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Co-Determination, Efficiency, and Productivity

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ABSTRACT

Co-Determination, Efficiency, and Productivity*

We present the first panel estimates of the productivity effects of the unique German institution of parity, board-level co-determination. Although our data span two severe recessions when labour hoarding costs of co-determination are probably highest, and the panel is too short to capture the likely long run benefits in terms of human capital formation and job satisfaction, we find positive productivity effects of the 1976 extension to parity co-determination in large firms.

JEL Classification: D2, J5, L2

Keywords: co-determination, employee involvement, productivity

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1 Introduction

While human resource management (HRM) has long been an important topic in business management and organisational studies, it is only fairly recently that rigorous economic analysis in this area has emerged (Bailey et al, 2001; Kleiner et al, 2002). In a sample of British firms, Patterson et al (1998) find good HRM practices including employee involvement and team working to have much greater influence on productivity and profitability than traditionally predominant concerns such as business strategy, advanced technology and R & D. The authors also find HRM to be "one of the most neglected areas" of management practice in their sample.

Ichniowski et al (1997) show that combining various aspects of HRM into an integrated employee involvement package was much more effective in raising productivity in the US steel industry than any single component , such as team work, job-enlargement, or profit sharing on its own. These findings confirm earlier results summarised by Levine and Tyson (1990), who also argue that low national levels of training, employee involvement and other HMR measures could constitute a Nash equilibrium for most firms in the American (or British) institutional framework. Acemoglu and Pischke (1999) argue that compression of the wage distribution in Germany encourages firms to provide more general training and supports a high - skill equilibrium.

In contrast to the high proportion of unqualified workers in the Anglo-American economies, HRM has been given central importance by legislation on co-determination in the German economy. This legislation includes the right to elect plant level works councils with far reaching powers in most personnel related decisions, and since 1976 a requirement that half of the supervisory board (**Aufsichtsrat**) of firms with over 2000 employees, consist of labour representatives (though the chair, appointed by owners, retains a second vote in case of a tie).

The supervisory board, essentially the non-executive and outside directors, appoints the top management board (**Vorstand**) and has to approve major strategic decisions such as acquisitions or plant closure. Ultimate corporate power thus resides with the supervisory board, and directly sharing this power with labour and union representatives represents a radical break with the neoclassical model, where maximizing shareholder value should be the (only) goal of the firm, and owners or their representatives hold residual power to safeguard their residual income.

Surprisingly, perhaps, board level co-determination has received little attention from economists. In contrast to the much analysed effects of works councils, there is

not much econometric evidence on codetermination. While simple comparisons of mean values before and after the 1976 legislation by Benelli et al (1987) and Gurdon and Rai (1990) were inconclusive, most German commentators argue that board level co-determination complements plant level councils and has helped to maintain co-operative labour relations, with greater weight given to HRM than in other countries (Streeck, 1984, 1995). Business representatives, however, have strongly opposed the various extensions of co-determination, particularly in 1976, as primarily redistributive measures. The employers' attempt to block the 1976 law as "unconstitutional" was finally defeated in the Federal Constitutional Court in 1979. Interest has also been raised by the 1997 report "European Systems of Worker Involvement" of a European Commission group chaired by former vice-President Etienne Davignon, which recommends that one fifth of seats in supervisory boards of European public companies formed from merging companies from at least two different countries should be held by employee representatives.

The EC report explicitly rejected the idea of giving labour representatives on the supervisory board only consultative - rather than voting-rights, although the practice of co-determination generally seems to lead to unanimous agreement after extensive discussion. The suggested level of involvement emerged as compromise to fit the widely differing levels of employee involvement in EU countries. Though offering much less empowerment than the German system of almost-parity, it still represents a substantial departure from neo-classical and Anglo-American notions of exclusive residual control for capital owners or their representatives.

