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Analyzing the Effects of Labor Standards on U. S. Export Performance. A Time Series Approach with Structural Change*

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Abstract

We analyze the empirical effects of different measures of labor standards on the export performance of the United States using annual data for the period 1950-1998, applying a time series approach based on the structural change literature. Hence, we estimate a model with endogenous breaks following the methodology proposed by Bai and Perron (1998). The results show that the labor standards, represented by the number of hours worked, the rate of occupational injuries and the unionization rate, are all very important to explain the behavior of exports for the United States. In particular, we find that low labor standards may both improve or lead to a deterioration in export performance.

Keywords: Exports, Labor Standards, Stationarity, Structural Change, Unit Roots.

JEL: F1, C1, C4

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1 Introduction

There is still a lot of debate as to whether labor standards should be uniform (or harmonized) across countries, and this debate will very likely continue to be part of future trade talks as it remains a means for countries with high labor standards to attempt to maintain a protectionist policy against imports from low-standard countries (Anderson, 1995). Indeed, both the United States (henceforth *US*) and France have unsuccessfully attempted to introduce the issue of labor standards during the Uruguay Round, and in the last ten years, the *US* has included a worker rights clause into many of its trade agreements. Efforts by some governments, albeit extensive, to include core labor standards in the World Trade Organization (*WTO*) have also led to the acceptance that this issue should be raised, but only at *WTO* preparatory meetings (Aggarwal, 1995).[1] Many, on the contrary, have argued that only the International Labor Organization (*ILO*), whose main role is to improve labor standards worldwide, should deal with the issue of labor standards. However, proponents of the latter argument have to also acknowledge the lack of enforcement mechanisms by the *ILO* to ensure compliance with labor standards across countries.

Several studies have examined the relationship between labor standards and export performance. Aggarwal (1995) argues that the imposition of core labor standards in developing countries is likely to increase production costs and, therefore, investigates whether labor standards are being suppressed by developing countries so that they can reduce production costs and encourage exports. Her examination of the export patterns of ten developing countries to the *US* for 1994 shows that sectors with the weakest labor standards are not the only or primary share of these countries' exports. She does not find any clear evidence to support the argument that low labor standards are a boost to export performance. An *OECD* (1996) study, considering both *OECD* and non-*OECD* countries, finds no evidence that countries with low labor standards, as measured by freedom of association and collective

[1] The proposition that trade agreements should not include labor standards has lost ground over the last few years. Indeed, the *WTO* has already considered intellectual property rights and environmental matters, and indicated that it will also have to deal with labor standards in some manner.

bargaining rights, can achieve a better export performance than countries with high labor standards. The study, which is unfortunately essentially based on graphical analysis, also finds that countries that have liberalized trade do not necessarily face a worsening of their labor standards, in that case equated with freedom of association rights.

Mah (1997) also examines the relationship between core labor standards and export performance of developing countries. His cross-sectional study is, however, limited by the fact that he uses the ratification of core conventions as an indicator for labor standards. Indeed, the fact that countries ratify conventions does not necessarily mean that the standards that these conventions represent are being enforced.[2] Despite this major shortcoming, Mah's results overall indicate that higher labor standards have a negative impact on export performance. In another study, Rodrik (1996) uses different proxies for labor standards[3] and looks at their effects on labor costs and comparative advantage. His results show that labor standards are significant determinants of labor costs when one controls for productivity; but they are not important determinants of comparative advantage, the latter being determined mostly by factor endowments.

Van Beers (1998) considers the relationship between labor standards and trade flows of $OECD$ countries, using an indicator for labor standards based on actual labor regulations. The indicator is a synthetic index constructed by the $OECD$ and which takes into account the enforcement of various government regulations such as working time, employment contracts, minimum wages and workers' rights. The author extends a gravity model, which considers bilateral trade flows, with variables that represent the strictness of labor regulations, and tests the hypothesis whether labor standards have a detrimental effect on exports due to a fall in competitiveness. His results do not show any significant impact of labor standards stringency on exports of

[2] In fact, we have shown in another paper (Samy, 2000) that the impact on export performance is largely dependent on the choice of indicators for labor standards, and that Mah's model is very sensitive to the type of specification used.

[3] He in fact uses ILO conventions ratified, Freedom House indicators of civil liberties and political rights, an indicator of the incidence of child labor, statutory hours in a normal week in manufacturing and construction, days of paid annual leave in manufacturing, and percentage of the labor force that is unionized.

labor-intensive commodities. However, when a distinction is made in terms of skill-intensities, both the exports of labor-intensive and capital-intensive commodities, which are produced with relatively high-skilled labor, deteriorate with an increase in the strictness of labor standards.

Most of the existing studies related to the issue of trade and labor standards assume that firms, industries or countries can boost their competitiveness by lowering labor standards since the latter involves a cost. However, in our view, this is not always true, and it is quite possible that the lowering of standards may have adverse effects on productivity and efficiency, and hence on export performance. In other words, weaker labor standards, instead of providing a competitive advantage, may raise costs. Maskus (1997) and Martin and Maskus (1999) have actually shown, through a series of simple partial equilibrium models, that the lack of enforcement of core labor standards reduces an economy's efficiency and alters its comparative advantage.

