HOW DO ECONOMIC REFORMS AFFECT THE STRUCTURE OF WAGES? THE CASE OF BRAZILIAN MANUFACTURING, 1984–1996

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Abstract

One of the established findings in the literature on inter-industry wage differentials is the long-term stability of the wage structure. In this paper, we examine how market-oriented and other economic reforms undertaken by an industrialising country affect the dispersion and structure of wages. Using a large, individual-level dataset, we find that the labour market is highly responsive to the economic reforms undertaken in Brazil in the early 1990s. Wage dispersion falls dramatically just after the implementation of economic reforms and we find evidence that the wage structure is under transition after the changes. Human capital variables gain importance in the explanation of wage differentials, while industry affiliation and institutional characteristics become less important. This finding is consistent with the labour market being flexible enough to adapt to the new economic conditions, and becoming more competitive as a result of the economic reforms.

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1. Introduction

Throughout the last 20 years, industrialising countries have experienced many economic changes. Following the international economic instability of the late 1970s and early 1980s, they launched structural adjustment programmes aimed at solving external account imbalances and controlling high inflation rates. From the mid-1980s, many of these countries undertook unprecedented economic reforms involving trade liberalisation, privatisation of state companies, and the deregulation of labour, financial and goods markets. These reforms have been reshaping many developing economies very rapidly and are giving rise to profound economic transformations.

With respect to the labour market, the empirical literature has revealed substantial changes in wages and employment in developing countries (Gonda and Gomez, 1994; Pessino, 1994; Collado *et al.* 1994; Paus and Robinson, 1997; Revenga, 1997; Milner and Wright, 1998). The pattern of wage dispersion has also attracted much attention, and the available evidence has identified the tradable *vs.* non-tradable sectors issue as an important source of wage variance (Revenga, 1997; Curie and Harrison, 1997; Milner and Wright, 1998).[1] This kind of analysis tends to rely on testing the validity of the Hecksher-Ohlin model under the institutional and economic constraints of the industrialising countries. While this model predicts increasing wage dispersion in *developed* economies, wage dispersion should shrink in *developing* countries since they have an excess supply of unskilled workers and,

[1] For a theoretical discussion on this topic, see Edwards (1988).

thus, tend to specialise in the production of labour-intensive goods.[2] Within this theoretical framework, trade liberalisation in a developing country can affect wage dispersion in two possible ways as a consequence of the shifts in the structure of labour demand: (i) wage dispersion may decrease, suggesting that the labour market is competitive; or (ii) wage dispersion may not change or may even increase, implying that the labour market is segmented.

Few attempts have been made to investigate whether and how the recent economic reforms in developing countries have influenced the dispersion and structure of wages from the perspective of industry affiliation. On the one hand, in a context of price and wage flexibility and free factor mobility, the expected outcome of macroeconomic adjustment and economic reforms is a change in relative prices, which would shift resources between sectors, thus affecting the dispersion and structure of wages. On the other hand, the literature has long established the pervasiveness and stability of the structure of wages over very long periods for countries with different institutional and structural characteristics. Industry affiliation appears to be the fundamental source of the explanation of the phenomenon (Slichter, 1950; Krueger and Summers, 1987, 1988; Gittleman and Wolff, 1993; Katz and Summers, 1989, *inter alia*). If industry affiliation is a key factor in determining labour market segmentation, then structural adjustments and market-oriented economic reforms may have little impact on wage dispersion and structure of wages. In the presence of elements of labour market segmentation such as efficiency wages, the level of average wages and the pattern of the inter-industry wage structure may only be moderately affected by economic reforms.[3] The effects of industry

[2] For a discussion of the controversy on the effects of globalisation on wage dispersion in developed countries, see Wood (1994), Sachs and Shatz (1994), Desjonqueres *et al*. (1997), Machin (1996), and Berman *et al.* (1994).

[3] Gaston and Trefler (1994) investigate the effects of trade and protection on individual wages and inter-industry wage structure in the US in 1983 – a period of rapid change in American international trade – and conclude that trade does not affect the structure of wages. Freeman

affiliation associated with efficiency wages are that wages tend to be rigid. In this case, only microeconomic measures that influence the efficiency wage frontier will affect the wage structure. In these circumstances, neither restrictive structural adjustment policies, nor openness to international trade will be enough to generate greater wage flexibility. As a consequence, the labour market may retain a stable inter-industry wage structure and similar wage dispersion, but with higher levels of unemployment, at least in the short-term (Riveros and Boulton, 1994).

Understanding the responses of the inter-industry wage structure and the dispersion of wages in industrialising countries during periods of radical economic reforms helps to shed light on the links between labour market adjustment during episodes of stabilisation programmes and increasing global competition. As such, this may be valuable for the design of economic policies, especially those concerned with income inequality. The aim of this paper is to investigate the pattern of inter-industry wage differentials in Brazil before and after the economic reforms implemented at the beginning of the 1990s, and to track the impact of industry affiliation on wage dispersion and structure of wages. The case of Brazil is of interest not only because of the neo-liberal nature of reforms adopted from 1990 onwards, but also because of the pace of the changes, which can provide useful lessons to many other industrialising countries which have embarked on similar reforms.

The economic reforms undertook in Brazil in the 1990s – trade liberalisation, deregulation and privatisation of state-owned enterprises – have affected the whole of the economy, but especially the manufacturing sector which is more exposed to international

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and Abowd (1991) find that changes in trade and domestic demand in the US over the period 1958-84 affect the structure of manufacturing wages. They estimate that a 10 percent decrease in industry revenues from increased import penetration reduces an industry's relative wage for production workers by only 0.5 percent, thus suggesting a relatively low response from wages to international trade.

competition.[4] Hay (1998) finds a sharp fall in profits, and a marked increase in the efficiency of large manufacturing firms as a result of trade liberalisation. Salm *et al*. (1997) find a remarkable increase in productivity in the manufacturing sector from 1990, and highlight the rapid adoption of modern managerial methods. Profound changes in the labour market have also been identified. Barros *et al*. (1996) and Moreira and Najberg (1997) find a decline in manufacturing employment as a consequence of the increased openness to international trade. Using Granger causality tests, Arbache (1996) identifies productivity as the key variable driving the labour market from 1990 onwards. He also reports structural changes in unemployment and manufacturing real wages at the beginning of the decade. Gonzaga (1996) presents evidence of increasing elasticity of the manufacturing labour demand curve with respect to labour costs and output in the 1990-96 period.

This paper is organised as follows: the next section presents an overview of the Brazilian economy for the last two decades. Section 3 discusses the methodology used to estimate wage differentials and describes the database utilised. Section 4 reports the main econometric results on wage dispersion, and section 5 assesses the impact of industry affiliation on wage structure. Section 6 concludes.

