The Rise of the Skill Premium in Mexican Maquiladoras

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Abstract: Mexican maquiladoras are global production sharing firms subject to high speed of technology diffusion. Estimating how the skill premium between skilled and unskilled workers responds to technical progress, we substitute the time trend for a capital-expenditure share. Employing monthly data over 1990-2006, the benchmark model yields plausible elasticities of substitution (σ) varying from 1.73 to 2.42. This supports theoretical models in which the skill premium increases in the long-run under strong technology effects. With the capital-expenditure share, however, σ is estimated less precisely. Error correction models confirm fast adjustment to long-run equilibrium, lasting around 4 months.

Running Title: Skill Premium in Mexican Maquiladoras.

Keywords: Capital Equipment Expenditures, Skill Premium, SBTC, Time Trend.

JEL Classification Number: F16.

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1. Introduction

The Mexican maquiladora industry has been referred to as an example of global production sharing [e.g., Hummels et al. (2001), Hanson (2002), Helg and Tajoli (2005)], whose importance to the Mexican economy is not only remarkable but growing. Multidisciplinary work in Samstad and Pipkin (2005) acknowledges an upscaling process within the maquiladora sector, in which firms increase their competitiveness through training workers to become more productive.

The behavior of the maquiladoras' *skill premium* (defined as the relative wage between administrative workers and manual workers) amid trade openness, economic crisis and stabilization has been much less discussed. We conjecture that the speed of technology diffusion is high in the industry and estimate, in particular, how the skill premium between skilled and unskilled workers responds to technical progress.

Figure 1 shows that from 1980 to 2006 the share of maquiladora exports to total Mexican exports has risen by between 4 and 5 times; similar figures for imports lie between 3 and 4 times initial levels. From 1990 to 2006 (consistent data are not available for the 1980s), Figure 2 shows that the skill premium has risen by about 20%, while the skilled employment share (lsratio) has moved up only more recently from 9% to 11%. While such skilled labor share in the maquiladoras is lower than those reported for Mexican manufacturing (see below), it is in line with widespread belief suggesting the industry is very intensive in unskilled workers.

[Figures 1 and 2 here]

Studies of changing wage inequality rely on several explanations, ranging from labor supply shifts in Katz and Murphy (1992) to trade in Freeman and Katz (1991) and to skill-bias technical change (SBTC) in Berman et al. (1994). Marjit et al. (2004) discuss the increasing wage gap following a more open trade and investment regime on Latin America, in contrast to East Asian countries. Empirical evidence of the trade hypothesis has grown lately for developing economies and for Mexico, in particular, since at least when the country opened to foreign trade and slashed tariffs and quotas across all sectors from the mid to late 1980s.[1]

This paper looks at the share of capital equipment expenditures to total maquiladora expenses, which can be interpreted as a measure of international technology diffusion mentioned by Keller (2004). Our aim to enhance our understanding of how relative wages at home respond to increases in world production sharing in at least two ways.

First, there is the theoretical motivation. Using monthly data from Mexican maquiladoras's from 1990 to 2006, our research strategy compares the basic Katz and Murphy (1992) model to Acemoglu (2002) who takes into account the complementarity between skilled labor and capital. The main idea dates to Greenwood et al. (1997) through the introduction of new and more efficient capital goods as an important source of U.S.

[1] The following is a far from exhaustive list on Mexico. For the maquiladora industry, Coronado et al. (2004) and Fullerton and Torres-Ruiz (2004) document statistically significant payroll responses in Mexico's Tijuana and Chihuahua's maquiladoras, respectively, to economic forces and Mollick and Wvalle-Vásquez (2006) report how maquiladora employment reacts to wage effects allowing for Chinese competition effects. More generally for Mexico, Tybout and Westbrook (1995) estimate cost functions between 1984 and 1990 and the efficiency effects caused by greater openness. Feenstra and Hanson (1997) explore FDI inflows into Mexico and the role of outsourcing by Northern multinationals. Individual characteristics of workers are handled by Cragg and Epelbaum (1996). One explanation for the skill premium in Mexico growing after 1985 with little change in the relative employment of skilled labor, 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 period from 1984 to 1990. Evidence on technology effects on the skill premium is perhaps confined to Esquivel and Rodríguez-López (2003), who 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. Robertson (2004) also examines changes in relative output prices.

productivity change. Technological advances have made equipment less expensive, triggering increases in the accumulation of equipment both in the short and long run. Krusell et al. (2000) hypothesize that capital-skill complementarity may be important for understanding wage inequality, because the stock of equipment has been growing at a very high rate in recent years. They find that changes in observed factor inputs can account for most of the variation in the skill premium over the last 30 years. Goldin and Katz (1998) observe a strong positive association between changes in capital intensity and the nonproduction worker wage bill from 1909-1919. Autor et al. (1998) find a substantial and growing wage premium associated with computer use despite a large increase in the supply of workers with computer skills. For Topel (1997), 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.

Second, this research reports plausible values of parameters under several estimation methods. In the model below the skill premium is explained as a function of 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 w falling could be compensated by changes in technology. Hitherto, the only application of this idea to developing economies is perhaps limited to Sánchez-Páramo and Schady (2003), who argue that in Argentina, Brazil, Chile, Colombia, and Mexico very implausible values of the elasticity of substitution between skilled and unskilled workers (σ) are found. Inspired by the conjecture in Acemoglu (2002) that capital intensity may better capture technology than the time trend, we check the model sensitivity to our measure of technology diffusion.

To the best of our knowledge, no empirical study has verified the possibility of unit roots in the skill premium and relative labor supplies. This should be more than a statistical exercise since error correction modeling may enlighten how the skill premium responds in the short-run to changes in the relative supply. Acemoglu (1998), for example, argues that the impact of an increase in the supply of skills on the skill premium is determined by two competing forces: *the substitution effect*, which makes the economy move along a downward sloping relative demand curve; and a *directed technology effect*, which shifts the relative demand curve for skills. A large increase in skilled labor first moves the economy along a short-run relative demand curve (in which technology is constant), reducing the wage premium. The second effect is later felt on. If the technology effect is strong the model predicts that the skill premium should increase in the long-run. In Kiley (1999), an increase in the supply of skilled labor leads to a temporary fall in the skill premium, followed by an expanding gap between the wages of skilled and unskilled workers as technologies adjust towards a more skill-intensive mix.

Our major results are as follows. The benchmark model borrowed from Acemoglu (2002) finds support by Granger causality tests. We find unidirectional support from lagged relative labor supplies to wage premiums for the Mexican maquiladoras, implying that wages respond to quantities, as discussed in Slaughter (2001). The elasticity of substitution (σ) between skilled and unskilled workers is estimated in the range (1.73-2.42), not far from the consensus of σ between 1 and 2 in Johnson (1997). Allowing for unit roots in the data, we reinforce the estimated σ at around 2.11-2.12 (by DOLS) or at 1.85 (by Johansen). A very fast adjustment to long-run equilibrium is found: between 23% and 29% of the deviations are adjusted within a month. The short-run adjustment supports both positive and negative effects of changes in relative labor supply on changes in the wage premium. Yet the long-run result is always consistent with the strong *directed technology effect*.