In the present paper, we are able to consider data for labor standards for a period extending over almost fifty years in order to analyze the effect of labor standards on export performance for the US. Unlike the general approach in the literature, which is based on cross-section analysis (see, for example, Rodrik (1996) and Mah (1997)), we use a time-series approach based on the structural change literature. This allows us, not only to identify structural breaks in the data, but also to see how the effects of labor standards on export performance evolve over time. The structure of the paper is as follows. Section 2 presents the model that will be tested. Section 3 describes the indicators for labor standards which are considered in the empirical analysis. Section 4 examines the stationarity properties of the series being considered, describes the methodology which is used, and reports the empirical results as well as their implications. Section 5 concludes.

2 The Model

The model that we seek to estimate is derived from the perfect and imperfect substitutes models, which are used to analyze the time-series behavior of exports and imports. In these models, prices and incomes are what essentially determine exports and imports (see Goldstein and Khan, 1985). The perfect substitutes model is based on the assumption that homogeneous goods are traded on international markets at a common price. It can be represented by the following set of equations for a representative country i:

$$
\begin{aligned}
D_i &= f(P_i, Y_i), \qquad f_1 < 0, \quad f_2 > 0 & (1)\\
S_i &= g(P_i, Y_i), \qquad g_1 > 0, \quad g_2 < 0 & (2)
\end{aligned}
$$

where D_i is the quantity of goods demanded in country i; S_i is the supply of goods produced in country i; P_i is the domestic price of traded goods; and Y_i and F_i are money income and factor costs. The model also considers the following equations:

$$
\begin{aligned}
I_i &= D_i - S_i & (3)\\
X_i &= S_i - D_i & (4)\\
PI_i &= PX_i = eP_w & (5)
\end{aligned}
$$

Equations (3) and (4) show the quantity of country i's imports and exports. PI_i, PX_i and $e.P_w$ are the import, export and world prices respectively. Equation (5) tells us that there exists only one traded goods price, abstracting from transport costs and other trade barriers, determined by the interaction of world demand and world supply. Replacing equations (1) and (2) in equation (5), we obtain the following:

$$X_i = g(P_i, Y_i) - f(P_i, Y_i) \qquad (6)$$

Equation (6), therefore, implies that a country's export volume depends on domestic prices, money income, and factor costs within the country.

The main assumption of the imperfect substitutes model is that imports and exports are not perfect substitutes for domestic goods.[4] If this were not true, then, contrary to what is often observed, a given country should not import or export the same traded good. The main difference between the imperfect substitutes model and the perfect substitutes model is that equation (5) does not hold any more.

[4] For a comprehensive treatment of the imperfect substitutes model, refer to Goldstein and Khan (1985).

In order to analyze the impact of labor standards on export performance, the basic equation, which is an application of the perfect substitutes model described above (see also Mah (1997)), is as follows:

$$\log(exp_t/gdp_t) = \alpha_0 + \sum_{i=1}^{h} \alpha_i labstd_{it} + \epsilon_t \tag{7}$$

where, $\log(exp_t/gdp_t)$ is the logarithm of the ratio exp_t/gdp_t at time t. In this case, exp_t is the export value in US dollars and the variable gdp_t is gross domestic product in US dollars. The variable $labstd_t$ is a measure of a particular labor standard. Here, we use three different measures ($h = 3$); and ϵ_t is the disturbance term which is assumed to be $i.i.d.$ $(0, \sigma_\epsilon^2)$.

In the imperfect substitutes model, prices have to adjust in each time period to maintain the equality between demand and supply. Due to the existence of costs to changing prices in imperfectly competitive markets, firms have to weigh the costs of changing prices against adjustment costs such as changes in inventories or unfilled orders. The basic equation to analyze the effects of labor standards on export performance in the case of the imperfect substitutes model will therefore be as follows:

$$\log(exp_t/gdp_t) = \alpha_0 + \sum_{i=1}^{h} \alpha_i labstd_{it} + \gamma rin_t + e_t \tag{8}$$

where, rin_t is defined as the lending rate minus the rate of consumer price increase and the other terms are as indicated in equation (7). The variable rin_t is included to take account of the additional costs of adjustment mentioned above.

Equations (7) and (8) in effect both assume that the export performance of a country is determined by its price competitiveness. The effect of a labor standard is to lead to an increase in labor costs and hence cause a deterioration in the price competitiveness of a country's exports. In order to interpret the results in the next section of the paper, the null hypothesis is that the coefficient of $labstd_t$ (that is, α_i) is zero such that labor standards do not influence export performance. As far as the alternative hypothesis goes, it is assumed that the sign of α_i is different from zero. Even though one would expect α_i to be negative since the labor standards lead to a rise in labor costs, it is quite possible, as explained at the end of section 1 of the paper, that they may have the opposite effect.[5] The coefficient γ is expected to have a negative sign because an increase in the real interest rate can raise the capital cost and hence lead to a deterioration of price competitiveness.