2. The Brazilian Economy: 1980-1996

If one phrase is able to summarise the Brazilian economy since the beginning of the last decade, this phrase is ‘macroeconomic instability’. Since the second international oil shock in 1979, policy-makers have moved from an import-substitution industrialisation policy

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[4] The stabilisation of the inflation rate post-1994 is another remarkable economic change of the period. However, a study of its impact on the structure of manufacturing wages is beyond the scope of this paper.

to a pragmatic economic policy pattern seeking to solve both the balance of payments problem and to control the accelerating rate of inflation. In practice, the government has alternated measures based on the restriction of demand aimed at containing inflation and reducing external deficits, with expansionist policies seeking to boost the rate of growth. The lack of clear macroeconomic targets is seen as a source of uncertainty and instability, and the accelerating rate of inflation, especially in the end 1980s and early 1990s, can be seen to be one outcome of this process (Amadeo, 1994a).

Table 1 presents some key macroeconomic indicators for the Brazilian economy for 1980 to 1996. GDP growth was highly volatile, while the inflation rate increased almost steadily, reaching dramatic levels in 1990 and 1994. The unemployment rate has increased steadily since 1989. Between 1986 and 1994, six stabilisation plans based on freezing or controlling prices and wages were put forward in attempting to halt the accelerating rate of inflation, but only the 1994 plan – the *Plano Real* – succeeded in breaking the indexation pattern and keeping inflation at comparatively low levels.

Prior to 1990, the Brazilian economy was highly protected (Pinheiro and Almeida, 1994), regulated, and public sector companies dominated a variety of infra-structure activities, among other industries. From 1990, the efforts to contain inflation were combined with vigorous trade liberalisation, economic deregulation and the privatisation of public companies. Up to the middle of 1993, a complex and bureaucratic set of non-tariff barriers was removed, and a new tariff structure was imposed, which substantially reduced the degree of protectionism. The general average effective tariff dropped from 45.5 percent in 1990, to around 20 percent in 1993 (Kume, 1996). As a result of the new economic policies, from 1990 to 1996, imports increased by 257 percent, while exports increased by 151 percent. Most of the effects of trade liberalisation on the trade balance were, however, postponed to 1994

onwards due to the 1990-92 economic stagnation and revaluation of exchange rate in 1991-92 (see columns 6 and 7). In 1994, the combination of new liberalising measures seeking to discipline domestic prices and overvaluation of the exchange rate affected the trade accounts in such a way that the trade balance started to have increasing deficits after a long period of surplus (see column 4 in Table 1).

<<Table 1 here>>

Table 2 shows the effective protection in manufacturing, which takes into account tariff and non-tariff barriers by sector. The average effective protection suggests that there was no significant change in trade policy from 1980 to 1991, despite the lower tariff dispersion in this year (see the last two rows of Table 2). By 1993, however, the mean effective protection had fallen from 40 percent in 1980-91, to only 17 percent, and the standard deviation dropped from 30.5 percent in 1991 to under 10 percent by 1993. These statistics suggest both a change in the trade regime and significant tariff convergence in the manufacturing sector.

<<Table 2 here>>

Contrary to what occurred in many other developing countries in which trade reforms were gradually introduced and had a moderate impact (Adriamananjara and Nash, 1997), the change in trade policy in Brazil seems to have impacted the economy in a more radical manner. This is confirmed by the very rapid rise in the import penetration ratio in the manufacturing sector illustrated in Figure 1.[5] There is a clear shift in the import pattern from 1990 onwards. By 1996, the import penetration ratio had reached 11.5 - more than twice the figure for 1990. There are, however, large differences in import penetration by industry. Beverages, Food, Furniture, and Wood, for example (the typical non-tradable and labour-

[5] The data on import penetration, $M_i / (Y_i - X_i + M_i)$, where M is imports, Y is production, and X is exports of industry i, are from Haguenauer et al. (1998).

intensive sectors), experience a level of import penetration which reaches at most 6 percent of the apparent domestic demand in each industry. In contrast, industries such as Chemical and Electronic (the typical tradable and capital-intensive industries), experience very large changes in import penetration (as much as from 8 percent to 30 percent). As a consequence, trade liberalisation can be expected to have differential impacts on wages and employment within different manufacturing industries.

<<Figure 1 here>>

With regard to the deregulation of markets, the restrictive rules and laws which prevented internal and external competition in many sectors for a considerable period of time began to be removed in 1990. This stimulated lively competition, especially in manufacturing industries and in some service sectors. By 1995, most of the steel and chemical sectors had already been privatised, and many other sectors such as transportation, power generation and distribution, mineral companies and banks were undergoing privatisation.

In the labour market, important changes are identifiable in the manufacturing sector during the period under analysis. Total employment in manufacturing fell by 25 percent from 1990 to 1996, according to the Monthly Employment Survey, and about one million jobs were lost. The proportion of workers employed without the coverage of labour regulation increased steadily after the turn of the decade. In the 1980s, around 20 percent of the full-time workforce in manufacturing were not covered by formal labour contracts. By 1992, this figure had risen to 23 percent, and by 1996 it had reached 26.5 percent. Trade unions become stronger in the 1980s, and important rent-sharing effects arose in wage formation (Amadeo, 1994b; Carneiro, 1998). However, the union-non-union wage gap began to fall during the 1990s, suggesting that unions began to lose their importance in wage setting (Arbache and Carneiro, 1998).

Brazil also experienced important institutional changes in the 1980-96 period. In 1985, the 21-year dictatorship era finished and, as a consequence, several social and political demands started to flourish all over the country. In 1993, after a long corruption trial, President Collor was removed from office, which, once again, caused new social demands, affecting the spirits and expectations of economic agents.

The effects of the economic and institutional changes described above on the structure and dispersion of wages are not easily identified, and depend on the interactions among the economic variables, and on these interactions with the institutional changes. At most, only partial analysis can be carried out unambiguously. On the basis of the Hecksher-Ohlin model, for example, the increased openness to international trade and deregulation of markets in a developing country tend to reduce the wage dispersion and to impact the structure of wages in such a way that it reflects the abundance of capital, technology and quality of labour. The effect of privatisation of state companies on the labour market is towards a higher competitiveness, since it reduces an important source of segmentation in developing countries' labour markets (Rama, 1995; Teal, 1996). Presumably, the wage structure and dispersion will better reflect market forces. The impact of unionism on wage determination in Brazil, however, is to increase, rather than to decrease wage dispersion, as shown by Arbache (1998a) and Arbache and Carneiro (1998).