Previous econometric studies of board level codetermination by FitzRoy and Kraft (1993) and Gorton and Schmid (2000) used cross sectional data and thus could not control for firm-specific effects. They found no evidence of productivity benefits. To overcome the problems inherent in cross sectional data, in the present paper we use panel data for 179 firms, from 1972-76 and from 1981-85, allowing for adjustment to the 1976 law. Both periods contain severe recessions, as well as more normal years, so the sample is still, if anything, biased towards exaggerating the costs of co-determination, though obviously less so than the single recession year cross sections. The empirical results suggest a significant, though small, positive influence on productivity from the 1976 strengthening of co-determination law. These results reject the critique of co-determination made by some business and academic observers and provide support for the Davignon Report. In view of the likelihood of recessionary labour hoarding in our samples, the results do seem to provide indirect evidence of compensating positive productivity effects, even in the relatively short

run. However, as we argue below, the legal framework of co-determination complements other characteristics of the environment in which it operates so extension to different environments needs careful consideration.

The plan of the paper is to outline the economic and institutional framework in section 2, and provide an overview of the data used in section 3. Section 4 then presents the empirical results and conclusions are summarised in a final section 5.

2 Economic and Institutional Framework

Traditionally economists see little value in state intervention into a functioning market and are therefore sceptical about the efficiency effects of codetermination. The property rights theory, in general, dislikes intervention by government into the decision rights of firms. The argument is that if it were efficient, then it would emerge in an evolutionary way by itself (see among others Jensen and Meckling 1979). Clearly this argument is relevant, although there are situations possible, where the market mechanism does not work. However, even if the result is a cooperative one, decisions will certainly take longer if a consensus is needed; compromises will in part limit the interests of the capital owners and flexibility is necessarily reduced. It is clearly possible that such an environment reduces the innovativeness of an organization. Co-determination may well lead to a tendency to maintain the status quo in order to avoid any conflict. Additional support for a skeptical view concerning this law might come from a political-economic perspective. The co-determination law of 1976 was passed under the SPD/FDP (social democrats and liberals) government. Traditionally, the social democrats are associated with unions and one could argue that this government introduced this law in order to do the unions a favor, even if the whole economy does not benefit from it. This has some logic, if one believes in rent-seeking.

Participation theorists, on the other hand, argue that the market process does not necessarily lead to a first best solution. Co-determination enables the use of information from employees, which would otherwise be lost. Furthermore, it will lead to a more cooperative solution, and the conflict between capital owners and workers is reduced if not solved. According to this view, productivity will increase as a result, and such firms will be successful on the market.

In a world of imperfect information and incomplete contracts the potential benefits of employee involvement or ‘voice’ in decision-making have long been recognized (Hirschman, 1970; Steinherr ,1977; McCain,1980; Sertel ,1982). However, involvement may also divert effort to internal rent seeking and these ‘influence costs’ could in some circumstances outweigh efficiency gains (Milgrom and Roberts, 1992). Even when worker participation generates net efficiency gains, employees' bargaining power over distribution of the enterprise surplus or rent is likely to increase, so that the share remaining for owners or managers may decline, thus explaining the widespread opposition to employee involvement by business leaders. As Freeman and Lazear (1995), and others, have emphasized, the distributional effect will probably lead employees to demand a degree of empowerment that

exceeds the social optimum, while employers will maximise their share of the surplus when employee involvement is below the socially optimal level.

The designers of co-determination seem to have been aware of these problems, because collective bargaining in Germany is formally quite separate from all aspects of co-determination. To summarise the institutional set-up, industry unions bargain with employers' associations over regional wages, and neither works councils nor employee board representatives are formally involved at any stage. However, works councillors are usually union members, and board representatives are usually also works councillors, though some outside board members are appointed by the relevant union, so there are close personnel links between the three institutions. Furthermore, the works councils are closely involved in decisions over working conditions which may affect workers' skill and pay classification. Supplementary wage and working time agreements at plant or enterprise level typically do involve works councils, and have been increasing in recent years. Works councillors are not allowed to call a strike, though they have been involved in unofficial strikes.