3 Indicators for Labor Standards

The choice of appropriate indicators for labor standards entails certain difficulties. Ideally, one would first agree on the labor standards to be considered and then use them to assess their effects on export performance. Following $OECD$ (1996), we take labor standards as being norms, rules and conventions that govern working conditions and industrial relations. For example, the ratification of ILO conventions, the number of hours worked, and occupational health and safety standards at the workplace would fit that definition.

Several authors, among which Stern (1998), Aggarwal (1995) and Fields (1995), have discussed these definitional issues at length. The biggest difficulty, it seems, is to distinguish labor rights (that is, standards that should apply to everybody regardless of a country's stage of development, also known as core labor standards) from other labor standards, which are more dependent on national characteristics. There is now, however, a growing consensus that core labor standards should include freedom of association, collective bargaining rights, elimination of exploitative forms of child labor, prohibition of forced labor, and non-discrimination in employment among genders.

What is more important in our case is to find indicators that can encapsulate rules related to the conditions of work and this leads us to another difficulty, which is the availability of data. Because of the nature of our analysis, namely the time-series approach, we have to find data for labor standards that go back far enough in time. As a result, the number of indicators that we use for labor standards is limited to the ones described below. Time series data is collected for the US for the period 1950 to 1998, and we

[5] Swinnerton (1997) also provides a detailed explanation of these efficiency effects.

define the labor standards that we considered in the next paragraph.[6]

We consider the actual number of weekly hours worked in the manufacturing sector. The numbers are obtained from various issues of the *Yearbook of Labor Statistics* of the *ILO* and generally represent the average hours worked by wage earners. They vary from roughly 37 to 42 hours per week for the period considered. This variable is denoted by $hour_t$ and we use the logarithm of this variable ($lhour_t$). We consider the occupational injuries per thousand people employed in manufacturing industries and it can be interpreted as an indicator of safety at the workplace. We are able to construct the series from various issues of the *Yearbook of Labor Statistics* of the *ILO*. This variable is denoted as inj_t. Finally, we consider the union membership as a percentage of the non-agricultural paid workers. The data is obtained from the *Datapedia of the United States* (1994) and the *World Labor Report* (1997). This variable is denoted as $union_t$.

4 Empirical Analysis

In this section, the empirical analysis applied to the *US* data is presented. The data is annual and it spans the period 1950 to 1998. The section is divided into two parts. First, the stationarity properties of each variable are presented. Second, the methodology used to estimate the models and the results obtained from these estimations, as well as the implications thereof, are presented.

4.1 Stationarity Analysis

In order to investigate the univariate process of each of the time series that we are considering, three tests are applied. The first and second statistics are the augmented Dickey-Fuller statistic (hereafter *ADF*) proposed by Dickey and Fuller (1979) and extended by Said and Dickey (1984) to the case of data having an ARMA structure; and the *ADF* statistic based on the Generalized Least Squares (*GLS*) detrending procedure proposed by El-

[6] Given the methodology we use, we cannot use the ratification of ILO Conventions (represented by dummy variables) as one of our indicators (as used by Rodrik (1996) and Mah (1997), for instance).

liot, Rothenberg and Stock (1996) (hereafter ERS). This statistic is labeled as ADF^{GLS}. The third statistic allows for a more complex dynamic by introducing the possibility that the series are best represented by a shifting mean, a broken trend, or both simultaneously. Originally, this kind of test was proposed by Perron (1989) using a known break point. After that, Zivot and Andrews (1992), Banerjee, Lumsdaine and Stock (1992), Perron (1997) and Perron and Rodríguez (2002), have proposed similar tests but considering an unknown break point. The version of Perron (hereafter $P97$) is used here.

The class of process considered can be described as follows. We denote the relevant data series by y_t (in our case, this variable represents one of the following time series $log(exp/gdp)_t$, inj_t, $lhour_t$, rin_t, $union_t$) and write:

$$\begin{aligned} y_t &= d_t + u_t \\ A(L)u_t &= B(L)e_t \end{aligned} \tag{9}$$

where $A(L)$ is a p^{th} order autoregressive polynomial in the lag operator L (defined such that $Ly_t = y_{t-1}$) defined by $A(L) = 1 + \phi_1 L + +\phi_p L^p$. Similarly, $B(L)$ is a q^{th} order moving-average polynomial defined by $B(L) = 1 + \theta_1 L + +\theta_q L^q$. The errors $\{e_t\}$ are assumed to be martingale differences. The system (9) simply describes a process that is the sum of a deterministic time trend (d_t) and a noise function modeled as an $ARMA$ process. The null hypothesis is that one root of the autoregressive polynomial is unity. The term d_t is equal to $\psi' z_t$ where z_t is the set of deterministic components in the series. Hence, we can have $z_t = \{1\}$ or $z_t = \{1, t\}$.

The ADF statistic is based on the idea that a stationary and invertible $ARMA$ process can be approximated by an autoregression. Hence, the relevant autoregression estimated by OLS is:

$$y_t = d_t + \alpha y_{t-1} + \sum_{j=1}^{k} c_j \Delta y_{t-j} + v_t \tag{10}$$

Here, the parameterization (10) is such that the coefficient α is the sum of the autoregressive coefficients. Hence, the null hypothesis can be tested using the t-statistic constructed for $\alpha = 1$.