2. Data Description

This study examines relative wages, employment, and capital intensity for Mexican maquiladoras. Monthly data are used and the 1990-2006 period is chosen because it is the longest possible with detailed information on relative labor types and maquiladora expenditures. The period includes NAFTA taking into effect in 1994, which "gave Mexico a considerable preferential tariff advantage" according to Kose et al. (2005), as long as U.S. tariffs on imports from non-NAFTA partners were much higher than those on imports from Mexico.

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). For the Mexican case, Robertson (2004) looks at additional data sources and shows that production workers have less education in every industry than non-production workers and concludes that the two-tier distinction to classify skill intensity seems valid in Mexico.

All labor data are from Mexico's INEGI in the "Banco de Información Económica" (BIE) section (http://www.inegi.gob.mx). The subsections on "IME" (industria maquiladora de exportación) for the maquiladoras and on "Sector Manufacturero" for manufacturing are explored. It is worth emphasizing that the maquiladora data contrast to Mexican manufacturing data, primarily for one methodological reason. The manufacturing data are based on a sampling methodology, including the Monthly Industrial Survey

(Encuesta Industrial Mensual - EIM - in Spanish) and the Yearly Industrial Survey (Encuesta Industrial Anual - EIA). The EIM and EIA surveys started in 1964 and 1963, respectively, with 29 groups of economic activity. They were later expanded in 1976 to 57 groups and in 1987 to 129 groups. The particular data we work on below to compare to the maquiladora data are from the further expansion in 1993, which is now in place, and counts 205 groups of economic activity. Neither the EIM nor the EIA data include information on the maquiladora firms. The manufacturing data for the 205 groups start only in 1994 and are based on samplings and not on the total number of firms.

The maquiladora data are based on national coverage through surveys to *all firms* which were awarded the authorization to operate as a “maquiladora firm”. For the skilled labor premium in a month, we take: **Wu** from “obreros” or unskilled workers; and **Ws** from “empleados administrativos”, management employees or skilled workers, both in pesos. The deflator is Mexico’s CPI (the INPC), with third and fourth weeks of June 2002 = 100.

Data on hours worked are incomplete. Mexico’s INEGI provides only data on hours worked for unskilled workers and for total workers, which comprises all three types of workers in the maquiladoras: management personnel, technicians, and laborers. Looking at the available data on hours worked for unskilled workers and total workers, one infers that hours worked of unskilled workers are from 21% to 25% higher in a month than hours worked of total workers from 1990: 1 until 1996:12. Then suddenly the gap becomes very close to zero from 1997 onwards. This is an indication that measurement errors possibly exist with the data on hours worked, and we choose to leave aside this piece of data.

A few observations are worth mentioning. First, in Figure 2 the skill premium in the maquiladoras (**wprem**) moves from about 4 to 5 in the period, suggesting an upward trend of relative wages. If one had defined instead the skilled premium as the ratio between wages of technicians with respect to unskilled workers the skill premium fluctuates within the 2.50-3.00 range over the sample, with no clear upward trend. The relative participation of non-production workers in the total workforce (**lsratio** or **H/L ratio**) has fluctuated but not by much, between 9% and 11%. Second, while such skilled labor share in the maquiladoras is much lower than those reported for Mexican manufacturing (**lsratioman**) in Figure 3 (between 41% and 46%), it matches the belief that the maquiladora industry is very intensive in unskilled workers. The share of skilled workers in employment appears to rise after 2001, a result also observed for the ratio of wages between technicians and unskilled workers (not reported). This upward trend in skill intensity has not always been present in related studies. Gera et al. (2001), for example, report no strong evidence that skill intensity increased at the aggregate level in Canada over 1981-1994. Similar to the Mexican maquiladoras, the relative labor supply of skilled workers in manufacturing rise after 2001. The wage premium in manufacturing (**wpremman**), however, does not move much over the time period.

[Figure 3 here]

Third, we investigate the share of capital expenditures with respect to all monthly expenses of maquiladora firms. The intensity of capital and equipment employed in the maquiladoras rises from a little over 0.5% in 1990 to about 2% in 2006, as can be verified by the K1 measure in Figure 4. We believe this is the closest measure possible to equipment expenditures in the SBTC literature such as in Krusell et al. (2000) and

Acemoglu (2002). Since no data exist on the price of capital equipment for the maquiladoras, we implicitly use the fact that more capital is being used, perhaps due to lower prices and more diffusion of technology. Keller (2004) reviews several approaches to measure technology and refers to the “diffusion of technology”, which includes market transactions and externalities and to the fact that data on the former is straightforward, such as firms making royalty payments for the use of patents, licenses and copyrights. A broader measure takes into account buildings and land rents and acquisitions by maquiladoras (K2), which has started at about 10% of total expenses and then moved up to 12% during the Mexican peso collapse to finally stabilize at below 10% at the end of the sample. The broader measure K3 is obtained by K1 + K2. Figure 4 contains these three measures.

[Figure 4 here]

3. The Model

Suppose maquiladoras hire only two types of workers: skilled (H) and unskilled (L), who are imperfect substitutes and supply labor inelastically. Suppose workers are risk-neutral, maximize the present value of labor income, and labor markets are competitive. Under a constant elasticity of substitution (CES) production function 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:

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

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

Combining (2) and (3), the skill premium becomes:

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

which turns out to be, in logarithmic form:

$$\ln w = [(\sigma\text{-}1)/\sigma] \ln (A_H/A_L) - (1/\sigma) \ln (H/L) \qquad (5),$$

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) in the vertical axis against the relative supply of skills (H/L) in the horizontal axis. An increase in the (vertical) relative supply in the (w, H/L) locus moves the equilibrium point along the downward sloping relative demand curve, reducing the skill premium w.

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. This would be a shift out of the relative demand curve. The conventional wisdom is that w increases when skilled workers become relatively more productive, consistent with $\sigma > 1$.

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

It is possible to explore the idea that technical change has taken placed steadily. As suggested by Acemoglu (2002), this hypothesis can be captured by:

$$\ln (A_{Ht}/A_{Lt}) = \gamma_0 + \gamma_1 \text{trend} \qquad (6),$$

where trend is calendar time. Substituting (6) into (5) immediately yields:

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

$$\ln w = \beta_0 + \beta_1 \text{trend} + \beta_2 \ln (H/L) \qquad (8),$$

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. When H/L grows faster than the rate of SBTC, w will fall. When H/L falls short of this rate, w will increase. Estimating a value of –0.709 for β_2 in (9) with $R^2 = 0.52$, Katz and Murphy (1992) argue that $\sigma = 1.41$ for the U.S. labor market under 25 annual data: 1963-1987.