The design of the German system of co-determination has thus at least attempted to ameliorate the conflict between efficiency and distributional goals. Participants are legally enjoined to promote enterprise success and not just their own constituents' goals in internal bargaining. However the system is now under strain as increasing numbers of employers have opted out of the industry employers' federation to conduct enterprise or plant level bargaining in which works councils are likely to be at least informally involved.

Since displaced workers usually suffer substantial losses of future income, employee organisations have an incentive to oppose restructuring and downsizing that would be optimal for owners. Parity co-determination may have strengthened the already strong Federal labour laws and the role of works councils in limiting employers' flexibility in Germany, but also encouraged investment in firm specific human capital and HRM. Case studies reveal wide variation in the implementation of co-determination, ranging from management hostility and deliberate restriction of information flows, to enthusiastic co-operation (Nagel, 1996). Of course, a similar range of management-labour relations is also found in firms and countries without co-determination. However, the very high average levels of blue collar skills and training in Germany also complement the institutions of mandatory co-determination, allowing the most skilled workers (**Facharbeiter**) to hold extensive responsibilities in production, of a kind that would be reserved for supervisory, white collar employees in most other countries. Case studies of matched British and German plants again show that human

capital deployment is the main reason for the superior productivity of the German plants (Prais, 1990).

Gorton and Schmid (2000) investigate the effects of parity co-determination on measures of firm performance relevant for capital owners, such as return on assets and equity, with cross-sectional data. Although they control for the influence of capital (total assets) with semi-parametric regression, their finding of generally negative co-determination coefficients is subject to the usual problem of interpretation in cross-section analysis where firm specific effects cannot be controlled for. Since only the largest firms in terms of employment are co-determined, they may be picking up a size effect rather than any direct effect of co-determination. Gorton and Schmid note that co-determination may be a way of protecting non-contractible investments in firm specific human capital through a transfer of power to employee stakeholders, and that evidence for such a transfer is not sufficient to judge the social benefits of co-determination.

The 1976 co-determination law gives the chair of the supervisory board, who is appointed by owners, an extra vote to break a tie between the equal numbers of labour and owners' representatives on the board of companies with more than 2000 employees. This situation of 'almost parity' is considered by most observers to have substantially strengthened labour's position compared to the earlier one-third representation (for all firms with over 500 employees) and further strengthened the role of HRM. Most board decisions appear to be by unanimous vote. Parallels can also be drawn between the organisation of large Japanese firms, with no formal co-determination but extensive involvement, and the effects of co-determination combined with a highly trained work force in Germany (FitzRoy, Acs, Gerlowski, 1998).

Addison, Schnabel and Wagner (2004, 2001), and Huebler and Jirjahn (2003) survey the evidence on both the effects and determinants of works councils, which are the most studied institution of codetermination. Positive or negligible effects seem to dominate the literature. In 1994 between 98-100% of all firms with 2000 employees or more had a works council, in contrast to only 17-20%. of all firms. Hence the large codetermined firms will almost surely also have a works council, while this is unclear for the smaller units. Unfortunately we have no information concerning the existence of works councils in our sample.

While Freeman and Lazear (1995) and most other writers on this topic neglect the interaction between collective bargaining and co-determination, Huebler and Jirjahn (2003) follow McCain (1980) to develop a model of collective bargaining and works councils. They

find that councils raise productivity only in firms under the industry collective bargain, while the main wage effect is in firms that have opted out of the industry bargain to negotiate independently with the union. However their data does not include capital, so their results on productivity have to be interpreted with caution. FitzRoy and Kraft (1995) find a positive interaction between works councils and profit sharing.

The 1976 move from one third- to almost – parity, board level co-determination in large firms is agreed by most observers to have strengthened employee bargaining power, (Rogers and Streeck, 1995). More protection for non-contractible human capital investment in the form of reduced employment fluctuation would tend to lower the return to capital (as found by Gorton and Schmid, 2001), unless offset by greater efficiency. Since the large firms affected by board-level co-determination in our sample almost all had works councils, and were likely to be covered by the industry wage bargain, the question we address is: what effects do different kinds of board representation, if any, have on productivity? Most theoretical models agree that there is likely to be an optimal degree of co-determination in terms of efficiency, and 'too much' participation or employee representation will be counterproductive. There is thus no particular presumption that the two degrees of parity should have similar-or even same signed- effects, and since we have a panel of firms that changed from one third -to almost-parity after 1976 we can provide the first estimates of the *additional* effect of this controversial move, while controlling for firm specific effects.