To apply the ADF^{GLS} statistic, the approach of ERS consists of first locally removing the deterministic components of $\{y_t\}$ via GLS. Denoting $y_t^{\overline{\alpha}}$ and $z_t^{\overline{\alpha}}$ as:

$$\left\{ \begin{array}{ll} y_t^{\overline{\alpha}} = y_1, (1-\overline{\alpha}L)y_t & t = 2,3,...,T \\ z_t^{\overline{\alpha}} = z_1, (1-\overline{\alpha}L)z_t & t = 2,3,...,T \end{array} \right\} \tag{11}$$

where $\overline{\alpha} = 1+\overline{c}/T$, with $\overline{c} = -7.0$ for the case where $z_t = \{1\}$, and $\overline{c} = -13.5$ when $z_t = \{1, t\}$. We define $\widetilde{\psi}$ as the estimator which minimizes the squared sum of residuals:

$$S(\psi, \alpha) = \sum_{t=0}^{T} (y_t^{\overline{\alpha}} - \psi' z_t^{\overline{\alpha}})^2. \tag{12}$$

Then we use $\widetilde{\psi}$ to construct the detrended series $\widetilde{y}_t = y_t - \widetilde{\psi}' z_t$ and to apply an ADF test on $\widetilde{y}_t$:

$$\widetilde{y}_t = \alpha \widetilde{y}_{t-1} + \sum_{j=1}^{k} c_j \Delta \widetilde{y}_{t-j} + v_t \tag{13}$$

An issue of empirical importance is the choice of the order of the autoregression k in (10) and (13). Following Campbell and Perron (1991) and Ng and Perron (1995), a data-dependent method based on a general to specific recursive procedure is used. Starting from a maximal order of k (say, $kmax$), the method tests if the last lag included is significant[7], and if not, the order of the autoregression is decreased by one and the coefficient of the last lag again examined. This is repeated until a rejection occurs or the lower bound zero is reached. In our case, $kmax = int(10 * (T/100)^{0.25})$ is considered.

For the unit root test in the presence of structural change, three possible cases are considered according to the deterministic components. In the first model, $z_t = \{1, t, DU_t\}$ with $DU_t = 1(t > T_B)$, where $1(.)$ is the indicator function. This model considers the possibility that there is a break in the intercept, which in the terminology of Perron (1989) corresponds to the "crash model". The second model considers $z_t = \{1, t, DT_t\}$, where $DT_t = 1(t > T_B)(t - T_B)$. In this model, a change in the slope of the trend

[7] According to the recommendation of Campbell and Perron (1991), a significance level of 90.0% is used.

function is allowed. Finally, a third model where $z_t = \{1, t, DU_t, DT_t\}$ is also considered. In this model both changes are allowed. In all cases, T_B is the date of the break point, which is also defined as $\lambda = T_B/T$.

In order to choose T_B, we estimate the ADF statistic above through $t = kmax + 1, ..., T-1$, and we select the date, $\widehat{T}_B$, as the point associated to the infimum value of the $t_{\widehat{\alpha}}$.

Results are presented in Table 1. For all variables, except for inj_t, an intercept and a time trend in the specification of the deterministic components of the time series are included. For the statistic $P97$, only the case where a rejection was obtained is presented. When no rejection is found, results for the three models considered for the analysis of stationarity are presented.

In general, the ADF statistic presents evidence of non-stationarity. The only exception is the dependent variable. When the ADF^{GLS} statistic is applied[8], we found evidence of stationarity for most of our variables. When the test with one break of Perron (1997) is applied, a more clear evidence in favor of stationarity is obtained. In the case of the variable $log(exp/gdp)_t$, a rejection of the hypothesis of a unit root is obtained with the model where only a break in the intercept is allowed. We found a break point at the year 1970. The variable rin_t also shows evidence of stationarity with the year 1979 as the date of the break point. The logarithm of the worked hours ($lhour_t$) is also stationary using the model where there is only a break in the slope of the trend function. In this case, the break point found corresponds to 1981. Finally, the variable $union_t$ is stationary using the model where a break is allowed only in the intercept. In this case, the break point corresponds to the year 1982. It is interesting to note that the break points obtained for the variables rin_t, $lhour_t$ and $union_t$ correspond to the second oil-price shock of 1979 to 1982, and the ensuing worldwide recession, the effects of which were also felt in the U. S..

[8] Finite critical values from ERS for $T = 50$ when $z_t = \{1, t\}$ were used.

4.2 Estimating the Model with Multiple Structural Breaks

According to the results from the stationarity analysis, the variables are stationary and possibly with a broken trend. Hence, we will estimate a model with endogenous breaks that enables us to have different coefficients for each regime. This method to estimate multiple structural change models was proposed by Bai and Perron (1998).

