under its control), under regular compensation schemes. It includes regular workers under temporary leave, excluding temporary and retired workers;

*Hours Worked:* number of hours, regular and extra, worked during a given month, excluding suspended time due to stops, strikes, vacation, illness, or to natural factors;

*Non-Production Workers ("Empleados"):* workers engaged in planning, supervision, accounting, sales, research, and advertising, including the owner;

*Production Workers ("Obreros"):* workers engaged in manual work, operating machines and equipment, and in cleaning, repair, storing, packing, and so on;

*Wages paid to Non-Production Workers ("Empleados"):* the cash amount, before any deduction, paid in a given month to such workers. It includes productivity bonuses, incentives, vacational pay, an extra month of pay at the end of the year ("aguinaldos"), commissions on sales and temporary leave under payment. It excludes pensions to retired workers, payment of "honorarios" to workers not under the payroll and termination pay; and

*Wages paid to Production Workers ("Obreros"):* defined as in the preceding item.

Exports by industry are in U.S. dollars and shipments (or sales) by industry are in pesos. In order to define the *export intensity ratio* (**xint**), we divide mnufacturing exports by shipments, employing the nominal exchange rate (the interbank average bid rate from INEGI) to convert the denominator into U.S. dollars[3]. The export-intensity ratio for the aggregate of Mexican manufacturing, shown in Figure 2a, more than doubles over the 9-year period, especially after the peso devaluation of December 1994. The same upward adjustment is also observed in Figure 2b for the later sample.

[Figures 2a and 2b here]

The real exchange rate in Figures 3a and 3b show the one-shot like adjustment at the time of the currency crises of December 1994 and early 1995. For the earlier sample, the trend for the real exchange rate was of sustained appreciation before the collapse of the fixed exchange rate regime. After that, a gradual appreciation of the Mexican peso against major currencies once again follows suit.

[Figures 3a and 3b here]

## 3. The Models

Suppose firms hire only two types of workers: skilled (H) and unskilled (L), who are imperfect substitutes. There are $L_t$ unskilled (or low-education) workers and $H_t$ (high-education) workers, supplying labor inelastically at time t. Suppose workers are risk-neutral, maximizing the present value of labor income, and that labor markets are competitive. We proceed under a constant elasticity of substitution (CES) production function for the aggregate economy as in Acemoglu (2002):

$$Y_t = [(A_{Lt} L_t)^{\rho} + (A_{Ht} H_t)^{\rho}]^{1/\rho} \qquad (1),$$

where $\rho \leq 1$ and $A_{Lt}$ and $A_{Ht}$ are factor-augmenting technology terms. The elasticity of substitution between the two types of workers is $\sigma \equiv 1/(1-\rho)$. The two types of workers are gross substitutes when $\sigma > 1$ (or $\rho > 0$) and gross complements when $\sigma < 1$ (or $\rho < 0$).

Special cases include Leontieff fixed proportions ($\sigma \rightarrow 0$ or $\rho \rightarrow -\infty$), perfect substitution ($\sigma \rightarrow \infty$), and the CES function collapsing to the Cobb-Douglas case ($\sigma \rightarrow 1$).

Under competitive labor markets, the unskilled wage is (we omit the t subscript):

$$W_L = \partial Y/\partial L = A_L^{\rho} \, [A_L^{\rho} + A_H^{\rho} \, (H/L)^{\rho}]^{(1-\rho)/\rho} \qquad (2),$$

implying that as the fraction of skilled workers in the labor force increases, the unskilled wage should increase. Similarly, the optimal skilled wage is:

$$W_H = \partial Y/\partial H = A_H^{\rho} \, [A_L^{\rho}(H/L)^{-\rho} + A_H^{\rho}]^{(1-\rho)/\rho} \qquad (3),$$

implying that, with the rise in skilled workers, their wages must fall.

Combining (2) and (3), the skill premium (defined by the ratio $W_H/W_L$) becomes:

$$rwhr = (W_H/W_L) = (A_H/A_L)^{\rho} \, (H/L)^{-(1-\rho)} = (A_H/A_L)^{(\sigma-1)/\sigma} \, (H/L)^{-1/\sigma} \qquad (4),$$

which turns out to be, in logarithmic form:

$$\ln (rwhr) = [(\sigma-1)/\sigma] \ln (A_H/A_L) - (1/\sigma) \ln (H/L) \qquad (5),$$

where one can see that the skill premium increases when skilled workers become more scarce: the partial derivative is $-(1/\sigma) < 0$. This is the substitution effect showing that – for given skill bias of technology (the term $A_H/A_L$) – the relative demand curve is downward sloping with elasticity $1/\sigma = (1-\rho)$. A figure could plot the relative demand for skills in (5) on the vertical axis against the relative supply of skills (H/L) on the horizontal axis. An increase in the (vertical) relative supply in the (rwhr, H/L) locus moves the equilibrium point along the downward sloping relative demand curve, reducing the skill premium.

One can also see how the skill premium reacts to technology, and this will again depend on the elasticity of substitution: $[(\sigma-1)/\sigma]$. If $\sigma > 1$, then improvements in the skill-complementary technology increase w, a shift out of the relative demand curve. If, however, $\sigma < 1$, an improvement in $(A_H/A_L)$ shifts the demand curve inwards, reducing rwhr. The conventional wisdom is that w increases when skilled workers become relatively more productive, consistent with $\sigma > 1$. This is usually, but not always, supported in the data.

Note that as H/L increases, rwhr should fall, which would represent graphically a rightward shift in the relative vertical labor supply to the right. But this tendency of falling rwhr could be compensated by changes in technology in the first term of (RHS) of (5). Recent U.S. experience has witnessed a rapid increase in the supply of skills (H/L) but no corresponding fall in rwhr. This suggests the demand for skills must have increased to prevent the relative wages of skilled workers from declining. SBTC is a natural explanation since the relative productivity of skilled workers must have increased.

It is possible to explore the idea that technical change has taken placed steadily through a time trend. Bernard and Jensen (1997) have found, however, that increases in employment of exporters account for almost all of the increase in the wage gap. Supposing a greater impetus of exports to skilled-labor saving technologies than to that of unskilled labor saving, we modify the equation suggested by Acemoglu (2002) by:

$$\ln (A_{Ht}/A_{Lt}) = \gamma_0 + \gamma_1 t + \gamma_2 \ln (x) \qquad (6),$$

where: t is calendar time and x can be either xint (the degree of export intensity measured by exports/shipments in the industry) or rer (the real exchange rate). Both export intensity and real exchange rate terms can be viewed as shift factors to the labor demand curve in the Katz and Murphy (1992) model. Several studies have addressed the incorporation of trade related variables on wages. At the individual worker level, a particular set of works follows the two-stage framework of Gaston and Trefler (1994), including Pizer (2000) for the wage premium between the oligopolistic and competitive sectors in U.S. industries; the impact of trade reforms on wages by Attanasio et al. (2004) for Colombia and Pavcnick et al. (2004) for Brazil; and Milner and Tandrayen (2004) for the impact of trade reforms in manufacturing firms of six Sub-Saharan countries.[4] While there are comparatively less studies on the real exchange rate, Robertson (2003) discusses the reasons why the impact of real exchange rate movements on labor demand varies with the share of exported output, the amount of imported intermediate inputs, and the degree of complementarity between imported inputs and labor. Robertson (2003) finds, for Mexico's monthly industrial surveys, that real exchange rate appreciation has a large

positive effect on the wage premium. See also Verhoogen (2007) for quality upgrading induced by the exchange rate shock increasing wage premiums in Mexico.

