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THE INTER-INDUSTRY WAGE STRUCTURE: EMPIRICAL EVIDENCE FOR GERMANY AND A COMPARISON WITH THE U.S. AND SWEDEN*

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February 1992

Abstract

This paper investigates the inter-industry wage structure in the 1984 wave of the German Socioeconomic Panel and compares the main findings with those available for the U.S. and Sweden. In contrast with what emerges from aggregate data, empirical evidence based on individual data emphasizes cross-country differences. Although industry differentials appear significant, labour quality and other compensating factors have a major impact in explaining the wage structure, thus suggesting the possibility that industry differences just reflect the effect of unobservable characteristics. The degree of centralization of wage bargaining may play a role in accounting for the observed pattern of industry wages.

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"A few years ago we hired a new secretary in my department. She was smart and efficient and we were pleased to have her. Much to our dismay, after just a few months she was offered and accepted a job from an IBM facility in a nearby city. She told me that she had been on a waiting list there for a year or so, and would be a fool to turn IBM down since they paid so much more than any of the other local employers. I wondered at the time whether her marginal product typing IBM interoffice memos could be that much higher than it would be typing manuscripts and referee reports, and/or why IBM should find it profitable to pay much more than the going wage. (...). These observations seem to violate the law of one price, a fundamental component of the theory of competitive markets." (Thaler, 1989, pp.181-182).

1. Introduction

Several empirical studies of the inter-industry wage structure show results which do not seem consistent with the competitive model of the labour market. On the one hand, at a micro data level of analysis, they suggest that there exist unexplained wage differentials which depend on industry affiliation. Even after controlling for varying human capital factors and working conditions, industry wage differences appear to remain substantial (Krueger and Summers, 1987 and 1988; Dickens and Katz, 1987a and 1987b; Katz and Summers, 1989). On the other hand, using aggregate industry wage data, some authors claim that the structure of inter-industry wage differences appears to be stable over time and across different countries, a result that may call for the introduction of non-competitive considerations to be rationalized (Krueger and Summers, 1987; Katz and Summers, 1989).

The existing empirical literature on the determinants of industry wage differentials using individual data essentially concerns the U.S. labour market. Some evidence is also available for the Swedish labour market (Edin and Zetterberg, 1989). In this paper I present evidence for Germany, based on the individual data available from the German Socioeconomic Panel.

The approach here adopted consists in the estimation of a wage equation which includes measures of human capital and working conditions, as well as industry affiliation controls, as explanatory variables. This method permits a better evaluation of competitive and non-competitive influences on the process of wage determination and the resulting inter-industry structure of relative wages. As we will see, industry differentials appear to a certain extent significant. However, differently from what emerges for the U.S. case, labour quality and other competitive factors have a major impact in explaining the wage structure, thus suggesting the possibility that industry differences just reflect the effect of unobservable characteristics.

Empirical analyses of the stability of wage structures across countries have been based so far on aggregate industry wage data. Here I consider cross-country comparisons of the inter-industry wage structure as emerging from micro data, rather than from average industry data. I will contrast my results with those presented in two similar studies, the first on the U.S. by Krueger and Summers (1988) and the second on Sweden by Edin and Zetterberg (1989). All the papers make use of individual level data in a similar cross-sectional approach for 1984. My interest in comparing the wage dispersion across industries in these three countries derives from the fact that they are usually considered as characterized by different labour market institutional structures and in particular, by a different degree of centralization of wage bargaining. In contrast with what emerges from aggregate data, empirical evidence based on individual data emphasizes cross-country differences and the degree of centralization of wage bargaining may play a role in accounting for the observed pattern of industry wages.

The rest of the paper is structured as follows. In section 2 I illustrate theories of the labour market which provide explanations for the existence of inter-industry wage differentials. In section 3 I present several criteria for evaluating the degree of centralization of wage setting and compare the different institutional conditions for wage bargaining in the U.S., Germany, and Sweden; these countries span the range of the degree of centralization in OECD labour markets. In section 4 I show and discuss the limits of empirical evidence for the inter-industry wage structure based on aggregate industry wage data as presented in the recent literature. In section 5 I describe the main results of the empirical analysis of the German case obtained with the data from the Socioeconomic Panel. In section 6 I compare my empirical evidence with that available for the U.S. and Sweden. Section 7 contains some concluding remarks.

2. Theories of the Labour Market and Inter-industry Wage Differentials

Competitive theories of the labour market suggest that job characteristics which have no direct influence on the utility of workers should not affect the level of wages. Workers are paid a wage equal to their opportunity cost, which depends on accumulated human capital and their working conditions. In other words, equally skilled workers should be compensated in a way that guarantees equal levels of utility. In this context, inter-industry wage differentials might simply reflect labour quality differences which vary systematically across industries and/or compensating differentials for some aspects of the working conditions in the various industries of employment.

Non-competitive theories, instead, predict that job attributes having no effect on workers' utility should systematically affect the wage structure, as far as they have an influence on the optimal solution to the firms' maximization problem. Equally skilled workers are paid differently according to features like industry affiliation. Possible explanations for persistent inter-industry wage differentials for equally productive workers in the class of non-competitive theories are offered by efficiency wage models (Krueger and Summers, 1988) and the insider-outsider model (Lindbeck and Snower, 1988).

The efficiency wage model in all its variants - shirking, labour turnover, adverse selection, gift exchange models (surveyed by Stiglitz, 1986 and Katz, 1986) - provides a rationale for the fact that firms' attributes not affecting workers' utility can indeed affect the structure of relative wages. The key issue is firms' imperfect information: if firms differ in their ability to face turnover costs, to measure the productivity of workers, or to monitor them, either because of differences in management capacity, or because of differences in the technology of production, then this may be reflected in substantial wage differentials for similar workers. The optimal wage will be different among firms and, in particular, some firms will find it profitable to pay non-competitive wages above the going rate, thus inducing involuntary unemployment. And if some forms of imperfect information - like those giving rise to turnover costs or monitoring costs - depend on industry specific characteristics, such as the technology of production, then firm-level differentials will be reflected, at an aggregate level, in significant inter-industry wage differentials, consistently with the efficiency wage hypothesis (Krueger and Summers, 1988).

Another possible explanation of the inter-industry wage structure is the one arising from the insider-outsider theory. According to the insider-outsider model, wages are the outcome of a bargaining process whereby firms and their 'insiders' share the economic rent from insider employment. In this context, the insiders' wages will be higher, the more their firms stand to lose from a breakdown in wage negotiations. The degree of vulnerability to the exercise of the insider power is in turn a function of industry specific conditions, like profit opportunities, the technology of production, the concentration ratio and workers market power in their industries. Inter-industry wage differentials may therefore reflect differences in these sectoral conditions (Lindbeck and Snower, 1988).

Recent literature has also suggested a more general explanation for the observed pattern of inter-industry wage differentials, based on the relationship between the degree of wage dispersion across industries and the type of labour market institutions (Bell and Freeman, 1986; Freeman, 1988). Labour market institutions have received increasing attention as a determinant of the labour market performance of the advanced economies in terms of employment and unemployment rates and, more generally, of the macroeconomic performance of a country (Bruno and Sachs, 1985, ch.11; Newell and Symons, 1987; Calmfors and Driffill, 1988; Freeman, 1988; Layard et al., 1991). Among the aspects characterizing the structure of labour markets, the focus is on the degree of centralization of wage bargaining. Countries with very high or very low wage dispersion are supposed to be the countries with highly decentralized or highly centralized wage setting procedures respectively, while intermediate degrees of centralization tend to be associated with an intermediate wage dispersion across industries (Freeman, 1988).

The explanation of industry wage differentials in terms of degree of centralization in wage setting procedures - and in general of labour market institutions - may be seen as a more general approach when compared with the efficiency wage and insider-outsider models. I believe that it is not in contrast with these two theories and that, in a sense, it encompasses both of them. Efficiency wage and insider-outsider theories are not independent of the institutional framework and can be considered as alternative microfoundations for wage determination, which become relevant under different labour market institutions. According to the efficiency wage models, firms are assumed to make both wage and employment decisions unilaterally. This assumption seems more realistic in a labour market characterized by a high degree of decentralization of wage bargaining, where workers have a limited market power in wage negotiations. In the insider-outsider model, the market power of the insiders in the bargaining process is certainly increased by some degree of unionization, but at the same time unions are assumed not to take into account the interest of the outsiders. These conditions seem consistent with an intermediate degree of centralization of wage setting, where unions can exert some market power but are led to ignore the unemployment consequences of their actions. If we test the nature of the relationship between the degree of centralization of wage bargaining and the degree of wage dispersion, we therefore implicitly test also the potential relevance of efficiency wage and insider-outsider models in explaining the observed pattern of industry wage differentials and we ultimately evaluate their importance in accounting for involuntary unemployment.

Despite the general agreement about the significant role played by labour market institutions from a theoretical point of view, existing empirical analyses of the nature of the relationship between cross-country differences in the degree of centralization and of wage dispersion and the differing labour market performance across countries have offered conflicting interpretations. Some studies (Bean, Layard and Nickell, 1986; Newell and Symons, 1987; Layard et al., 1991) relate good outcomes with respect to employment and unemployment to centralized labour markets with low wage dispersion, where large and all-encompassing trade unions take into account the unemployment effects of wage determination. Others (OECD, 1985 and 1987) associate labour market success with decentralization of wage setting and greater cross-industry wage dispersion, since a higher flexibility of relative wages allows greater scope for industry-specific factors. Finally, some authors (Calmfors and Driffill, 1988; Freeman, 1988) postulate the existence of a non-monotonic relationship between centralization/dispersion and the labour market performance, with countries having highly centralized institutions and low wage dispersion and countries with highly decentralized bargaining and high wage dispersion both showing better employment outcomes than countries with intermediate types of wage structure and labour market institutions.

These divergences in the conclusions about the relation between wage dispersion, institutions and economic performance suggest the danger of generalizing from results based on aggregate data, as it is the case in the studies previously mentioned. On the one hand, countries exhibiting a similar degree of wage dispersion at an aggregate level are characterized by very different institutional frameworks and wage bargaining procedures. For example, the U.S. and Austria, which are generally regarded as extreme cases of decentralization and centralization respectively, are both classified by Freeman (1988) as having a high and increasing degree of wage dispersion within OECD countries over the period 1970-86; similarly for Switzerland and Germany, as less extreme examples. On the other hand, countries forming natural pairs in institutional terms - such as Sweden and Norway, the U.K. and Ireland - present, again according to Freeman (1988), rather different degrees of dispersion in the inter-industry wage structure. The same sort of ambiguity can be detected within some of the above mentioned studies with respect to the relationship between indicators of the labour market structure and labour market performance, since countries with similar institutions may perform very differently - for example, the U.S. and Japan, Sweden and Denmark (Calmfors and Driffill, 1988; Freeman, 1988).

If cross-industry wage dispersion is used as the most important indicator of the underlying labour market institutions to be related to employment performances, average industry wages at an aggregate level are not a proper magnitude for evaluating the inter-industry structure of relative compensations. In fact, they may reflect many factors other than institutional characteristics which present wide variations across industries - like differences in technology and productivity - and which are perfectly consistent with the competitive theories of the labour market.

For a better evaluation of the relationship between the degree of centralization and industry wage dispersion, at least two problems deserve some attention. Firstly, we need an accurate definition of centralization, in order to classify the various countries unambiguously. The following section 3 discusses this subject. Secondly, inter-industry wage differentials of a non-competitive nature need to be properly measured. This major point is treated extensively in section 4, where existing evidence based on aggregate industry wage data is presented and critically assessed, and in sections 5 and 6, where empirical results based on micro data are illustrated and opposed to those obtained with aggregate data.

3. Institutional Conditions for Wage Bargaining in the U.S., Germany, and Sweden

Following Tarantelli (1986), I define the degree of centralization in wage bargaining as one of the main dimensions of a labour market characteristic termed "corporatism" by political scientists, the others being the degree of cooperation between trade unions and employers' representatives in wage bargaining and the system of regulation of industrial conflicts. According to Bruno and Sachs (1985), corporatism is in turn one of the two important dimensions of labour market flexibility, together with nominal wage responsiveness to changing labour market conditions. It is therefore clear that the degree of centralization represents only one of the aspects of the broader concept of labour market flexibility, but hopefully one which is more easily measurable.

In their attempt to classify different countries according to the degree of centralization of wage bargaining, various authors have considered several definitions of centralization, each of them putting emphasis on different factors characterizing the process of wage setting. Bruno and Sachs (1985) focus on the level at which wage negotiations proceed and on the extent of coordination within trade unions and employer associations. According to them, the key feature is voting on collectively bargained agreements. A similar but more limited aspect is emphasized in the definition provided by Calmfors and Driffill (1988). They define it as the extent of inter-union and inter-employer cooperation in wage bargaining. The focus is therefore on the extent to which coalitions are formed among unions and employers respectively, that is on the behavioural content of wage setting rather than on the formal one, as it is the case when we consider the level at which bargaining occurs. Other authors - reported and discussed by Calmfors and Driffill (1988) - define centralization according to similar criteria. Schmitter (1981) considers only the union side. Cameron (1984) takes into account only the union side, but in addition he considers the extent of unionization, in an attempt to measure cooperation among workers in general rather than among unions only. Blyth (1979) relies on two criteria: the extent to which coalitions are formed within unions and within employers in actual wage bargaining and the level at which bargaining takes place. A much more comprehensive definition of centralization is proposed by Tarantelli (1986). According to him, a centralized system of industrial relations is characterized by bargaining mainly taking place at the national and/or industrial or regional level, rather than at the company and plant level, by a high degree of coordination within the organizational structure of trade unions, by a few key contracts influencing directly or indirectly a high percentage of the labour force, and by bargaining taking place at close - for example, one year - and non-overlapping or synchronous intervals. This last condition is particularly interesting, since it attempts to measure the possibility of fine-tuning wage agreements preserving the structure of relative compensations in the presence of changing economic conditions. Finally, Freeman (1988) suggests the use of union density as a very simple indicator of the degree of centralization. The main problem with this approach is to judge the extent to which differing unionization rates actually reflect differences in the formal or informal coverage of union contracts (Calmfors and Driffill, 1988). Moreover, differences in national definition and measurement are likely to affect the comparability across countries. Freeman himself concludes from his empirical analysis that after controlling for other indicators of centralization, union density does not play a significant additional role (Freeman, 1988).

In the light of the various definitions of centralization that I have provided, I can now turn to examine the principal institutional conditions for wage setting in the three countries considered in the present study - the U.S., Germany, and Sweden - in order to try to classify their degree of centralization of wage bargaining. I will especially consider the situation as it appeared in the first half of the 1980s, which is the relevant time period for my later empirical analysis.

3.1 The U.S.

The U.S. labour market is usually qualified as one of the most decentralized among Western economies. Union density in 1984/5 was only 18 percent of non-agricultural wage and salaried employees and fell sharply during both 1970-79 (-6%) and 1979-85 (-7%) (Freeman, 1988). Wage negotiations occur predominantly at the enterprise and plant level. There is no traditional involvement by central organizations in bargaining: the main U.S. labour confederation, the AFL-CIO, does not bargain for its affiliated unions and therefore has never signed a wage contract; no national employer federation is engaged in the collective bargaining process. As a general practice, a large proportion of collectively bargained agreements must be ratified by individual union members. The U.S. system exhibits a largely unstable and complex network of pattern bargaining, with 195,000 collective agreements affecting about 25% of the labour force (at the end of the 1970s). Synchronization of contract renewals is very low and contracts have a long duration - often three years (Bruno and Sachs, 1985; Tarantelli, 1986; Calmfors and Driffill, 1988).

3.2 Germany

Germany presented in 1984/5 a moderate degree of union density, 42 percent of industrial employees, which exhibited an increase in the period 1970-79 (+5%) and remained constant in the period 1979-85 (Freeman, 1988). Collective wage agreements are normally struck within industries on a regional basis, but regional negotiations are closely coordinated by national trade unions and employer associations. Wage bargaining in the metal industry provides the guide-lines for negotiations in all the other industrial and service sectors. Trade unions and employer associations have comparatively centralized and encompassing organizational structures. There exists one central union confederation, the DGB, and single unions are organized on a branch basis - 17 branch trade unions associated with the DGB. A large proportion of employers is organized in associations, the most important of which - on a national basis - is the BDA. Central associations, however, are not usually involved in actual collective bargaining. Contracts signed in the metal sector by the metal industry union, the IGM, represent a benchmark for all the other contracts, thus influencing as a matter of fact a high percentage of the labour force. The duration of contracts is typically one year and negotiations take place throughout the year, though again the metal sector is considered to be the key industry for setting the contract renewals pattern, which guarantees a fair degree of synchronization (Flanagan et al., 1983; Bruno and Sachs, 1985; Calmfors and Driffill, 1988).

A distinctive feature of German collective bargaining is that the relationship between the trade unions and the employer associations is characterized by a relatively high level of cooperation and willingness to compromise. During the 1980s, trade unions accepted wage agreements which led - for the first time in the post-war era - to a considerable reduction of net real wages. In the spring of 1984, after the biggest strike in the history of West Germany, employers agreed to a reduction of the working week from forty to thirty-eight and a half hours and unions - with IGM as the pacemaker - consented in return to an increase in the degree of flexibility of working time. These agreements also reflect a general tendency toward the decentralization of industrial relations, implying a shift of competence in collective bargaining from an industry to a company level and a delegation of decision power to the management and work councils under the system of "co-determination", which is the peculiar institution of German industrial relations (Streeck, 1987; Jacobi and Mueller-Jentsch, 1990).