4.2.1 The Methodology

Following Bai and Perron (1998) and using similar notation, the following multiple linear regression with m breaks ($m+1$ regimes) is considered:

$$\left\{ \begin{array}{cc} i_t = z_t'\gamma_1 + x_t'\beta + \epsilon_t, & t = 1, 2, ..., T_1 \\ i_t = z_t'\gamma_2 + x_t'\beta + \epsilon_t, & t = T_1 + 1, ..., T_2 \\ . & \\ . & \\ . & \\ i_t = z_t'\gamma_{m+1} + x_t'\beta + \epsilon_t, & t = T_m + 1, ..., T \end{array} \right\} \tag{14}$$

where i_t is the observed dependent variable at time t; z_t $(q \times 1)$ and x_t $(p \times 1)$ are vectors of covariates and γ_j $(j = 1, 2, ..., m+1)$ and β are the corresponding vectors of coefficients; and ϵ_t is the disturbance term at time t. In our case, $i_t = log(exp/gdp)_t$ and the other variables are included either in vectors x_t or z_t. If the analyst has reasons to believe that one or more coefficients are stable through time, then we have a model where both $q > 0$ and $p > 0$, which is known as a "partial structural change model" since the coefficients contained in β are not allowed to shift. Whereas if $p = 0$, we then have a "pure structural change model" and all the coefficients are shifting. This is the model that we consider here. The indices $(T_1, ..., T_m)$, or the break points, are treated as unknown. Hence, this methodology enables us to estimate the unknown regression coefficients together with the break points when T observations on (i_t, x_t, z_t) are available.

The method of estimation is based on the least squares principle. For each $m-$partition $(T_1, ..., T_m)$, the associated leas-squares estimates of β and γ_j are obtained by minimizing the sum of squared residuals $S_T(.) =$

$\sum_{i=1}^{m+1}\sum_{t=T_{i-1}+1}^{T_i}[i_t - x_t'\beta - z_t'\gamma_i]^2$. Then we have $\hat{\beta}(\{\hat{T}_{B_j}\})$ and $\hat{\gamma}_j(\{\hat{T}_{B_j}\})$ which can be considered as the resulting estimates based on the given $m-$partition $(T_1, ..., T_m)$ denoted by $\{T_j\}$. Substituting these estimates in the objective function, the estimated break points $(\widehat{T}_1, ..., \widehat{T}_m)$ are such that $(\widehat{T}_1, ..., \widehat{T}_m) = \arg\min_{\{T_1,...,T_m\}} S_T(T_1, ..., T_m)$. Then, the break-point estimators are global minimizers of the objective function. Finally, the regression parameter estimates are obtained using the associated least-squares estimates at the estimated $m-$partition $\{\widehat{T}_j\}$, i.e. $\widehat{\beta} = \widehat{\beta}(\{\widehat{T}_j\}), \widehat{\gamma} = \widehat{\gamma}(\{\widehat{T}_j\})$.

Before to be able to estimate the model, we have to select the number of break points. For that, Bai and Perron (1998) propose various statistics. The first statistic is the *sup* F_T type test of the null hypothesis of no structural break $(m = 0)$ versus the alternative hypothesis that there are $m = k$ breaks. Let R be the conventional matrix such that $(R\gamma) = (\gamma_1' - \gamma_2', ..., \gamma_k' - \gamma_{k+1}')$. The *sup* F_T test proposed is asymptotically equivalent and it uses the estimates of the break dates obtained from the global minimization of the sum of squared residuals. Denoting these estimates by $\widehat{\lambda}_i = \widehat{T}_i/T$ for $i = 1, ..., k$, the test is then $\sup_{(\lambda_1,...,\lambda_m)\in\Lambda_\epsilon} F_T(k; q) = F_T(\widehat{\lambda}_1, ..., \widehat{\lambda}_k; q)$, where $\widehat{\lambda}_1, ..., \widehat{\lambda}_k$ are the arguments that maximize the following F_T-statistic:

$$F_T(\lambda_1, ..., \lambda_k; q) = \frac{1}{T}(\frac{T - (k+1)q - p}{kq})\widehat{\gamma}' R'(R\widetilde{V}(\widehat{\gamma})R')^{-1}R\widehat{\gamma}, \qquad (15)$$

and $\widetilde{V}(\widehat{\gamma}) = (\frac{\overline{Z}' M_X \overline{Z}}{T})^{-1}$, is the covariance matrix of $\widehat{\gamma}$ assuming spherical errors. Maximizing this F_T-statistic is equivalent to minimizing the global sum of squared residuals. As is showed in Bai and Perron (1998), this procedure is asymptotically equivalent since the break dates are consistent even in the presence of serial correlation.[9]

When the investigator wishes not to pre-specify a particular number of breaks to make inference, Bai and Perron (1998) propose two tests of the null hypothesis of no structural break against an unknown number of breaks given some upper bound M. These are called the double maximum tests. The first statistic is an equal weighted version; while the second statistics

[9] The asymptotic distribution still depends on the specification of the set Λ_ϵ via the imposition of the minimal lenght of a segment. Hence, $\epsilon = h/T$.