The elasticity of substitution parameter is admittedly difficult to estimate since it refers to an elasticity that combines substitution between and within industries. 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 since we have the whole of the maquiladora industry. See Slaughter (2001) for individual firms facing perfectly elastic labor supplies when employing 4-digit data since industry labor supply is closer to perfectly elastic than to perfect inelastic.

As discussed in the Introduction, for Acemoglu (1998), the impact of an increase in the supply of skills on the skill premium is determined by two competing forces: *the substitution effect* and the *directed technology effect*. Krusell et al. (2000) hypothesize that capital-skill complementarity may be important for understanding wage inequality, because the stock of equipment has been growing at a very high rate in recent years. Acemoglu (2002) introduces this idea to U.S. data and conjectures that if the decline in the relative price of equipment capital is related to the increase in the demand for skills, it should proxy in (8) for the demand for skills and perform better than a linear trend. Adding the log of the relative price of capital to (8), Acemoglu (2002) does not find statistically significant capital terms when the time trend is included. When the time trend is excluded, however, the fit of the regression (measured by the adjusted R^2) is worse than (8). Acemoglu (2002, p. 29) concludes that "while this evidence may simply reflect the fact that the relative price of equipment is measured with error, it casts some doubt on the view that the relative price of equipment capital is directly linked to the demand for skills."

Introducing this "observable", the modified equation becomes:

$$\ln w = \beta_0 + \beta_1 \text{trend} + \beta_2 \ln (H/L) + \beta_3 \ln (K1) + \varepsilon \qquad (9),$$

where: K1 is the share of equipment capital rents relative to total maquiladora expenses. Linearity is assumed for convenience and follows Acemoglu (2002). In Krusell et al. (2000), the skill premium is derived as the sum of three ratios: capital equipment to skilled work, hours worked between unskilled and skilled workers, and quality per hour of the type of worker. The latter is an unmeasured series and hours worked for the maquiladora industry is incomplete as discussed in Section 2. Beaudry and Green (2005) extend the empirical framework above to total factor productivity, which requires physical capital stock data, also unavailable for maquiladoras. For these reasons, we follow linearity in (9). The essence of the argument is to see whether the observable on capital expenditures better capture technology trend than the time trend, upon which evidence is provided below.

4. Results

Table 1 reports unit root tests for the series in levels and in first differences. The tests suggest all series are non-stationary in levels but achieve stationarity after first-differencing. Usually the ADF tests do not reject the unit root in levels but do reject at standard confidence levels in first-differences. The only exception is the ratio of maquiladora capital equipment rents to total expenses (K1), when the ADF test indicates stationary series in levels (at the 10% level), a result reversed under the KPSS test. The KPSS test overwhelmingly confirms the I (1) decision, rejecting the null of stationary series in levels and not rejecting it in first-differences.

[Table 1 about here]

The existence of a unit root in relative wages could be a matter of concern since it would be difficult for an economic model to predict that relative wages grow without bound. If this reasoning is correct, the evidence in favor of a unit root would more likely mean that, in the transition period when Mexico suffers a currency crisis (1994-1995), there is so much fluctuation in relative wages or in relative labor supplies. These abnormal events would make it more difficult to reject the null of unit roots and the evidence below is robust to the presence of unit roots in the data.

An accompanying table available upon request reports the sample correlation coefficients in levels and in first-differences. In levels, there are high correlation coefficients between capital share and the wage premium (0.741), followed by the correlation coefficient between capital share and relative labor supply (0.516). The wage premium appears to be only weakly correlated with H/L ratio (0.217). In first-differences, the respective correlation coefficients are 0.170, 0.010, and -0.246.

As a preliminary route to check the plausibility of the model in (8), we check for Granger causality tests for several lag-lengths. Longer lags yielded not statistically significant coefficients and specifications with shorter lags suffered from misspecification problems. Table 2 contains these results for: the maquiladoras; the manufacturing sector; and for the interaction between the maquiladoras and the manufacturing sector. As explained in Section 2, data constraints lead us to adopt different time periods for each sector. When we combine both sectors, we run the regressions with the smaller time period (from 1994 to 2006) when we have simultaneous data for both sectors.

The model represented by (8) finds preliminary support in the upper panel of Table 2, as we reject the null that lagged relative labor supplies do not influence wage premiums for the Mexican maquiladoras. On the other hand, when relative labor supply is the dependent variable we find no statistically significant effects of wage premiums. This result is reversed for the Mexican manufacturing in the middle panel, in which the direction of causality runs contrary to that specified in the theoretical model. The most likely explanation is that, not capturing the whole of manufacturing violates the assumption discussed earlier that wages should respond to quantities as in Slaugther (2001). When we combine both sectors, the lower panel of results does not indicate substantial effects other than the one for 5 lags (yet not significant at the 5% level) when the wage premium of manufacturing is responding to the labor supply of the maquiladora sector. Therefore, despite the hypothesis that a low wage premium in manufacturing may induce a shift of skilled labor into the more dynamic maquiladora sector, the evidence is weak that there is much of this mechanism.

[Table 2 about here]

Table 3 reproduces OLS estimates of specifications (8) and (9) under different assumptions for the deterministic terms (constant and trend). We focus on the parameter associated with the relative supply of workers and leave the econometric fit of the estimates for further analysis below. The elasticity of substitution between the two labor types is derived by computing $(-1/\sigma) = \beta_2$ on ln (H/L) in (8) or (9). Table 3 shows that specifications without the time trend in the first two columns yield very implausible values of σ. It also shows that the implied elasticity of substitution varies between 1.730 and 2.353, under specification (8) with the time trend. This suggests 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 β_2-coefficient equals -0.578 in the specification without constant term in the estimation, labeled (8)t. When capital share is included, the implied elasticity of substitution varies between 1.727 and 2.415 under specifications (9)t and (9)ct. The magnitudes for both models appear reasonable since recent empirical estimates suggest σ 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).

[Table 3 about here]

When the time trend is included together with the constant, the implied σ becomes greater than 2. If correct, the value of σ suggests 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, as derived in Acemoglu (2002, pp. 38-39).

Support for SBTC is obtained in the estimation of the time trend, which is always statistically significant and positive. Its value lies between 0.001 and 0.002, in agreement with the expected positive value if technical change is skill biased. Katz and Murphy

(1992) report an overall value for the U.S. labor market over 1963-1987 of a 0.033 trend coefficient for annual data. Our monthly estimates are a little off this value, if annualized, but still provide support for the SBTC hypothesis.