3 The Data

We use published data for 179 public companies (not all of them traded in the stock market) from all sectors of the manufacturing industry except steel, coal mining and brewing in (former) West Germany. Steel and mining firms are excluded because of their stronger codetermination rights and particular structural problems related to these industries. Brewing is not considered because of its special circumstances, which are not representative for the manufacturing sector in general. Overall beer consumption is declining, the smaller breweries make frequent losses, there are many mergers, large firms are gaining market share and the whole industry is subject to severe structural change. The criteria for selection of the other firms is simply the availability of the necessary data from companies that publish this information for the years in question. We include all information that is available to us.

The data form two unbalanced panels, with at least 3 and at most 5 annual observations in the two periods, 1972-1976 and 1981-1985, before and after the 1976 Co-determination Act. The 65 largest co-determined companies all switch from one-third labour representation before 1976 to ‘almost parity’ after 1976. This is the strongest form of co-determination which we use as a benchmark for productivity comparison, so that our ‘parity co-determined’ sub-sample consists of all those firms with employment exceeding 2000 after 1976 (by a generally large margin). The rest of our sample consists of firms which are all much smaller, mostly with less than 500 employees and thus not covered by the first (1952) law on co-determination which mandated 'one third parity' or board representation by labour. We denote these the ‘non co-determined’ firms. Since we wish to compare the effects of parity co-determination (post 1976) with the effects of weak co-determination (one third parity) for the *same* sub-sample, thus controlling for firm effects, in order to identify any influence of the 1976 legislation, we include a few firms with one-third parity in the 114 ‘non-codetermined’ pre-1976 sub-sample. We consider one-third parity separately below. None of our firms has changed its legal status, has reduced employment below the 2000 employees limit or split up in legally separated firms in order to escape from the codetermination law.

The following Table 1 lists definitions of variables used and presents descriptive statistics for the six sub-samples. We use sales as our gross output measure, and include materials and intermediate goods as an input in addition to labour and capital thus avoiding the problems of bias from published value-added data in an imperfectly competitive environment (Basu and Fernald, 1995). Estimations of production functions are frequently plagued by the reverse causality running from output to factor demand. In order to avoid a possible simultaneous equation bias, we use lagged values of the production factors labour, capital and materials. The production factors are in this case predetermined and can be used as a simple way to avoid a simultaneous equation bias. However none of our results depend on the use of lagged variables. The sales volume and materials are divided by the price index of the relevant industry and are therefore real values. Since our three factors are all related to firm size, we are also less likely to conflate co-determination effects with firm size effects on productivity. Since size is a major determinant of works council election (Addison, Schnabel and Wagner, 1997), our large co-determined firms almost certainly all have works councils.

Our general strategy is it to consider the relative performance of non-codetermined firms compared with the codetermined ones before and after the 1976 legislation. One possible reason for firm size effects not captured by the production factors, is that our non-co-

determinated, smaller firms may not all have works councils, though we have no data on this. There is thus a possibility that any cross-sectional productivity differences may also be affected by the presence or absence of works councils. Since councils are rarely disbanded or newly elected, we focus on the change in productivity measured before and after the 1976 law. Nevertheless, as we and most other students of these institutions have emphasized, plant- and board-level co-determination are likely to be complementary.