Substituting (6) into (5) immediately yields:

$$\ln (rwhr) = [(\sigma-1)/\sigma]\, \gamma_0 + [(\sigma-1)/\sigma]\, \gamma_1 t - (1/\sigma)\, [\ln (H/L)] + \gamma_2 \ln (x) \qquad (7), \text{ or}$$

$$\ln (rwhr) = \beta_0 + \beta_1 t - \beta_2 \ln (H/L) + \beta_3 \ln (x) + \varepsilon \qquad (8),$$

where: $\varepsilon$ are random errors and (8) is a testable equation carrying the idea that technological developments occur at a constant rate, but the supply of skilled workers could grow at different rates, and the degree of export intensity have a positive effect on the wage premium. Ignoring trade-related considerations, when H/L grows faster than the rate of SBTC rwhr will fall. When H/L falls short of this rate, rwhr will increase. Estimating a value of -0.709 for $\beta_2$ in (8) with $R^2 = 0.52$, without trade factors, Katz and Murphy (1992) argue that $\sigma = 1.41$ for the U.S. labor market using 25 years of annual data from 1963 to 1987.

Estimation of (8) is conducted in this paper at the manufacturing level. Two problems arise with the time series estimation of (8), however. First, there is the possibility of specification errors. Omitted variables problem, such as capital stock, may be present. While (8) takes into account the substantial increase in exports at the industry level

observed in Mexican manufacturing, data availability precludes this strategy for capital stock at the sector level. Second, the series in (8) may be non-stationary.

Banerjee et al. (1998) propose an error correction modeling (ECM) test based upon the OLS coefficient of the lagged dependent variable in an autoregressive distributed lag model augmented with leads of the regressors. The test, embedded in a single equation framework, has its limit distribution not depending upon nuisance parameters but depending on the dimension of the system, is performed as follows:

$$\gamma(L)\Delta\ln(rwhr)_t = \beta_0 + \beta_1 t + \alpha(L)\Delta\ln(x)_t + \beta\ln(rwhr)_{t-1} + \phi\ln(x)_{t-1} + \Sigma\, \varphi_i\, \Delta\ln(x)_{t+j} + \varepsilon_t \quad (9a),$$

where: $x_t$ = (H/L, xint)$_t$ or (H/L, rer)$_t$, $\gamma(L)$ is a scalar polynomial and $\alpha(L)$ is a vector polynomial suitable chosen by information criteria. Lags and leads are chosen by information criteria, while j is recommended by Banerjee et al. (1998, p. 275) to be 1 or 2 for N=100.

In addition to the cointegration procedure outline above, we conduct the Pesaran et al. (2001) bounds test approach given the uncertain nature of the order of integration of the variables. Doing so, the calculated F- and t-statistics test the significance of the lagged levels of the variables in the error-correction mechanism. Since the asymptotic distributions of these statistics are non-standard under the null of no-cointegration, we compare the calculated values to the tables in Pesaran et al. (2001). The general ECM-type cointegration equation is:

$$\Delta(\ln(rwhr)_t) = c_0 + c_1 t + c_2 dum95{:}01 + \beta_1 \ln(rwhr)_{t-1} + \beta_2 \ln(H/L)_{t-1} + \beta_3 \ln(x)_{t-1} +$$

$$\Sigma\, \phi_i\, \Delta \ln (z)_{t-i} + \delta_1\, \Delta(\ln(H/L))_t + \delta_2\, \Delta(\ln(x))_t + \varepsilon_t \qquad (9b),$$

where: $x_t$ = xint or rer; and $z_t$ = (rwhr, H/L, xint)$_t$ or (rwhr, H/L, rer)$_t$ and the appropriate test is a t-test on $\beta_1$ or a F-test on $\beta_1 = \beta_2 = \beta_3 = 0$. As before, lags of differenced variables are chosen according to statistic criteria. The general-to-specific methodology chooses the lag-length of the differenced terms in the right hand side of (9b).

## 4. Results

Table 1 contains standard unit root tests, such as the sequential Augmented Dickey-Fuller (ADF) methodology recommended by Ng and Perron (1995) and the KPSS (1992) tests under the stationary null. Details on lag selection are provided at the notes to Table 1. In general, the standard tests are supportive of unit roots. The existence of a unit root in relative wages could be, however, a matter of concern since it would be difficult for an economic model to predict that relative wages grow without bound.

Further examination is pursued. The sequential unit root tests by Ayat and Burridge (2000) in Tables 2 and 3 also reject the unit root null (the $\rho - 1$ coefficient must be different from zero for the rejection), other than a few specifications with only the constant term, which are less preferable to specifications allowing for trend terms. One must recall the trend-like pattern in wages, exports, or the real exchange rate discussed in Section 2. In addition to these sequential tests in general suggesting rejections of the unit root null and the fact that export intensity is subject to level shifts, we perform the Zivot-Andrews

(1992) test referred to as the "crash model". The test correctly identifies the 1995:01 breakpoint on the two samples and suggest stationary series when allowing for the break.

[Tables 1, 2, and 3 here]

Since the evidence on unit roots is mixed, we proceed with estimations of both standard OLS as well as cointegration-based methods. In the estimations that follow, the elasticity of substitution is important for the behavior of wage premium when labor supply changes. Let the basic assumption be that the industry level labor supply is perfectly inelastic, implying that - *given exogenous quantities* - *wages are endogenously determined*. This assumption has its major support on the level of data aggregation used: industries at the 1-digit level are presumably closer to industry (or to the entire nation) than to the firm level.[5] It is also consistent with the theoretical model above.

Table 4 reproduces OLS estimates of (8) for the benchmark and augmented specifications. All standard errors in the estimates are computed with the Newey-West method for correction for heteroskedascity and serial correlation. We focus on the parameter associated with the relative supply of workers, as in Katz and Murphy (1992), and leave the econometric fit of the estimates for further analysis below. We handle first the estimates of the elasticity of substitution between the two labor types derived by computing $(-1/\sigma) = \beta_2$ associated with ln (H/L) in (8).

According to Table 4, for the benchmark model the implied elasticity of substitution is 1.03 for the 1987-1995 sample, implying that the wage premium moves in tandem with the relative supply of skilled workers. As relative supply of skilled workers rise, their relative wage falls: the $\beta_2$-coefficient equals -0.97. This magnitude appears

reasonable since recent empirical estimates suggest $\sigma$ is likely to be between 1 and 2 with an emerging consensus "best guess" estimate at around 1.4 to 1.5 in Johnson (1997). For the 1994-2007 sample, the $\beta_2$-coefficient is estimated to be -0.584, implying a higher elasticity of substitution of 1.712.

[Table 4 here]

Support for SBTC is obtained by the estimation of the time trend. The time trend coefficient for the basic model in Table 4 is always statistically significant and positive for the 1987-1995 sample and negative for the 1994-2007 sample. Positive values of the term trend coefficient are in agreement with skill biased technical change. For the latter sample, however, the SBTC result is reversed. Katz and Murphy (1992) report an overall value for the U.S. labor market over 1963-1987 of a $\beta_1$-coefficient of 0.033 for annual data.

It is interesting to explore how the basic model changes with the modifications to (6). For the 1994-2007 sample, the elasticity of substitution is relatively high at 2.506 for the model with export intensity. This value of $\sigma$ implies that a rise in H/L encourages so much SBTC that the demand for skills increases more than enough to offset the potential increase in the supply of skills.[6]

Overall, specification issues suggest problems with (8) in its original form. One finds, for example, widespread serial correlation in the estimates of Table 4 as documented by complementary Breusch-Pagan Lagrange Multiplier tests and Q-statistics. In all cases, the null of no serial correlation is rejected at standard significance levels. The serial correlation problems may be caused by either model misspecification ("omitted

variables"), the presence of unit roots ("spurious regressions"), or by a combination of both.