The benchmark model with technology trend and relative labor supply in Table 3 explains around three quarters of the overall movements in relative wages. All standard errors in the estimates are computed with the Newey-West method for correction for heteroskedascity and serial correlation. There is, in fact, widespread serial correlation in the estimates of Table 3 as documented by complementary Breusch-Pagan Lagrange Multiplier tests (not reported). In all cases, the null of no serial correlation is rejected at standard significance levels, which suggests further work, either due to model misspecification (“omitted variables”) or to the presence of unit roots (“spurious regressions”).[2]

Given the unit root in the data, we thus check the existence of cointegration in the basic model by applying a standard ADF test on the residuals. If the residuals are stationary, the variables are cointegrated and there exists an ECM incorporating lagged deviations from the stationary equilibrium. Since ADF tests are known to have low power in finite samples, we report also KPSS tests on the residuals. For the latter, the null hypothesis is that the series is stationary. In Table 3, the row below the β_3-coefficient presents evidence on both sets of tests: the ADF statistics in the upper row and the KPSS in the lower row. For both ADF and KPSS tests, there is cointegration at the 5% level

[2] We also apply first differences to the benchmark model, thus eliminating the deterministic terms. Details are omitted for space constraints but this procedure suggests very close β_2-coefficients for models (8) and (9) of -0.46 (not far from those already reported in Table 3) and the estimated effect of capital changes on the wage premium becomes a very strong 2.61 capital-coefficient. The corresponding elasticities of substitution vary only from 2.14 to 2.16 in this case. Serial correlation problems are not severe, at the cost of worsening the overall fit of the model as expected.

(stationary residuals) for both specifications (8) and (9).[3] Evidence in favor of stationary residuals is absent for the model without the time trend. However, the KPSS does not reject the null of stationary residuals in any case. We thus conclude the evidence on cointegration is particularly strong when the time trend is present.

In order to confirm the values of the elasticity of substitution, we apply next the Stock and Watson (1993) dynamic OLS (DOLS) procedure allowing for contemporaneous, as well as leads and lags ($k = 2$) of first differenced terms. The results, displayed in Table 4, are in line with those in Table 3: clear rejection for the model without trend, and cointegration for the models with trend. The elasticity of substitution in the latter cases varies between 2.114 and 2.123.

[Table 4 about here]

Applying the ECM methodology to the equations in Tables 3 and 4 that passed cointegration [namely, (9)t and (9)ct in Table 3 and (8)ct and (9)ct in Table 4; other specifications did not change the basic results], negative and significant values of the β_1 coefficients are found throughout in Table 5. Since ε measures the deviations of the wage premium from its long-run value (estimated by the model), a positive term in the previous period implies a reduction in the current wage premium. We estimate a model for variations of w in first-differences, under a general-to-specific methodology (with 14 maximum lags) in choosing the lag-length of the differenced terms in the right hand side. The results do not change with respect to the benchmark or modified model: between 23%

[3] The sequential ADF method recommended by Ng and Perron (1995), starting with maximum lag length kmax = 24, is applied on the residuals of the cointegrating equations. The KPSS tests are employed with Newey-West bandwidth and window at 4, with test results unchanged to the specific window used.

and 29% of the deviations from the long-run equilibrium are corrected in the following month.

It is also possible to see the (short-run) elasticity of the wage premium with respect to relative labor supply through the estimation of the γ-parameters. Table 5 reports positive and statistically significant γ-parameters when the residuals come from the OLS model with only the time trend: between 0.420 and 0.458 in column (1), which suggests that, away from the steady-state, increases in H/L bring about increases in the wage premium, consistent with the SBTC hypothesis and a strong *directed technology effect* even in the short-run. Note, however, that medium run parameters (γ_3 and γ_4-parameters) are negative and statistically significant for the other three specifications. This implies that *the substitution effect*, which makes the economy move along a downward sloping relative demand curve, also has some role in the short-run estimations.

[Table 5 about here]

Contrary to the serial correlation found in the original model in levels, the Breusch-Pagan Lagrange Multiplier tests do not detect any in Table 5. In all cases, the null of no serial correlation is not rejected at standard significance levels, which suggests a good specification. The explanatory fit of the ECM (by the R^2) varies between 46% and 48% for the OLS-based models and at around 45% for the DOLS-based models.

It is well known that no single cointegration test is found to dominate. As Gregory et al. (2004) put, the choice of which test to apply is dictated by the nature of the investigation. If interest is on a particular relation or variable (H/L on w, for example), single equation tests are likely used. If, however, we have a truly multivariate setting (H/L combined with K1 to explain w), a system approach should be adopted.

We apply the Johansen cointegration method next in case the true relationship is given by (w, H/L, K1). In the first-stage the Johansen cointegration method is used for estimation of the long-run vector. The method of estimation is the vector error correction model with eleven seasonal dummies. In the second stage, lagged one period residuals from the first-stage are used in differenced form in the ECM. Table 6 contains VECM estimations of equations (8) and (9). Finding partial evidence of cointegration at the bottom of Table 6, column (1) reports a positive time trend for the model without capital and a β_1-coefficient of -0.541, implying σ = 1.848, a value very much close to those reported in Tables 3 and 4 under different cointegration procedures.

[Table 6 about here]

The ECM term is statistically significant and strong (-0.235) in the first column, suggesting a fast adjustment to equilibrium. The magnitude is about the same as those β_1-coefficients reported under residual-based methods in Table 5. Similar to the previous residual-based results, positive and statistically significant γ_1-parameter is found in the first column: 0.286 (standard error = 0.121); becomes less significant in the second column at 0.211 (standard error = 0.123); and not significant at all in the third column at 0.175 (standard error = 0.123). Away from the steady-state, increases in H/L bring about increases in the wage premium, consistent with the SBTC hypothesis.

When the capital-expenditure share (K1) is introduced, however, the β_2-coefficient becomes weaker: -0.334 without the time trend (statistically significant at the 5% level) and -0.186 with the time trend (not statistically significant). There are positive coefficients for the capital-expenditure share (varying from 0.206 to 0.324) but the augmented model in columns (2) and (3) fails to display cointegration. In this respect, our findings are similar to Acemoglu (2002), who reported that the relative price of equipment capital is not significant in such regressions. Once capital terms were entered simultaneously with the time trend however, the time trend was found to be significant but there was no evidence in Acemoglu (2002) that the relative price of capital matters for the demand for skills.

The serial correlation properties of the underlying VARs are different too. While the benchmark model fails to reject the null of serial correlation at standard significance levels, the augmented model fails to do so for 12 lags as reported by the LM test. In the VECM context, Bruggerman et al. (2006) argue that the LM variant has the most favorable size properties. Therefore, the benchmark model (8) in a systems-framework performs best.