Table 1a: Descriptive Statistics for 1972-1976
(Mean Values, Standard Deviation in parenthesis)

VARIABLE	**Codetermined Firms (N = 65)**	**Third Parity-Codetermined Firms (N = 61)**	**Non-Codetermined Firms (N=53)**
COD (=1 if the employment is >=2000)	1.	.0	.0
LABOUR Number of Employees	17586 (26926)	1137 (547)	321 (126)
CAPITAL Total Capital Stock in TDM	1320188 (2187755)	65906 (40753)	16949 (12092)
MAT Material Input in TDM	963477 (1631943)	46217 (51503)	14674 (10618)
SALES Total Sales Volume in TDM	1879424 (3096734)	95906 (64853)	28248 (16677)
OVER Average Number of Overtime Hours per Worker at Industry Level	2.30 (0.94)	2.44 (1.03)	2.57 (1.20)
CONC Percentage Market Share of the Six Largest Firms – Industry Level	33.58 (18.88)	22.80 (15.06)	21.65 (16.72)
IMP (Value of Imported Goods/Total Production) – Industry Level	0.17 (0.11)	0.19 (0.12)	0.22 (0.17)
EXP (Value of Exported Goods/Total Sales Volume) – Firm Level	0.33 (0.10)	0.30 (0.06)	0.30 (0.12)

Table 1b: Descriptive Statistics for 1981-1985

(Mean Values, Standard Deviation in parenthesis)

VARIABLE	**Codetermined Firms (N = 65)**	**Third Parity-Codetermined Firms (N = 61)**	**Non-Codetermined Firms (N=53)**
COD (=1 if the employment is >=2000)	1.	.0	.0
LABOUR Number of Employees	18918 (37834)	1100 (450)	315 (117)
CAPITAL Total Capital Stock in TDM	2246018 (3865700)	117990 (89117)	28904 (22076)
MAT Material Input in TDM	2082503 (3862749)	90468 (98639)	30746 (43039)
SALES Total Sales Volume in TDM	3729622 (6706128)	175726 (130640)	53270 (48942)
OVER Average Number of Overtime Hours per Worker at Industry Level	1.43 (0.59)	1.57 (0.63)	1.71 (0.79)
CONC Percentage Market Share of the Six Largest Firms – Industry Level	33.43 (18.65)	23.55 (15.37)	23.30 (16.95)
IMP (Value of Imported Goods/Total Production) – Industry Level	0.25 (0.16)	0.28 (0.19)	0.34 (0.22)
EXP (Value of Exported Goods/Total Sales Volume) – Firm Level	0.44 (0.17)	0.42 (0.18)	0.42 (0.18)

4 Empirical Results

Given the limitations of the data, our empirical strategy is to estimate the effects of co-determination on total factor productivity in the pooled data of the pre-1976 and post-1976 panels. In view of the major size differences between co-determined and non-codetermined firms our main interest is in testing for a change in the effect of co-determination on large firms as a result of moving from one third parity to almost parity in 1976. Opponents of co-determination argue that this move, supported by the social democratic West German government of the time, went beyond efficient involvement in order to redistribute more surplus to labour. However, as argued above, redistribution does not exclude efficiency gains as well, through improved co-operation and human capital formation.

By focusing on the change in board representation in a given subset of (65) firms, our co-determination dummy variable is obviously defined as time invariant. The usual firm fixed effects thus cannot be distinguished from co-determination effects, and the fixed effects or 'within' estimator cannot be used. We shall thus use the solution suggested by Hausman and Taylor (1981) for such a situation. This model allows that some of the explanatory variables are related to the firm-specific effects, while others are not. In particular they consider four groups of explanatory variables. The variables in the first group are time-varying and uncorrelated with the firm-specific effects, the second group considers time-varying variables, that are related to the specific effects, in the third category the variables are time-invariant and uncorrelated with the fixed effects and finally there are variables, which are time-invariant and correlated with the fixed effects. The model is rather complicated and perhaps this is the reason for the few applications of this very useful solution to a frequent problem. Our specification is accepted by the appropriate Hausman (1978) test on a possible correlation between the fixed effects and our explanatory variables. We explain the econometric model to the interested reader in more detail in an appendix. For other readers we just report and discuss the results.