3.3 <u>Sweden</u>

The Swedish system is built around nearly universal union participation and this puts Sweden among the countries with the largest extent of union coverage. In 1984/5, 95 percent of blue-collar workers were represented in the national trade union confederation, the LO, and approximately 75 percent of white-collar workers were represented in two other union organizations, TCO and SACO-SR. The overall union density experienced a sharp rise both during 1970-79 (+10%) and during 1979-85 (+6%) (Bruno and Sachs, 1985; Freeman, 1988). The level of negotiation is highly centralized: detailed wage bargaining typically takes place at the industry level and then is further refined at the local level, but national level agreements serve as an essential guide-post for negotiations at the industry and firm level. The union confederation LO is organized along branch lines - 24 branch federations associated with the central organization. Employers are almost universally represented in the national employer confederation, the SAF. Central collective agreements are negotiated, without exception, between LO and SAF. Negotiators at the branch level have the power to reach agreements which are binding for all the members of the branch and individual union members voting on these agreements is virtually nonexistent. As a result, national and a few branch-level agreements affect almost the entire economy's wage setting. Synchronization of contract renewals is high and contracts duration is normally one-two years (Bruno and Sachs, 1985; Calmfors and Driffill, 1988).

Between the 1950s and the mid-1970s, a wide national consensus was reached in Sweden concerning the so-called "Scandinavian model". Its norm was that labour should maintain a constant share of national income and the scope for overall real wage increases was measured on this rule. Despite the changes in general economic conditions since then, there remained, in the 1980s, a basic commitment to a constant-share strategy of wage setting (Bruno and Sachs, 1985).

Wage bargaining in Sweden has also been crucially influenced by the "solidaristic wage policy". Initially conceived in 1936 and fully elaborated by LO in the 1950s with both growth and equity objectives, it developed along with government labour market programmes about training and labour mobility. The basic principle was "equal pay for equal workers": workers performing the same job should receive the same wage, irrespective of inter-industry differences in productivity and profitability. The principle has been implemented by raising the relative wages of workers in low-productivity sectors and by not fully exercising bargaining power in sectors with the greatest ability to pay. Lacking exact criteria for comparing jobs in different industries, the solidaristic policy has given way, since the 1960s, to a strictly egalitarian narrowing of wage differences between workers in different occupations (Flanagan, 1987).

The Swedish case seems, in several respects, to be representative of a highly centralized economy. There are however some caveats against the traditional classification of Sweden as one of the most centralized countries within OECD (Calmfors and Nymoen, 1990). Contrary to the frequently held view, centralized bargaining never concerns the whole economy. There is no coordination between the private and public sector unions within central organizations and both private and public employers bargain with different associations for blue-collar and white-collar workers. More relevantly, bargaining never occurs at the central level only. Because of the "wage drift" - wage increases at the local level in excess of the central agreements - which has become more important with real wage moderation after 1983, it is not clear to what extent the final outcome is controlled by central wage bargaining. Between 1971 and 1984, the wage drift accounted for 40 to 50 percent of the hourly earnings increases of blue-collar workers (40% in 1984) and for 20 to more than 50 percent of the hourly earnings increases of white-collar workers in private industry (61% in 1984) (Flanagan, 1987). Some authors (Bruno and Sachs, 1985) support the hypothesis that central bargaining may be able to offset variations in wage drift, but this view is challenged by others (Flanagan, 1987), and the question remains unsettled.

From these various institutional conditions characterizing the process of wage setting, I believe we may conclude that the U.S. have an extremely decentralized system of wage bargaining, Germany represents an intermediate case, and Sweden has one of the highest degrees of centralization of wage bargaining among Western countries. This view about the ranking of the three countries is shared with several authors (Blyth, 1979; Schmitter, 1981; Cameron, 1984; Calmfors and Driffill, 1988; Freeman, 1988).

Differences in the degree of centralization of wage setting in the U.S., Germany, and Sweden may be reflected in a different inter-industry wage structure in the three countries. Swedish labour market policy should imply relatively small wage differentials among workers in different industries, while the U.S. should represent the opposite extreme, having no central bargaining or policy restraint. In Germany the industry specific character of negotiations and policies, somehow mitigated by a tendency to the diffusion of institutional innovations across sectors, should create an intermediate situation. To what extent different degrees of centralization are associated with a different degree of wage dispersion is an empirical issue that I will address in the following sections.

4. The Inter-industry Wage Structure with Aggregate Industry Wage Data

One of the findings from the U.S. literature which is often put forward as supportive of non-competitive explanations for industry wage differences is the stability of the wage structure across time and countries (Krueger and Summers, 1987; Katz and Summers, 1989).

The results that I discuss in this section are derived from a study by Krueger and Summers, where the authors present, among other things, some empirical evidence for inter-industry wage differentials based on aggregate industry wage data (Krueger and Summers, 1987, pp.21-28).

The authors start considering the stability of wage differentials over time in the U.S., by comparing aggregate log average annual earnings of full-time equivalent employees in nine major industries for selected years between 1984 and 1900. The industries are agriculture, manufacturing, mining, construction, transportation, communications, wholesale and retail trade, FIRE (finance, insurance and real estate), and services. Their results are reported in the second column of Table I. Simple correlations

TABLE I

Industry wage structure through time in the U.S.: estimated correlation coefficients[a], respective standard errors[b], and 99% confidence intervals for log average annual earnings of full-time equivalent employees in nine major industries[c]

Year	Correlation with 1984[a] r	Standard error[b] s_r	99% Confidence interval[d]
1984	1.000	0.000	1.000 - 1.000
1980	0.984	0.016	0.942 - 1.026
1975	0.961	0.039	0.861 - 1.061
1970	0.909	0.084	0.693 - 1.125
1965	0.898	0.093	0.659 - 1.137
1960	0.893	0.097	0.644 - 1.142
1955	0.893	0.097	0.644 - 1.142
1950	0.866	0.117	0.565 - 1.167
1945	0.891	0.098	0.638 - 1.144
1940	0.836	0.137	0.482 - 1.190
1935	0.793	0.164	0.370 - 1.216
1930	0.761	0.182	0.291 - 1.231
1925	0.801	0.159	0.390 - 1.212
1920	0.807	0.156	0.406 - 1.208
1915	0.627	0.243	-0.001 - 1.255
1910	0.604	0.252	-0.046 - 1.254
1905	0.636	0.240	0.017 - 1.255
1900	0.616	0.248	-0.023 - 1.255

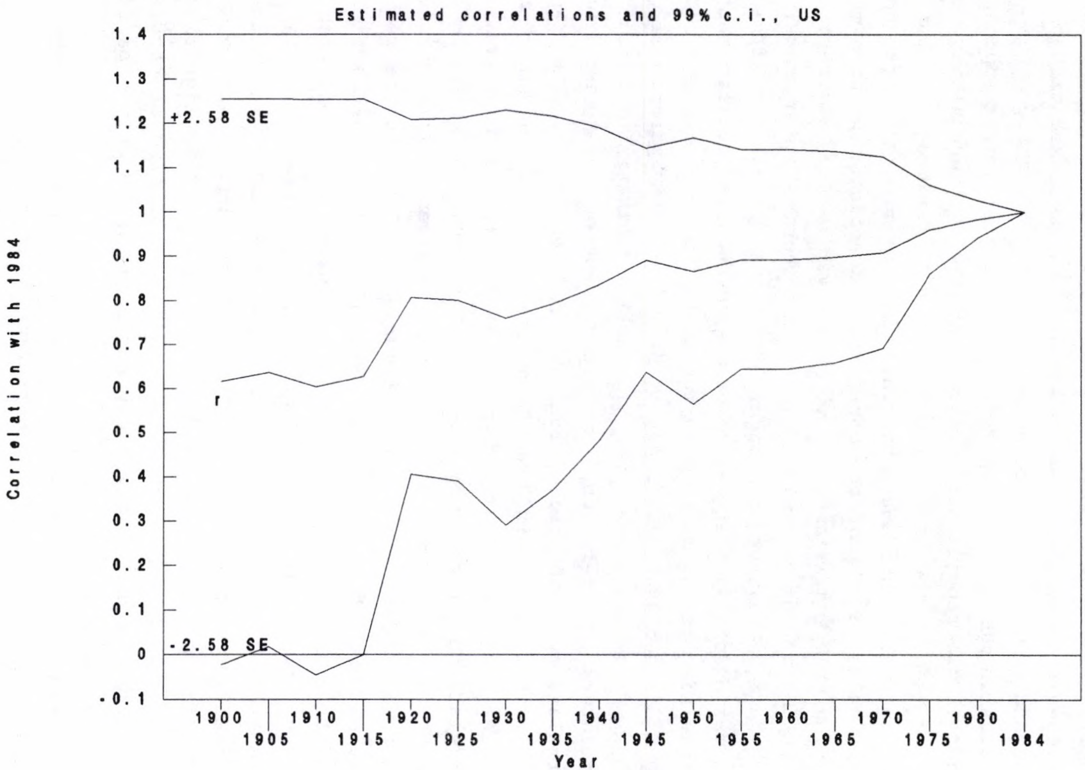
[a] <u>Source</u>: Krueger and Summers (1987, p.24, Table 2.2).
[b] See Appendix A for details about the estimate of the standard error of the correlation coefficient s_r with the formula suggested by Hotelling (1953).
[c] The sample size for the estimate of the correlation coefficients is n=9. Industries include: agriculture, manufacturing, mining, construction, transportation, communications, wholesale and retail trade, FIRE (finance, insurance and real estate), and services.
[d] Approximate 2-sided 99% confidence interval, under the assumption that $(r-\rho)/s_r$ follows an asymptotically standard normal distribution for the case $\rho=0$.

of log average annual earnings are proposed by Krueger and Summers to suggest that the industry wage structure for all industries has remained remarkably stable since 1920, with correlations with the wage structure in 1984 ranging from 0.76 to 0.98 (Krueger and Summers, 1987, Table 2.2). Before 1920 the pattern of industry wages appears less similar to the 1984 industry wage structure, but the correlations are still above 0.60 for all years up to 1900. Taking into account changes in industry definitions and sampling errors, this seems to imply, according to the authors, that the structure of relative industry wages changed only moderately over a very long time interval.

As is well known, however, the significance of simple correlation coefficients is sensitive to the sample size (Atkinson et al., 1990, pp.83-89). Given the fact that the correlations presented by Krueger and Summers are based on a very small sample of only nine pairs of observations - one for each industry aggregate - I want to construct a rigorous test for their statistical significance. The sampling distribution of the correlation coefficient is not well-behaved and several methods have been suggested to approximate it.

As a first way to address this issue, I construct two-sided 99% confidence intervals for the calculated correlation coefficients using an estimator of the approximation to the standard error of the correlation proposed by Hotelling (1953, p.212). The relevant formulae and other details about the calculations are given in Appendix A. The results obtained applying this method are reported in the last two columns of Table I, together with the respective estimated correlations, and graphed in Figure I. We can immediately notice the extreme width of the confidence intervals in many cases. For example, with a sample of only 9 observations, we have a 99% confidence interval of about 0.64 for a correlation coefficient equal to 0.616 and a 99% confidence interval of about 0.35 for a correlation coefficient equal to 0.836. Since the estimated standard error is a monotone decreasing function of the sample correlation coefficient - for positive values of the sample correlation - it tends to zero as the correlation approaches 1 in 1984. We also

FIGURE I Wage structure over time



observe that for some values of the correlation coefficient we are unable to reject the null hypothesis of a population correlation equal to zero.

This method, however, is not appropriate for accurate hypothesis testing with small samples. The approximation to the standard error is a consistent but biased estimator of its population counterpart and the correlation coefficient approaches a normal distribution only asymptotically and very slowly. This can be seen in the last column of Table I from the fact that the two-sided 99% confidence intervals thus derived exceed the upper limit value of +1 for the correlation coefficient, which means that, with a small sample of nine observations, only a very poor approximation to a typically skew distribution can be obtained.

An alternative approach, which avoids the problems of estimating the sampling error of the correlation coefficient and of claiming the normality of its distribution, is suggested by Kendall and Stuart (1977, p.416, §16.28). In the particular case in which we want to test the null hypothesis $\rho=0$ for the population correlation, the distribution of the sample correlation coefficient can be reduced to a "Student's t-distribution". This transformation can be used for a more precise hypothesis testing of the significance of the correlations presented by Krueger and Summers (1987). The details about the construction of the tests are given in Appendix A. In the third and fourth columns of Table II, I report the results obtained with the t-transformation of the correlation coefficient for the 1-tailed tests of the null hypothesis $\rho=0$ against the alternative hypothesis $\rho>0$, which is the relevant one in this case since stability of the industry wage structure over time is claimed. Outcomes are provided both at the conventional 1% significance level and at the more rigorous 0.5% significance level, in order to take into account the problem of precision that arises from having a very small sample of only 9 pairs of observations.

From Table II we observe that, with this method, the null hypothesis of a population correlation equal to zero cannot be rejected in four of the eighteen cases at the 1% significance

TABLE II

Industry wage structure through time in the U.S.: estimated correlation coefficients[a] and tests of their statistical significance[b] for log average annual earnings of full-time equivalent employees in nine major industries[c]

Year	Correlation with 1984[a] r	t-transformation Reject H_0: $\rho=0$ at significance level: 1%[d]	.5%[d]	z-transformation Reject H_0: $\rho=0$ at significance level: 1%[d]	.5%[d]
1984	1.000	yes	yes	yes	yes
1980	0.984	yes	yes	yes	yes
1975	0.961	yes	yes	yes	yes
1970	0.909	yes	yes	yes	yes
1965	0.898	yes	yes	yes	yes
1960	0.893	yes	yes	yes	yes
1955	0.893	yes	yes	yes	yes
1950	0.866	yes	yes	yes	yes
1945	0.891	yes	yes	yes	yes
1940	0.836	yes	yes	yes	yes
1935	0.793	yes	no	yes	no
1930	0.761	yes	no	yes	no
1925	0.801	yes	yes	yes	yes
1920	0.807	yes	yes	yes	yes
1915	0.627	no	no	no	no
1910	0.604	no	no	no	no
1905	0.636	no	no	no	no
1900	0.616	no	no	no	no

[a] Source: Krueger and Summers (1987, p.24, Table 2.2).
[b] See Appendix A for details about the construction of the tests with the t-transformation and the z-transformation of the correlation coefficient.
[c] See note c to Table I.
[d] 1-tailed tests of the null hypothesis H_0: $\rho=0$, against the alternative hypothesis H_1: $\rho>0$. The relevant claim is, in fact, that the industry wage structure is stable over time. Only positive correlations are therefore expected.

level and in six cases at the 0.5% significance level.

This approach based on the t-transformation overcomes the difficulty of directly estimating the sampling error of the correlation coefficient, but it still relies in some degree on the asymptotic properties of its distribution in large samples. I therefore consider also a different transformation of the correlation coefficient first introduced by Fisher (1921) and subsequently improved by Hotelling (1953), which has been extensively used in statistical literature for testing the significance of observed correlations and for setting confidence limits (see Hotelling, 1953). The so-called "Fisher's z-transformation" provides a function of the correlation coefficient which tends to normality very much faster than the correlation coefficient itself and with a standard error almost independent of the population correlation. For these reasons, it seems particularly appropriate for hypothesis testing with very small samples. Appendix A gives the formulae used for the z-transformation and the details about the construction of the tests. Table II presents in the last two columns the outcomes for the 1-tailed tests of the null hypothesis $\rho=0$ against the alternative hypothesis $\rho>0$, at the 1% and 0.5% significance levels. We can see that, also with this method, some of the correlations do not appear statistically significant and that the results obtained with the t-transformation are confirmed by the z-transformation.

To summarize all the previous findings, it seems that Krueger and Summers' assertion about the extreme stability of the U.S. wage structure over time is somewhat overstated. If we take into account the smallness of the sample they use to compute their correlations, conclusions are more ambiguous than what they claim.

Further evidence of the regularities characterizing the inter-industry wage structure is proposed by Krueger and Summers through international wage structure comparisons. The authors suggest that a regular pattern in the wage structure for diverse countries is evidence that some common aspects of labour markets, and not country specific institutions, are responsible for the observed wage differentials. Their correlations of industry wage structures between nations refer to average wages in manufacturing industries for the various countries in 1982 (Krueger and Summers, 1987, Table 2.3). Data are drawn from the International Labor Organization's (ILO) Yearbook of Labor Statistics. The classification of manufacturing industries, the earnings measure and the type of workers covered by these data differ somehow across countries. Krueger and Summers describe in their data appendix the number of industries available for each country (Krueger and Summers, 1987, Table 2.A.1), but they do not specify the exact size of the samples actually used to compute simple correlations. In the analysis that follows, I assumed the sample size to be equal, for each correlation, to the smaller of the two numbers of industry sectors available for the pair of countries involved in the correlation coefficient (see notes c of Table III and of Table A.II in Appendix A). In the first part of Table III, I present a sub-set of the results published by the authors, limiting my attention to the major industrialized Western countries and including the three countries analysed in the present study, Germany, Sweden, and the U.S..