In sum, there is evidence for cointegration in the benchmark model under Tables 3 to 5, which is corroborated under Table 6 for system-cointegration. On the other hand, Johansen tests cast doubt on a stable relationship under the augmented model with K1. What this research documents is a fairly similar value of the elasticity of substitution. In fact, the several estimated values of σ in this paper are close to plausible values whenever a long-run relationship is shown to exist and can be compared to other works. Sánchez-Páramo and Schady (2003), for example, have argued that in Argentina, Brazil, Chile, Colombia, and Mexico their estimates of σ are very imprecise with t-statistics of one or lower; in Chile, for example, their estimate of σ is higher than 10 but the t-value is 0.2. Recent work in 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 more recently: their estimate of σ is 0.22, with standard error of 0.36 for the 1976 to 2000 period.

5. Final Remarks

We borrow from Acemoglu (2002) the idea that an "observable" related to the level of capital intensity may better capture technology than the time trend in the Mexican maquiladora industry. Granger causality tests find unidirectional support from lagged relative labor supplies to wage premiums for the Mexican maquiladoras; yet do not suggest cross-effects with the manufacturing sector. The elasticity of substitution (σ) between skilled and unskilled workers is estimated in the range (1.73-2.42), not far from the consensus of σ between 1 and 2 in Johnson (1997).

While support for SBTC is overwhelming, measured by the significance of the time trend in the wage premium equations, we allow for unit roots in the series, a point overlooked in the literature. In addition to reinforcing the estimated σ at around 2.11-2.12 (by DOLS) or at 1.85 (by the Johansen method) we offer two findings. First, system methods suggest (partial) cointegration for the benchmark model but not for the specifications with the capital-expenditure share. Second, a very fast adjustment to long-run equilibrium is found: between 23% and 29% of the deviations are adjusted within a month. The short-run adjustment in this paper supports both positive [the *directed technology effect* in Acemoglu (1998)] and negative (the *substitution effect*) effects of changes in relative labor supply on changes in the wage premium. But the long-run result remains consistent with a strong *directed technology effect* found herein.

We leave for further research extensions of the Katz and Murphy (1992) model through TFP in Beaudry and Green (2005) or allowing for differences in hours worked in Krusel et al. (2000). Fluctuations in the price of oil may also have had a role in Mexico as

SBTC should change with trends towards unskilled workers and shrinking wage premiums.

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Figure 1. Shares of Mexican Maquiladoras Exports and Imports over Total of Exports and Imports: 1980:1-2006:3.

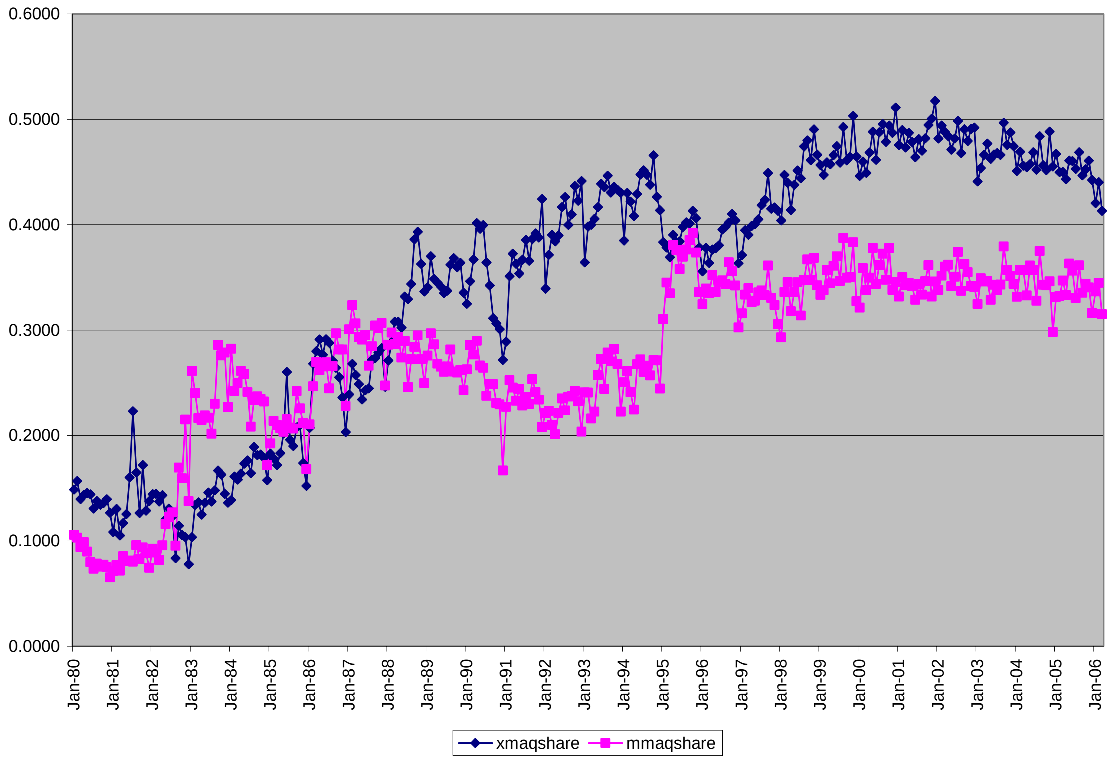

Figure 2. Wage Premium (wprem), at left, and Ratio of Skilled to Unskilled Workers (lsratio), at right, in Mexican Maquiladoras: 1990:1-2006:3.