In order to handle these problems, our modified framework has trade and SBTC with complementary roles in a new theory of wage dispersion. The rationale for the estimations herein is that increased trade must cause a change in the path of technological progress. We then modify the basic model for trade-related characteristics, of which export intensity is explicitly employed in (8b) and the real exchange rate appears in (8c). Table 4 contains these results, in which the export intensity and real exchange rate coefficients are negative for the 1987-1995 sample: -0.089 and -0.097, respectively. The interpretation is as follows: a higher degree of export intensity leads to a fall in the wage premium, consistent with the notion that higher levels of trade lead to a higher (external) demand for the relatively abundant labor type (unskilled labor) in Mexico. Similarly, a depreciation of the real exchange rate makes the wage premium to shrink, tending to benefit more unskilled workers. One possible channel is that if skilled workers are more complementary with imported intermediate inputs, an appreciation of the real exchange rate should favor skilled workers, as in Robertson (2003).

Compared to the benchmark model, despite finding some improvement in the specification measured by the Durbin-Watson statistics, the values of the not reported Q-statistics at 36 lags are still large enough to suggest specification problems.[7] In other words, accounting for omitted variables alleviates serial correlation problems but does not solve them completely. Also, for the later sample, the results are reversed with positive coefficients estimated for the trade-related variables: 0.072 and 0.048, respectively, casting doubt on the interpretation above.

We next check the existence of cointegration in the modified model by the Banerjee et al. (1998) methodology associated with (9a), which has good finite-sample properties compared to residual-based cointegration models. An assumption in the OLS estimation of (9a) is that the x-vector is weakly exogenous. In order to check the plausibility of this assumption in the present context, we perform the general LM test for weak exogeneity developed by Engle (1984). For all three models and two samples it was not possible to reject the weak exogeneity null.[8]

Referring then to the evidence from the Banerjee et al. (1998) procedure, we report in Table 4 the estimated coefficients for the model with 1 lead of the "x-vector". The results with 2 leads yield the similar results and are reported right below the BDM t-statistics. For space considerations, the full results of the BDM model are available upon request. It follows from these tests that the variables appear to share a long-run relationship for the more recent sample but not for the earlier one, which has only weak evidence of cointegration for the model with export intensity in column (2).

Given that the series are believed to be either I (0) or I (1) by the unit root tests, F- and t-statistics as well as the coefficients based on the Pesaran et al. (2001) methodology are provided in Table 5 with these results. Consistent with the theoretical discussion above, the higher the relative supply the lower the relative wage in favor of skilled workers: -0.717 for the benchmark model in the first sample and -0.266 in the second sample.

Note the (short-run) elasticity of the wage premium with respect to export intensity. For instance, increases in export intensity yield either a decrease in the wage premium between skilled and unskilled workers in Mexican manufacturing in the 1987-1995 sample or no effect in the 1994-2007 sample. A similar conclusion stands for real exchange rate

effects. Reinforcing the OLS results, for the earlier sample there is a negative and statistically significant relationship between the wage premium and: export intensity (-0.054) as well as the real exchange rate (-0.132). The tests for cointegration suggest very robust patterns for the 1994-2007 sample and a long-run relationship only for the model with export intensity in column (2) of Table 5 for the 1987-1995 sample. For the real exchange rate in column (3), only the t-tests support long-run relationships.

[Table 5 here]

Contrary to the original model, formal tests do not detect specification problems in the results based on Pesaran et al. (2001). In all cases, the null of no serial correlation is not rejected at standard significance levels, which suggests a good specification. Stability tests indicate excellent performance as well and the explanatory fit of the ECM in (9b) varies from adjusted $R^2$ of 46% or 49% in the earlier sample to 56% or 58% in the later sample.

Overall, the role of trade in driving relative wages can be interpreted as follows. Under both OLS and cointegration procedures, the impact of higher export intensity is to mitigate the wage premium for the earlier sample. This is consistent with goods exports being on average relatively intensive in unskilled labor. This matches the results from Li and Xu (2003) for China who found, conditioning for technology, that export intensity reduces wage inequality in exporting firms. For the real exchange rate, if skilled workers are more complementary with imported intermediate inputs, an appreciation of the real exchange rate should favor skilled workers, as in Robertson (2003). We find that trade and, to a lesser extent, real exchange rate mechanisms are operative in our earlier sample *at the very start of the trade liberalization.*

## 5. Final Remarks

This paper explores the trade-based hypothesis by Freeman and Katz (1991) and Revenga (1992) on Mexican manufacturing wage premiums. Two sample periods are studied with remarkable differences, ranging from the institutional trade framework (GATT and NAFTA), to the exchange rate regime (fixed or floating peso) and manufacturing wage premiums (55% rise versus 11% fall). In contrast to outsourcing in Feenstra and Hanson (1997) or to Stolper-Samuelson arguments in Hanson and Harrison (1999), this paper employs *relative supply shifts* as the major driving force of the wage premiums.

We borrow from Acemoglu (2002) the theory in which SBTC accelerates due to trade opening, motivated by the finding in Bernard and Jensen (1997) that increases in employment of exporters account for almost all of the increase in the wage gap. The elasticity of substitution ($\sigma$) between skilled and unskilled workers is estimated at 1.03 (earlier sample) or 1.71 (recent sample). Allowing trade considerations to modify the factor augmenting technology ratio, we examine the impact of export intensity or the real exchange rate on the wage premium. For the earlier sample, higher export intensity and a weaker peso are associated with decreases in the wage gap.

These effects are not observed in the later sample. Overall, the model with only relative supply shifts seems to be more appropriate for the more recent sample from 1994 to 2007 when the higher relative supply of skilled workers leads to lower wage premiums. Further work at the breakdown of the manufacturing sector may shed further light on these contrasting patterns.

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**Table 1. Unit Root Tests for Mexican Manufacturing.**

| Series | Deterministic Terms? | ADF (k) Series in Levels | ADF (k) Series in First Diffs. | KPSS (4) Series Levels | KPSS (4) Series in First-Diffs. |
|---|---|---|---|---|---|
| **1987-1995 Sample** | | | | | |
| rwhr | constant | -3.360(11)** | -6.198(10)*** | 2.189*** | 0.034 |
| | constant and trend | -1.796(11) | | 0.327*** | |
| H/L | constant | -1.003(12) | -2.662(11)* | 1.507*** | 0.156 |
| | constant and trend | -2.240(12) | | 0.272*** | |
| rer | constant | -2.658(10)* | -2.655(12)* | 1.008*** | 0.355* |
| | constant and trend | -2.619(10) | | 0.338*** | |
| xint | constant | 1.205(11) | -13.258(0)*** | 1.778*** | 0.271 |
| | constant and trend | -1.332(11) | | 0.357*** | |
| | Zivot-Andrews (1992) | -5.579(1)*** | | | |
| | chosen breakpoint | [1995:01] | | | |
| **1994-2007 Sample** | | | | | |
| rwhr | constant | 0.141(14) | -3.651(13)*** | 2.659*** | 0.059 |
| | constant and trend | -3.006(14) | | 0.352*** | |
| H/L | constant | -1.123(13) | -3.102(14)** | 1.922*** | 0.185 |
| | constant and trend | -2.428(13) | | 0.539*** | |
| rer | constant | -1.743(8) | -4.040(7)*** | 1.481*** | 0.107 |
| | constant and trend | -2.701(8) | | 0.320*** | |
| xint | constant | -1.160(13) | -5.448(12)*** | 2.448*** | 0.196 |
| | constant and trend | -3.831(13)** | | 0.296*** | |
| | Zivot-Andrews (1992) | -7.847(2)*** | | | |
| | chosen breakpoint | [1995:01] | | | |