applies weights to the individual tests such the marginal p-values are equal across values of m. It implies weights that depend on q and the significance level of the test, say α. Both statistics are defined by (in their asymptotically equivalent version):

$$
\begin{aligned}
UD\max\ F_T(M,q) &= \max_{1\leqslant m\leqslant M} F_T(\hat{\lambda}_1,...,\hat{\lambda}_m;q), && (16)\\
WD\max\ F_T(M,q) &= \max_{1\leqslant m\leqslant M} \frac{c(q,\alpha,1)}{c(q,\alpha,m)} F_T(\hat{\lambda}_1,...,\hat{\lambda}_m;q), && (17)
\end{aligned}
$$

where, as before, $\widehat{\lambda}_i = \widehat{T}_i/T$ $(i = 1,...,m)$ are the estimates of the break points obtained using the global minimization of the sum of squared residuals and $c(q,\alpha,m)$ is the asymptotic critical value of the test $\sup_{(\lambda_1,...,\lambda_m)\in\Lambda_\epsilon} F(\lambda_1,...,\lambda_m;q)$ for a significance level α.

Other proposed statistic works in a sequential way. This test is named *sup* $F_T(l+1|l)$. The method amounts to the application of $(l+1)$ tests of the null hypothesis of no structural change versus the alternative hypothesis of a single change. The test is applied to each segment containing the observation $\widehat{T}_{i-1}$ to $\widehat{T}_i$ $(i = 1,...,l+1)$. In this case, the estimates $\widehat{T}_i$ need not be the global minimizers of the sum of squared residuals, all that is required is that the break fractions $\widehat{\lambda}_i = \widehat{T}_i/T$ converge to their true value at rate T. We conclude for a rejection in favor of a model with $(l+1)$ breaks if the overall minimal value of the sum of squared residuals (over all segments where an additional break is included) is sufficiently smaller than the sum of squared residuals from the l breaks model. The break date is selected as the one associated with this overall minimum.

Finally, other criteria to select for the number of breaks are the BIC and LWZ methods. They are defined, respectively, by

$$
\begin{aligned}
BIC(m) &= \ln\hat{\sigma}(m) + g\ln(T)/T && (18)\\
LWZ(m) &= \ln\left(\frac{S_T(\hat{T}_{B_1},...,\hat{T}_{B_m})}{T-g}\right) + \left(\frac{g}{T}\times c_0(\ln(T))^{2+\delta_0}\right) && (19)
\end{aligned}
$$

where $g = (m+1)q+m+p$, and $\hat{\sigma}^2(m) = T^{-1}S_T(\hat{T}_1,...,\hat{T}_2)$ and, according to Liu *et al.* (1997), $c_0 = 0.299$ and $\delta_0 = 0.1$.

4.2.2 The Results

All the results were obtained using the Gauss program constructed by Bai and Perron (1998). In the configuration of the parameter of the model, a trimming of 15% ($\epsilon = 0.15$) is considered, which corresponds to an $h = 7$. We have $q = 5$ or $q = 6$, depending if equation (7) or (8) is estimated, respectively. As we know, parameter q, specifies the number of regressors that change in the time period. Given that we have the lagged dependent variable as another explanatory variable, we do not use automatic correction for autocorrelation. Finally, the maximum number of breaks allowed in the estimation was $M = 5$.

The results of the estimation of the models are presented in Tables 2-4. Table 2 presents the results from the different statistics to select break points for equation (7) and (8), as is showed in second and third columns, respectively. Firstly, consider the results for equation (7). The statistics $UDmax$ and $WDmax$ show (at 1.0% level of significance) that there exists at least one break in the model. A similar conclusion is obtained when the statistic $sup\ F_T$ is used. In this case, we always reject the null hypothesis that there is no break in the model. The sequential application of the test $sup\ F_T$ rejects the null hypothesis of one break in favor of two breaks using 1.0% level of significance. The next null hypothesis (of only two breaks) is also rejected in favor of three breaks but this time using 2.5% level of significance. Finally, the result using BIC also shows that there is evidence in favor of three breaks. Notice also that the sequential procedure found only one break when 1.0% level of significance is used. Three breaks are found using other significance levels.

Using the previous results, Table 3 shows the results from the estimation of the model with three breaks and one break, respectively. Let consider the case with 4 regimes. In the first regime (1950-56), the variables $lhour_t$ and $union_t$ are significant at the 1% level and explain the behavior of $log(exp/gdp)_t$. In the second regime (1957-1971), it appears that only $lhour_t$ is important (significant at the 5% level). However, while the negative sign of variable $union_t$ (in both regimes one and two) implies that low labor standards are a boost to export performance, the negative sign of variable $lhour_t$ implies the contrary. In the third and fourth regimes (1972-1981 and 1982-1998 respectively), inj_t also appears to explain the behavior of the dependent variable, and the positive sign on this variable implies that lower labor standards provide a comparative advantage in trade by leading to a higher export performance. The variable $union_t$ is also significant at the 5% level in the fourth regime and is again negative. The dependent variable lagged one period is also significant in all but one regime (1957-71), which implies that past export performance is important in explaining current export performance.