3. Measuring Wage Dispersion and the Structure of Wages

The strategy we adopt to estimate the degree of wage dispersion and to examine the structure of wages is through the estimation of inter-industry wage differentials. We follow the approach proposed by Haisken-DeNew and Schmidt (1997), which improves the standard

procedure popularised by Krueger and Summers (1988).[6] The wage equations are estimated in the following form:

$$\ln w_{ij} = \alpha + \beta X_i + \varphi Z_j + \varepsilon_{ij}, \tag{1}$$

where $\ln w_{ij}$ is the natural logarithm of the hourly real wage of worker i in industry j, X_i is a vector of personal characteristics, occupations and regions, Z_j is a vector of industry dummies which includes *all* industries, α is the intercept term, ε_{ij} is a random disturbance term reflecting unobserved characteristics and the inherent randomness of earnings statistics, and β and φ are the vectors of parameters to be estimated. Since in this model the cross-product matrix of the regressors is not of full rank, a linear restriction is imposed on the φs as follows,

$$\sum_j \varphi_j n_j = 0, \tag{2}$$

where n_j is the employment share in industry j. The reported coefficients can then be interpreted as the proportionate difference in wages between a worker in industry j and the average worker from the whole of manufacturing industry, after controlling for the other factors which determine wages. The formulation given by (1) and (2) provides both economically sensible coefficients and their correct standard errors in a single regression step.

The standard deviation of wage differentials is computed as

$$SD(\varphi) = \sqrt{n'(\overline{H}(\varphi_j))\varphi_j - n'\overline{D}(V(\varphi_j))}. \tag{3}$$

$SD(\varphi)$ gives the weighted and adjusted standard deviation of coefficients, $\overline{H}(.)$ transforms a column vector into a diagonal matrix whose diagonal is given by the column vector itself, $\overline{D}$

[6] For a detailed discussion on inter-industry wage differentials methodologies, see Arbache (1998b).

denotes the column vector formed by the diagonal elements of a matrix, and *V* is the variance-covariance matrix (see Haisken-DeNew and Schmidt, 1997, for further details).

The cross-sectional micro data we use to estimate inter-industry wage dispersion are from the PNAD, conducted by the Brazilian Institute of Geography and Statistics (IBGE). Each survey contains data on about a hundred thousand randomly selected households throughout the country, fully describing a substantial number of individual and household characteristics for labour market analyses. The sample data we analyse were filtered in the following way: active individuals, non-employers, between the ages of 18 and 65, working 21 or more hours a week in the main job,[7] and affiliated to any of the 22 industries (gêneros da indústria - IBGE) that comprise the manufacturing sector. This classification corresponds to the 2-digit SIC.

The data set consists of hourly wages, industry affiliation, education (years), experience (years), 5 region dummies, gender dummy, race dummy, over time worked dummy, metropolitan area dummy, head of family status dummy, 7 occupation dummies, union membership status dummy, and work-card dummy. The reported wages are monthly wages in the main occupation that individuals held during the period of interview, and do not include other income or non-wage compensation. Thus, we avoid any bias from self-employment, second jobs, or any other income source in our estimation of wage differentials. In order to obtain hourly wages, we divided the monthly wage by the product of usual weekly hours in the main job multiplied by 4. Wages were deflated to the CPI of December 1995 and transformed to natural logarithms. Summary statistics are presented in Table A1.

[7] Working journeys more than 20 hours per week are defined as full-time according to the Brazilian legislation.

4. Wage Dispersion: 1984-1996

Results of raw cross-section regressions are reported in Table 3. The coefficients show the proportionate difference in wages between an employee in a given industry and the weighted average worker in manufacturing. The figures indicate that, for instance, in 1996 a worker in the Apparel industry earned about 30.2 percent less than the average manufacturing wage, while a worker in the Pharmaceutical industry earned about 95 percent above the mean wage.[8] The weighted and adjusted wage dispersion is about 34 percent in 1984, 36.6 percent in 1988, and then falls to 30 percent in 1992, and to 28.4 percent in 1996. The wage premium dispersion is sizeable for all years, suggesting a high inter-industry wage variability. There is, perhaps, evidence of a slight convergence in the dispersion of wages in the 1990s.

In order to examine the robustness of these results, we compare them to alternative estimates of wage differentials based on aggregate data. We calculate the Spearman rank correlation coefficient between the figures in Table 3, and the proportionate difference between the wage of a worker in an industry, and the weighted average wage in all industries, estimated with aggregate data from Annual Report of Social Information (RAIS), for two of the four years under analysis (see Table A3). The coefficients for 1988 and 1992 are, respectively, 0.8645 (p=0.000) and 0.8688 (p=0.000),[9] suggesting that the estimated wage structures based on these two different data sets are quite similar. This is an indication that differences in average wages seem to be, at least in part, a result of the impact of industry characteristics on the structure of labour demand.

<<Table 3 here>>

[8] In order to derive percentage values from the dummy variable coefficiencts, the following formula is adopted: $(e^{\beta_i} - 1) * 100$ (see Halvorsen and Palmquist, 1980).

[9] The sample sizes of RAIS are 6,026,359 for 1988, and 4,828,707 for 1992. The RAIS CD-ROM provides average wages by industry, region, race etc., but not individual wages. The RAIS for 1984 is not published, and the RAIS for 1996 is not yet available.

Table 4 reports the wage differentials when control variables are included in the wage equations[10]. The results show that most of the estimated industry coefficients decline in magnitude as compared to the figures in Table 3, suggesting that the control variables have strong explanatory power in wage determination. As a consequence, the weighted and adjusted inter-industry wage dispersion falls in all years. In the 1980s, wage dispersion remained around 14 percent. However, in 1992, it dropped to 8.4 percent, and fell to only 4.3 percent in 1996. The huge effect of the control variables implies high and increasing returns to productive attainments across individuals. More significant than the small magnitude of the industry wage dispersion is its reduction over time which implies a convergence of the inter-industry wage differentials. Accordingly, the control variables explain 59 percent of the inter-industry wage dispersion in 1984 (calculated as $100\times(0.3393-0.1401)/0.3393$); 60 percent in 1988; 72 percent in 1992; and 85 percent in 1996. Over time, an increasing proportion of the inter-industry wage dispersion is due to differences in individual characteristics, rather than being due to (unexplained) industry characteristics. The convergence path identified is remarkable, and requires the labour market to be highly flexible and responsive to the economic changes. The evidence suggests that the labour market is becoming more competitive, which seems to be an unambiguous response to the economic reforms carried out in the 1990s.[11]

<<Table 4 here>>

The rank correlation coefficients between the wage structures estimated with and without control variables are 0.7483, 0.7489, 0.8588, and 0.7504 for, respectively, 1984,

[10] Coefficients on the control variables are reported in Table A2.

[11] The estimated wage differentials are robust to different specifications of the wage equation, and the presence or absence of particular variables in the model do not significantly change the ranking of inter-industry wage differentials.

1988, 1992, and 1996.[12] Thus the addition of control variables does little to affect the ranking of inter-industry wages. This implies that the differences in the average wages between industries are partially explained by differences in labour quality.