*Notes*: Data are of monthly frequency for the two different samples. ADF(k) refers to the Augmented Dickey-Fuller t-tests for unit roots, in which the null is that the series contains a unit root. The lag length (k) is chosen by the Campbell-Perron data dependent procedure, whose method is usually superior to a fixed k chosen a priori and to k chosen by the information criterion. See Ng and Perron (1995). The method starts with an upper bound, $k_{max}$=14, on k. If the last included lag is significant, choose k = $k_{max}$. If not, reduce k by one until the last lag becomes significant (we use the 5% value of the asymptotic normal distribution to assess significance of the last lag). If no lags are significant, then set k = 0. Next to the ADF critical t-value, in parenthesis is the selected lag length. The KPSS test follows Kwiatkowski et al. (1992), in which the null is that the series is stationary and k=4 is the lag truncation parameter. The Zivot-Andrews (1992) test refers to Model A "crash model". The reported statistic is for testing the $\alpha = 1$ hypothesis, in which the critical value for inf $t_\alpha$ ($\lambda$) at the 5% level is -4.80, and $\lambda = T_B/T$ is the location of the breakpoint $T_B$. The symbols * [**] (***) attached to the figures indicate rejection of the null at the 10%, 5%, and 1% levels, respectively.

**Table 2. Ayat and Burridge (2000) Sequential Unit Root Tests for Uncertainty in Trend: 1987-1995 Sample.**

$$\Delta y_t = (\rho\text{-}1)y_{t-1} + \alpha_1 + \alpha_2 t + \alpha_3 t^2 + \Sigma\, \phi_i \Delta y_{t-i} + \nu_t$$

| | **rwhr** | **rwhr** | **rwhr** | **H/L** | **H/L** | **H/L** |
|---|---|---|---|---|---|---|
| **1987-95** | | | | | | |
| $(\rho\text{-}1)$ | -0.016<br>(0.023) | -0.362***<br>(0.111) | -0.855***<br>(0.160) | -0.049***<br>(0.029) | -0.135***<br>(0.042) | -0.136***<br>(0.047) |
| $\alpha_1$ | 0.060<br>(0.060) | 0.758***<br>(0.227) | 1.653***<br>(0.307) | 0.015**<br>(0.009) | 0.040***<br>(0.012) | 0.040***<br>(0.014) |
| $\alpha_2$ | | 0.004***<br>(0.001) | 0.014***<br>(0.003) | | 0.00002***<br>(0.000006) | 0.00002<br>(0.00002) |
| $\alpha_3$ | | | -0.00005**<br>(0.00001) | | | 0.00000001<br>(0.0000002) |
| Number of $\phi_i$'s | 3 | 3 | 3 | 1 | 1 | 1 |
| N | 104 | 104 | 104 | 106 | 106 | 106 |
| DW | 1.979 | 1.986 | 1.969 | 2.019 | 2.001 | 2.002 |
| Adj. $R^2$ | 0.391 | 0.442 | 0.517 | 0.035 | 0.091 | 0.082 |
| | **xint** | **xint** | **xint** | **rer** | **rer** | **rer** |
| $(\rho\text{-}1)$ | -0.017<br>(0.020) | -0.050<br>(0.038) | -0.157***<br>(0.061) | -0.051**<br>(0.024) | -0.038<br>(0.031) | -0.121***<br>(0.045) |
| $\alpha_1$ | 0.002<br>(0.011) | -0.002<br>(0.011) | 0.050**<br>(0.026) | 4.841*<br>(2.453) | 2.792<br>(3.768) | 16.696**<br>(6.783) |
| $\alpha_2$ | | 0.0007<br>(0.0003) | -0.0009<br>(0.0008) | | 0.014<br>(0.019) | -0.237**<br>(0.104) |
| $\alpha_3$ | | | 0.000021**<br>(0.00001) | | | 0.0020**<br>(0.0008) |
| Number of $\phi_i$'s | 1 | 1 | 1 | 1 | 1 | 1 |
| N | 106 | 106 | 106 | 106 | 106 | 106 |
| DW | 1.894 | 1.897 | 1.885 | 2.035 | 2.022 | 2.036 |
| Adj. $R^2$ | 0.064 | 0.092 | 0.125 | 0.045 | 0.041 | 0.085 |

*Notes*: Data are of monthly frequency. The tests further extend those discussed in Table 1 when the actual trend term is unknown as explained in Ayat and Burridge (2000). The number of $\phi_i$'s is determined by Schwarz information criterion. The symbols * [**] (***) attached to the figure indicate rejection of the null at the 10%, 5%, and 1% levels, respectively.

**Table 3. Ayat and Burridge (2000) Sequential Unit Root Tests for Uncertainty in Trend: 1994-2007 Sample.**

$$\Delta y_t = (\rho\text{-}1)y_{t-1} + \alpha_1 + \alpha_2 t + \alpha_3 t^2 + \Sigma\, \phi_i \Delta y_{t-i} + \nu_t$$

| | **rwhr** | **rwhr** | **rwhr** | **H/L** | **H/L** | **H/L** |
|---|---|---|---|---|---|---|
| **1994-2007** | | | | | | |
| $(\rho\text{-}1)$ | -0.057<br>(0.048) | -0.349***<br>(0.088) | -0.576***<br>(0.111) | -0.014<br>(0.012) | -0.032**<br>(0.016) | -0.040**<br>(0.021) |
| $\alpha_1$ | 0.158<br>(0.136) | 1.055***<br>(0.264) | 1.701***<br>(0.327) | 0.004<br>(0.004) | 0.009**<br>(0.005) | 0.012*<br>(0.006) |
| $\alpha_2$ | | -0.0008***<br>(0.0002) | 0.00005<br>(0.0005) | | 0.0000044***<br>(0.0000026) | -0.0000009<br>(0.000009) |
| $\alpha_3$ | | | -0.000008**<br>(0.000003) | | | 0.00000004<br>(0.00000006) |
| Number of $\phi_i$'s | 2 | 2 | 3 | 1 | 1 | 1 |
| N | 154 | 154 | 153 | 155 | 155 | 155 |
| DW | 1.889 | 1.902 | 1.992 | 1.969 | 1.961 | 1.965 |
| Adj. $R^2$ | 0.433 | 0.482 | 0.511 | 0.101 | 0.112 | 0.109 |
| | **xint** | **xint** | **xint** | **rer** | **rer** | **rer** |
| $(\rho\text{-}1)$ | -0.086***<br>(0.029) | -0.206***<br>(0.056) | -0.227***<br>(0.064) | -0.043**<br>(0.021) | -0.076***<br>(0.026) | -0.105***<br>(0.028) |
| $\alpha_1$ | 0.079***<br>(0.025) | 0.146***<br>(0.036) | 0.152***<br>(0.037) | 3.417*<br>(1.730) | 7.633***<br>(2.611) | 12.605**<br>(3.137) |
| $\alpha_2$ | | 0.0004**<br>(0.0002) | 0.0007<br>(0.0005) | | -0.020**<br>(0.009) | -0.110***<br>(0.034) |
| $\alpha_3$ | | | -0.000002<br>(0.000002) | | | 0.0005***<br>(0.0002) |
| Number of $\phi_i$'s | 2 | 2 | 2 | 1 | 1 | 1 |
| N | 154 | 154 | 154 | 154 | 154 | 154 |
| DW | 1.950 | 1.944 | 1.942 | 2.030 | 2.037 | 2.039 |
| Adj. $R^2$ | 0.230 | 0.256 | 0.253 | 0.037 | 0.059 | 0.098 |

*Notes*: Data are of monthly frequency. The tests further extend those discussed in Table 1 when the actual trend term is unknown as explained in Ayat and Burridge (2000). The number of $\phi_i$'s is determined by Schwarz information criterion. The symbols * [**] (***) attached to the figure indicate rejection of the null at the 10%, 5%, and 1% levels, respectively.