As the model is already quite involved, we have not tried to use more complicated production functions like CES or translog, but use instead the simple Cobb-Douglas function. Therefore we start with the following log-linear specification:

$$(1) \quad \ln SALES = \beta_0 + \beta_1 COD + \beta_3 \ln LABOUR + \beta_4 \ln CAPITAL + \beta_5 \ln MAT + \beta_6 OVER + \beta_7 CONC + \beta_8 IMP + \beta_9 EXP + \text{time dummies} + \text{random firm effect} + \text{error}.$$

To identify the effects of parity in the COD subsample we now define an additional dummy variable COD80, which takes the value one for co-determined firms *only after 1980.* This variable thus measures the *additional effect* of moving from one third to almost parity after 1976, in the given subset of firms with over 2000 employees. The idea of this specification is to control for any possible size effect present at the 2000 employees limit, which is not captured by the production factors. If COD has a significant coefficient, there is such an effect. Wooldridge (2002, 129-130) shows that this specification is a difference- in-difference estimator. A combination of size effects with the impact of the codetermination law is now only possible for the situation, that something linked to size has *changed* from the seventies to the eighties. One possible reason might be different developments of the qualification structure. Data on this topic is very hard to obtain as it has to be differentiated according to firm size classes and cover the seventies and eighties. We obtained from the Bundesinstitut für Berufsbildung in Berlin information concerning apprenticeship and the change from 1980 to 1985. The share of firms with of apprentice programs rose for firms with 50-499 employees from 0.73 to 0,77. From the larger firms with 500 employees or more this ratio increased from 0.93 to 0.94.(We have no separate information for firms with 2000+ employees.) Hence according to these figures there was no dramatic change in apprentices and as the qualification structure changes rather slowly over time, we do not think that such a change could affect our results.

In addition to the co-determination dummies, we add industry overtime OVER, as an index of capacity utilization, and the 4-firm industry concentration ratio, CONC, as an index of market power. The role of international trade is represented by the share of sales from foreign firms divided by total sales volume, IMP, measured on the two digit industry level. The share of sales exported is calculated on the individual firm level and added as EXP. Time dummies are also included to control for inflation and macroeconomic conditions.

In terms of the Hausman -Taylor model, our group one variables are the industry variables OVER, CONC, IMP, EXP and the time dummies, in group two we have the production factors, in group three are the industry dummies and finally in the forth group we put the codetermination dummies. Hausman and Taylor (1981) suggest that a time invariant variable such as either of our co-determination dummies in the relevant years could be

instrumented using explanatory variables that are correlated with it, but not with firm effects. This suggestion is followed by instrumenting the codetermination variable by the individual mean values of IMP, EXP; OVER, CONC and industry dummies, which is then denoted CÔD (for the years 1972-1985) and CÔD80 (for the years 1981-1985). The additional productivity effect of codetermination is thus estimated by β_2, independently of the size effect that is captured by β_1. Our final specification is thus:

$$\text{(2)} \quad \ln SALES = \beta_0 + \beta_1 \, C\hat{O}D + \beta_2 \, C\hat{O}D80 + \beta_3 \ln LABOUR + \beta_4 \ln CAPITAL + \beta_5 \ln MAT + \beta_6 \, OVER + \beta_7 \, CONC + \beta_8 \, IMP + \beta_9 \, EXP + \text{time dummies} + \text{random firm effect} + \text{error.}$$

The results, presented in Table 2, show that parity co-determination does have a small, significant and positive additional effect on productivity after 1976, thus rejecting the hypothesis of purely redistributive effects. As the Hausman-Taylor approach is rather complex and involves a number of variable transformations, we also present simple random effects estimations without the many data purifications in order to check the robustness of the results. The basic results with respect to the codetermination variables remain unchanged. We also performed the Hausman-Taylor approach without the IV transformation of the COD variables. The basic results were not affected.