In order to assess the impact of industry affiliation on wage dispersion, and to investigate whether it has decreased its effect on wages (as suggested by the convergence of wage differentials), we estimate the share of wage variance due to industry affiliation using analysis of covariance techniques. Given an earnings function such as equation (1), this technique can partition the variance of wages, σ_w^2, into that due to industry affiliation and to covariates, as follows:

$$\sigma_w^2 = \sigma_\beta^2 + \sigma_\varphi^2 + \sigma_\varepsilon^2, \tag{4}$$

where σ_β^2 is the covariates wage variance, σ_φ^2 is the inter-industry wage variance, and σ_ε^2 is the variance due to the error term. The share of each variance component reflects its importance to the overall variance of wages. Here we are particularly interested in obtaining the proportion of wage variation due to industry affiliation. The total proportion of wage variation explained by industry affiliation and the covariates is given by the coefficient of determination of the earnings function. We estimate the lower and upper bound of wage dispersion due to industry affiliation as follow.[13] First, we take the coefficients of determination from the regression of log wages on industry dummies and covariates together, $R_{\beta,\varphi}^2$, on industry dummies alone, R_φ^2, and on covariates alone, R_β^2. Second, to estimate the lower bound of wage dispersion due to industry affiliation, we take the difference between $R_{\beta,\varphi}^2$ and R_β^2. This difference provides a conservative measure of the industry effects - the

[12] All coefficients are significant at the 1 percent level.
[13] This methodology follows Dickens and Katz (1987) and Groshen (1991).

correlation between the covariates and industry dummies (for example, due to occupational groups) implies that the lower bound of the industry effect is larger than this since the joint impact of industry affiliation and the covariates is attributed to the covariates only. The upper bound is given by R^2_{φ}.

The covariates we include in the regressions are education, experience, experience squared, occupation, region, work-card, gender, and head of family status.[14] The decomposition analysis is reported in Table 5. Our results show that, in the 1980s, the contribution of the vector of industry dummies remains broadly constant, and accounts for between 4.5 and 16 percent of the wage variation. However, in the 1990s, this falls significantly, to between 3.6 and 15.2 percent in 1992 and falls even further to between 2.5 and 12.8 percent in 1996. Not only are the bounds for the 1990s lower, but the interval range is also smaller (see Figure 2). An estimate of the variance of wages attributable to industry affiliation is 0.1206, 0.1384, 0.0972, and 0.0857 for, respectively, 1984, 1988, 1992, and 1996 (estimated as the proportion of sum of squares due to industry alone multiplied by the variance of earnings in the sample). The F-statistics for the hypothesis that industry effects are jointly zero are always large, thus rejecting the hypothesis that industry affiliation does not contribute to wage dispersion. The smaller industry effects in the 1990s imply that wage determination is less affected by industry segmentation, therefore supporting the previous evidence that the labour market is becoming more competitive.

<<Table 5 and Figure 2 here>>

[14] We do not include union membership in the analysis because this variable is not available for 1984. Race is also excluded from the earnings function because of a large number of missing values for 1984. The lack of these two variables may result in the overestimation of the importance of industry affiliation.

An alternative examination of the effects of industry affiliation on wages is provided by the computation of the change in the standard error of the regression once industry dummies are removed from the wage equation. The greater the fall in the standard error, the higher the marginal contribution of industry affiliation. We also compute the marginal contribution of other control variables in order to provide a comparative analysis, and to assess their importance to the explanation of wage variation.

Table 6 reports the results from this exercise and supplies a ranking of variables according to their contribution to the variance of wages. The results are striking. Industry dummies rank third in the 1980s, but fall to fourth in 1992, and to seventh in 1996, always ranking behind human capital and occupation variables. For all years, education is by far the most important variable affecting wage variance, an outcome consistent with previous results for Brazil (Lan and Levison, 1992; Bonelli and Ramos, 1995). The relatively higher importance of the human capital variables in the 1990s is evidence that skills and individual attributes are being increasingly rewarded, while the high rank of occupation suggests that skills embodied in occupation, or specific skills, are also highly rewarded. These results are consistent with Cragg and Epelbaum's (1996) findings for Mexico during its increasing openness to international trade in the 1980s. The decrease in the marginal contribution of industry dummies indicates the decreasing importance of industry specific-skills plus industry specific-characteristics in the determination of wages.

<<Table 6 here>>

Our results on inter-industry wage differentials differ from previous studies for Brazil (Pinheiro and Ramos, 1994; Gatica *et al.*, 1995) in at least two important ways. Firstly, depending on the year of analysis, our wage equations can explain at least 60% and up to 81% of the inter-industry wage differentials. This contrasts with the two previous studies which

could only account for about 50% of the raw wage dispersion. Secondly, our analysis covers a longer period, which encompasses the main institutional and political changes Brazil undertook during the last two decades, as well as the rise and fall of the inflation era, the several stabilisation plans, the increasing openness to international trade, privatisation and deregulation. Perhaps of greater importance than the longer time span is the fact that the period we analyse is particularly dynamic, and thus we can provide a more comprehensive assessment of the responses of the labour market to the changing economic, institutional, political and social environment. Our results are not directly comparable to the previous estimates due to differences in sampling and the nature of the data, the period of analysis, and model specification. Moreover, the two studies cited above both use the Krueger and Summers (1988) methodology, while we have adopted the Haisken-DeNew and Schmidt (1997) procedure. Thus, any comparisons should be made with caution.

5. Wage Structure: 1984-1996

Despite the increasing importance of human capital, occupation and other control variables, the regression results suggest that manufacturing wages are still influenced by industry characteristics. *Ceteris paribus*, workers affiliated to Chemical, Pharmaceutical, Transportation, Electronic and Mechanic (the traditionally capital-intensive industries), are the most highly paid within manufacturing, while workers affiliated to Wood, Leather, Apparel, Food and Beverages (the typical labour-intensive industries), are the paid the least. While this evidence is in line with the literature on the effects of industry affiliation on the wage structure, it is also clear that industries have distinct evaluations of workers' productive attributes, and that the marginal productivity of workers with similar attainments may vary among industries.

The inter-industry wage structure would be expected to have remained stable in the 1990s - a period of very rapid economic changes - if:

(i) elements of wage rigidity such as efficiency wages and rent-sharing dominated the wage determination process - this seems unlikely, since the impact of industry affiliation on wages has been falling over time, thus reducing the effects of industry characteristics on wages;

(ii) the structure of trade barriers remained stable over time, or the effect of the reduction of trade barriers was approximately equal across industries (that is, it affected the level but not the structure of tariffs and other forms of protection);[15]

(iii) the market-oriented reforms 'uniformly' affected the structure of demand across manufacturing industries - which seems improbable;[16]

(iv) other factors such as institutional changes, deflation, unionism, tax reforms, tastes, technology, etc. affected all manufacturing industries in a similar manner - which also seems fairly unlikely. With regard to unionism, the diminishing importance of unions in wage determination during the 1990s (Arbache and Carneiro, 1998) suggests that rent-sharing and wage bargaining are unlikely to have a sizeable effect on the determination of the wage structure in the period.

[15] This does not seem to have happened. The Spearman rank correlation coefficients for effective trade protection in manufacturing industries fall steadily as time goes by (see Table A4), suggesting the structure of protection is changing within manufacturing.