Krueger and Summers argue that these correlations show how the pattern of relative wages is remarkably similar across countries. Their correlations are regarded by them as quite high, ranging from about 0.7 to 0.9. Particularly strong seems the correlation between log average wages at an aggregate industry level in Germany and in the U.S., equal to 0.85, in Germany and Sweden, equal to 0.84, in Sweden and in the U.S., equal to 0.82.

In order to assess rigorously the statistical significance of the correlation coefficients appearing in Table III, I construct hypothesis tests following the approaches previously described for both the t-transformation and the z-transformation of the correlation coefficient. Details for these transformations and the tests are presented again in Appendix A. The results obtained for the 1-tailed tests at the 1% and 0.5% significance levels are reported in the second part of Table III.

In this case we see that the null hypothesis of a population correlation equal to zero can be rejected for all the observed

TABLE III

International wage structure: estimated correlation coefficients[a] and tests of their statistical significance[b] for log average manufacturing wages among countries[c], 1982

	Can	Fra	Jap	US	Ger	UK	Nor	Swe
Canada	1.00	0.85	0.82	0.92	0.83	0.88	0.67	0.79
France		1.00	0.95	0.90	0.87	0.93	0.80	0.84
Japan			1.00	0.89	0.86	0.93	0.80	0.81
US				1.00	0.85	0.95	0.67	0.82
Germany					1.00	0.90	0.74	0.84
UK						1.00	0.70	0.83
Norway							1.00	0.74
Sweden								1.00

Countries	Correlation[a] r	t-transformation Reject H_0: $\rho=0$ at significance level: 1%[d]	.5%[d]	z-transformation Reject H_0: $\rho=0$ at significance level: 1%[d]	.5%[d]
Can,Fra	0.85	yes	yes	yes	yes
Can,Jap	0.82	yes	yes	yes	yes
Can,US	0.92	yes	yes	yes	yes
Can,Ger	0.83	yes	yes	yes	yes
Can,UK	0.88	yes	yes	yes	yes
Can,Nor	0.67	yes	yes	yes	yes
Can,Swe	0.79	yes	yes	yes	yes
Fra,Jap	0.95	yes	yes	yes	yes
Fra,US	0.90	yes	yes	yes	yes
Fra,Ger	0.87	yes	yes	yes	yes
Fra,UK	0.93	yes	yes	yes	yes
Fra,Nor	0.80	yes	yes	yes	yes
Fra,Swe	0.84	yes	yes	yes	yes
Jap,US	0.89	yes	yes	yes	yes
Jap,Ger	0.86	yes	yes	yes	yes
Jap,UK	0.93	yes	yes	yes	yes
Jap,Nor	0.80	yes	yes	yes	yes

(continued)

TABLE III (Continued)

Countries	Correlation[a] r	t-transformation Reject H_0: $\rho=0$ at significance level: 1%[d]	.5%[d]	z-transformation Reject H_0: $\rho=0$ at significance level: 1%[d]	.5%[d]
Jap,Swe	0.81	yes	yes	yes	yes
US,Ger	0.85	yes	yes	yes	yes
US,UK	0.95	yes	yes	yes	yes
US,Nor	0.67	yes	yes	yes	yes
US,Swe	0.82	yes	yes	yes	yes
Ger,UK	0.90	yes	yes	yes	yes
Ger,Nor	0.74	yes	yes	yes	yes
Ger,Sew	0.84	yes	yes	yes	yes
UK,Nor	0.70	yes	yes	yes	yes
UK,Swe	0.83	yes	yes	yes	yes
Nor,Swe	0.74	yes	yes	yes	yes

[a] Source: Krueger and Summers (1987, p.26, Table 2.3).

[b] See Appendix A for details about the construction of the tests with the t-transformation and the z-transformation of the correlation coefficient.

[c] For each country, average manufacturing wages are available for the following numbers of industries (see Krueger and Summers, 1987, p.45, Table 2.A.1):

Country	Number of industries
Canada	21
France	20
Japan	21
US	17
Germany	24
UK	21
Norway	27
Sweden	26

Data for international wage comparisons are reported in the ILO Yearbook of Labor Statistics (1983).

[d] 1-tailed tests of the null hypothesis H_0: $\rho=0$, against the alternative hypothesis H_1: $\rho>0$. The relevant claim is, in fact, that the international wage structure is stable across countries. Only positive correlations are therefore expected.

correlations. However, it must be recalled that I have made a strong assumption about the size of the samples actually considered by Krueger and Summers for their correlation coefficients. This assumption is likely to give an overestimate of the real sample sizes, given the heterogeneity of the criteria used in different countries to classify industry sectors, which may affect the degree of comparability[1]. As I have already remarked, the smallness of the samples used to compute correlation coefficients may critically reduce their precision. For example, if a correlation of 0.67 - as in the case of the U.S. and Norway - were in fact computed with a sample of 12 pairs of observations, rather than 17, it would not be significantly different from zero at the 0.5% significance level[2]. I therefore believe that some caution should be used in drawing conclusions based on this sort of simple statistics.

Krueger and Summers conclude their analysis of the regularities in the inter-industry wage structure asserting:

> "The evidence (...) indicates the presence of pervasive regularities in the wage structure. A similar industrial pattern of wages recurs in different eras and different places (...). Such a uniform pattern ought to be explicable without resort to highly idiosyncratic factors specific to particular workers, industries, times or places. (...) [This] cannot plausibly be rationalized without the introduction of non-competitive considerations or additional constraints (...)." (Krueger and Summers, 1987, p.37).

Two objections can be raised about these remarks. First, as I have tried to suggest in this section, care needs to be employed when interpreting the type of evidence provided by the authors. The accuracy of the simple correlations obtained by them

[1] It seems, for example, that Krueger and Summers (see their Figure 2.2 on p.27, 1987) used a sample of only 13 pairs of observations - instead of the 17 theoretically available - for the correlation between log average manufacturing wages in the U.S. and Japan.

[2] Test based on the t-transformation.

with the smallest samples of data is questionable in several cases. The lack of accuracy may be seen in terms of the large sampling errors associated with their estimates and this casts some doubts about the claimed over-all stability of the wage structure across time and space. Second, empirical evidence based on average aggregate data may indeed be consistent with competitive explanations for inter-industry wage differentials. Before the observed wage structure can be regarded as supportive of the non-competitive theories, plausible competitive rationalizations such as compensating differentials and differences in labour quality and productivity - which could well lead to a stable pattern of wages across time and countries - must be ruled out. A more rigorous attempt to demonstrate the existence and measure the importance of industry wage differentials of an actually non-competitive nature requires a different approach: an empirical analysis based on micro data to test the relevance of industry affiliation in explaining relative wages after controlling for individual human capital characteristics and working conditions. As we will see in the following sections, this technique leads to conclusions which contradict those emerging from aggregate industry wage data.

5. Empirical Evidence for Germany

5.1 Methodology and Model Specification

My empirical approach is to estimate a standard cross-section wage equation in the framework of the earnings function of human capital theory, enriched by demographic and working conditions variables and by industry dummy variables. In order to examine the importance of industry affiliation in explaining relative wages, I want to evaluate the effects of industry dummy variables after controlling for human capital, demographic background, and working conditions as well as possible. Under the hypothesis of a competitive model - if the list of controls is complete - the estimated coefficients of industry dummy variables would not be significantly different from zero. The general structure of the wage regression model is of the following form:

$$\ln w_{ij} = \mathbf{x}_i'\boldsymbol{\beta} + \mathbf{y}_i'\boldsymbol{\gamma} + \mathbf{d}_j'\boldsymbol{\delta} + u_{ij} \tag{1}$$

where w_{ij} is the wage of individual i in industry j, $\mathbf{x}_i$ is a vector of human capital variables for individual i, $\mathbf{y}_i$ is a vector of demographic and working conditions variables for individual i, $\mathbf{d}_j$ is a vector of industry dummy variables for industry j affiliation, and u_{ij} is a random disturbance term assumed to be normally distributed with zero mean and constant variance σ_u^2. The regression parameters $\boldsymbol{\beta}$, $\boldsymbol{\gamma}$ and $\boldsymbol{\delta}$ are estimated with the OLS method.

The existence of statistically significant industry effects in a wage regression like (1), however, is not a definite proof in support of non-competitive theories of wage determination. Unmeasured labour quality differences - such as ability and motivation - which might vary systematically across industries and unmeasured differences in industry specific working conditions which necessitate compensating wage differentials may indeed induce biased estimates of the coefficients of the industry dummy variables. This may lead to an overestimate of the pure industry affiliation effect in explaining the observed wage structure. It is therefore crucial to incorporate the whole information available from the data in the set of control variables included in the wage regression, in the attempt to - at least - minimize the bias due to omitted variables.

5.2 Data, Sample Characteristics, and Construction of the Variables

The empirical analysis of industry wage differentials in Germany is based on individual cross-sectional data from the 1984 wave of the German Socioeconomic Panel (SOEP) provided by the Deutsches Institut fuer Wirtschaftsforschung (Berlin). The panel is derived from a nationwide representative survey using a sample of about 5,000 randomly selected households for each year from 1984 onwards. The survey relies on two different types of questionnaire: the first collects information about the household as a whole; the second, which is the relevant one for the purpose of my study, is addressed to each individual household member 16 years old or older and contains detailed questions about participation in gainful employment. Both Germans and foreigners living in Germany are represented in the survey[3].

The wage variable I want to use in my regression analysis is a measure of standard hourly earnings for salaried workers regularly employed in agricultural, industrial and service firms, both in the private and in the public sector. The measure of earnings available in the data set is not independent from the hours of work, which may vary across individuals, firms, and sectors. Usual hourly earnings provide therefore a measure of the wage rate which is more properly comparable across industries. Moreover, I consider only regularly salaried workers, since the focus of the present study is on the behaviour of employers and their employees in the process of wage determination. I also decided to concentrate on a sample of male workers only, to avoid the problem of self selection connected with female labour supply.

The initial sample of all male individuals 16 years old or older contains 6,007 observations. The sub-sample I analyse is composed of German and foreign male employees selected according to the following criteria:

i) individuals younger than 65 years, the normal age of retirement in Germany - this selection reduces the sample to

[3] The number of foreigners included in the survey sample is over-representative of the actual proportion of foreigners in the population. In the sub-sample of workers used in the present study, foreigners are about 39% of the whole sub-sample. This is due to the purpose of the SOEP of providing a specific sample of foreigners which is by itself representative of the entire universe of foreign residents in Germany. In order to take this peculiarity of the survey into account, I construct, for my regression analysis, special variables apt to capture at least a part of the distinctive aspects of foreign workers human capital - such as the variables for modelling foreign education and, more generally, dummy variables for foreign nationalities.

5,498 observations (about 92% of the initial sample);

ii) full-time or regular part-time private and public employees, excluding individuals not in the labour force, unemployed and irregularly employed workers - this selection reduces the sample to 4,141 observations (about 69% of the initial sample, with a reduction of -25% with respect to the previous sub-sample);

iii) blue- and white-collar workers, excluding professional men, self-employed workers, trainees and civil servants qualified as "Beamten", who being public officials are subject to peculiar regulations affecting their position in the labour market (e.g. clerical officers, judges, career military personnel) - this selection reduces the sample to 3,368 observations (about 56% of the initial sample, with a reduction of -19% with respect to the previous sub-sample);

iv) employees working a fixed number of hours per week, for whom the measure of usual hourly earnings is more exactly determined and less subject to measurement errors - this selection reduces the sample to 3,240 observations (about 54% of the initial sample, with a reduction of -4% with respect to the previous sub-sample).

For the empirical construction of the wage rate variable, a measure of earnings and a measure of hours of work are required. The relevant information on earnings given in the questionnaire refers to gross earnings in the month preceding that of the interview, excluding special payments - such as vacation bonuses or back pay - but including pay for overtime. Two different measures of hours worked are given: hours actually worked weekly, on average, in the month preceding that of the interview, including overtime, and normal hours worked weekly, excluding overtime. Both measures include transitory components, such as sick time or vacation days, as normal work time. There is a possible problem in computing standard hourly earnings that arises due to the non-linearity of total earnings as a function of hours actually worked when overtime work is done[4]. I

[4] Even if a constant wage rate is earned over all hours normally worked each week, this need not imply that the wage rate obtained by working overtime hours equals the individual's

therefore consider only a sub-sample of straight-time employees, defined as individuals who did not work overtime hours in the relevant period[5]. The wage rate variable is thus constructed as gross earnings in the last month divided by four times the normal hours worked weekly, where neither the earnings measure nor the hours of work measure are affected by overtime work. I first eliminate from the sample all the observations for which information on earnings and/or hours of work are missing. This reduces the sample to 2,945 observations (about 49% of the initial sample, with a reduction of -9% with respect to the previous sub-sample). Then I eliminate individuals who report normal hourly earnings less than 1 DM, considered to be outliers. This reduces the sample by only 1 observation. The subsequent selection of the sub-sample of straight-time employees leads to a final sample size of 2,072 observations (about 34% of the initial sample, with a reduction of -30% with respect to the previous sub-sample).

The exclusion of overtime workers from the sub-sample, however, may raise two problems related to the possible non-randomness of the selected sample. First, it might be the case that overtime workers are not proportionally distributed across industries. A higher wage rate might be systematically associated with more compensated overtime work[6] or, on the contrary, the usual wage rate might be relatively lower for individuals who do work overtime[7]. In both cases, the exclusion

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average wage rate. It may therefore be that additional hours worked overtime would change the marginal and the average wage rate. Unfortunately, the SOEP does not provide information on this issue, but only a measure of total earnings including payments for overtime work.

[5] There is an additional reason to prefer an analysis of the straight-time sample. Earnings due to overtime work and especially hours actually worked weekly on average in the last month - including overtime hours - are likely to be seriously affected by measurement errors.

[6] This implies a positive wage elasticity of overtime labour supply in face of short-run adjustments on the demand side.

[7] Some indirect evidence of this negative wage elasticity of overtime labour supply is provided by Borjas (1980).

of employees doing overtime work may disproportionately reduce the number of observations only in certain sectors - only in high wage sectors or low wage sectors respectively - thus reducing the precision of the respective estimated industry differentials. Second, such segmentation of the sample raises a serious risk of selection bias. The sample selection rule that determines the availability of data for the wage regression - only a sub-sample of straight-time workers is considered - may have critical consequences on the estimated coefficients. If observations are not excluded randomly, the wage function estimated on the selected sample confounds the parameters of interest with the parameters of the function determining the probability of entrance into the sample - i.e., the probability of not working overtime hours in the relevant period (Heckman, 1979). Under these circumstances, the OLS method gives biased estimates of the parameters of the wage equation and, among these, of the parameters relative to the industry dummy variables, which may lead to an incorrect evaluation of the relevance of industry affiliation in explaining relative wages.

A way to address the first problem is to construct a chi-square test to compare the distribution of employees across industries for the selected sub-sample with the distribution of employees across industries for a sample including both straight-time and overtime employees. The test gives the following result[8]:

$$\chi^2(35) = 16.996 < \chi^2_{.01}(35) \approx 57.3$$

We cannot therefore reject the null hypothesis of identical distributions and this can be interpreted as a signal that the overtime workers excluded from the sub-sample are proportionally

[8] The classification of industry sectors considered for this chi-square test is the original one provided by the SOEP, which consists of 36 different industries. In my later regression analysis I use a different classification, an aggregation of the original into 26 sectors only.

distributed across industry sectors.

As far as the second problem is concerned, given the possibly serious effects arising from a sample selection bias, I adopt a rigorous treatment using the technique suggested by Heckman (1979). Heckman's two-stage estimator is consistent even when non-randomly selected samples are used to estimate behavioural relationships with simple regression methods, like least squares. The model that I in fact consider is not simply represented by equation (1), but by the following two equation system with a selection criterion equation:

$$\ln w_{ij} = \mathbf{x}_i'\boldsymbol{\beta} + \mathbf{y}_i'\boldsymbol{\gamma} + \mathbf{d}_j'\boldsymbol{\delta} + u_{ij} \tag{1}$$

$$ot_i^* = \mathbf{v}_i'\boldsymbol{\alpha} + e_i \tag{2}$$

where equation (1) is defined as above, while ot_i^* is an unobserved index variable indicating the desired amount of overtime work done in the relevant period by individual i, $\mathbf{v}_i$ is a vector of explanatory variables for the overtime work of individual i, and e_i is a random disturbance term normally distributed with zero mean and constant variance σ_e^2. The joint distribution of u_{ij} and e_i is a bivariate normal and their covariance is σ_{ue}. Observations on the wage rate w_{ij} are included in the sub-sample if $ot_i^*=0$, while if $ot_i^*>0$ they are excluded.