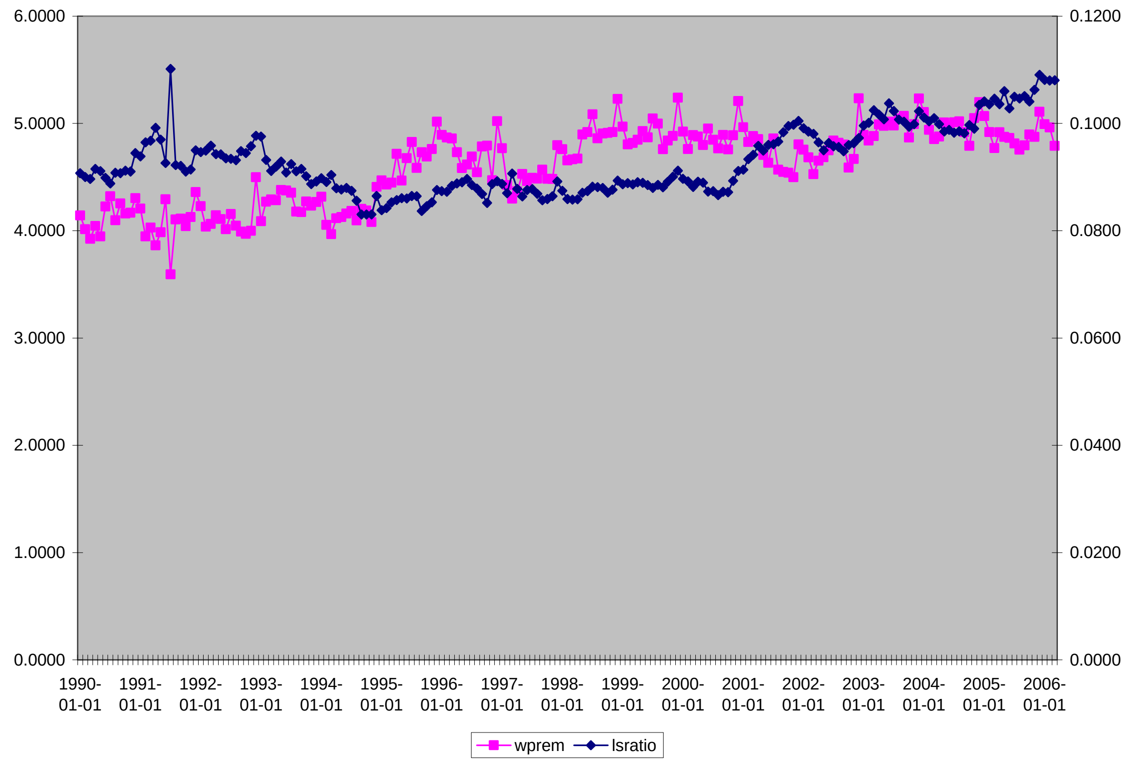

Figure 3. Wage Premium (wpremman), at left, and Ratio of Skilled to Unskilled Workers (lsratioman), at right, in Mexican Manufacturing: 1994:1-2006:3.

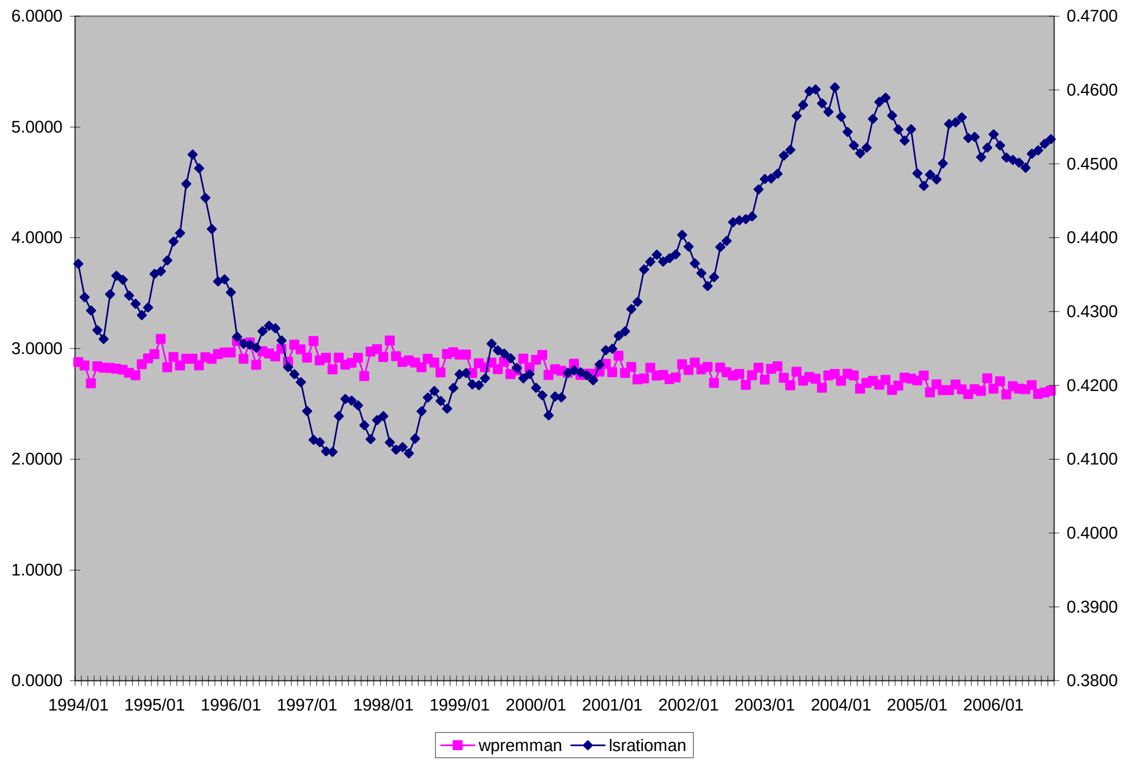

Figure 4. The Ratios between Capital and Equipment Expenses (K1) to Total Expenses, between Buildings and Land Purchases to Total Expenses (K2), and between (K1 + K2 = K3) to Total Expenses in Mexican Maquiladoras: 1990:1-2006:3.

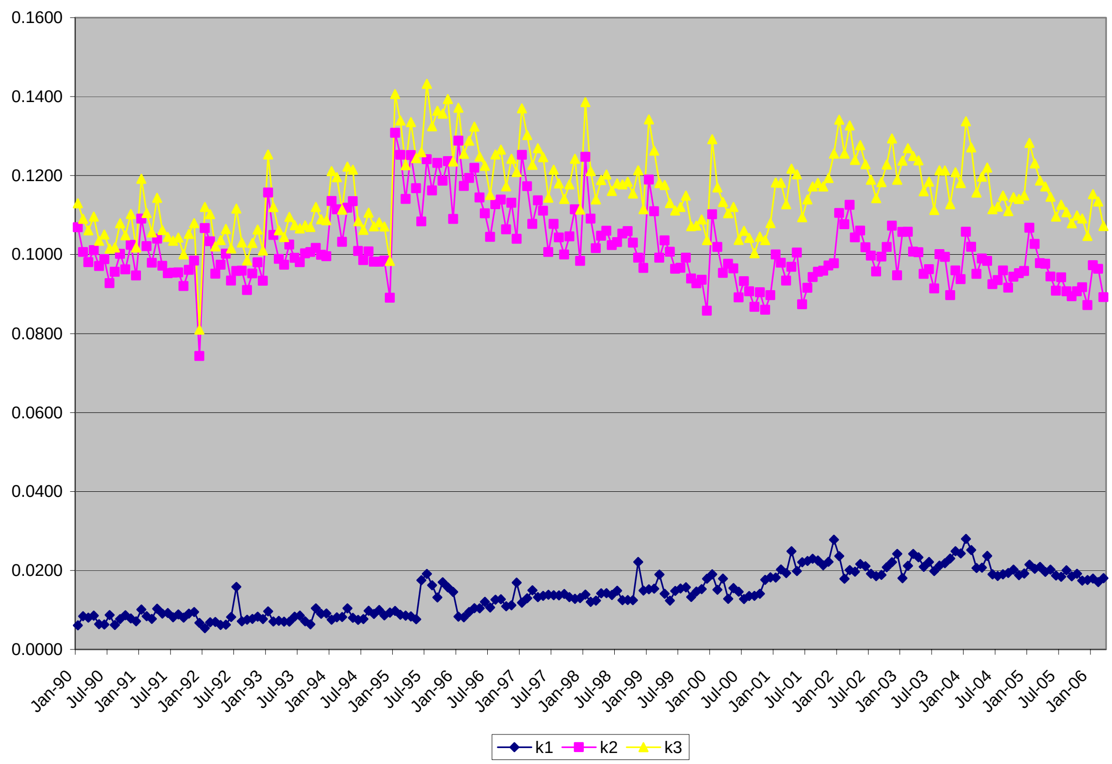

Table 1. Unit Root Tests on Monthly Data for Mexican Maquiladoras: 1990:1 to 2006:3.