[16] Tables A5 and A6 report the Spearman rank correlation coefficients of import penetration and the export ratio. The results suggest a change in the structure of imports and exports in the manufacturing sector.

However, one could expect changes in the inter-industry wage structure in the 1990s as compared to the 1980s to have reflected the reallocation of factors of production and technological changes, and to have matched the changes in the structure of imports, exports and trade protection.[17]

In order to assess the evolution of the wage structure, we compute the Spearman rank correlation coefficients for industrial wage differentials with and without control variables included in the wage equations. Table 7 reports the results for non-controlled wage differential vectors, and shows a high degree of stability of the wage structure, even after the economic changes of the 1990s. This is in accordance with the literature which stresses the high stability of the rank of inter-industry wage premia. Based on these results, one could conclude that the inter-industry wage structure in Brazilian manufacturing is not explained by short-run factors.

<<Table 7 here>>

Table 8, however, reveals a rather different story. Once control variables are included in the wage equations, the wage structure can be seen to significantly change after the turn of the decade (although the computed correlation coefficients remain significant at the 1 percent level).[18] Column 1 shows that, over time, the wage structure experiences a quite dramatic change. The observed changes reflect shifts in the structure of labour quality, changes in the composition of labour demand, and a more efficient use and allocation of labour. Some industries seem to be more enthusiastic than others to attract higher quality labour, thus changing the distribution of labour skills within the manufacturing sector. These results are supported by the increasing importance of productive traits in wage determination as shown

[17] Gaston and Trefler (1994) argue that the temporal stability of the inter-industry wage differentials in the US manufacturing is associated with the temporal stability of the structure of imports, exports and trade barriers.

[18] The difference between the figures shown in Tables 7 and 8 emphasises the effect of the control variables on wages in the 1990s.

above, and might, at least partially, be one consequence of the market-oriented economic reforms.

The key implication of these results is that the Brazilian labour market seems to be flexible enough to deal effectively with the reallocation of labour necessitated by the economic reforms and structural changes during the period, and also to allow for changes in relative wages. Thus, in the 1990s, the wage structure in Brazilian manufacturing industry does seem to have been affected by short-run changes. This is a strong result in the light of the stylised fact that the inter-industry wage structure is stable and is little affected by short-term factors and business fluctuations.

<<Table 8 here>>

6. Summary and Conclusions

This paper has investigated the impact of economic changes in the dispersion and structure of wages in an industrialising country experiencing rapid market-oriented and other economic reforms. The stylised fact in the literature - that the structure of wages is stable over time and is affected little by business cycles and other economic changes - does not seem to apply to the manufacturing industry in Brazil. The inter-industry wage differentials in the 1990s – the period in which the economic reforms were introduced – show a significant change when compared to the 1980s. The dispersion of the wage premia, which had remained stable in the 1980s, fell dramatically in the 1990s, to a level which is only around one third of that in 1988. While human capital and occupation variables proved to be the main factors determining wages, industry affiliation – the traditional factor accounting for wage segmentation – has decreased in importance after the turn of the decade. In the 1980s, the marginal contribution of industry affiliation to the wage was ranked in third place. By 1996, it

had dropped to seventh place, ranking behind all human capital variables, occupation, as well as other variables such as gender and over-time worked. With regard to the structure of wages, the evidence reveals that there was a clear change in the 1990s, in contrast with the stylised fact that the wage structure of comparable workers is stable over time. The shifts in the structure of labour demand seem to be associated with the economic changes of the period.

Contrary to the established evidence in the literature on inter-industry wage differentials, our results suggest that wages in Brazilian manufacturing are sensitive to short-run economic changes. The responsiveness of the wage structure and the wage premia suggest that the labour market is flexible enough to adjust quite rapidly to the new economic environment. To the extent that human capital variables are the main factors determining wages, and that the wage dispersion converges over time, the labour market seems to have become more competitive. Since the wage premia fall over time, productive gains derived from the better efficiency and employment of better labour quality may have been transferred to prices.

Of course, the convergence of wages and the higher competitiveness in the labour market might not be exclusively due to the market-orientated reforms of the period, but also to other factors. In this paper, we have presented some evidence that the reforms appear to be associated with the observed changes. However, other factors such as the stabilisation of inflation, foreign direct investment, political stability, decreasing union strength, among others, might also have contributed to the results.

It is worth noting that the *pace* of economic change observed in Brazil in the 1990s may be an important contributory factor in the explanation of the changes we have identified in the labour market. Thus, the Brazilian experience should be compared with caution to the results on the stability of the wage structure previously found for developed countries,

although our results may be informative for a number of developing countries which are embarking on similar reforms, and may yield insights into the links between labour market adjustment, increasing global competition, and the effects of market-oriented reforms in these countries. However, there are still many gaps in our understanding of labour market dynamics in developing countries. In particular, research is needed into the impact of privatisation of state enterprises, deregulation, trade liberalisation, unionism, and inflation stabilisation on wage determination.

Two important policy implications can be drawn from our results. First, to the extent that elements of labour market rigidities lose importance over time, macroeconomic policies would appear to have improved the effectiveness of the operation of the labour market. Second, the increasing relevance of human capital and occupational variables suggests that policies seeking to tackle income inequality and the social exclusion problem might be directed towards broad educational and vocational training programmes that improve the quality and, more importantly, the access to skills acquisition.

Table 1

Macroeconomic Indicators for the Brazilian Economy: 1980-1996

Year	GDP % change	Industrial output % change	Unemployment rate[1]	Export - import balance[2]	GDP deflator % change[3]	Real exchange rate[3]	Effective exchange rate[4]
1980	9.2	9.2	7.2	-2.8	100.2	-	-
1981	-4.3	-8.8	7.9	1.2	109.9	100	-
1982	0.8	0.0	5.4	0.8	95.4	98.6	-
1983	-2.9	-5.9	6.7	6.5	154.5	128.3	-
1984	5.4	6.3	7.1	13.9	220.6	127.9	-
1985	7.8	8.3	5.3	12.5	225.5	131.0	-
1986	7.5	11.7	3.6	8.3	143.5	114.3	-
1987	3.5	1.0	3.7	11.2	223.2	90.8	-
1988	-0.1	-2.6	3.9	19.2	684.6	91.9	118.5
1989	3.2	2.9	3.4	16.1	1320	75.5	91.1
1990	-4.3	-8.2	4.3	10.7	2740	61.0	78.4
1991	0.3	-1.8	4.8	10.6	414.7	69.4	89.0
1992	-0.8	-3.8	5.8	15.2	991.4	71.7	100.0
1993	4.2	6.9	5.4	13.0	2103	66.4	94.7
1994	6.0	7.0	5.1	10.4	2407	48.4	83.0
1995	4.2	2.0	4.6	-3.5	67.5	39.5	69.9
1996	3.0	2.5	5.4	-5.6	11.1	38.8	65.3

Notes:

1. IBGE open unemployment rate. 1980 and 1981 is the average unemployment rate for São Paulo and Rio de Janeiro.
2. In nominal US$ billions.
3. IGP-DI Geral FGV, real exchange rate 1980=100.
4. 1992=100.