**Table 4. OLS Estimations of (8) and (9a) on Mexican Manufacturing:**

$$\text{OLS} \quad \ln(rwhr)_t = \beta_0 + \beta_1 t + \beta_2 \ln(H/L)_t + \varepsilon_t \qquad (8a)$$
$$\text{OLS} \quad \ln(rwhr)_t = \beta_0 + \beta_1 t + \beta_2 \ln(H/L)_t + \beta_3 \ln(xint)_t + \varepsilon_t \qquad (8b)$$
$$\text{OLS} \quad \ln(rwhr)_t = \beta_0 + \beta_1 t + \beta_2 \ln(H/L)_t + \beta_3 \ln(rer)_t + \varepsilon_t \qquad (8c)$$

$$\gamma(L)\,\Delta[\ln(rwhr)_t] = \beta_0 + \beta_1 t + \alpha(L)\,\Delta[\ln(x)_t] + \beta\ln(rwhr)_{t-1} + \phi\ln(x)_{t-1} + \Sigma\,\varphi_i\,\Delta\ln(x)_{t+j} + \varepsilon_t$$
$$\text{where } x_t = (H/L, xint)_t \text{ or } (H/L, rer)_t \qquad (9a)$$

| | Benchmark 1987-1995 | Trade 1987-1995 | RER 1987-1995 | Benchmark 1994-2007 | Trade 1994-2007 | RER 1994-2007 |
|---|---|---|---|---|---|---|
| **Constant** | -0.448<br>(0.531) | 0.343<br>(0.552) | -0.030<br>(0.472) | 0.387**<br>(0.175) | 0.647***<br>(0.167) | 0.023<br>(0.237) |
| **Trend** | 0.0044***<br>(0.0003) | 0.005***<br>(0.0004) | 0.004***<br>(0.0002) | -0.0005***<br>(0.0001) | -0.0008***<br>(0.0001) | -0.0004***<br>(0.0001) |
| **$\log(H/L)_{t-1}$** | -0.970**<br>(0.432) | -0.205<br>(0.466) | -1.005**<br>(0.409) | -0.584***<br>(0.141) | -0.399***<br>(0.134) | -0.705***<br>(0.136) |
| **$\log(xint)_{t-1}$** | | -0.089***<br>(0.025) | | | 0.072***<br>(0.016) | |
| **$\log(rer)_{t-1}$** | | | -0.097***<br>(0.026) | | | 0.048**<br>(0.021) |
| **Implied $\sigma$** | 1.031 | | 0.995 | 1.712 | 2.506 | 1.418 |
| **BDM Tests on $\beta$ of (9a) with 1 lead** | -3.125 | -3.970* | -3.382 | -9.275*** | -9.158*** | -7.890*** |
| **BDM Tests on $\beta$ of (9a) with 2 leads** | -3.013 | -3.908* | -3.135 | -6.966*** | -6.763*** | -6.810*** |
| **DW** | 1.038 | 1.274 | 1.281 | 1.578 | 1.882 | 1.718 |
| **Adj. $R^2$** | 0.924 | 0.940 | 0.937 | 0.685 | 0.727 | 0.704 |

*Notes:* The dependent variable is the skill premium: $\ln(rwhr)_t$. Data are of monthly frequency from 1987:01 to 1995:12 and from 1994:01 to 2007:01 in two different samples. The standard errors below the estimated coefficients are computed by the Newey-West correction of the variance-covariance matrix for heteroskedasticity and autocorrelation. The symbols * [**] (***) attached to the figure indicate rejection of the null of zero-value coefficients at the 10%, 5%, and 1% levels, respectively. The Banerjee, Dolado and Mestre (1998) tests are applied to the dynamic equation (9a), in which the relevant test is on $\beta \ln(rwhr)_{t-1}$ and either 1 or 2 leads of the "x" vector was assumed. Reported are the coefficients for the estimated model with one lead. The critical values of the t-ratio ECM test (with N=100) with constant and trend are: -3.75 for the benchmark case (with only relative labor supplies) and -3.98 for the augmented models with either trade or the real exchange rate with size = 0.05. The critical values from Banerjee, Dolado and Mestre (1998, p. 277) Table I (continued) are: -3.43 and -3.66, respectively, with size = 0.10. ** indicates rejection of the null hypothesis at 5%; * at 10%.

**Table 5. Pesaran et al. (2001) Estimations of (9) for Mexican Manufacturing:**

$$\Delta(\ln(rwhr)_t) = c_0 + c_1 t + c_2 dum95{:}01 + \beta_1 \ln(rwhr)_{t-1} + \beta_2 \ln(H/L)_{t-1} + \beta_3 \ln(x)_{t-1} + \Sigma\, \phi_i\, \Delta \ln(z)_{t-i} + \delta_1\, \Delta(\ln(H/L))_t + \delta_2\, \Delta(\ln(x))_t + \varepsilon_t \quad ,$$

$$\text{where } x_t = xint \text{ or } rer; \text{ and } z_t = (rwhr, H/L, xint)_t \text{ or } (rwhr, H/L, rer)_t \qquad (9b)$$

| | Benchmark 1987-1995 | Trade 1987-1995 | RER 1987-1995 | Benchmark 1994-2007 | Trade 1994-2007 | RER 1994-2007 |
|---|---|---|---|---|---|---|
| **Constant** | -0.601<br>(0.275) | -0.074<br>(0.459) | -0.305<br>(0.327) | 0.677***<br>(0.131) | 0.713***<br>(0.180) | 0.877***<br>(0.200) |
| **Trend** | 0.0018***<br>(0.0006) | 0.003**<br>(0.0006) | 0.0018****<br>(0.0005) | -0.0007***<br>(0.00008) | -0.0007***<br>(0.0002) | -0.0008***<br>(0.0001) |
| **Dum95:01** | -0.016<br>(0.014) | -0.002<br>(0.016) | 0.044*<br>(0.024) | 0.046***<br>(0.008) | 0.060**<br>(0.025) | 0.053***<br>(0.011) |
| **$\log(rwhr)_{t-1}$** | -0.371***<br>(0.116) | -0.485***<br>(0.112) | -0.524***<br>(0.110) | -0.941***<br>(0.090) | -0.991***<br>(0.101) | -0.927***<br>(0.103) |
| **$\log(H/L)_{t-1}$** | -0.717***<br>(0.255) | -0.285<br>(0.397) | -1.087***<br>(0.331) | -0.266**<br>(0.127) | -0.265*<br>(0.136) | -0.172<br>(0.127) |
| **$\log(xint)_{t-1}$** | | -0.054***<br>(0.022) | | | -0.026<br>(0.049) | |
| **$\log(rer)_{t-1}$** | | | -0.132***<br>(0.048) | | | -0.023<br>(0.015) |
| **Implied $\sigma$** | 1.395 | | 0.920 | 3.759 | 3.774 | |
| **Additional regressors** | 3 diff. lags | 3 diff. lags | 3 diff. lags | 3 diff. lags | 3 diff. lags | 3 diff. lags |
| **F-IV** $\alpha = \varphi = \Psi = 0$ | 3.957 | 6.248 | 6.575 | 31.028*** | 21.152*** | 17.704*** |
| **F-V** $\varphi = \Psi = 0$ | 5.815 | 7.903** | 7.923 | 62.019*** | 33.079*** | 28.767*** |
| **t-V** **t-ratio on $\varphi$** | -3.196 | -4.330** | -4.765** | -10.429*** | -9.773*** | -9.000*** |
| **DW** | 1.977 | 1.937 | 1.932 | 2.015 | 2.053 | 2.015 |
| **Adj. $R^2$** | 0.465 | 0.482 | 0.494 | 0.557 | 0.575 | 0.569 |