Evidence is similar when we consider results with only one break point. In this case, variable inj_t explains the behavior of the dependent variable for the two regimes (1950-1971 and 1972-1998). The variable $union_t$ appears to be useful in the first regime. However, and once again, the signs of the variables inj_t and $union_t$ imply that labor standards can both increase or reduce export performance respectively. In fact, the sign of variable $union_t$ has changed and past export performance is significant (at the 1% level) in the second regime but not in the first one.

Last column of the Table 2 presents evidence of the presence of breaks for equation (8). The $UDmax$ and $WDmax$ tests show a clear evidence (at 1.0% level of significance) of the existence of breaks in the model. A similar conclusion (with similar level of significance) is found when the statistic $sup\ F_T$ is used. The sequential application of the $sup\ F_T$ shows evidence in favor of three breaks and this result is confirmed by the BIC.

Table 4 presents the results of the estimations with 3 breaks. Firstly, the variable rin_t is important for all regimes except for the last one (1982-1998). However, even though the sign of rin_t was expected to be negative, this is so only for the period 1957-1971. The variable $lhour_t$ is significant for the first three regimes (1950-1956, 1957-1971, 1972-1981). The variable $union_t$ appears to be important for all regimes. Finally, the variable inj_t is significant for the first, third and fourth regimes. The signs of variables inj_t and $union_t$ indicate that low labor standards are a source of comparative advantage since they lead to an improvement in export performance; variable $lhour_t$, however, shows the opposite. We do have evidence of a change in the sign of the labor standards variables from one regime to another (for instance variable inj_t is positive in the first regime and negative in the second one), but we do not obtain a change in sign from one significant variable to another significant one. Past export performance (the dependent variable lagged one period) is also significant for all regimes in explaining current export performance.

5 Conclusions

This paper has analyzed the effects of labor standards on the export performance of the US using annual data for the period 1950 to 1998. An important issue is the choice of indicators for labor standards, and in this paper we are able to consider three indicators, namely actual weekly hours worked, occupational injuries, and the rate of unionization. Unlike the general approach in the literature, which is based on cross-section analysis, this paper considers a time series approach based on the structural change literature. Such an approach is in our view more useful since it yields some benefits when compared with cross-section estimates such as those of Rodrik (1996) and Mah (1997). First, by identifying structural breaks in the data, we can look at the effects of labor standards on export performance under different regimes and also examine the evolution of the direction of these effects. Second, and because of the nature of the data, past export performance is included as an additional explanatory variable to control for current export performance, and it turns out to be a significant factor in many cases.

As far as the actual results are concerned, the evidence assembled in this paper is mixed. For example, the unionization rate and the rate of occupational injuries ($union_t$ and inj_t), are significant determinants of export performance and they support the general view that low labor standards can improve export performance. The possibility that lower labor standards may instead have an adverse effect on productivity and efficiency, and hence raise costs, is also confirmed by some of the results in the paper (for instance with regards to variable $lhour_t$). What is more important is the fact that we obtain evidence that labor standards do matter for export performance, and even in the case of a developed country such as the US. As was mentioned in the beginning of the paper, the emphasis on a North-South framework to analyze the issue of trade and labor standards has overlooked the importance of this issue among developed countries. This paper has thus filled some of that gap.

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Table 1. Unit Root Tests*

Variables	Statistic	$\widehat{\alpha}$	$t_{\widehat{\alpha}}$	k^*	$\widehat{T}_B$
$log(exp/gdp)_t$	ADF^{GLS}	0.65	-3.77^a	2	
	ADF	0.62	-4.04^a	6	
	$P97$	0.53	-6.38^a	8	1970
rin_t	ADF^{GLS}	0.80	-2.90^d	0	
	ADF	0.61	-2.90	3	
	$P97$	-0.05	-6.31^a	8	1979
$lhour_t$	ADF^{GLS}	0.66	-3.06^d	0	
	ADF	0.65	-3.09	0	
	$P97$	0.42	-4.35^c	8	1981
$union_t$	ADF^{GLS}	0.80	-2.31	4	
	ADF	0.69	-2.34	4	
	$P97$	-0.54	-5.95^a	8	1982
inj_t	ADF^{GLS}	0.90	-1.90^d	1	
	ADF	0.83	-2.63	1	
	$P97 - Slope$	0.62	-3.31	8	1970
	$P97 - Intercept, Slope$	0.70	-4.24	8	1976
	$P97 - Intercept$	0.63	-4.10	8	1978

* a, b, c, d denotes significance levels at the 1%, 2.5%, 5% and 10%, respectively.

Table 2. Statistics from the Equations (7) and (8)*

Statistics	Equation (7)**	Equation (8)***
$UD\max$	44.78^a	73.64^a
$WD\max(1\%)$	65.22^a	133.97^a
$\sup F_T(1)$	43.31^a	35.79^a
$\sup F_T(2)$	35.19^a	37.49^a
$\sup F_T(3)$	44.78^a	69.82^a
$\sup F_T(4)$	41.43^a	66.53^a
$\sup F_T(5)$	35.34^a	73.64^a
$\sup F_T(2\|1)$	23.73^a	66.04^a
$\sup F_T(3\|2)$	23.73^b	66.04^a
$\sup F_T(4\|3)$	16.08	19.85
$\sup F_T(5\|4)$	6.42	19.55
BIC	3	3
LWZ	0	0
$Sequential$	$3^{b,c,d}$, 1^a	$3^{a,b,c,d}$

* a, b, c, d denotes significance levels at the 1%, 2.5%, 5% and 10%, respectively.