In the case of independence between u_{ij} and e_i ($\sigma_{ue}=0$), so that the observations on w_{ij} would be randomly excluded from the sub-sample, least squares estimators might be used to estimate $\boldsymbol{\beta}$, $\boldsymbol{\gamma}$ and $\boldsymbol{\delta}$ on the selected sample and the only cost of having an incomplete sample would be a loss in efficiency. But in the case of overtime work as a selection criterion, the probability of including observations on w_{ij} in the sub-sample may vary with its value, or with the values of variables affecting w_{ij}; then the probability of observing w_{ij} will depend on u_{ij}, σ_{ue} will be different from zero and the sub-sample wage regression function will depend not only on $\mathbf{x}_i$, $\mathbf{y}_i$, and $\mathbf{d}_j$, but also on $\mathbf{v}_i$. Under these circumstances, least squares estimators of $\boldsymbol{\beta}$, $\boldsymbol{\gamma}$ and $\boldsymbol{\delta}$ in equation (1) estimated on the selected sample will be biased, as in an ordinary problem of omitted variables (Heckman, 1979)[9]. An important source of selection bias is represented by the omission of variables in $\mathbf{v}_i$ not contained in $\mathbf{x}_i$, $\mathbf{y}_i$, or $\mathbf{d}_j$ - variables affecting the probability of observing w_{ij}, but not w_{ij} directly - but correlated with these included variables. A symptom of selection bias is in fact that variables that do not belong to the true structural wage equation - variables in $\mathbf{v}_i$ not in $\mathbf{x}_i$, $\mathbf{y}_i$, or $\mathbf{d}_j$ - may appear to be statistically significant determinants of w_{ij} when they are incorrectly included as regressors in the wage equation and the wage regression is fitted on the selected sample (Heckman, 1979, p.155).

The selection rule implies that observations are excluded from the sub-sample if any overtime work is done, independently of its amount. In the context of my sample selection model, I am therefore interested in the probability that employees work any positive amount of overtime hours. I then build, for ot_i^*, the counterpart binary variable ot_i according to the following criterion:

$$ot_i = 1 \text{ if } ot_i^* > 0$$
$$ot_i = 0 \text{ if } ot_i^* \leq 0$$

which substituted in equation (2) gives the binary probit selection equation:

$$ot_i = \mathbf{v}_i' \boldsymbol{\alpha} + e_i \tag{2'}$$

Individuals then enter the sub-sample used to estimate equation

[9] Moreover, if $\sigma_{ue} \neq 0$, the usual formulae for standard errors of least squares coefficients are not appropriate: they understate the true standard errors and overstate estimated significance levels (Heckman, 1979, pp.157-58).

(1) when ot_i is equal to 0, while they are eliminated when ot_i is equal to 1.

The solution proposed by Heckman (1979) consists of the following three steps:

i) apply probit analysis to equation (2′) for the full sample, to estimate the parameters of the probability that $ot_i=1$ - the probability that any amount of overtime work is done - i.e., to estimate $\boldsymbol{\alpha}/\sigma_e$;

ii) for each observation, estimate the Heckman's $\boldsymbol{\lambda}$ in the form which is appropriate to the case of selection on the value 0 for ot_i:

$$\boldsymbol{\lambda}_i = \frac{-\phi\left(\dfrac{v_i'\boldsymbol{\alpha}}{\sigma_e}\right)}{1-\Phi\left(\dfrac{v_i'\boldsymbol{\alpha}}{\sigma_e}\right)}$$

using the probit estimated coefficients for $\boldsymbol{\alpha}/\sigma_e$; all of these estimators are consistent;

iii) estimate equation (1) with OLS for the selected sub-sample, regressing $\ln w_{ij}$ on $\mathbf{x}_i$, $\mathbf{y}_i$, $\mathbf{d}_j$, and the estimated value of $\boldsymbol{\lambda}_i$; regression estimators of equation (1) are consistent for $\boldsymbol{\beta}$, $\boldsymbol{\gamma}$, $\boldsymbol{\delta}$, and σ_{ue}/σ_e - the coefficients of $\mathbf{x}_i$, $\mathbf{y}_i$, $\mathbf{d}_j$, and $\hat{\boldsymbol{\lambda}}_i$ respectively.

The sample selection bias introduced by eliminating employees working overtime hours is significant only if the coefficient for $\hat{\boldsymbol{\lambda}}_i$ in the wage regression (σ_{ue}/σ_e) is significantly different from zero, since this implies a significant covariance ($\sigma_{ue}\neq 0$) between the wage regression and the selection equation disturbances u_{ij} and e_i.

As far as the choice of explanatory variables in equations (1) and (2′) is concerned, the set of human capital, demographic background, and working conditions controls used in the wage equation ($\mathbf{x}_i$ and $\mathbf{y}_i$) includes: age, age squared, tenure in the current job (years), tenure squared, 5 dummies for German education, 5 dummies for foreign education, 9 skill dummies, the number of nights spent in a hospital in the previous year as a measure of health conditions, 4 marital status dummies (married, living with spouse; married, permanently separated; divorced; widowed), the degree of satisfaction with the current job in a scale from 1 to 10[10], and 3 dummy variables for the size of the firm of current employment. The vector of variables that permits evaluation of the relevance of industry affiliation in explaining relative wages ($\mathbf{d}_j$) includes 25 industry dummies.

The set of explanatory variables used in the overtime work probability equation ($\mathbf{v}_i$) consists both of variables not included in the wage equation - assumed to affect the probability of working overtime hours but not affecting the wage rate directly - and of variables also included in the wage equation - assumed to influence simultaneously the probability of working overtime hours and the wage rate. The sub-set of non-overlapping variables is: the number of children under the age of 16 years living in the household, a dummy variable for a second house/apartment in the Federal Republic of Germany, a dummy variable for mortgages on the house/apartment which is the main residence of the household, and 5 nationality dummies. The sub-set of overlapping variables includes: age, age squared, 9 skill dummies, the number of nights spent in a hospital in the previous year, the degree of satisfaction with the current job, 3 dummy variables for the size of the firm of current employment, and the 25 industry dummies.

Other variables are used both in the wage equation and in the overtime work probability equation to deal with the problem of missing values. Instead of excluding observations with missing values in any of the explanatory variables - that would reduce further on the sample size (for example, missing values for industry affiliation are 191, 9.2% of the observations in the selected sub-sample) - I prefer to introduce a separate dummy variable for missing data about education, tenure, marital status, nationality, and industry affiliation. Dummy variables

[10] On the use of job satisfaction as an economic variable in labour market analysis, see Freeman (1978).

for missing values of education, tenure, and marital status are later eliminated from the model because their estimated coefficients are not significantly different from zero - that is, the effect of education, tenure, and marital status for individuals with missing values is not statistically different from the effect of the same variables for individuals whose characteristics define the base group of each dummy variable - and because their omission do not affect the other estimated coefficients.

Two different specifications of the sample selection model expressed by equations (1) and (2') are estimated: the first is the general model, which includes both control variables ($\mathbf{x}_i$ and $\mathbf{y}_i$) and industry dummies ($\mathbf{d}_j$) in the wage equation (1); the second is a restricted model, which involves only industry dummy variables ($\mathbf{d}_j$) in the wage equation (1). In both specifications, equation (2') has the same form. With respect to the selection bias problem, the estimates give the following results: the coefficient for $\boldsymbol{\lambda}_i$ in the wage equation fit on the selected sample for the general model is not significantly different from zero (estimated σ_{ue}/σ_e is -0.070, with a t-statistic of -0.90), indicating that I fail to reject the null hypothesis of no sample selection bias induced by the exclusion of employees working overtime hours; the coefficient for $\boldsymbol{\lambda}_i$ in the wage equation for the restricted model is significantly different from zero at the 1% level (estimated σ_{ue}/σ_e is -0.164, with a t-statistic of -2.59), which implies that, in this case, I reject the null hypothesis of no sample selection bias. This last outcome is not unexpected: in the restricted model I exclude variables from the wage equation (the controls $\mathbf{x}_i$ and $\mathbf{y}_i$) which, as we will see, are statistically significant determinants of the wage rate and many of these controls (the sub-set of overlapping variables in $\mathbf{v}_i$) enter as explanatory variables the overtime work probability equation; $\boldsymbol{\lambda}_i$, as a function of $\mathbf{v}_i$, proxies the effect on wages of these controls and hence its coefficient appears statistically significant in the wage equation merely because of their exclusion from the equation. Also in the case of the restricted model, however, Heckman's two-stage method guarantees consistent estimates of the parameters of equation (1). The detailed results for the estimated probit overtime work equation and for the estimated wage equations, including $\hat{\lambda}_i$ among the regressors, are presented in Appendix B.

5.3 Basic Results

In Table IV I report the results of cross-section estimates of inter-industry wage differentials in a sample selection model for an aggregation of three-digit industries according to the German industry classification, which is nearly comparable with the two-digit classification used by the other authors. The dependent variable is the logarithm of usual hourly earnings in the month of reference. As suggested by Krueger and Summers (1988), the estimated industry wage differentials are reported as deviations from the employment-weighted mean differential; that is, I calculate the employment-weighted average of wage differentials for all industries from the wage regression - treating the omitted industry variable as having zero effect on wages - and report the difference between the industry differentials and the weighted average differential. The resulting statistics therefore are the proportionate difference in wages between an employee in a given industry and the average employee in the whole economy. Again following Krueger and Summers (1988, p.267), to summarize the overall variability in industry wages I present the employment-weighted adjusted standard deviation of the industry wage differentials[11].

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[11] The adjustment is required because the estimated industry wage differentials include a least squares sampling error, which leads to an upward bias in the standard deviation of wage differentials (Krueger and Summers, 1988, p.267). The standard deviation is adjusted using the formula:

$$SD(\hat{\delta}) \approx \sqrt{var(\hat{\delta}) - \sum_{j=1}^{K} \hat{\sigma}_j^2 / K}$$

where $\hat{\delta}_j$ is the estimate of the true wage differential δ_j of industry j=1,...,K, and $\hat{\sigma}_j$ is the standard error of $\hat{\delta}_j$. Like

<u>TABLE IV</u>

Estimated wage differentials in a sample selection model for two-digit industries, 1984: deviations from the employment-weighted mean differential (unadjusted OLS t-statistics in parentheses)

Industry	I Without controls	II With controls[a]
1. Energy, Water and Mining	.131	.113
	(5.32)	(3.66)
2. Chemical	.136	.085
	(5.78)	(3.39)
3. Rubber	-.095	-.005
	(2.42)	(1.65)
4. Stone, Clay and Glass	-.009	.037
	(3.48)	(2.37)
5. Iron and Steel	-.053	.021
	(3.53)	(2.47)
6. Machinery, Excl. Elec.	.049	.061
	(4.92)	(3.21)
7. Electrical Machinery	.062	.030
	(4.86)	(2.50)
8. Lumber, Wood, Paper and Printing	-.092	.038
	(2.71)	(2.60)
9. Textile and Apparel	-.046	.014
	(3.20)	(2.09)
10. Food, Beverages and Tobacco	-.132	-.101
	(2.15)	(.14)
11. Construction	-.035	.044
	(3.77)	(2.91)
12. Wholesale Trade	.021	-.078
	(3.43)	(.44)
13. Retail Trade	-.186	-.104
	(1.48)	(.09)
14. Railroads	-.072	-.070
	(2.22)	(.50)
15. Mail Service	-.014	.011
	(2.68)	(1.51)
16. Other Transport and Communications	.002	-.011
	(3.79)	(1.67)
17. Banking	.078	-.094
	(4.12)	(.22)
18. Insurance	.386	.163
	(6.28)	(3.37)
19. Personal Services	-.524	-.323
	(-2.36)	(-3.20)

(continued)

TABLE IV (Continued)

Industry	I Without controls	II With controls[a]
20. Entertainment	.237	-.016
	(6.22)	(1.45)
21. Health Services	.037	-.112
	(3.77)	(-.05)
22. Legal and Business Services	.277	.064
	(5.42)	(2.19)
23. Non-profit Organizations and Private Households	.092	-.061
	(4.25)	(.70)
24. Local Collective Organizations	.023	-.047
	(4.23)	(1.08)
25. Social Security	.095	-.059
	(3.30)	(.57)
Weighted Adjusted Standard Deviation of Differentials[b]	.146	.072
F-statistics for No Industry Effect	9.768**	5.618**
Adjusted R^2	0.100	0.515
Sample Size	2,072	2,072

[a] Controls include age and its square, tenure and its square, five German education dummies, five foreign education dummies, nine skill dummies, four marital status dummies, a measure of health conditions, degree of satisfaction, and three firm-size dummies.

[b] Weights are employment shares for each industry.

** F test that industry wage differentials jointly equal 0 rejects at the 1% level. The 1% critical points are $F_{.01}(26,2044) \approx 1.76$ for the regression without control variables and $F_{.01}(26,2012) \approx 1.76$ for the regression with control variables. The number of restrictions (degrees of freedom for the numerator) refers to the coefficients of 25 dummy variables for industrial sectors and the coefficient of a dummy variable for missing industry affiliation.

The results presented in Table IV are obtained using as weights employment shares by industry derived from national census statistics for the whole economy (Statistisches Bundesamt, 1984). I also tried as weights employment shares by industry as resulting from within the sample used for my regression analysis. With these two sets of weights, I obtained only minor differences in the levels of industry wage differentials as deviations from the employment-weighted mean differential[12] and nearly identical employment-weighted adjusted standard deviations of industry wage differentials[13]. This seems to confirm that the sample used in the present study is representative of the underlying population in terms of distribution of employees across industries.

Column I of Table IV presents raw wage differentials, that is industry wage differentials estimated without controlling for human capital, demographic and working conditions. The industry dummy variables are jointly statistically significant (the appropriate F-statistic is 9.768, significant at the 1% level being the critical 1% point $F_{.01}(26,2044) \approx 1.76$). Industry dummies are also statistically significant individually in 24 of the 25 cases at the 5% level and in 20 cases also at the 1% level. Estimates of the industry wage differentials range from -52 percent in the personal services sector, to +39 percent in the insurance sector. The employment-weighted adjusted standard deviation of raw industry wage differentials is 14.6 percent. This result is consistent with the findings of Krueger and Summers (1987, Table 2.4), where inter-country comparisons based

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Krueger and Summers, I neglect the covariance terms.

[12] For the differentials of column I, the use of sample employment shares as weights gives wage differentials that exceed those obtained with population employment shares by 0.007; for the differentials of column II, the use of population employment shares gives wage differentials that exceed those obtained with sample employment shares by 0.010.

[13] For the differentials of column I, the use of population and sample employment shares as weights gives employment-weighted adjusted standard deviations equal to 0.146 and 0.147 respectively; for the differentials of column II, the employment-weighted adjusted standard deviations are 0.072 and 0.075 respectively.

on industry aggregate data show for Germany a standard deviation of log average earnings of 14.1 percent in 1982[14].

Column II of Table IV presents estimated industry wage differentials when control variables for human capital and working conditions are introduced in the wage equation (see Appendix B, Table B.II for the parameter estimates of the control variables in the wage regression). With respect to raw industry differentials, 17 of the 25 differentials decrease in absolute size, with a mean relative reduction of about 48%; however, the other 8 differentials exhibit a considerable increase in absolute size, so that the overall mean relative change due to the introduction of controls is a growth of differentials of about 24%. In 14 of the 25 cases the differentials change in sign and only 5 of these remain significant in the wage regression with the controls. Estimates of industry differentials range from -32 percent in the personal services sector, to +16 percent in the insurance sector, the same sectors as in the case of raw differentials. It seems therefore that the addition of controls alters to a certain extent the pattern of wage differences. Industry dummies are statistically significant in 12 out of 25 industries at the 5% level and in 7 cases also at the 1% level. They are also jointly significant at the 1% level (the appropriate F-statistic is 5.618, whereas the critical 1% point is $F_{.01}(26,2012) \approx 1.76$). Controlling for worker characteristics also reduces significantly the estimated inter-industry wage dispersion. The employment-weighted adjusted standard deviation of industry wage differentials falls to 7.2 percent.

We can observe that, differently from the U.S. case

[14] Krueger and Summers (1987, Table 2.4) provide the standard deviation of log average earnings among 24 manufacturing industries computed with aggregate annual data by industry. The earnings measure is earnings per hour, where earnings include wages and all wage supplements. Data are derived from the ILO Yearbook of Labor Statistics (1983). Their standard deviation of log earnings based on aggregate data - as a measure of wage dispersion - is comparable with my employment-weighted adjusted standard deviation of raw earnings differentials estimated in a regression based on individual data, where the dependent variable is log hourly earnings and the set of regressors includes only industry dummies.

illustrated by Krueger and Summers (1987 and 1988), raw industry wage differentials are not a very satisfactory approximation of the differentials obtained when control variables are introduced. This can be seen in Figure II, where the plot of wage differentials estimated with controls against wage differentials estimated without controls does not show a very strong positive linear relationship. The Pearson correlation between the estimated wage differentials in column I and II of Table IV - a measure used by Krueger and Summers (1987, p.19) to claim the stability of the pattern of industry wages with respect to controls[15] - is 0.73, significantly different from zero at the 0.005% level. An alternative way to verify whether controls have an impact on the ranking of industry wage differences is to compute the Spearman rank correlation coefficient. For the German case, the rank correlation of the differentials estimated with and without controls is 0.45, which is not significantly different from zero at the 1% level[16]. Differences in observed labour quality and working conditions seem to explain a considerable part of the variability of wages among industries. When human capital and working conditions controls are introduced in the wage regression, the standard error of the regression is reduced by 30 percent (from 0.338 to 0.237), the adjusted R^2 increases from 10 percent to 52 percent and the employment-weighted adjusted standard deviation of differentials is reduced from 15 to 7 percent.