*Notes*: The standard errors below the estimated coefficients are computed by the Newey-West correction of the variance-covariance matrix for heteroskedasticity and autocorrelation. Below the coefficients of the table are F (Wald-type) and t-tests for the existence of long-run relationships. If the values fall outside the critical value bounds, a conclusive inference can be drawn without needing to know the integration or cointegration status of the underlying regressors. ** indicates rejection of the null hypothesis at 5%; * at 10%. The bounds of the critical value of the F-test for the benchmark model are 6.56 - 7.30 (5%) and 5.59 - 6.26 (10%) and, for the t-test, -3.41 to -3.69 (5%) and -3.13 to -3.40 (10%). The bounds of the critical value of the F-test for the augmented model (2 independent variables) are 4.87 - 5.85 (5%) and 4.19 - 5.06 (10%) and, for the t-test, -3.41 to -3.95 (5%) and -3.13 to -3.63 (10%). All critical values are taken from the tables in Pesaran et al. (2001).

Figure 1a. Relative Wage Ratio (rwhr) and Relative Labor Supply Ratio (lsratio) in Mexican Manufacturing: 1987-1995.

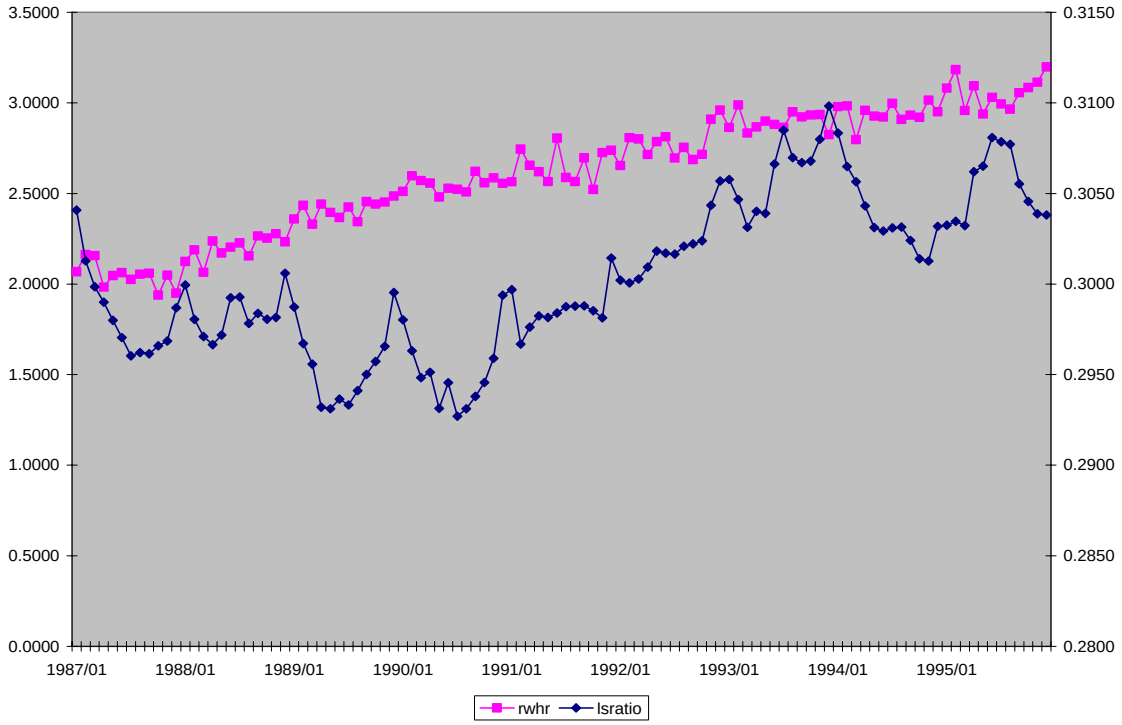



Figure 1b. Relative Wage Ratio (rwhr) and Relative Labor Supply Ratio (lsratio) in Mexican Manufacturing: 1994-2007.

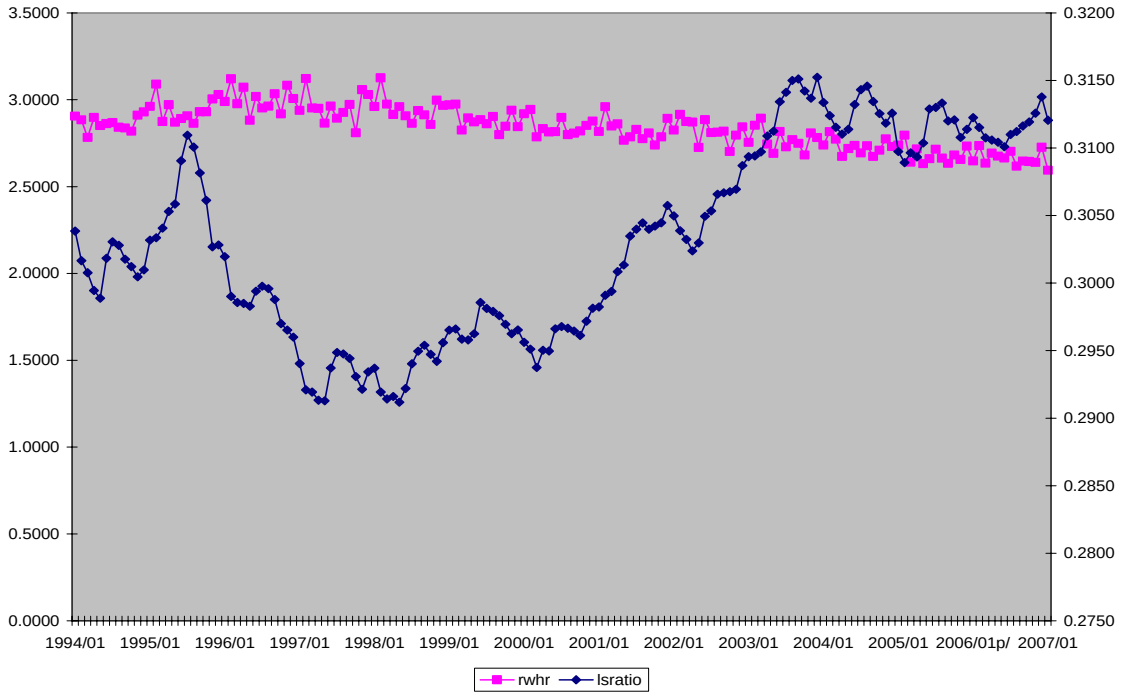


Figure 2a. Export Intensity Ratio in Mexican Manufacturing: 1987-1995.