Series	Trend?	ADF (k)	KPSS (4)
w	**Yes**	-2.77 (13)	0.40***
Δ (w)	**No**	-3.85 (12)**	0.04
H/L	**Yes**	-1.65 (1)	0.83***
Δ (H/L)	**No**	-13.48 (1)***	0.17
K1	**Yes**	-3.14 (3)*	0.18**
Δ(K1)	**No**	-6.92 (7)***	0.04

Notes: Data are of monthly frequency from 1990:1 to 2006:3. **w** refers to relative (real) wages between skilled and unskilled workers at Mexican maquiladoras; **H/L** refers to to relative labor supply between skilled and unskilled workers at Mexican maquiladoras; and **K1** is the share of capital equipment expenditures to total maquiladora expenditures. The symbol Δ refers to the first-difference of the original series. We include the deterministic trend only when testing in levels as suggested from graph inspection. 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) for ADF tests is chosen by the Campbell-Perron data dependent procedure, whose method is usually superior to k chosen by the information criterion, according to 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 at the 5% level. If no lags are significant, then set k = 0. Next to the reported calculated 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 used lag truncation parameter. The symbols * [**] (***) attached to the figure indicate rejection of the null at the 10%, 5%, and 1% levels, respectively.

Table 2. Granger Causality Tests between Wage Premiums and Relative Labor Supplies across Sectors: Varying Time Periods.

Sector	Lsratio ⇒ wprem	wprem ⇒ Lsratio
Maquiladoras, 1990 to 2006		
with 6 lags	2.602**	1.194
	[0.020]	[0.312]
with 5 lags	2.069*	1.372
	[0.072]	[0.237]
with 4 lags	2.667**	1.715
	[0.034]	[0.149]
Manufacturing, 1994 to 2006	**Lsratioman ⇒ wpremman**	**wpremman ⇒ Lsratioman**
with 6 lags	0.718	2.208**
	[0.636]	[0.047]
with 5 lags	0.384	3.836***
	[0.859]	[0.003]
with 4 lags	0.741	3.281**
	[0.566]	[0.014]
Maquiladoras and Manufacturing, 1994 to 2006	**Lsratio ⇒ wpremman**	**wpremman ⇒ Lsratio**
with 6 lags	1.759	1.455
	[0.114]	[0.200]
with 5 lags	2.081*	0.862
	[0.072]	[0.509]
with 4 lags	1.972	0.226
	[0.103]	[0.924]

Notes: Data are of monthly frequency from 1990:1 to 2006:3 for the maquiladoras and from 1994:1 to 2006:3 for the manufacturing sector. Reported are F-statistics of Granger causality tests, allowing for constant, time trend and eleven seasonal dummies. When the wage premium is the dependent variable, the null hypothesis is that the group of coefficients associated with lagged relative labor supplies is statistically significant to zero. The similar reasoning holds for when relative labor supply is the dependent variable. P-values are in brackets. The symbols * [**] (***) attached to the figure indicate rejection of the null at the 10%, 5%, and 1% levels, respectively.

Table 3. OLS Estimations of the Skill Premium for Mexican Maquiladoras.

$$\ln (w)_t = \beta_0 + \beta_1 trend + \beta_2 \ln (H/L)_t + \varepsilon_t \qquad (8)$$
$$\ln (w)_t = \beta_0 + \beta_1 trend + \beta_2 \ln (H/L)_t + \beta_3 \ln (K1)_t + \varepsilon_t \qquad (9)$$

	(8)c	(9)c	(8)t	(8)ct	(9)t	(9)ct
β_0	2.095*** (0.340)	1.720*** (0.228)		0.375* (0.192)		0.558** (0.243)
β_1			0.002*** (0.00007)	0.001*** (0.00007)	0.002*** (0.00009)	0.001*** (0.0001)
β_2	0.240* (0.143)	-0.217** (0.104)	-0.578*** (0.004)	-0.425*** (0.081)	-0.579*** (0.030)	-0.414*** (0.082)
Implied σ	**-4.167**	**4.608**	**1.730**	**2.353**	**1.727**	**2.415**
β_3		0.165*** (0.014)			0.0005 (0.015)	0.031* (0.018)
ADF (k) and KPSS on ε_t	**-1.45(11) 2.66*** **	**-3.01(12)** 0.51** **	**-5.56(14)*** 0.086**	**-5.28(14)*** 0.077**	**-5.56(14)*** 0.087**	**-5.29(14)*** 0.068**
Stationary Residuals?	**No**	**No**	**Yes**	**Yes**	**Yes**	**Yes**
Diagnostics						
DW	0.262	0.823	0.840	0.876	0.748	0.879
Adj. R^2	0.034	0.626	0.749	0.759	0.839	0.764

Notes: Data are of monthly frequency from 1990:1 to 2006:3. The dependent variable is the logarithm of the skill premium (w). 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 row "implied elasticity of substitution" reports the negative of the reciprocal of the parameter estimated in ln $(H/L)_t$; see section 3 for further explanation. We highlight in bold the cases when the elasticity is statistically significant and as expected. In the ADF (k) procedure the optimal k is selected by the sequential method proposed by Ng and Perron (1995), with kmax = 14. In the KPSS test we use the k=4 recommended lag length, in which the null hypothesis is of stationarity. Both tests have different null hypothesis as explained in the text.