The general conclusion seems to be that although the size, significance, and dispersion of inter-industry wage differentials may cast some doubts on the standard competitive model of the

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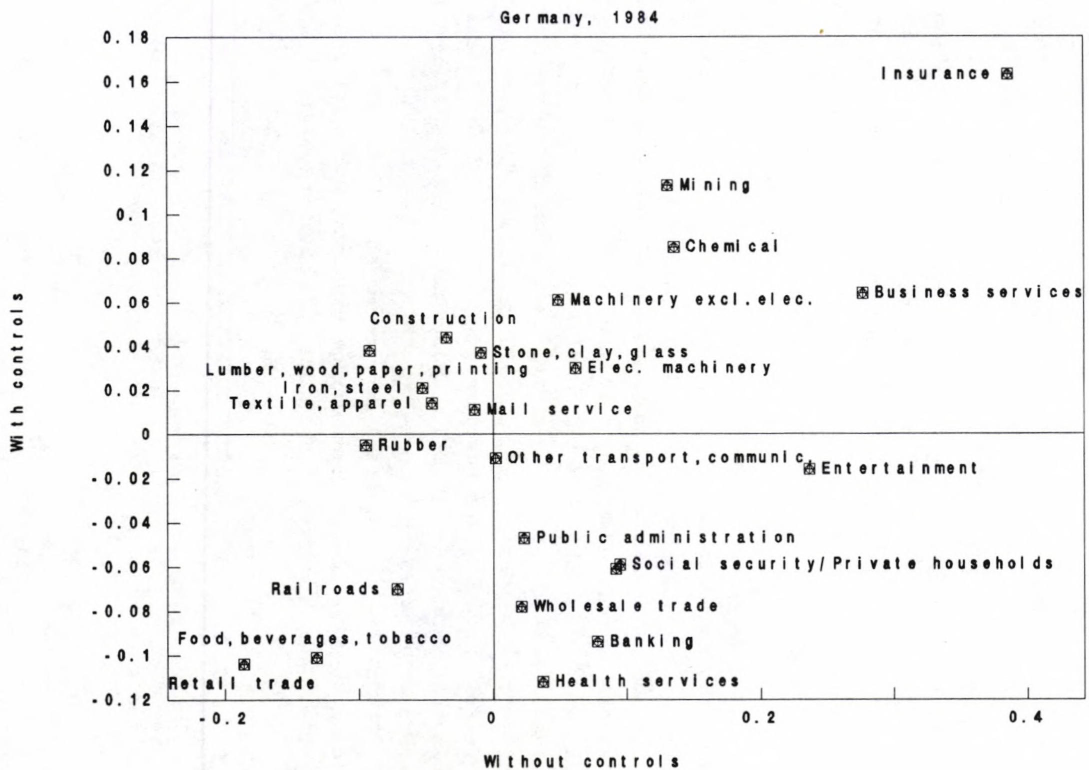
[15] "It is clear that the addition of these controls barely alters the ranking of industry wage differences. Indeed the correlation of the industry wage differentials estimated with and without controls is 0.95." (Krueger and Summers, 1987, p.19). I verified that this correlation is significantly different from zero at the 0.005% level.

[16] Using the results by Krueger and Summers (1987, Table 2.1), I computed the rank correlation for the U.S. case. This is also equal to 0.95, significantly different from zero at the 0.005% level. The claim of no impact of controls on the pattern of industry wages is therefore more justifiable in the U.S. case.

FIGURE II Estimated wage differentials



labour market, human capital and working conditions factors play a crucial role in explaining the observed wage structure in Germany.

Two important caveats should be also taken into account in evaluating my findings. On the one hand, estimated industry differentials may appear smaller than the true differentials of non-competitive nature because of the inclusion of firm size variables among the controls in the wage equation. Both efficiency wage and insider-outsider theories, in fact, predict a positive relationship between firm size and wages: in the context of the efficiency wage model, turnover and monitoring costs may be higher in larger than in smaller firms and thus the efficiency wage may increase with firm size (Salop, 1973); in the context of the insider-outsider model, labour is expected to be better organized in large firms and thus insiders may be able to obtain a larger profit share (Weiss, 1966). There is some evidence for Germany supporting non-competitive explanations of the observed firm size-wage effect (Schmidt and Zimmermann, 1990). Whatever the underlying theoretical reason, firm size variables may "pick-up" some aspects of a non-competitive process of wage determination and therefore reduce the estimated industry affiliation effect.

On the other hand, estimated industry differentials may appear larger than the true differentials of non-competitive nature because of unobservable labour characteristics. In an approach based on the estimate of an earnings function like equation (1), unobservable characteristics which vary systematically across industries may produce upward biased estimates of the coefficients of the industry dummy variables, thus overstating the actual importance of industry affiliation in explaining the structure of wages. This problem seems even more serious in the German case than in the U.S. case, since here observable characteristics do have a substantial impact on relative wages[17].

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[17] Krueger and Summers (1987) find for the U.S. case that controlling for observable characteristics of workers does not change the pattern of wage differences and argue: "Unless

Inter-country comparisons of the sort presented in section 4 to claim stability of the wage structure rely on the hypothesis that average wage differentials as emerging from aggregate data are a good approximation of the wage differentials that would result when all compensating differentials for labour quality and working conditions are controlled for. As I have already mentioned, there is some evidence supporting this hypothesis for the U.S. case. This leads Krueger and Summers (1987) to conclude:

> "The finding (...) allows for the comparison of industry wages over time and across countries with aggregate industry wage data since it is unlikely that controls would change the pattern of industry wages in these data." (Krueger and Summers, 1987, p.20).

However, as we have seen in the present section, this hypothesis does not seem a realistic one for the German case. Even in an approach that cannot exclude compensating differentials for unobservable human capital characteristics and working conditions, observable labour quality controls explain a large amount of the variability of wages across industries and modify the pattern of inter-industry wage differentials.

In the following section, I will therefore consider inter-country comparisons based on results obtained with micro data and contrast the emerging conclusions with those drawn from evidence based on aggregate industry wage data.

6. Comparisons with the U.S. and Swedish Evidence Based on Micro Data

In this section I will compare my results based on micro data for Germany with those of analogous work for the U.S. (Krueger and Summers, 1987 and 1988) and for Sweden (Edin and

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unmeasured aspects of labor quality are only weakly correlated with tenure, age and education, and are far more important than measurable aspects, it is hard to see how they could account for inter-industry wage differences." (Krueger and Summers, 1987, p.38).

Zetterberg, 1989). When moving from aggregate to individual level data, we will see that differences between countries, rather than similarities, tend to emerge.

The three empirical studies use a similar approach: a wage equation - the logarithm of usual hourly earnings as a function of controls for human capital and working conditions and of industry dummies - estimated from a cross-sectional regression on individual data for 1984 with the OLS method. However, the degree of comparability between countries is affected by several differences in the definition of the samples of interest and of the dependent and explanatory variables, as well as in the statistical methodology applied for the estimate of the model.

The U.S. data are derived from the Current Population Survey (CPS), collected by the Bureau of Census for May 1984. CPS contains labour force data for members of the sampled households who are 14 years old or older. The sub-sample used by Krueger and Summers (1987 and 1988) in their regression analysis consists of full and part-time private non-agricultural employees 16 years old or older. Both male and female workers are included. Employees who report hourly earnings smaller than $1.00 or greater than $250.00 are considered to be outliers and eliminated from the sample. The authors obtain in this way a sub-sample of 11,512 individuals from a nationwide representative sample (10,289 individuals for the estimates of raw industry differentials in Krueger and Summers, 1987). The dependent earnings variable is defined as usual weekly earnings in the relevant month (May 1984) divided by usual weekly hours of work in the same time period.

Data for Sweden are obtained by Edin and Zetterberg (1989) from the HUS-Survey for 1984, which contains labour force and work place related data for members of about 1,500 households. A sub-sample is selected for full and part-time employees in public and private sectors. Both male and female workers are included. Observations with missing values in any of the dependent or explanatory variables are eliminated from the sample. A sub-sample of 1,298 individuals is thus obtained from a nationwide representative sample. The dependent variable is the log of the hourly wage rate, calculated as usual weekly (or other time units) earnings divided by usual hours of work for the corresponding time unit.

The other authors do not provide more information about the procedure of sample selection and the construction of the dependent variable. In particular, they do not specify anything about the treatment of overtime workers and ignore the possible consequences in terms of selection bias of any of the selection criteria. Moreover, the choice of including both male and female employees in the selected samples and a dummy variable for sex among the control variables does not seem particularly appropriate, because it neglects the problem of self selection connected with female labour supply.

With respect to the sets of control variables[18], the major differences between the German and the U.S. studies are represented by the inclusion among the controls for the U.S. case of a dummy variable for union membership and by the exclusion of firm size variables and of any other control for working conditions. Union membership and firm size variables are both likely to "pick-up" non-competitive influences on industry wages. With respect to the Swedish case, the main difference consists in a richer set of working conditions and work-place related controls, likely to "pick-up" both competitive and non-competitive aspects of the process of wage determination. Other minor differences in the sets of control variables mainly reflect peculiar institutional conditions characterizing the

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[18] The set of explanatory variables for the U.S. case includes: education and its square, 6 age dummies, 8 occupation dummies, 3 region dummies, sex, race, central city, union membership, marital status, veteran status dummies, several interaction terms, and 42 industry dummies.

The set of explanatory variables used in the Swedish case includes: education (years of schooling), experience and its square, tenure, age, sex, white-collar, native language not Swedish dummies, plant size, logarithm of regional unemployment rate, 6 shift dummies (3-shift, 2-shift, working weekends, working nights, irregular shifts, other shifts), 4 wage-form dummies (individual/group/mixed piece-rates, other piece-rates), and 26 industry dummies.

For the analysis of the German case, I refer to the list of controls presented in section 5.2.

labour markets of the three countries.

Finally, as already noticed, no treatment for sample selection bias is contemplated in the U.S. and Swedish studies: the simple OLS method is used to provide estimates for the wage regressions on the selected samples.

Taking into proper account the limit to the degree of comparability arising from all these differences - the most serious probably being in the use of a sub-sample of male workers only for my analysis of the German case - I will proceed with inter-country comparisons considering some aspects of the empirical evidence available from the different studies. The estimates of raw industry differentials for the three countries are summarized in Table V, while Table VI presents industry wage differentials estimated with controls for human capital and working conditions. A further limit in the actual comparison is represented by the differences in the industry sectors classifications used in the U.S., Germany, and Sweden. I therefore preferred to consider, for each country, a sub-set of industries which are fairly similar according to the definition of each sector, even at the cost of oversimplifying the comparison to a certain extent.

A first relevant difference between Germany and the U.S. and between Germany and Sweden is the magnitude of industry wage differentials. German evidence shows industry differentials which, in absolute size, are smaller than the differentials in the U.S. and larger than those in Sweden in the majority of industries, both in terms of raw differentials and when control variables are introduced in the wage regression. Germany therefore represents an intermediate case between the U.S. and Sweden.

The three countries also differ in terms of statistical significance of individual industry dummies. For the U.S., industry dummy variables remain generally statistically significant when human capital and working conditions controls are introduced in the wage regression. For Germany, almost a half of industry dummies (12 out of 25) are statistically significant at the 5% level in the regression including other control

<u>TABLE V</u>

Estimated wage differentials <u>without</u> controls for human capital and working conditions for Germany, U.S., and Sweden, 1984: deviations from the employment-weighted mean differential (unadjusted OLS standard errors in parentheses)

Industry	Germany	U.S.[a]	Sweden[b]
1. Mining	.131	.404	.036
	(.082)	(.043)	(.086)
2. Chemical	.136	.362	-.004
	(.076)	(.041)	(.064)
3. Rubber	-.095	.038	——
	(.087)	(.051)	——
4. Stone, Clay, Glass	-.009	.357	-.009
	(.085)	(.061)	(.094)
5. Iron, Steel	-.053	.357	.017
	(.071)	(.048)	(.064)
6. Machinery, Excl. Elec.	.049	.335	——
	(.072)	(.028)	——
7. Electrical Machinery	.062	.185	——
	(.076)	(.030)	——
9. Textile, Apparel	-.046	——	-.231
	(.081)	——	(.066)
10. Food, Bever., Tobacco	-.132	——	.034
	(.080)	——	(.061)
11. Construction	-.035	.216	.066
	(.072)	(.024)	(.039)
12. Wholesale Trade	.021	.171	.073
	(.095)	(.026)	(.047)
13. Retail Trade	-.186	——	-.097
	(.080)	——	(.033)
17. Banking	.078	.084	.129
	(.093)	(.026)	(.060)
18. Insurance	.386	.105	.048
	(.110)	(.026)	(.059)
19. Personal Services	-.524	-.329	——
	(.093)	(.030)	——
20. Entertainment	.237	-.181	-.042
	(.087)	(.043)	(.064)
21. Health Services	.037	-.183	-.076
	(.091)	(.026)	(.090)
22. Business Services	.277	.027	.181
	(.107)	(.027)	(.041)
23. Private Households	.092	-.776	——
	(.093)	(.038)	——
24. Public Administration	.023	——	.048
	(.078)	——	(.036)

(continued)

TABLE V (Continued)

Industry	Germany	U.S.[a]	Sweden[b]
25. Social Security	.095 (.121)	-.194 (.032)	-.024 (.015)
Weighted Adjusted Standard Deviation of Differentials[c]	.146	.240	.071
F-statistic for No Industry Effect	9.547**	N.A.	4.05**
Sample Size	2,072	10,289	1,298

[a] Source: Krueger and Summers (1987, pp.20-21, Table 2.1).
[b] Source: Edin and Zetterberg (1989, pp.8-9, Table 1).
[c] Weights are employment shares for each industry.
** F test that industry wage differentials jointly equal 0 rejects at the 1% level.

TABLE VI

Estimated wage differentials with controls for human capital and working conditions for Germany, U.S., and Sweden, 1984: deviations from the employment-weighted mean differential (unadjusted OLS standard errors in parentheses)

Industry	Germany	U.S.[a]	Sweden[b]
1. Mining	.113	.241	.024
	(.061)	(.033)	(.071)
2. Chemical	.085	.221	.054
	(.057)	(.033)	(.053)
3. Rubber	-.005	.054	—
	(.063)	(.041)	—
4. Stone, Clay, Glass	.037	.085	.008
	(.062)	(.044)	(.077)
5. Iron, Steel	.021	.162	.011
	(.053)	(.037)	(.056)
6. Machinery, Excl. Elec.	.061	.185	—
	(.053)	(.024)	—
7. Electrical Machinery	.030	.107	—
	(.056)	(.025)	—
9. Textile, Apparel	.041	—	-.074
	(.059)	—	(.056)
10. Food, Bever., Tobacco	-.101	—	.020
	(.057)	—	(.051)
11. Construction	.044	.126	.069
	(.053)	(.020)	(.034)
12. Wholesale Trade	-.078	.047	.044
	(.070)	(.020)	(.039)
13. Retail Trade	-.104	—	-.054
	(.056)	—	(.028)
17. Banking	-.094	.064	.028
	(.068)	(.022)	(.048)
18. Insurance	.163	.071	.038
	(.081)	(.021)	(.048)
19. Personal Services	-.323	-.154	—
	(.067)	(.025)	—
20. Entertainment	-.016	-.141	-.127
	(.064)	(.034)	(.053)
21. Health Services	-.112	-.082	-.002
	(.060)	(.023)	(.074)
22. Business Services	.064	.000	.099
	(.079)	(.023)	(.034)
23. Private Households	-.061	-.366	—
	(.069)	(.033)	—
24. Public Administration	-.047	—	.028
	(.057)	—	(.030)

(continued)

TABLE VI (Continued)

Industry	Germany	U.S.[a]	Sweden[b]
25. Social Security	-.059 (.088)	-.246 (.027)	-.030 (.059)
Weighted Adjusted Standard Deviation of Differentials[c]	.072	.140	.012[d]
F-statistic for No Industry Effect	5.618**	N.A.**	1.86**
Sample Size	2,072	11,512	1,298

[a] Source: Krueger and Summers (1988, pp.265-266, Table II).
[b] Source: Edin and Zetterberg (1989, pp.8-9, Table 1).
[c] Weights are employment shares for each industry.
[d] Not computable with the formula in footnote 9, since the variance of estimated residuals is less than the average standard error. This result is obtained accounting for covariance terms, which are elsewhere neglected.
** F test that industry wage differentials jointly equal 0 rejects at the 1% level.

variables. For Sweden, only three individual dummies are significant, even at the 5% level, when a large number of individually related and work-place related controls are introduced in the wage regression. However, they remain jointly statistically significant. It is worth noting that differences in the level of significance may simply reflect differences in the sample sizes. The standard error of least squares coefficients is in fact an increasing function of the inherent variability of the dependent variable and a decreasing function of the sample size and of the variability of each explanatory variable. Thus, other things being equal, a larger sample size leads to more accurate estimates and higher significance levels. The size of the sample used in the U.S. study (11,512) is about 6 times the sample size in my analysis of the German case (2,072) and about 9 times that of the Swedish study (1,298).