Table 2

Effective trade protection in manufacturing sector

	1980	1991	Jul-1993
Apparel	46.7	55.9	23.8
Beverage	-1.1	108	24.8
Chemical	86.4	11.4	8.2
Electronics	129.3	44.8	25.5
Food	26.1	27.5	16.4
Furniture	52.7	42.7	26.4
Leather	13.9	13.3	8.4
Mechanic	93.3	34.7	23.1
Metallurgic	34.2	27.4	16.9
Mineral	-4.2	1.4	-0.4
Non-metallic	-19.6	17.3	8.8
Other	171.7	43.2	0.4
Paper	-18.5	13.1	9.2
Perfumes	91.6	64.8	26.1
Pharmaceuticals	116.3	19.1	13.1
Plastics	28.3	46.1	22.4
Publishing	31.9	10.8	8.4
Rubber	-21.4	44.6	15.7
Textiles	36.7	47.4	19.2
Tobacco	5.7	117.6	23.9
Transport	-6.5	72.6	39.7
Wood	17.7	10.8	10
Mean	41.42	39.75	16.82
Standard deviation	52.69	30.45	9.71

Note:
Table constructed based on Table 6 of Pinheiro and Almeida (1994).

Table 3

Inter-industry wage differentials – no controls included

	1984		1988		1992		1996	
Industry	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.
Apparel	-0.3875	(.0172)	-0.4485	(.0208)	-0.3024	(.0167)	-0.3602	(.0173)
Beverage	-0.0462	(.0387)	-0.0250	(.0481)	0.0168	(.0428)	0.0837	(.0467)
Chemical	0.6442	(.0239)	0.6106	(.0323)	0.4059	(.0277)	0.3381	(.0308)
Electronics	0.4035	(.0261)	0.4440	(.0298)	0.5024	(.0330)	0.3382	(.0335)
Food	-0.3260	(.0136)	-0.3711	(.0180)	-0.2592	(.0142)	-0.2511	(.0147)
Furniture	-0.2304	(.0261)	-0.2722	(.0324)	-0.2166	(.0270)	-0.1313	(.0260)
Leather	-0.2093	(.0644)	-0.2312	(.0751)	-0.1265	(.0593)	-0.2024	(.0800)
Mechanic	0.4159	(.0250)	0.5096	(.0321)	0.3139	(.0287)	0.3291	(.0286)
Metallurgic	0.2630	(.0153)	0.2721	(.0195)	0.2850	(.0175)	0.2734	(.0184)
Mineral	0.1300	(.0210)	0.1398	(.0267)	-0.0918	(.0302)	-0.0081	(.0382)
Non-metallic	-0.3511	(.0215)	-0.3123	(.0268)	-0.3097	(.0229)	-0.2941	(.0252)
Others	-0.1398	(.0339)	-0.1192	(.0428)	-0.1296	(.0389)	-0.0103	(.0391)
Paper	0.0846	(.0397)	0.3011	(.0555)	0.3617	(.0457)	0.1560	(.0484)
Perfume	0.2442	(.0715)	-0.2291	(.0971)	0.0695	(.0737)	0.1309	(.0745)
Pharmaceuticals	0.5966	(.0551)	0.5905	(.0694)	0.5348	(.0664)	0.6685	(.0678)
Plastics	0.0184	(.0407)	0.2203	(.0491)	0.1376	(.0411)	-0.0079	(.0408)
Publishing	0.2374	(.0292)	0.2302	(.0369)	0.2350	(.0304)	0.3433	(.0336)
Rubber	0.2755	(.0574)	0.1999	(.0622)	0.1531	(.0653)	0.3570	(.0722)
Textile	-0.2251	(.0227)	-0.1742	(.0272)	0.0098	(.0268)	-0.1434	(.0304)
Tobacco	0.1828	(.0750)	-0.0349	(.0971)	0.3789	(.0926)	0.3240	(.1087)
Transport	0.5245	(.0226)	0.5880	(.0291)	0.6347	(.0286)	0.5445	(.0292)
Wood	-0.3830	(.0210)	-0.4510	(.0277)	-0.3949	(.0262)	-0.2919	(.0257)
R^2	0.1626		0.1606		0.1518		0.1281	
F-value	187.96		135.38		119.16		96.79	
SD	0.3393		0.3662		0.3006		0.2837	
n	20,244		14,876		13,996		13,846	

Note:
SD is the weighted and adjusted standard deviation of inter-industry wage dispersion calculated using equation 3.

Table 4

Inter-industry wage differentials

	1984[2]		1988		1992		1996	
Industry	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.
Apparel	-0.0717	(.0136)	-0.1734	(.0162)	-0.1242	(.0140)	-0.1224	(.0138)
Beverage	-0.1518	(.0275)	-0.0946	(.0339)	-0.0416	(.0320)	0.0238	(.0334)
Chemical	0.2840	(.0172)	0.2271	(.0229)	0.2064	(.0208)	0.1284	(.0221)
Electronics	0.1750	(.0188)	0.1824	(.0216)	0.1187	(.0252)	0.0626	(.0245)
Food	-0.1819	(.0100)	-0.1791	(.0131)	-0.1165	(.0112)	-0.0979	(.0111)
Furniture	-0.0890	(.0190)	-0.0637	(.0233)	-0.0395	(.0208)	0.0393	(.0193)
Leather	-0.1032	(.0457)	-0.0756	(.0528)	-0.1544	(.0445)	-0.2222	(.0571)
Mechanic	0.0844	(.0182)	0.1235	(.0230)	0.0329	(.0217)	0.0600	(.0207)
Metallurgic	0.0764	(.0111)	0.0405	(.0140)	0.0822	(.0133)	0.0787	(.0134)
Mineral	0.2970	(.0161)	0.3605	(.0202)	0.1420	(.0237)	0.1190	(.0281)
Non-metallic	-0.1593	(.0155)	-0.0800	(.0192)	-0.0534	(.0177)	-0.0355	(.0187)
Others	-0.0886	(.0242)	-0.1173	(.0303)	-0.0866	(.0292)	-0.0303	(.0279)
Paper	0.0551	(.0282)	0.1286	(.0389)	0.1661	(.0342)	-0.0425	(.0345)
Perfume	0.0415	(.0508)	-0.2333	(.0681)	0.0580	(.0549)	-0.0448	(.0531)
Pharmaceuticals	0.1190	(.0395)	0.1309	(.0490)	0.1667	(.0498)	0.2777	(.0487)
Plastics	-0.0246	(.0289)	0.0280	(.0345)	0.0077	(.0307)	-0.0240	(.0292)
Publishing	-0.1643	(.0212)	-0.1377	(.0264)	-0.0306	(.0232)	0.0035	(.0245)
Rubber	0.0999	(.0407)	0.0750	(.0436)	0.0123	(.0487)	0.1252	(.0515)
Textile	-0.0725	(.0164)	-0.0694	(.0195)	-0.0029	(.0203)	-0.0457	(.0219)
Tobacco	0.0700	(.0533)	-0.0137	(.0682)	0.2394	(.0691)	-0.0064	(.0776)
Transport	0.2225	(.0166)	0.2186	(.0210)	0.2732	(.0219)	0.1801	(.0215)
Wood	-0.0988	(.0160)	-0.0756	(.0205)	-0.0629	(.0206)	0.0020	(.0194)
R^2	0.5908		0.5904		0.5313		0.5599	
F-value	717.92		521.71		385.86		428.34	
SD	0.1401		0.1457		0.0839		0.0427	
N	20,244		14,876		13,996		13,846	