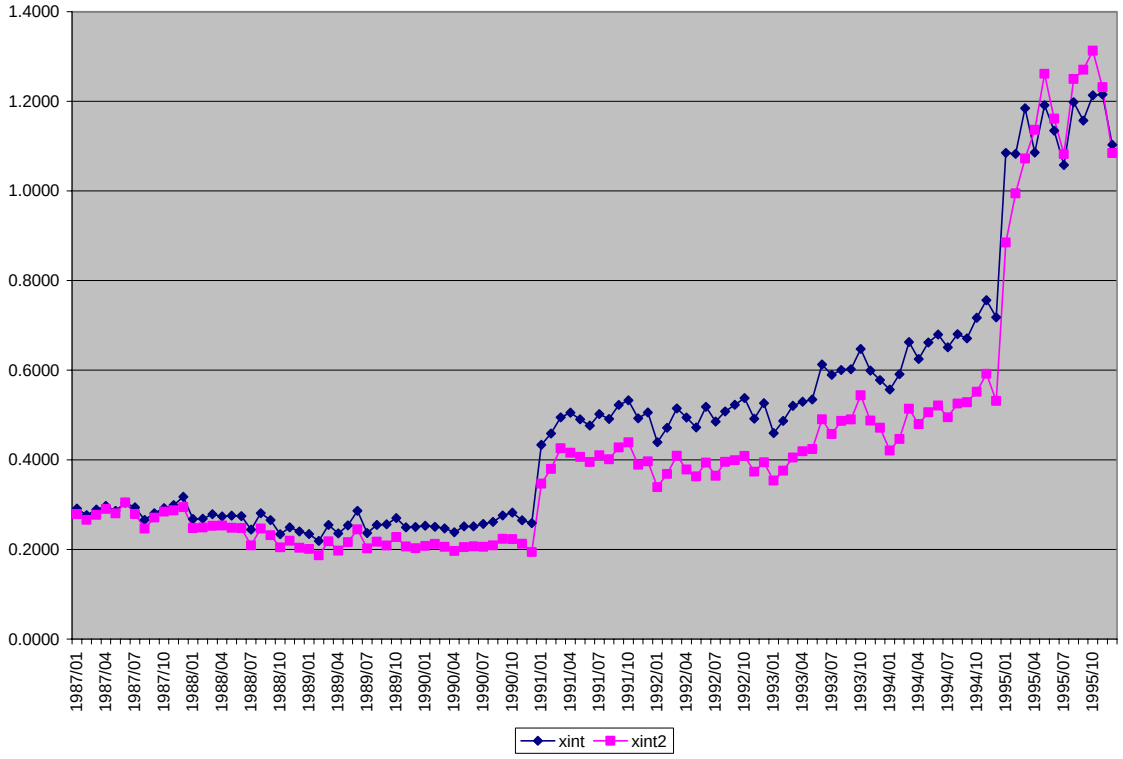



Figure 2b. Export Intensity Ratio in Mexican Manufacturing: 1994-2007.

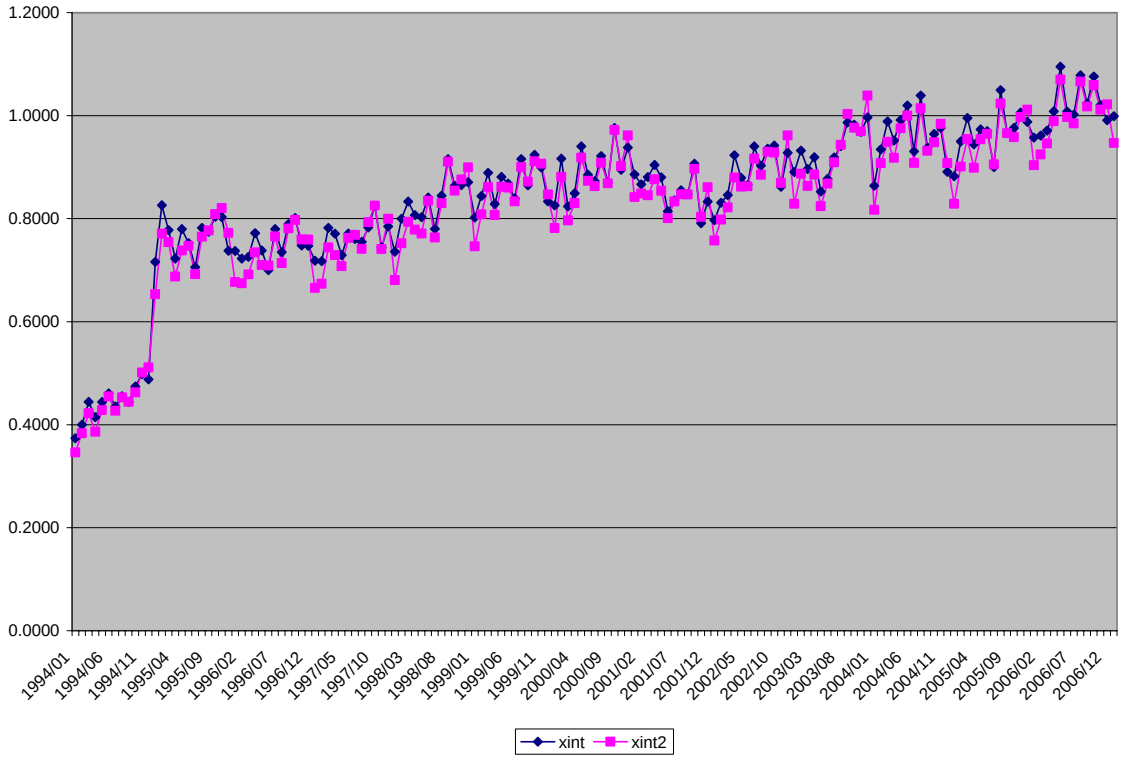


Figure 3a. Real Exchange Rate: 1987-1995.