The most striking difference among Germany, the U.S. and Sweden is in terms of the variability of wages across industries as measured by the employment-weighted adjusted standard deviation of industry differentials. Referring to the regression where a large number of controls are added, the figures for the three countries are respectively 7 percent, 14 percent and approximately zero. Germany again is between two extreme situations.

A further difference is the relative importance of human capital variables and industry variables in the wage equations for the three countries. In the U.S. industry variables are very important in explaining variations in individual wages. The standard error of the regression is reduced by 4.3 percent when industry variables are added to a regression that already controls for occupation, human capital, and demographic factors. In comparison, the standard error falls by 5.1 percent when human capital controls are added to the same regression. For Germany, when industry variables are introduced into a regression that already controls for a number of human capital and working conditions the standard error of the regression is reduced by 3.9 percent, while when human capital controls are added to the same regression the standard error falls by 4.5 percent. A totally different result is instead obtained for Sweden, where the reduction in the standard error of the regression due to the introduction of industry variables is 0.2 percent, while that due to human capital controls is 2.2 percent, more than ten times the reduction due to industry variables. This means that in the U.S. and in Germany industry variables and human capital controls are of nearly equal importance in explaining wage variations, with Germany exhibiting a slightly smaller impact of industry variables than the U.S., while in Sweden human capital variables are relatively much more important, with industry variables having almost no effect in the estimated wage equation.

Finally, differences emerge also in the comparison of industry wage structures. To get a first idea of the degree of stability of the industry wage pattern across countries, Figures III, IV, V, and VI present plots of wage differentials in the U.S and Germany and in Sweden and Germany, estimated both without and with controls for human capital and working conditions. The plots do not show the existence of strong positive relationships.

I have also computed the Pearson product-moment correlations and the Spearman rank correlations for the industry wage differentials appearing in Tables V and VI. For Germany and the U.S., the Pearson correlation between raw differentials using 17 industries is 0.14 (one-sided p-value 0.29) and the rank correlation is -0.03 (one-sided p-value 0.54); for Germany and Sweden, the Pearson correlation between raw differentials using 16 industries is 0.39 (one-sided p-value 0.07) and the rank correlation is 0.31 (one-sided p-value 0.12). None of these correlations are significantly different from zero at the 1% level. When wage differentials estimated with controls are considered, I obtain opposite results with respect to those presented by Edin and Zetterberg (1989), that is correlation coefficients tend to rise. For Germany and the U.S., the Pearson correlation is 0.59 (one-sided p-value 0.007) and the rank correlation is 0.70 (one-sided p-value 0.001); for Germany and Sweden, the Pearson correlation is 0.32 (one-sided p-value 0.11) and the rank correlation is 0.46 (one-sided p-value 0.04).

Without overvaluing the reliability of such comparisons, it

Figure III Estimated wage differentials

Without controls, U.S and Germany

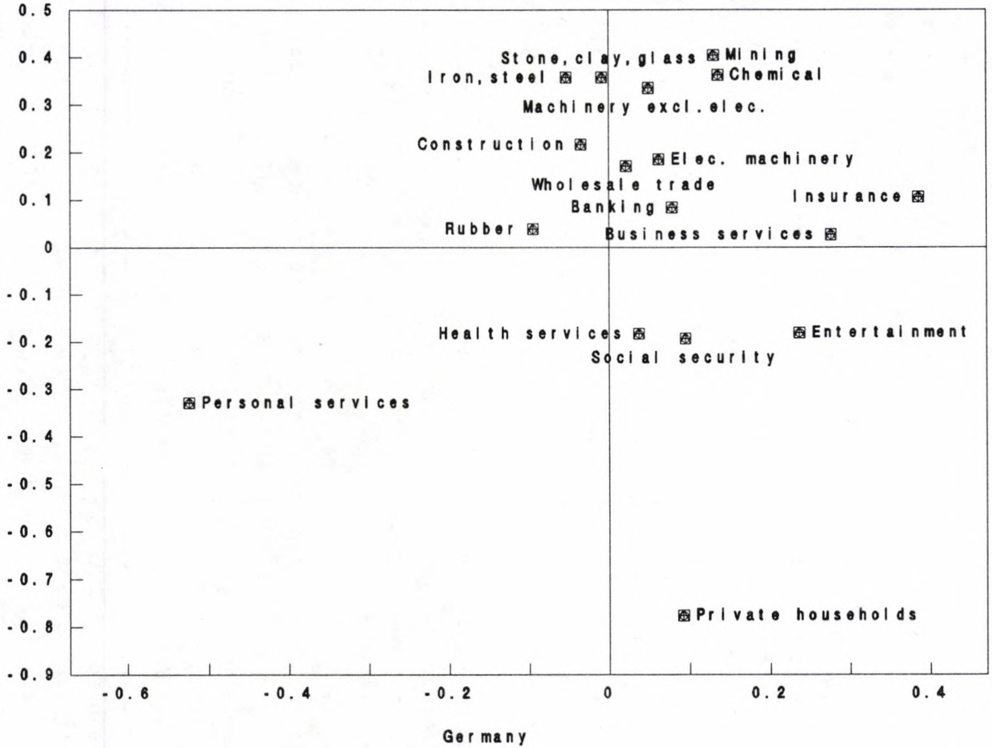


Figure IV Estimated wage differentials

Without controls, Sweden and Germany

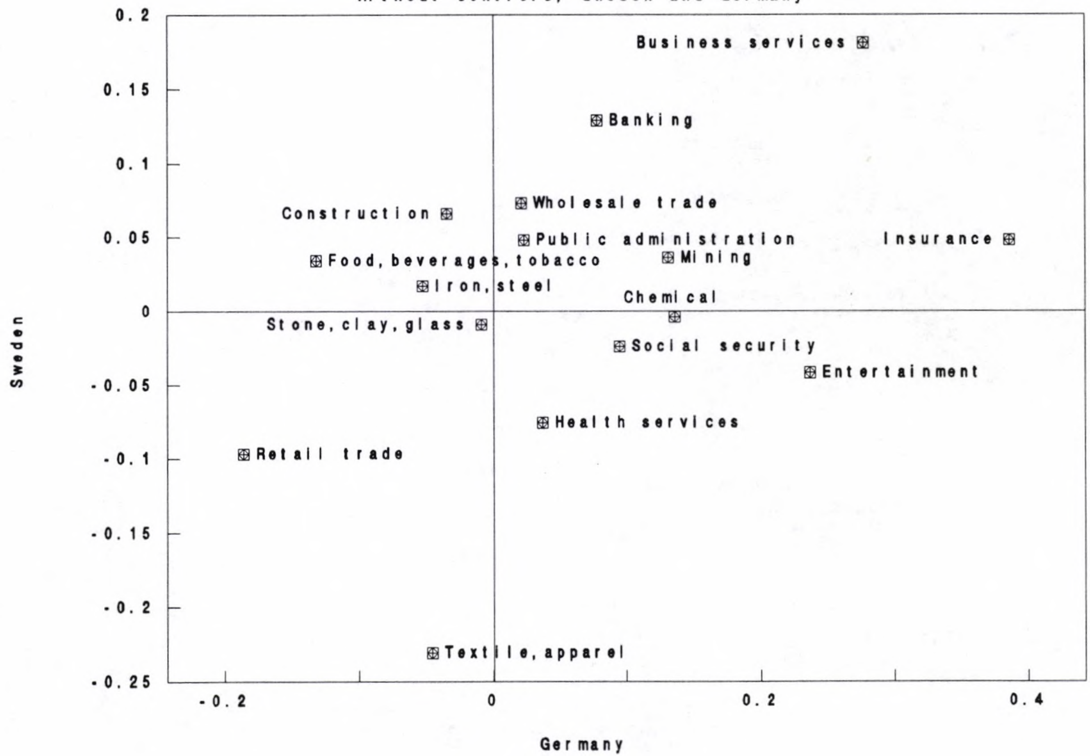


Figure V Estimated wage differentials

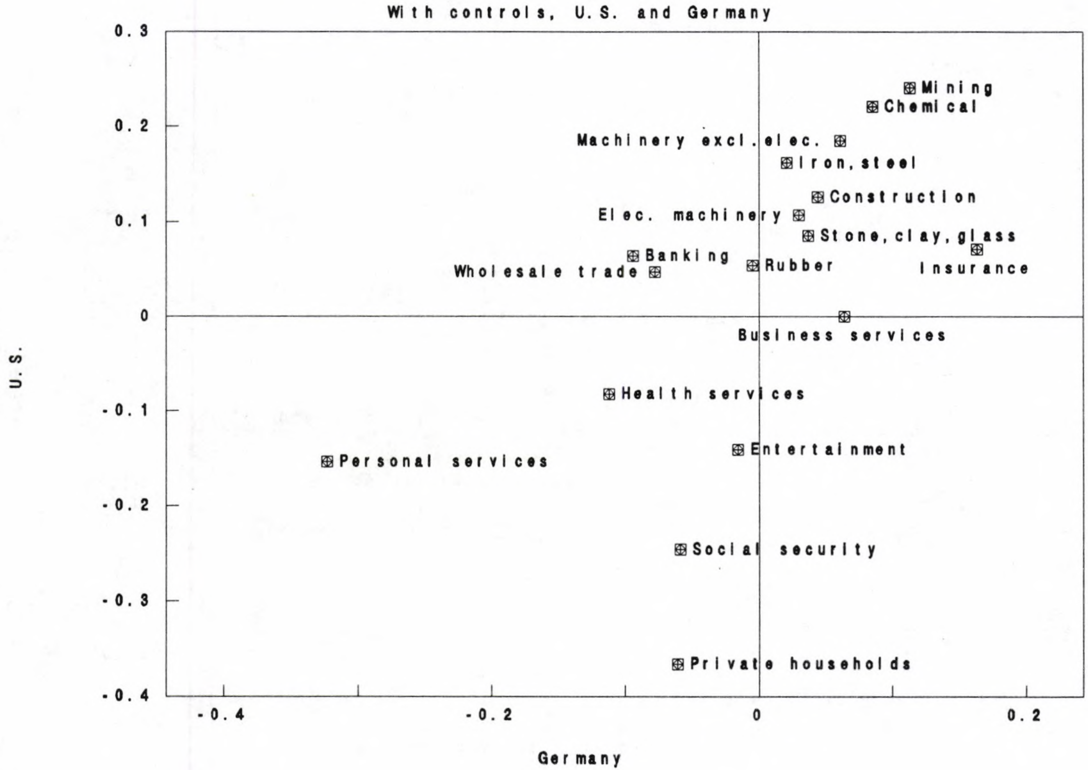
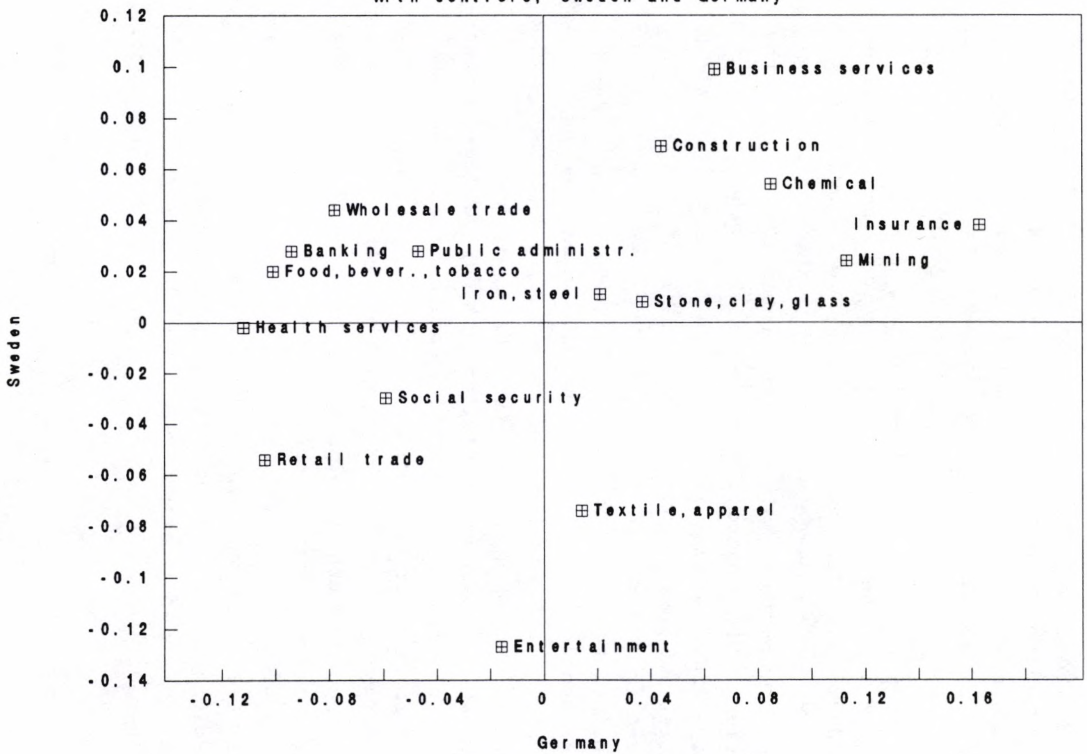


Figure VI   Estimated wage differentials

With controls, Sweden and Germany



is worth noting that we are very far from the results obtained by Krueger and Summers (1987) with aggregate data, which show correlations between log average industry wages equal to 0.85 and 0.84 for Germany and the U.S. and for Germany and Sweden respectively (see Table III in section 4). This indicates that the magnitude of correlations between countries may be overstated in aggregate data and that a substantial proportion of similarities may be due to the correlation of the industry distribution of observable and unobservable labour quality and job attributes across countries.

Within the limits of the actual comparability of empirical evidence for Germany, the U.S. and Sweden, the fact that differences between countries, rather than similarities, tend to emerge seems to suggest that institutional aspects of the labour market, and in particular the degree of centralization of wage bargaining, may play an important role in explaining the observed pattern of inter-industry wage differentials. The ranking of the three countries in terms of degree of centralization is consistent with the empirical results for industry wage differences estimated with individual data in the framework of human capital earnings functions. Union policies aiming at a reduction of wage differentials have been successful in Sweden and partly in Germany. On the other hand, the lack of centralized policies in wage negotiations appears to lead the U.S. towards a labour market characterized by wide inter-sectoral differentials. These findings are confirmed by additional evidence provided by Holmlund and Zetterberg (1991). Using panel data on aggregate industry wages for five countries including Germany, the U.S. and Sweden, they find that industry wages in the U.S. are substantially affected by industry specific conditions, while in Sweden these effects are negligible. Germany again plays an intermediate role, with substantial industry wage differentials but modest wage response to sectoral conditions.

7. Conclusions

The aims of this paper have been to present some empirical evidence based on micro data of inter-industry wage dispersion for Germany, to compare my results with those provided in two similar studies on the U.S. and Sweden, and to attempt an explanation of my findings in terms of competitive and non-competitive theories of the labour market.

The main conclusions may be summarized as follows. First, evidence for Germany shows that workers' quality and other compensating factors have an important impact on the observed wage structure, although the size, significance, and dispersion of inter-industry wage differentials cast some doubts on the standard competitive model of the labour market.

Second, comparisons with U.S. and Swedish evidence suggest that results obtained with individual data in a regression approach highlight differences among countries rather than similarities, in contrast to what emerges with aggregate data.

Third, in the class of non-competitive theories of the labour market, institutional conditions of wage bargaining and in particular the degree of centralization seem to play a relevant role in explaining the pattern of inter-industry wage differences.

These conclusions are obviously affected by differences in the methodology used in the various studies and by the fact that only three countries are considered. More evidence for other countries obtained with similar approaches would be very useful for generalizing these results.

APPENDIX A

Testing the Statistical Significance of the Correlation Coefficient in Small Samples

A variety of approaches have been suggested in the literature to approximate the properties of the correlation coefficient in samples from a bivariate normal population, since it has not a well-behaved sampling distribution (see, for example, Kendall and Stuart, 1977).

Hotelling (1953, p.212) proposes the following expression for the standard error of the sample correlation coefficient in finite samples:

$$\sigma_r = \frac{1-\rho^2}{\sqrt{n-1}}\left(1 + \frac{11\rho^2}{4(n-1)} + \frac{-192\rho^2+479\rho^4}{32(n-1)^2} + O(n^{-3})\right)$$

where ρ is the population correlation coefficient and n is the sample size. The terms of order n^{-3} and higher order become negligibly small even for a very small sample size and can therefore be ignored.

Following Atkinson et al. (1990, p.87), from Hotelling's formula I derive an expression for the estimated standard error of the correlation coefficient s_r, replacing the population ρ by its sample estimator r:

$$s_r = \frac{1-r^2}{\sqrt{n-1}}\left(1 + \frac{11r^2}{4(n-1)} + \frac{-192r^2+479r^4}{32(n-1)^2}\right).$$

Given asymptotic normality of r (Hotelling, 1953), I also construct two-sided 99% confidence intervals for the correlation coefficients as:

$$r - 2.58\,s_r < \rho < r + 2.58\,s_r.$$

Since r is a consistent estimator of ρ, it follows that s_r is also a consistent estimator of σ_r, which means that the derivation of s_r from Hotelling's expression is valid, but only asymptotically. For small samples - like in my case, where n=9 - the use of this formula to compute s_r may entail a considerable amount of bias. Moreover, the distribution of the correlation coefficient approaches normality very slowly for large samples (Kendall and Stuart, 1977, p.416). This may invalidate the construction of confidence intervals and hypothesis testing based on the estimator s_r and on the normal distribution for small samples.