Notes:

1. Control variables are 5 region dummies, 7 occupation categories dummies, education, experience, experience square, head of family, gender, work-card, race, metropolitan area, union, over time worked.
2. Union variable is not available, and race variable has many missing values for the 1984 data. Therefore, these variables are not included for this year.
3. *SD* is the weighted and adjusted standard deviation of inter-industry wage dispersion calculated using equation 3.

Table 5

Analysis of sources of wage variation

	Share of total sum of squares *(F-test of industry factor in parentheses)*							
Source of variation	1984		1988		1992		1996	
(1) Covariates and industry	0.533	(795.0)	0.524	(536.8)	0.470	(426.3)	0.486	(450.2)
(2) Covariates alone	0.488	(2,415.1)	0.484	(1743.9)	0.434	(1,338.1)	0.461	(1,480.2)
(3) Industry alone	0.163	(187.0)	0.161	(135.4)	0.152	(119.2)	0.128	(96.8)
Industry effects								
Lower bound (1) - (2)	0.045		0.040		0.036		0.025	
Upper bound (3)	0.163		0.161		0.152		0.128	
Total sum of squares	15,026		12,902		8,940		9,416	
Variance of ln(wage)	0.74		0.86		0.64		0.67	
Mean of ln(wage)	3.08		3.04		2.57		2.84	
Cases	20,244		14,876		13,996		13,846	

Notes:
1. Factor variable is industry.
2. The dependent variable is ln(wage).
3. Covariates are education, experience, experience square, gender, head of family status, 5 regions, work-card, and occupation.
4. All F-tests are significant at well above the 1 percent level.

Table 6

Rank of marginal contribution of explanatory variables

Change in standard deviation of regression after dropping exogenous variables – in %							
1984		1988		1992		1996	
Variable dropped		Variable dropped		Variable dropped		Variable dropped	
Education	10.87	Education	10.17	Education	7.29	Education	9.58
Occupation	4.66	Occupation	3.86	Occupation	3.37	Occupation	5.18
Industry	3.49	Industry	3.06	Region	2.60	Region	3.64
Experience	2.62	Experience	2.60	Industry	2.00	Experience	2.59
Gender	1.18	Region	2.00	Work-card	1.93	Gender	1.60
Head	1.07	Over time	1.47	Experience	1.80	Over time	1.58
Region	1.07	Work-card	1.29	Over time	1.37	Industry	1.17
Work-card	0.93	Gender	1.21	Gender	1.01	Head	0.85
Over time	0.75	Head	0.96	Head	0.63	Metropolitan	0.27
Metropolitan	0.20	Metropolitan	0.13	Metropolitan	0.18	Work-card	0.16

Table 7

Spearman rank correlation coefficients between non-controlled wage differentials

	1984	1988	1992	1996
1984	1			
1988	0.9119	1		
1992	0.9097	0.9176	1	
1996	0.9356	0.8780	0.9176	1

Note:
All coefficients are significant at the 1 percent level. Correlates are from Table 3.

Table 8

Spearman rank correlation coefficients between controlled wage differentials

	1984	1988	1992	1996
1984	1			
1988	0.8483	1		
1992	0.8182	0.7619	1	
1996	0.6894	0.7416	0.6680	1

Note:
All coefficients are significant at the 1 percent level. Correlates are from Table 4.

Figure 1: Import penetration ratio - manufacturing

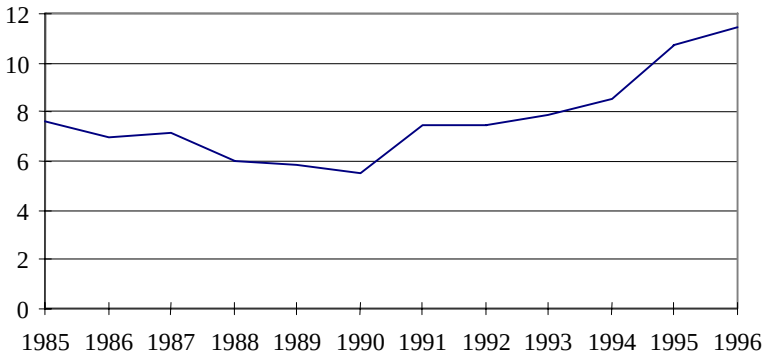
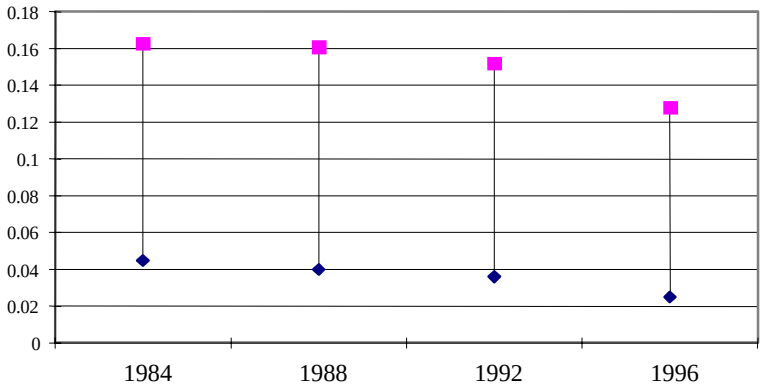