** The specifications are $M = 5$, $q = 5$ and $h = 7$.

*** The specifications are $M = 5$, $q = 6$ and $h = 7$.

Table 3. Estimation with Endogenous Breaks; Equation (7)[*,**]

Parameters	Equation (7)			
	Estimation with 3 Breaks		Estimation with 1 Break	
	Coefficients	t-stats	Coefficients	t-stats
$\gamma_{1,1}$	8.406^{c}	2.92	-1.253	-0.86
$\gamma_{2,1}$	0.566	1.12	1.100^{a}	2.71
$\gamma_{3,1}$	-2.495^{c}	-3.30	-0.116	-0.29
$\gamma_{4,1}$	-0.033^{c}	-3.46	0.007^{a}	3.31
$\gamma_{5,1}$	-0.352	-2.16	0.057	0.35
$\gamma_{1,2}$	1.138	0.87	-2.567	-1.49
$\gamma_{2,2}$	0.018	0.03	1.974^{a}	4.09
$\gamma_{3,2}$	-0.658^{c}	-2.05	0.588	1.27
$\gamma_{4,2}$	-0.005	-1.26	-0.000	-0.22
$\gamma_{5,2}$	-0.039	-0.25	0.855^{a}	11.18
$\gamma_{1,3}$	9.865^{c}	3.15		
$\gamma_{2,3}$	10.571^{a}	5.32		
$\gamma_{3,3}$	-2.704^{c}	-3.18		
$\gamma_{4,3}$	-0.047^{a}	-4.69		
$\gamma_{5,3}$	1.095^{a}	6.52		
$\gamma_{1,4}$	-2.366	-1.07		
$\gamma_{2,4}$	1.112^{a}	3.09		
$\gamma_{3,4}$	0.553	0.96		
$\gamma_{4,4}$	-0.012^{c}	-2.06		
$\gamma_{5,4}$	0.671^{a}	8.04		
	Estimated Break Point	95.0% Confidence Interval		
$\widehat{T}_{B_1}$	1956	[1954-1958]	1971	[1969-1973]
$\widehat{T}_{B_2}$	1971	[1969-1973]		
$\widehat{T}_{B_3}$	1981	[1979-1982]		
$\overline{R}^2$	0.979		0.956	
F	108.54^{a}		103.48^{a}	

[*] In all cases, the specifications are $M = 5$, $q = 5$ and $h = 7$.

[**] a, b, c, denotes significance levels at the 1%, 5% and 10%, respectively using freedon degree-adjusted critical values at two-tails of the t-student.

Table 4. Estimation with Endogenous 3 breaks; Equation (8)[*,**]

Parameters	Estimated Coefficients	t-statistics
$\gamma_{1,1}$	32.588^{b}	7.19
$\gamma_{2,1}$	0.032^{b}	5.55
$\gamma_{3,1}$	5.835^{b}	5.99
$\gamma_{4,1}$	-9.215^{b}	-7.35
$\gamma_{5,1}$	-0.091^{b}	-8.09
$\gamma_{6,1}$	-1.747^{b}	-6.70
$\gamma_{1,2}$	0.202	0.18
$\gamma_{2,2}$	-0.017^{b}	-2.86
$\gamma_{3,2}$	-0.622	-1.25
$\gamma_{4,2}$	-0.517^{c}	-1.97
$\gamma_{5,2}$	-0.008^{b}	-2.20
$\gamma_{6,2}$	-0.479^{b}	-2.39
$\gamma_{1,3}$	12.130^{a}	4.11
$\gamma_{2,3}$	0.005	1.88
$\gamma_{3,3}$	13.188^{a}	5.99
$\gamma_{4,3}$	-3.330^{a}	-4.15
$\gamma_{5,3}$	-0.050^{a}	-5.67
$\gamma_{6,3}$	1.308^{a}	7.14
$\gamma_{1,4}$	-2.613	-1.22
$\gamma_{2,4}$	0.003	1.07
$\gamma_{3,4}$	1.412^{a}	3.16
$\gamma_{4,4}$	0.607	1.09
$\gamma_{5,4}$	-0.012^{b}	-2.11
$\gamma_{6,4}$	0.681^{a}	8.39
	Estimated Break Point	95.0% Confidence Interval
$\widehat{T}_{B_1}$	1956	[1954-1957]
$\widehat{T}_{B_2}$	1971	[1969-1972]
$\widehat{T}_{B_3}$	1981	[1979-1982]
$\overline{R}^2$	0.983	
F	111.16^{a}	

[*] In all cases, the specifications are $M = 5$, $q = 6$ and $h = 7$.

[**] a, b, c denotes significance levels at the 1%, 5% and 10%, respectively using freedon degree-adjusted critical values at two-tails of the t-student.