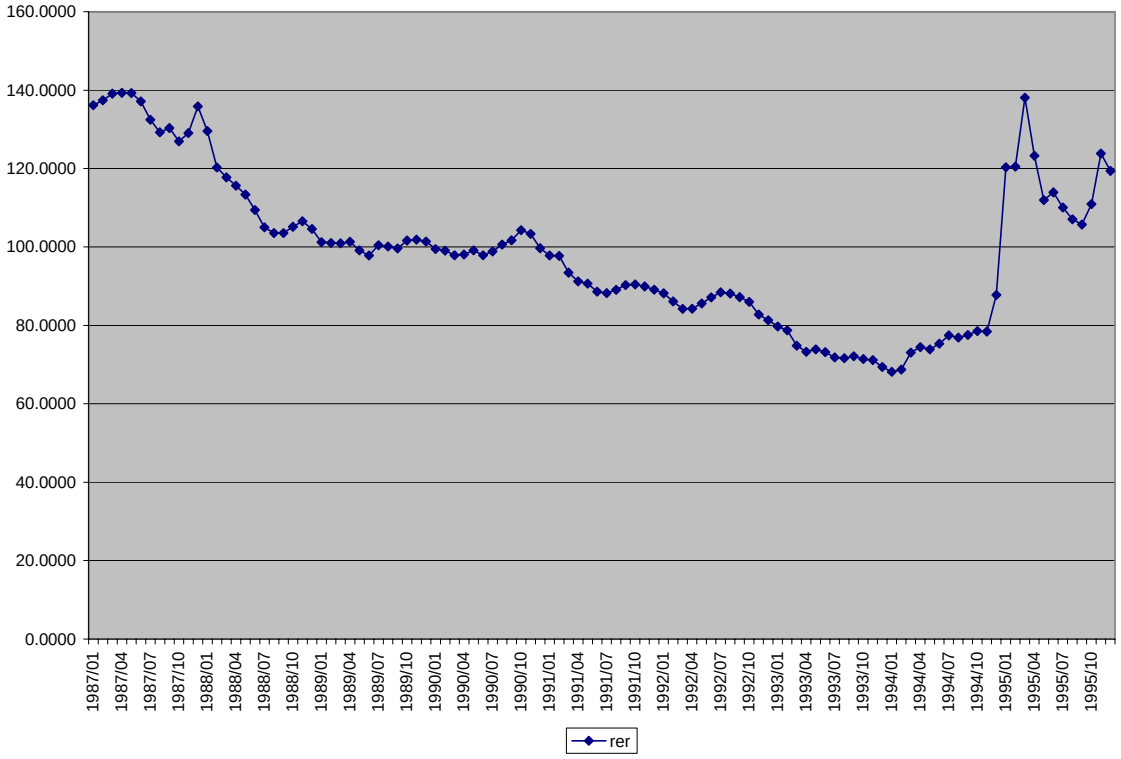
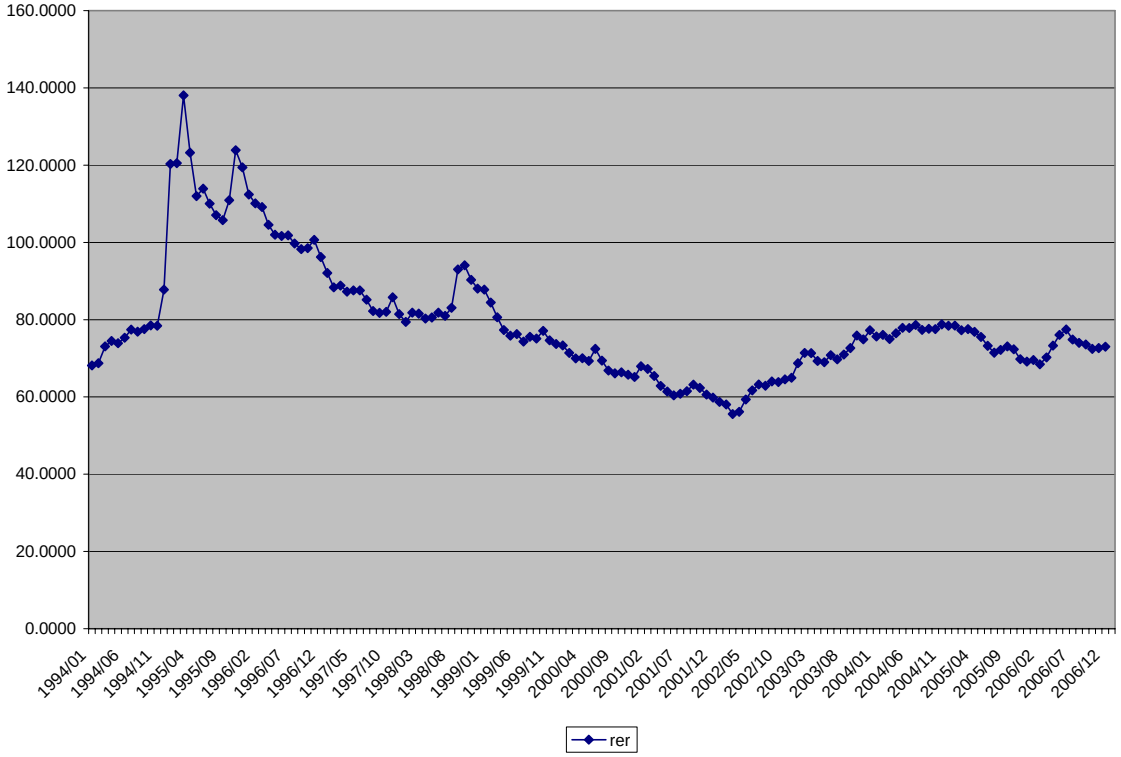



Figure 3b. Real Exchange Rate: 1994-2007.



[1] Export intensity and import penetration ratios, controlled for demand, do not help explain the rise in relative wages of white to blue-collar workers during 1987-1995 in Mexico, according to Mollick (2002). For Argentina, in industries where import penetration increased mostly, Galiani and Sanguinetti (2003) find that wage inequality widened more in favor of skilled workers, although trade does not explain a large portion of the rise in the skilled premium. Also, a burgeoning literature follows the two-stage framework of Gaston and Trefler (1994) at the individual worker level, such as: Pizer (2000) for U.S. industries; the impact of Colombian trade reforms on wages by Attanasio et al. (2004); Pavcnick et al. (2004) for Brazil; and Milner and Tandrayen (2004) for Sub-Saharan countries.

[2] Topel (1997) discusses certain changes in labor supply (immigration and increased female labor force participation) exacerbating inequality, while there is also the likelihood that human capital investment mitigates wage inequality. Will increased supply of college graduates reduce the relative wages? The answer depends on how well different skill groups substitute for one another in production, through the elasticity of substitution: $\sigma$. When the demand for college-education labor is fairly inelastic (high school and college are poor substitutes), increased supply of college graduates will reduce their relative wage.

[3] This definition of xint is very close to the alternative (xint2) in which exports are divided by the sum of imports, domestic sales, minus exports, as employed by Pizer (2000). See Figures 2a and 2b for the visual trend of both measures. With the 1994 devaluation, exports surge and imports fall, contracting the denominator and leading the figure slightly over 1 for the earlier sample. A different definition which bounds export intensity to lie within 0 and 1 divides exports by the sum of domestic sales plus exports. Qualitatively the results are the same under this new measure but we would not have a measure that is commonly used as xint or xint2 have been. Revenga (1992) has another import share definition as well.

[4] See also Revenga (1992) for the U.S. and Li and Xu (2003) for export intensity's role in explaining the skilled labor share in China. After estimating a Mincerian equation for wages in the first-stage, a second stage estimation regresses the industry wage premiums on a vector of trade related industry characteristics, of which export intensity is a regressor. For example, Attanasio et al. (2004) and Pavcnick et al. (2004) use lagged exports; Pizer (2000) uses lagged export intensity; and Milner and Tandrayen (2004) employ a dummy variable for exporting firms.

[5] See Slaughter (2001) for individual firms facing perfectly elastic labor supplies when employing 4-digit data. In his case, industry labor supply is closer to perfectly elastic than to perfect inelastic. An alternative identification strategy to the one above is to use instrumental variables, which are, however, difficult to obtain in practice. See Revenga (1992) and Mollick (2002) for examples in this context.

[6] This theoretical possibility is derived fully in Acemoglu (2002, pp. 38-39). It is possible to show that rwhr = $(p_H N_H / p_L N_L) = (H/L)^{(2\rho - 1)/(1-\rho)} = (H/L)^{\sigma - 2}$, where N is the number of machines used by each type of worker. If $\sigma > 2$ (or $\rho > 0.5$), the skill premium is an increasing function of the relative supply of skills.

[7] We also lag export intensity in order to allow for the possible endogeneity of trade flows. The lagged one period procedure was adopted by Attanasio et al. (2004) and Pavcnick et al. (2004) but did not change the qualitative nature of our findings. Rather, the magnitude of the $\beta_3$ coefficient decreased marginally. For the statistically significant coefficients, $\beta_3$ was found to be (std. error in parenthesis): -0.083 (0.024).

[8] The weak exogeneity methodology adopted is as follows. We first regress relative wages on labor supplies and save these first residuals. We then regress labor supplies on its own lagged terms as well as lagged wages and save these second residuals. Finally, we regress the first residuals on a constant, labor supplies and the second residuals. We repeat the procedure for the other two models, in which export intensity and real exchange rates are present in turn in the (RHS). Under the null of weak exogeneity, $nR^2$ from this regression is asymptotically distributed as a $\chi^2$ distribution. The null is rejected if $nR^2$ is greater than a critical value. It was never possible to reject the weak exogeneity null. The results for the $nR^2$ statistic were as follows: for the 1987-1995 sample, 0.0047 for the benchmark model, 2.4503 for the model with exports, and 0.8346 for the model with the real exchange rate; and for the 1994-2007 sample, 0.0003 for the benchmark model, 0.6552 for the model with exports, and 0.0005 for the model with the real exchange rate.