Table 4. DOLS Estimations of the Skill Premium for Mexican Maquiladoras.

$$\ln (w)_t = \beta_0 + \beta_1 trend + \beta_2 \ln (H/L)_t + \Sigma_{k=0\ to\ 2}\, \gamma_k\, \Delta \ln(H/L)_{t-k} + \varepsilon_t \qquad (8)$$

$$\ln (w)_t = \beta_0 + \beta_1 trend + \beta_2 \ln (H/L)_t + \beta_3 \ln (K1)_t + \Sigma_{k=0\ to\ 2}\, \gamma_k\, \Delta \ln(H/L)_{t-k} + \Sigma_{k=0\ to\ 2}\, \varphi_k\, \Delta \ln(K1)_{t-k} + \varepsilon_t \qquad (9)$$

	(8)c	**(8)ct**	**(9)t**	**(9)ct**
β_0	1.992*** (0.338)	0.259 (0.197)	1.511*** (0.208)	0.382 (0.307)
β_1		0.0014*** (0.00007)		0.0013*** (0.0002)
β_2	0.198 (0.141)	-0.473*** (0.082)	-0.335*** (0.096)	-0.471*** (0.086)
Implied σ	**-5.051**	**2.114**	**2.985**	**2.123**
β_3			0.181*** (0.016)	0.024 (0.033)
ADF (k) and KPSS on ε_t	**-2.47(12)** **2.55*** **	**-4.17(12)***** **0.06**	**-3.49(12)***** **0.02**	**-4.30(12)***** **0.06**
Stationary Residuals?	**No**	**Yes**	**Yes**	**Yes**
Diagnostics				
DW	0.240	0.940	0.731	0.927
Adj. R^2	0.071	0.775	0.691	0.775

Notes: Data are of monthly frequency from 1990:1 to 2006:3. The dependent variable is the logarithm of the skill premium (w). The method of estimation is dynamic OLS (DOLS) as in Sock and Watson (1993), allowing for contemporaneous, leads and lags (k = 2) of first differenced terms, which are not reported. 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 row "implied elasticity of substitution" reports the negative of the reciprocal of the parameter estimated in ln $(H/L)_t$; see section 3 for further explanation. We highlight in bold the cases when the elasticity is statistically significant and as expected. In the ADF (k) procedure the optimal k is selected by the sequential method proposed by Ng and Perron (1995), with kmax = 14. In the KPSS test we use the k=4 recommended lag length, in which the null hypothesis is of stationarity. Both tests have different null hypothesis as explained in the text.

Table 5. ECM Estimations for Mexican Maquiladoras.

$$\Delta \ln(w)_t = \beta_0 + \beta_1[\varepsilon]_{t-1} + \Sigma\, \gamma_i\, \Delta \ln(H/L)_{t-i} + \Sigma\, \phi_i\, \Delta \ln(w)_{t-i} + \Sigma\, \varphi_i\, \Delta \ln(K1)_{t-i} + \nu_t$$

	ols (9)t	**ols (9)ct**	**dols (8)ct**	**dols (9)ct**
β_1	-0.293***	-0.273***	-0.230***	-0.228***
	(0.054)	(0.054)	(0.063)	(0.063)
γ_1	0.420***	0.389***	0.354***	0.352***
	(0.095)	(0.083)	(0.086)	(0.085)
γ_3		-0.232***	-0.283***	-0.264***
		(0.084)	(0.089)	(0.101)
γ_4		-0.184***	-0.178***	-0.161**
		(0.072)	(0.071)	(0.081)
γ_{12}	0.458***	0.452***	0.329***	0.463***
	(0.070)	(0.073)	(0.058)	(0.070)
ϕ_1	-0.215***	-0.250***	-0.283***	-0.291***
	(0.083)	(0.076)	(0.089)	(0.087)
ϕ_3	-0.180***	-0.170***	-0.178**	-0.183**
	(0.077)	(0.066)	(0.071)	(0.070)
ϕ_{12}	0.330***	0.326***	0.329***	0.327***
	(0.059)	(0.059)	(0.058)	(0.059)
φ_1	0.023**	0.023**	0.023**	0.023**
	(0.009)	(0.009)	(0.009)	(0.009)
DW	2.159	2.118	2.106	2.100
Adj. R^2	0.464	0.475	0.455	0.454

Notes: Data are of monthly frequency from 1990:1 to 2006:3. The dependent variable is the first-difference of the skill premium: $\Delta \ln(w)_t$. The constant term is included in the estimation but is not reported. The method of estimation is OLS under the ECM framework, in which ε are the residuals from the cointegrating equation between wages and relative labor supplies in Table 2; and among wages, labor supplies and capital equipment share in Table 3. A general-to-specific methodology is adopted in selecting the lag-lengths for the γ, ϕ and φ terms, with maximum lags set at 14. We report in the Table only the statistically significant lags at the 5% level. 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.

Table 6. Vector ECM of Maquiladora Skill Premium: Base and Augmented Models.

$$\ln(w)_t = \beta_0 + \beta_1 \text{ trend} + \beta_2 \ln(H/L)_t + \varepsilon_t \quad (8)$$

$$\ln(w)_t = \beta_0 + \beta_2 \ln(H/L)_t + \beta_3 \log K1_t + \mu_t \quad (9a)$$

$$\ln(w)_t = \beta_0 + \beta_1 \text{ trend} + \beta_2 \ln(H/L)_t + \beta_3 \log K1_t + \mu_t \quad (9b)$$

	(8)	**(9a)**	**(9b)**
β_1	0.0014***		-0.0008
	(0.0002)		(0.0006)
β_2	-0.541***	-0.334**	-0.186
	(0.150)	(0.153)	(0.215)
β_3		0.206***	0.324***
		(0.024)	(0.081)
Diagnostics			
Lag-Length	3	4	4
(criteria)	[all 5]	[FPE, AIC]	[FPE, AIC]
LM-test h = 4	0.49	12.89	12.89
[P-value]	[0.97]	[0.17]	[0.17]
LM-test h = 8	0.37	5.64	5.64
[P-value]	[0.99]	[0.78]	[0.78]
LM-test h = 12	7.45	22.84**	22.84**
[P-value]	[0.11]	[0.01]	[0.01]
ECM in VECM	-0.235***	-0.165***	-0.092**
(std. error)	(0.063)	(0.053)	(0.039)
γ_i or φ_i in ECM	$\gamma_1 = 0.286$**	Various	Various
	(0.121)	φ's < 0	φ's < 0
Adj. R^2 in VECM	0.563	0.552	0.541
Null of No Coint. Vector	Trace: 24.59 < 25.87 Max. Eigenvalue: **21.09** > 19.39**	Trace: 26.71 < 29.80 Max. Eigenvalue: 19.22 < 21.13	Trace: 37.17 < 42.92 Max. Eigenvalue: 20.71 < 25.82

Notes: Data are of monthly frequency from 1990:1 to 2006:3. Standard errors are reported below the coefficients. The method of estimation is the vector error correction model with eleven seasonal dummies. In the first-stage the Johansen cointegration method is used and in the second stage residuals from the first-stage are used in differenced form. The LM t-stat. is a standard Lagrange Multiplier test on the residuals of the regression, calculated under the null hypothesis of no serial correlation. The trace and maximal eigenvalue tests associated with the null of no cointegration vectors in the Johansen test are reported at the bottom of the table. The symbols * [**] (***) attached to the figure indicate rejection of the null hypothesis of zero coefficients at the 10%, 5%, and 1% levels, respectively.