The results obtained applying this expression for s_r and using it to construct two-sided 99% confidence intervals for the correlation coefficients under the assumption of normality are presented in Table I and graphed in Figure I of section 4.

Kendall and Stuart (1977) suggest an alternative technique, which is more appropriate for testing if a correlation coefficient estimated with a small sample is significantly different from zero, since it does not involve the bias deriving from a direct estimate of the sampling error s_r. This is based on the transformation of the distribution function of the sample correlation coefficient into a "Student's t-distribution". Under the particular null hypothesis $\rho=0$, the distribution of the correlation coefficient may be obtained indirectly by putting:

$$t = \{(n-2)r^2/(1-r^2)\}^{\frac{1}{2}}$$

which follows a t-distribution with ν=n-2 degrees of freedom (Kendall and Stuart, 1977, p.416, §16.28). This t-transformation of the correlation coefficient is used to construct 1-tailed tests of the null hypothesis $\rho=0$ against the alternative hypothesis $\rho>0$.

Tables A.I and A.II present, together with the respective estimated correlations, the computed values of t and the outcomes for hypothesis testing at the 1% and 0.5% significance levels. Table A.I refers to correlations for the analysis of the stability of the U.S. wage structure over time and Table A.II

TABLE A.I

Industry wage structure through time in the U.S.: estimated correlation coefficients[a] and t-transformation[b] for log average annual earnings of full-time equivalent employees in nine major industries[c]

Year	Correlation with 1984[a] r	Degrees of freedom[d]	t^b	Reject H_0: $\rho=0$ at significance level: 1%[e]	.5%[e]
1984	1.000	7	∞	yes	yes
1980	0.984	7	14.612	yes	yes
1975	0.961	7	9.194	yes	yes
1970	0.909	7	5.770	yes	yes
1965	0.898	7	5.400	yes	yes
1960	0.893	7	5.250	yes	yes
1955	0.893	7	5.250	yes	yes
1950	0.866	7	4.582	yes	yes
1945	0.891	7	5.192	yes	yes
1940	0.836	7	4.031	yes	yes
1935	0.793	7	3.444	yes	no
1930	0.761	7	3.104	yes	no
1925	0.801	7	3.540	yes	yes
1920	0.807	7	3.615	yes	yes
1915	0.627	7	2.129	no	no
1910	0.604	7	2.005	no	no
1905	0.636	7	2.181	no	no
1900	0.616	7	2.069	no	no

[a] Source: Krueger and Summers (1987, p.24, Table 2.2).
[b] See Kendall and Stuart (1977, p.416, §16.28):

$$t = \{(n-2)r^2/(1-r^2)\}^{1/2}.$$

[c] The sample size for the estimate of the correlation coefficients is n=9. Industries include: agriculture, manufacturing, mining, construction, transportation, communications, wholesale and retail trade, FIRE (finance, insurance and real estate), and services.
[d] Degrees of freedom: $\nu = n-2$.
[e] 1-tailed tests of the null hypothesis H_0: $\rho=0$, against the alternative hypothesis H_1: $\rho>0$. The relevant claim is, in fact, that the industry wage structure is stable over time. Only positive correlations are therefore expected. The 1% critical point is $t_{.01}(7)=3.00$ and the .5% critical point is $t_{.005}(7)=3.50$.

TABLE A.II

International wage structure: estimated correlation coefficients[a] and t-transformation[b] for log average manufacturing wages among countries[c], 1982

Countries	Correlation[a] r	Degrees of freedom[d]	t[b]	Reject H_0: $\rho=0$ at significance level: 1%[e]	.5%[e]
Can,Fra	0.85	18	6.846	yes	yes
Can,Jap	0.82	19	6.245	yes	yes
Can,US	0.92	15	9.092	yes	yes
Can,Ger	0.83	19	6.486	yes	yes
Can,UK	0.88	19	8.076	yes	yes
Can,Nor	0.67	19	3.934	yes	yes
Can,Swe	0.79	19	5.617	yes	yes
Fra,Jap	0.95	18	12.908	yes	yes
Fra,US	0.90	15	7.997	yes	yes
Fra,Ger	0.87	18	7.486	yes	yes
Fra,UK	0.93	18	10.735	yes	yes
Fra,Nor	0.80	18	5.657	yes	yes
Fra,Swe	0.84	18	6.568	yes	yes
Jap,US	0.89	15	7.560	yes	yes
Jap,Ger	0.86	19	7.346	yes	yes
Jap,UK	0.93	19	11.029	yes	yes
Jap,Nor	0.80	19	5.812	yes	yes
Jap,Swe	0.81	19	6.021	yes	yes
US,Ger	0.85	15	6.249	yes	yes
US,UK	0.95	15	11.783	yes	yes
US,Nor	0.67	15	3.495	yes	yes
US,Swe	0.82	15	5.549	yes	yes
Ger,UK	0.90	19	9.000	yes	yes
Ger,Nor	0.74	22	5.160	yes	yes
Ger,Swe	0.84	22	7.261	yes	yes
UK,Nor	0.70	19	4.273	yes	yes
UK,Swe	0.83	19	6.486	yes	yes
Nor,Swe	0.74	24	5.390	yes	yes

[a] Source: Krueger and Summers (1987, p.26, Table 2.3).
[b] See Kendall and Stuart (1977, p.416, §16.28):

$$t = \{(n-2)r^2/(1-r^2)\}^{1/2}.$$

(continued)

(Continued)

[c] For each pair of countries, the sample size for the estimate of the correlation coefficient is assumed to be (see Krueger and Summers, 1987, p.45, Table 2.A.1):

Countries	n	Countries	n	Countries	n
Can,Fra	20	Fra,UK	20	US,Nor	17
Can,Jap	21	Fra,Nor	20	US,Swe	17
Can,US	17	Fra,Swe	20	Ger,UK	21
Can,Ger	21	Jap,US	17	Ger,Nor	24
Can,UK	21	Jap,Ger	21	Ger,Swe	24
Can,Nor	21	Jap,UK	21	UK,Nor	21
Can,Swe	21	Jap,Nor	21	UK,Swe	21
Fra,Jap	20	Jap,Swe	21	Nor,Swe	26
Fra,US	17	US,Ger	17		
Fra,Ger	20	US,UK	17		

Data for international wage comparisons are reported in the ILO Yearbook of Labor Statistics (1983).
[d] Degrees of freedom: ν = n-2.
[e] 1-tailed tests of the null hypothesis H_0: $\rho=0$, against the alternative hypothesis H_1: $\rho>0$. The relevant claim is, in fact, that the international wage structure is stable across countries. Only positive correlations are therefore expected. The 1% critical points for the various degrees of freedom are:

$$t_{.01}(15)=2.60$$
$$t_{.01}(18)=2.55$$
$$t_{.01}(19)=2.54$$
$$t_{.01}(22)=2.51$$
$$t_{.01}(24)=2.49$$

and the .5% critical points for the various degrees of freedom are:

$$t_{.005}(15)=2.95$$
$$t_{.005}(18)=2.88$$
$$t_{.005}(19)=2.86$$
$$t_{.005}(22)=2.82$$
$$t_{.005}(24)=2.80.$$

refers to correlations for international wage structure comparisons. The results obtained with this method are those reported in Table II and III of section 4 in the columns for the t-transformation.

A third approach to the problem of testing the significance of estimated correlations is based on another transformation of r, introduced by Fisher (1921):

$$z = \frac{1}{2} \log \frac{1+r}{1-r}.$$

The "Fisher's z-transformation" provides a function of the correlation coefficient having a distribution which approaches normality with great rapidity and a variance nearly independent of the population correlation ρ. This makes the transformation to z particularly useful in testing the significance of the deviation of observed correlations from hypothetical values and in constructing confidence intervals.

The normalizing and variance-stabilizing properties of z, however, are asymptotic ones and its use in very small samples may entail considerable errors (Hotelling, 1953)[1]. I therefore adopt the correction for bias in z proposed by Hotelling (1953, p.219) and compute the modified formula:

$$z' = z - \frac{r}{2n-5}.$$

An improved approximation for an unbiased estimate of the variance of z is also provided (Hotelling, 1953, p.220):

[1] In particular, for the variance of z Fisher recommends the formula:

$$\sigma_z^2 \approx \frac{1}{n-3}.$$

This simple formula is an approximation and its closeness depends on the actual value of ρ and on n (see Hotelling, 1953, p.218).

$$s_z^{2\prime} = \frac{1}{n - 3 + \frac{1}{2} r^2}.$$

Under the assumption of normality of z', I can construct one-sided 99% and 99.5% confidence intervals as, respectively:

$$\zeta > z' - 2.33\, s_z'$$
$$\zeta > z' - 2.58\, s_z'$$

where ζ is the population parameter corresponding to z'. These confidence intervals for z' are then transformed into the corresponding one-sided 99% and 99.5% confidence intervals for the correlation coefficient r, according to the inverse formula:

$$r = \tanh z'.$$

This is easily done using the special table of the transformation of z' to r provided by Fisher and Yates (1963, p.63). The intervals for r thus obtained are always within the range [-1,+1]. This means that the method of the z-transformation gives, for small samples, a better approximation to the sampling distribution of r than the one based on the direct estimate of its standard error s_r and on the assumption of its asymptotic normality - the first approach illustrated in this appendix. There we have, in fact, confidence intervals for r exceeding the upper limit value of +1, as shown in Table I of section 4.

Confidence intervals for r are finally used for the 1-tailed tests of the null hypothesis $\rho=0$ against the alternative hypothesis $\rho>0$, at the 1% and 0.5% significance levels.

Table A.III and A.IV present the values of z', s_z', the confidence intervals for z' and r and the outcomes for hypothesis testing at the 1% and 0.5% levels, for the wage structure estimated correlations over time and across countries respectively. The final results of the tests with the z-transformation are those reported in Table II and III of section 4.

TABLE A.III

Industry wage structure through time in the U.S.: estimated correlation coefficients[a] and z-transformation[b] for log average annual earnings of full-time equivalent employees in nine major industries[c]

Year	Correlation with 1984[a] r	z'[b]	$s_{z'}$[b]	99% one-sided confidence interval for z'	99% one-sided confidence interval for r	Reject H_0: $\rho=0$ at 1% significance level[d]
1984	1.000	∞	0.392	∞ - ∞	1.000 - 1.000	yes
1980	0.984	2.334	0.393	1.42 - ∞	0.890 - 1.000	yes
1975	0.961	1.885	0.393	0.97 - ∞	0.749 - 1.000	yes
1970	0.909	1.452	0.395	0.53 - ∞	0.485 - 1.000	yes
1965	0.898	1.393	0.395	0.47 - ∞	0.438 - 1.000	yes
1960	0.893	1.368	0.395	0.45 - ∞	0.422 - 1.000	yes
1955	0.893	1.368	0.395	0.45 - ∞	0.422 - 1.000	yes
1950	0.866	1.250	0.396	0.33 - ∞	0.319 - 1.000	yes
1945	0.891	1.358	0.395	0.44 - ∞	0.414 - 1.000	yes
1940	0.836	1.143	0.397	0.22 - ∞	0.217 - 1.000	yes
1935	0.793	1.018	0.398	0.09 - ∞	0.090 - 1.000	yes
1930	0.761	0.940	0.399	0.01 - ∞	0.010 - 1.000	yes
1925	0.801	1.040	0.398	0.11 - ∞	0.110 - 1.000	yes
1920	0.807	1.056	0.398	0.13 - ∞	0.129 - 1.000	yes
1915	0.627	0.688	0.402	-0.25 - ∞	-0.245 - 1.000	no
1910	0.604	0.653	0.402	-0.28 - ∞	-0.273 - 1.000	no
1905	0.636	0.703	0.402	-0.23 - ∞	-0.226 - 1.000	no
1900	0.616	0.671	0.402	-0.27 - ∞	-0.264 - 1.000	no

(continued)

TABLE A.III (Continued)

Year	Correlation with 1984[a] r	z'[b]	s_{z}'[b]	99.5% one-sided confidence interval for z'	99.5% one-sided confidence interval for r	Reject H_0: $\rho=0$ at .5% significance level[d]
1984	1.000	∞	0.392	∞ - ∞	1.000 - 1.000	yes
1980	0.984	2.334	0.393	1.32 - ∞	0.867 - 1.000	yes
1975	0.961	1.885	0.393	0.87 - ∞	0.701 - 1.000	yes
1970	0.909	1.452	0.395	0.43 - ∞	0.405 - 1.000	yes
1965	0.898	1.393	0.395	0.37 - ∞	0.354 - 1.000	yes
1960	0.893	1.368	0.395	0.35 - ∞	0.336 - 1.000	yes
1955	0.893	1.368	0.395	0.35 - ∞	0.336 - 1.000	yes
1950	0.866	1.250	0.396	0.23 - ∞	0.226 - 1.000	yes
1945	0.891	1.358	0.395	0.34 - ∞	0.328 - 1.000	yes
1940	0.836	1.143	0.397	0.12 - ∞	0.119 - 1.000	yes
1935	0.793	1.018	0.398	-0.01 - ∞	-0.010 - 1.000	no
1930	0.761	0.940	0.399	-0.09 - ∞	-0.090 - 1.000	no
1925	0.801	1.040	0.398	0.01 - ∞	0.010 - 1.000	yes
1920	0.807	1.056	0.398	0.03 - ∞	0.030 - 1.000	yes
1915	0.627	0.688	0.402	-0.35 - ∞	-0.336 - 1.000	no
1910	0.604	0.653	0.402	-0.38 - ∞	-0.363 - 1.000	no
1905	0.636	0.703	0.402	-0.33 - ∞	-0.319 - 1.000	no
1900	0.616	0.671	0.402	-0.37 - ∞	-0.354 - 1.000	no

[a] Source: Krueger and Summers (1987, p.24, Table 2.2).
[b] See Hotelling (1953, §9, p.217-221).
[c] See note c to Table A.I.
[d] 1-tailed tests of the null hypothesis H_0: $\rho=0$, against the alternative hypothesis H_1: $\rho>0$. The relevant claim is, in fact, that the industry wage structure is stable over time. Only positive correlations are therefore expected.

TABLE A.IV

International wage structure: estimated correlation coefficients[a] and z-transformation[b] for log average manufacturing wages among countries[c], 1982

Countries	Correlation[a] r	z'[b]	s_{z}'[b]	99% one-sided confidence interval for z'	99% one-sided confidence interval for r	Reject H_0: $\rho=0$ at 1% significance level[d]
Can,Fra	0.85	1.232	0.240	0.67 - ∞	0.585 - 1.000	yes
Can,Jap	0.82	1.135	0.234	0.59 - ∞	0.530 - 1.000	yes
Can,US	0.92	1.557	0.263	0.94 - ∞	0.735 - 1.000	yes
Can,Ger	0.83	1.166	0.233	0.62 - ∞	0.551 - 1.000	yes
Can,UK	0.88	1.352	0.233	0.81 - ∞	0.670 - 1.000	yes
Can,Nor	0.67	0.793	0.234	0.25 - ∞	0.245 - 1.000	yes
Can,Swe	0.79	1.050	0.234	0.51 - ∞	0.470 - 1.000	yes
Fra,Jap	0.95	1.805	0.239	1.25 - ∞	0.848 - 1.000	yes
Fra,US	0.90	1.441	0.263	0.83 - ∞	0.681 - 1.000	yes
Fra,Ger	0.87	1.308	0.240	0.75 - ∞	0.635 - 1.000	yes
Fra,UK	0.93	1.632	0.240	1.07 - ∞	0.790 - 1.000	yes
Fra,Nor	0.80	1.076	0.240	0.52 - ∞	0.478 - 1.000	yes
Fra,Swe	0.84	1.197	0.240	0.64 - ∞	0.565 - 1.000	yes
Jap,US	0.89	1.391	0.264	0.78 - ∞	0.653 - 1.000	yes
Jap,Ger	0.86	1.270	0.233	0.73 - ∞	0.623 - 1.000	yes
Jap,UK	0.93	1.633	0.233	1.09 - ∞	0.797 - 1.000	yes
Jap,Nor	0.80	1.077	0.234	0.53 - ∞	0.485 - 1.000	yes
Jap,Swe	0.81	1.105	0.234	0.56 - ∞	0.508 - 1.000	yes
US,Ger	0.85	1.227	0.264	0.61 - ∞	0.544 - 1.000	yes
US,UK	0.95	1.799	0.263	1.19 - ∞	0.831 - 1.000	yes
US,Nor	0.67	0.788	0.265	0.17 - ∞	0.168 - 1.000	yes
US,Swe	0.82	1.129	0.264	0.51 - ∞	0.470 - 1.000	yes
Ger,UK	0.90	1.448	0.233	0.90 - ∞	0.716 - 1.000	yes
Ger,Nor	0.74	0.933	0.217	0.43 - ∞	0.405 - 1.000	yes
Ger,Swe	0.84	1.202	0.216	0.70 - ∞	0.604 - 1.000	yes
UK,Nor	0.70	0.848	0.234	0.30 - ∞	0.291 - 1.000	yes
UK,Swe	0.83	1.166	0.233	0.62 - ∞	0.551 - 1.000	yes
Nor,Swe	0.74	0.935	0.207	0.45 - ∞	0.422 - 1.000	yes