Figure 2: Industry affiliation effect on wage



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APPENDIX

Table A1

Means and standard deviations

	1984		1988		1992		1996	
Variable	Mean	SD	Mean	SD	Mean	SD	Mean	SD
ln(wage)	30.08	0.8601	30.04	0.9301	20.57	0.8001	20.84	0.8202
Region								
Centre-West	0.0545	0.2271	0.0668	0.2496	0.0604	0.2382	0.0672	0.2505
North	0.0729	0.2599	0.0776	0.2675	0.0479	0.2137	0.0517	0.2215
South-East	0.4844	0.4998	0.4510	0.4976	0.4492	0.4974	0.4357	0.4959
South	0.2013	0.4010	0.2052	0.4039	0.2587	0.4380	0.2541	0.4354
North-East	0.1601	0.3602	0.2001	0.4005	0.1800	0.3902	0.1901	0.3903
Occupation								
Clerical	0.0830	0.2009	0.0820	0.2126	0.0794	0.2703	0.0651	0.2468
Manager	0.0570	0.2323	0.0551	0.2280	0.0612	0.2398	0.0631	0.2432
Sales	0.0390	0.1939	0.0401	0.1950	0.0433	0.2035	0.0498	0.2175
Semi-skilled	0.4780	0.4982	0.4620	0.4969	0.5116	0.4985	0.5259	0.4969
Skilled	0.1480	0.3300	0.1430	0.3255	0.1220	0.3273	0.1110	0.3214
Technician	0.0390	0.1933	0.0409	0.1971	0.0323	0.1768	0.0346	0.1828
Unskilled	0.1560	0.3902	0.1760	0.3801	0.1502	0.3602	0.1505	0.3504
Other variables								
Education (years)	5.73	3.96	6.02	3.92	6.09	3.79	6.57	3.81
Experience (years)	21.73	11.44	20.19	12.15	20.51	11.91	20.38	11.81
Gender (female=1)	0.2064	0.4048	0.2434	0.4292	0.2504	0.4332	0.2500	0.4331
Head of family (=1)	0.6539	0.4757	0.5483	0.4977	0.5953	0.4909	0.5884	0.4921
Metropolitan area (=1)	0.5744	0.4945	0.5383	0.4986	0.4993	0.5000	0.4779	0.4995
Over time worked	0.6143	0.4868	0.4881	0.4999	0.3288	0.4698	0.3619	0.4806
Race (white=1)	*na*	*na*	0.5781	0.4939	0.5877	0.4923	0.5978	0.4904
Union status (=1)	*na*	*na*	0.2520	0.4341	0.3039	0.4599	0.2668	0.4423
Work-card (signed=1)	0.8006	0.3995	0.7876	0.4090	0.7706	0.4204	0.7362	0.4407

Note:
The work-card variable is related to formal-informal labour contract. It captures the effect of labour legislation on wage determination. Skilled workers are defined as craftsmen, foremen and kindred workers. Semi-skilled workers are operatives and kindred workers. Unskilled workers are service workers.

Table A2

Coefficients of control variables

	1984		1988		1992		1996	
Industry	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.	Coef.	S.E.
Centre-West	0.1427	(.0197)	0.1680	(.0227)	0.2008	(.0223)	0.2702	(.0215)
North	0.2038	(.0176)	0.1910	(.0217)	0.1627	(.0246)	0.2287	(.0240)
South-East	0.2338	(.0113)	0.3037	(.0143)	0.3319	(.0137)	0.3664	(.0137)
South	0.1569	(.0131)	0.1516	(.0171)	0.2069	(.0159)	0.2761	(.0158)
Clerical	0.2879	(.0217)	0.2535	(.0257)	0.2985	(.0240)	0.2993	(.0253)
Semi-skilled	0.0317	(.0107)	0.0208	(.0132)	0.1528	(.0160)	0.1474	(.0163)
Skilled	0.3219	(.0148)	0.2822	(.0184)	0.3580	(.0211)	0.3728	(.0216)
Manager	0.7053	(.0202)	0.6212	(.0252)	0.6381	(.0255)	0.7505	(.0255)
Sales	0.3662	(.0226)	0.2607	(.0278)	0.2972	(.0280)	0.3242	(.0270)
Professional	0.6567	(.0235)	0.7111	(.0288)	0.6814	(.0322)	0.8395	(.0317)
Education	0.1061	(.0016)	0.1020	(.0019)	0.0771	(.0018)	0.0862	(.0017)
Experience	0.0518	(.0016)	0.0471	(.0017)	0.0349	(.0016)	0.0413	(.0016)
Experience^2	-0.0007	(.0000)	-0.0007	(.0000)	-0.0005	(.0000)	-0.0006	(.0000)
Gender	-0.2734	(.0125)	-0.2723	(.0142)	-0.2343	(.0133)	-0.2831	(.0132)
Head	0.2263	(.0108)	0.2156	(.0135)	0.1560	(.0122)	0.1766	(.0120)
Metropolitan	0.0786	(.0087)	0.0893	(.0109)	0.0841	(.0100)	0.0997	(.0100)
Overtime	-0.1492	(.0085)	-0.2226	(.0104)	-0.2004	(.0105)	-0.2109	(.0101)
Race	*na*	*na*	0.1480	(.0112)	0.1123	(.0107)	0.1576	(.0108)
Union	*na*	*na*	0.1561	(.0120)	0.1662	(.0109)	0.1513	(.0113)
Work-card	0.2210	(.0113)	0.2408	(.0138)	0.2539	(.0127)	0.0464	(.0120)

Table A3

RAIS Inter-industry wage differentials

Deviation from the weighted average wage (no control variables included)		
Industry	1988	1992
Apparel	-0.7448	-0.6988
Beverages	-0.1378	-0.2392
Chemical	0.5913	0.5541
Electronics	0.2985	0.3421
Food	-0.1870	-0.3898
Furniture	-0.5718	-0.6107
Leather	-0.6176	-0.5333
Mechanical	0.2521	0.2864
Metallurgic	0.1330	0.2267
Mineral	0.2659	0.2589
Non-metallic	-0.1261	-0.1339
Other	-0.0249	-0.0109
Paper	0.1095	0.1869
Perfumes	-0.0119	0.1016
Pharmaceuticals	0.3138	0.3766
Plastics	-0.1257	-0.0876
Printing	0.0156	0.0488
Rubber	0.0329	0.1384
Textiles	-0.2515	-0.2071
Tobacco	0.0876	0.4017
Transport	0.4611	0.5699
Wood	-0.6539	-0.7034
SD	0.3521	0.3836
n	6,026,359	4,828,707

Notes:
Parameters of the table: year; 22 industries; wage bands; workers employed in 31 December. *n* is the employment of the 22 industries.

Table A4

Spearman rank correlation of structure of effective protection

	1980	1991	1992	1993
1980	1			
1991	0.0813	1		
	(0.719)			
1992	0.1271	0.7929	1	
	(0.604)	(0.000)		
1993	0.1169	0.9873	0.8026	1
	(0.573)	(0.000)	(0.000)	

Note:
Effective protection estimated by Pinheiro and Almeida (1994), Table 6. *p*-value in parentheses.

Table A5

Spearman rank correlation of import penetration

	1985	1988	1992	1996
1985	1			
1988	0.9187	1		
1992	0.9255	0.9593	1	
1996	0.8103	0.869	0.9164	1

Note:
Estimated using the value of production from Haguenauer *et al*. (1998). All coefficients are significant at the 1 percent level.

Table A6

Spearman rank correlation of export ratio

	1985	1988	1992	1996
1985	1			
1988	0.8713	1		
1992	0.8159	0.9356	1	
1996	0.7899	0.9277	0.965	1

Note:
Estimated using the value of production from Haguenauer *et al*. (1998). All coefficients are significant at the 1 percent level.
Export ratio is exports/production.