(continued)

TABLE A.IV (Continued)

Countries	Correlation[a] r	z'[b]	s_{z}'[b]	99.5% one-sided confidence interval for z'	99.5% one-sided confidence interval for r	Reject H_0: $\rho=0$ at .5% significance level[d]
Can,Fra	0.85	1.232	0.240	0.61 - ∞	0.544 - 1.000	yes
Can,Jap	0.82	1.135	0.234	0.53 - ∞	0.485 - 1.000	yes
Can,US	0.92	1.557	0.263	0.88 - ∞	0.706 - 1.000	yes
Can,Ger	0.83	1.166	0.233	0.56 - ∞	0.508 - 1.000	yes
Can,UK	0.88	1.352	0.233	0.75 - ∞	0.635 - 1.000	yes
Can,Nor	0.67	0.793	0.234	0.19 - ∞	0.188 - 1.000	yes
Can,Swe	0.79	1.050	0.234	0.45 - ∞	0.422 - 1.000	yes
Fra,Jap	0.95	1.805	0.239	1.19 - ∞	0.831 - 1.000	yes
Fra,US	0.90	1.441	0.263	0.76 - ∞	0.641 - 1.000	yes
Fra,Ger	0.87	1.308	0.240	0.69 - ∞	0.598 - 1.000	yes
Fra,UK	0.93	1.632	0.240	1.01 - ∞	0.766 - 1.000	yes
Fra,Nor	0.80	1.076	0.240	0.46 - ∞	0.430 - 1.000	yes
Fra,Swe	0.84	1.197	0.240	0.58 - ∞	0.523 - 1.000	yes
Jap,US	0.89	1.391	0.264	0.71 - ∞	0.611 - 1.000	yes
Jap,Ger	0.86	1.270	0.233	0.67 - ∞	0.585 - 1.000	yes
Jap,UK	0.93	1.633	0.233	1.03 - ∞	0.774 - 1.000	yes
Jap,Nor	0.80	1.077	0.234	0.47 - ∞	0.438 - 1.000	yes
Jap,Swe	0.81	1.105	0.234	0.50 - ∞	0.462 - 1.000	yes
US,Ger	0.85	1.227	0.264	0.55 - ∞	0.501 - 1.000	yes
US,UK	0.95	1.799	0.263	1.12 - ∞	0.808 - 1.000	yes
US,Nor	0.67	0.788	0.265	0.10 - ∞	0.100 - 1.000	yes
US,Swe	0.82	1.129	0.264	0.45 - ∞	0.422 - 1.000	yes
Ger,UK	0.90	1.448	0.233	0.85 - ∞	0.691 - 1.000	yes
Ger,Nor	0.74	0.933	0.217	0.37 - ∞	0.354 - 1.000	yes
Ger,Swe	0.84	1.202	0.216	0.64 - ∞	0.565 - 1.000	yes
UK,Nor	0.70	0.848	0.234	0.24 - ∞	0.236 - 1.000	yes
UK,Swe	0.83	1.166	0.233	0.56 - ∞	0.508 - 1.000	yes
Nor,Swe	0.74	0.935	0.207	0.40 - ∞	0.380 - 1.000	yes

[a] Source: Krueger and Summers (1987, p.26, Table 2.3).
[b] See Hotelling (1953, §9, p.217-221).
[c] See note c to Table A.II.
[d] 1-tailed tests of the null hypothesis H_0: $\rho=0$, against the alternative hypothesis H_1: $\rho>0$. The relevant claim is, in fact, that the international wage structure is stable across countries. Only positive correlations are therefore expected.

APPENDIX B

Estimated Wage Equations in a Sample Selection Model

In this appendix I report the results for the estimated wage equations which were used to derive the industry wage differentials presented in section 5.

Table B.I gives the estimates of the probit overtime work equation used in our sample selection model. The probability that employees work overtime hours is estimated with the ML method as a function of a number of variables which explain individual overtime work. As far as the significance of each variable is concerned, we note that a relevant role is played by the number of children in the household, the nationality, the skill level, and the firm size.

In Table B.II I present the results for the regression of log hourly earnings on both industry variables and a set of controls for labour quality, demographic and working conditions in a sample selection model. The estimated industry wage differentials reported in this table were used to calculate the normalized industry differentials appearing in column II of Table IV. With respect to the raw estimates of Table B.III, we notice that the industry affiliation effects tend to decrease after controlling for human capital and working conditions. However, estimated industry differentials remain statistically significant both jointly and individually in most cases.

As far as other control variables are concerned, I defined several sets of dummy variables. Since the wage regression includes a constant, I omitted a dummy variable from each set and treated it as having a zero effect on wages. The reference groups for each set of dummy variables are the following: employees who did not receive any education or training at all for the education and training variables; unskilled blue-collar workers for the skill variables; single employees for the marital status variables; German nationality employees for the nationality variables; employees in firms with fewer than 20 employees for the firm size variables; employees working in the agriculture,

TABLE B.I

Estimated PROBIT overtime work equation in a sample selection model with controls for human capital and working conditions, 1984 (t-statistics in parentheses)

Variables	Parameter estimate	Mean of variable	Number of cases
Intercept	-.154 (-.40)	—	—
Age/10	-.099 (-.57)	3.963	2,944
Age squared/100	.003 (.12)	16.947	2,944
Skill variables:			
Unskilled worker with on-the-job training	.008 (.08)	.290	855
Trained worker	.068 (.68)	.274	807
Foreman	.195 (1.32)	.042	124
Master	.224 (1.10)	.018	54
White-collar industrial worker	-.014 (-.06)	.014	42
White-collar worker in basic positions	-.278 (-1.51)	.029	84
White-collar worker with advanced qualifications	-.084 (-.70)	.139	408
Highly trained white-collar worker	.030 (.23)	.084	248
White-collar worker with extensive leadership responsibilities	-.999 (-2.92)	.010	30
Nationality variables:			
Turkey	-.469 (-4.77)	.116	342
Yugoslavia	-.076 (-.74)	.072	213
Greece	-.189 (-1.54)	.052	154
Italy	-.357 (-3.46)	.082	240
Other nationality	-.472 (-4.20)	.068	199

(continued)

TABLE B.I (Continued)

Variables	Parameter estimate	Mean of variable	Number of cases
Dummy for missing nationality	-.796 (-1.33)	.002	7
Number of children under 16	.077 (3.02)	1.004	2,944
Dummy for second house	-.010 (-.06)	.025	73
Dummy for mortgage emprunt logement	-.072 (-1.08)	.195	573
Number of nights in hospital	-.00002 (-.01)	1.711	2,944
Degree of satisfaction with the current job (1-10)	.0009 (.09)	7.533	2,930
Firm size variables:			
20-200 employees	.210 (2.65)	.295	869
200-2000 employees	.136 (1.61)	.253	745
2000 or more employees	.167 (1.94)	.289	850
Industry variables:			
Energy, water and mining	-.225 (-.82)	.026	77
Chemical	-.352 (-1.37)	.047	137
Rubber	-.105 (-.37)	.019	56
Stone, clay and glass	-.083 (-.30)	.021	63
Iron and steel	-.142 (-.60)	.133	393
Machinery, excl. elec.	-.167 (-.71)	.123	362
Electrical machinery	-.274 (-1.09)	.054	159
Lumber, wood, paper and printing	.151 (.60)	.043	126
Textile and apparel	-.162 (-.60)	.029	85
Food, beverages and tobacco	-.059 (-.23)	.031	90

(continued)

TABLE B.I (Continued)

Variables	Parameter estimate	Mean of variable	Number of cases
Construction	-.236 (-1.01)	.135	398
Wholesale trade	.422 (1.46)	.017	51
Retail trade	-.350 (-1.30)	.031	91
Railroads	-.765 (-1.87)	.007	22
Mail service	-.341 (-.86)	.007	20
Other transport and communications	.053 (.20)	.032	94
Banking	.182 (.61)	.016	47
Insurance	-.245 (-.64)	.007	20
Personal services	.136 (.46)	.015	44
Entertainment	-.019 (-.07)	.019	56
Health services	.095 (.32)	.017	49
Legal and business services	-.168 (-.47)	.008	23
Non-profit organizations and private households	-.283 (-.88)	.013	37
Local collective organizations	-.170 (-.66)	.042	123
Social security	-.247 (-.59)	.005	15
Dummy for missing industry	-.140 (-.59)	.092	271
Log-Likelihood		-1727.3	
Restricted (Slopes=0) Log-L.		-1788.8	
Chi-squared (51)		122.95	
Significance Level		.000000009	
Sample size		2,944	

TABLE B.II

Estimated wage equation in a sample selection model with controls for human capital and working conditions, 1984 (t-statistics in parentheses)

Variables	Parameter estimate	Mean of variable	Number of cases
Intercept	1.570 (15.69)	—	—
Age/10	.380 (9.93)	4.012	2,072
Age squared/100	-.043 (-9.29)	17.356	2,072
Tenure/10	.064 (3.32)	1.159	1,971
Tenure squared/100	-.014 (-2.64)	2.108	1,971
German education variables:			
Short-course secondary school	.016 (.67)	.476	986
Intermediate type of secondary school	.071 (2.25)	.093	193
Technical high school or academically-oriented secondary school	.076 (1.86)	.027	56
Technical college, engineering school	.147 (3.49)	.029	60
College/University	.264 (5.96)	.031	64
Foreign education variables:			
Compulsory school without final examination	.071 (2.78)	.120	249
Compulsory school with final examination	.046 (1.83)	.167	346
Further schooling	.062 (1.64)	.026	53
Specialized professional school	.102 (3.13)	.044	92
College/University	-.159 (-1.93)	.004	9
Skill variables:			
Unskilled worker with on-the-job training	.017 (.91)	.298	618
Trained worker	.086 (4.14)	.267	552

(continued)

TABLE B.II (Continued)

Variables	Parameter estimate	Mean of variable	Number of cases
Foreman	.141 (4.14)	.038	79
Master	.328 (7.05)	.016	34
White-collar industrial worker	.247 (5.06)	.014	29
White-collar worker in basic positions	.005 (.14)	.032	66
White-collar worker with advanced qualifications	.219 (8.36)	.138	286
Highly trained white-collar worker	.448 (13.74)	.080	165
White-collar worker with extensive leadership responsibilities	.735 (12.09)	.013	27
Marital status variables:			
Married, living with spouse	.100 (5.36)	.448	928
Married, permanently separated	.450 (5.61)	.004	9
Divorced	.060 (1.40)	.017	35
Widowed	.133 (2.22)	.008	17
Number of nights in hospital	-.001 (-2.08)	1.751	2,072
Degree of satisfaction with the current job (1-10)	.005 (2.11)	7.529	2,060
Firm size variables:			
20-200 employees	.040 (2.17)	.282	585
200-2000 employees	.070 (3.80)	.254	527
2000 or more employees	.122 (6.38)	.293	606
Industry variables:			
Energy, water and mining	.222 (3.66)	.027	55
Chemical	.194 (3.39)	.051	105

(continued)

TABLE B.II (Continued)

Variables	Parameter estimate	Mean of variable	Number of cases
Rubber	.104 (1.65)	.019	39
Stone, clay and glass	.146 (2.37)	.021	44
Iron and steel	.130 (2.47)	.135	280
Machinery, excl. elec.	.170 (3.21)	.124	257
Electrical machinery	.139 (2.50)	.057	117
Lumber, wood, paper and printing	.147 (2.60)	.037	76
Textile and apparel	.123 (2.09)	.030	62
Food, beverages and tobacco	.008 (.14)	.029	60
Construction	.153 (2.91)	.140	289
Wholesale trade	.031 (.44)	.013	26
Retail trade	.005 (.09)	.034	71
Railroads	.039 (.50)	.009	19
Mail service	.120 (1.51)	.008	16
Other transport and communications	.098 (1.67)	.029	59
Banking	.015 (.22)	.014	28
Insurance	.272 (3.37)	.007	15
Personal services	-.214 (-3.20)	.014	28
Entertainment	.093 (1.45)	.018	38
Health services	-.003 (-.05)	.015	31
Legal and business services	.173 (2.19)	.008	16
Non-profit organizations and private households	.048 (.70)	.014	28
Local collective organizations	.062 (1.08)	.043	88

(continued)

TABLE B.II (Continued)

Variables	Parameter estimate	Mean of variable	Number of cases
Social security	.050 (.57)	.005	11
Dummy for missing industry	.091 (1.72)	.092	191
HECKMAN'S λ	-.070 (-.90)		
Correlation between regression and selection equation disturbances (ρ_{ue})		.297	
SE of the regression		.237[a]	
F-statistic (59,2012)		38.310[**]	
Adjusted R^2		.515	
Sample size		2,072	

[a] Selectivity corrected estimate of the standard error of the regression.
[**] F test that all the estimated slope coefficients jointly equal 0 rejects at the 1% level. The 1% critical point is $F_{.01}(59,2012) \approx 1.48$.

TABLE B.III

Estimated wage equation in a sample selection model <u>without</u> controls for human capital and working conditions, 1984 (t-statistics in parentheses)

Variables	Parameter estimate	Mean of variable	Number of cases
Intercept	2.481 (32.37)	——	——
Industry variables:			
Energy, water and mining	.436 (5.32)	.027	55
Chemical	.442 (5.78)	.051	105
Rubber	.210 (2.42)	.019	39
Stone, clay and glass	.296 (3.48)	.021	44
Iron and steel	.252 (3.53)	.135	280
Machinery, excl. elec.	.354 (4.92)	.124	257
Electrical machinery	.367 (4.86)	.056	117
Lumber, wood, paper and printing	.213 (2.71)	.037	76
Textile and apparel	.259 (3.20)	.030	62
Food, beverages and tobacco	.173 (2.15)	.029	60
Construction	.270 (3.77)	.139	289
Wholesale trade	.326 (3.43)	.013	26
Retail trade	.119 (1.48)	.034	71
Railroads	.233 (2.22)	.009	19
Mail service	.291 (2.68)	.008	16
Other transport and communications	.307 (3.79)	.028	59
Banking	.383 (4.12)	.014	28
Insurance	.691 (6.28)	.007	15

(continued)

TABLE B.III (Continued)

Variables	Parameter estimate	Mean of variable	Number of cases
Personal services	-.219 (-2.36)	.014	28
Entertainment	.542 (6.22)	.018	38
Health services	.342 (3.77)	.015	31
Legal and business services	.582 (5.42)	.008	16
Non-profit organizations and private households	.397 (4.25)	.014	28
Local collective organizations	.328 (4.23)	.042	88
Social security	.400 (3.30)	.005	11
Dummy for missing industry	.248 (3.40)	.092	191
HECKMAN'S λ	-.164 (-2.59)		
Correlation between regression and selection equation disturbances (ρ_{ue})		.486	
SE of the regression		.338[a]	
F-statistic (27,2044)		9.547**	
Adjusted R^2		.100	
Sample size		2,072	

[a] Selectivity corrected estimate of the standard error of the regression.
** F test that all the estimated slope coefficients jointly equal 0 rejects at the 1% level. The 1% critical point is $F_{.01}(27,2044) \approx 1.74$.

forestry and fishery sector for industry variables.

The choice of the reference groups - the omitted dummy variables - is completely arbitrary and does not affect the statistical properties of the model as a whole. However, if we do not take this arbitrariness into account, the interpretation of the estimated coefficients and of their t-statistics may be very misleading. The coefficients for the control dummy variables, in fact, must be interpreted as the relative wage differentials with respect to an individual characterized by the combination of all the aspects associated with each of the reference groups. For this reason in the main text we presented a "normalized" measure of industry differentials, which are the variables of our major concern in this study. Limiting therefore our consideration to the statistical significance of the other control variables, we notice that in the wage equation a relevant role is played by almost all variables: age, tenure in the current job, education, skill level, marital status, health conditions measured through the number of nights spent in a hospital, the degree of satisfaction with the current job, firm size and industry affiliation. Human capital and working conditions as a whole are very important in explaining variations in individual wages. When these controls are introduced in the wage regression, the standard error of the regression is reduced by 30 percentage points and the adjusted R^2 increases from 10 to 52 percent.

We can also note that the coefficient for the Heckman's λ in our sample selection model is not significantly different from zero, indicating that we fail to reject the null hypothesis of no sample selection bias introduced with the elimination of employees doing overtime work.

Table B.III gives the estimates of two-digit industry wage differentials from the regression of log hourly earnings on industry variables only in a sample selection approach. These estimates were used to calculate the normalized industry differentials reported in column I of Table IV. As already noticed, most of the estimated coefficients are statistically significant individually and they are also jointly statistically significant at the 1% level. We note that here the coefficient for the Heckman's λ is significantly different from zero at the 1% level, revealing a sample selection bias problem in this specification of the model. As already stressed, however, our estimates of the regression coefficients obtained with the Heckman's method are still consistent.

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