Table 2: Cobb-Douglas-Production Functions, with Pooled Data:
Dependent Variable Sales, (Mean Values, Standard Deviation in parenthesis)

Period VARIABLE	**Hausman-Taylor** **1972 – 1976,** **1981 –1985**	**Random Effects** **1972 – 1976,** **1981 –1985**
Constant	3.17 (11.12)	2.51 (20.42)
CÔD	0.24 (4.06)	0.005 (0.22)
CÔD80	0.06 (3.48)	0.04 (2.84)
InN	0.27 (6.49)	0.29 (14.11)
InK	0.27 (5.82)	0.27 (8.12)
InMAT	0.32 (5.51)	0.37 (8.48)
IMP	-0.13 (-1.15)	-0.21 (-2.33)
EXP	0.20 (2.54)	0.25 (3.25)
OVER	-0.0003 (-0.03)	0.006 (0.51)
CONC	-0.0003 (-0.10)	0.0001 (0.06)
$\mathbf{R^2}$		.998
$\mathbf{X^2}$ (p-value) Hausman test on Hausman-Taylor-RE versus Within Estimation of all time-varying variables	4.17 (.999)	-

Notes: t-values in brackets, n = 1630, 179 firms, time-dummies not reported

The coefficient of the CÔD80 dummy represents the productivity gain of the codetermined firms that switched from one third to 'almost parity', after 1976. The effect is rather small, but we interpret the significance of this coefficient as evidence against the hypothesis that codetermination is necessarily bad for efficiency. In our view this result has considerable policy relevance. In light of the recent controversial discussion on the introduction of a codetermination law on the EU level, according to our data and testing this will not disadvantage the European firms in question. Column one presents the results from a regression without the codetermination variables, which is made in order to test for the robustness of the results and to show the result of the Hausman-test.

The coefficients of the production factors are relatively small in particular if the results of the Hausman-Taylor model are considered. It is not uncommon that fixed effects estimations (or deviations from the mean) affect the magnitude of the coefficients of the production factors and reduce them. The coefficient of CÔD is much larger in the case of the Hausman-Taylor estimations and apparently a part of the size effect is captured by this variable and reduces the magnitudes of the coefficients of the production factors. Imports, and exports have a relatively strong impact.

It is also possible that the weaker, one third parity codetermination rights relevant for firms with 500 or more employees before 1976 had some impact. In order to test for this possibility we (re)define the following variables: COD5-20 is now a dummy for firms with more than 500 but less than 2000 workers. COD and COD80 are defined as above for firms with over 2000 employees. The small firms with less than 500 employees are the control group. The dummy-variables are again instrumented by the individual mean values of IMP, EXP; OVER, CONC and industry dummies. Aside of the results of the Hausman-Taylor model, simple random effects estimations are presented as well. All results are found in table 3 below.

This alternative estimation does not alter the conclusions. Almost -parity codetermination remains significant. The coefficient of one- third codetermination in the Hausman-Taylor specification also has a significant coefficient. However we are unable to carry out a before-after comparison, as the relevant legislation was introduced in 1952. As already noted, the co-determination that we consider is restricted to the supervisory board and this body does not determine wages (except the salaries of top management). Wages are set at

industry level. Therefore rent sharing can only be implemented by job – protection or labour hoarding, on which we have no empirical evidence.

Table 3: Cobb-Douglas-Production Function with Pooled Data, Dependent Variable Sales

Period VARIABLE	**Hausman-Taylor** **1972 - 1976,** **1981 –1985**	**Random Effects** **1972-1976,** **1981-1985**
Constant	3.18 (11.10)	2.48 (19.22)
CÔD	0.25 (4.03)	0.033 (1.14)
COD5-20	0.07 (2.02)	0.03 (1.65)
CÔD80	0.06 (3.45)	0.04 (2.80)
InN	0.27 (6.38)	0.29 (13.98)
InK	0.26 (5.71)	0.26 (8.05)
InMAT	0.32 (5.50)	0.37 (8.46)
IMP	-0.11 (-0.93)	-0.13 (-1.24)
EXP	0.17 (2.05)	0.21 (2.69)
OVER	-0.003 (-0.29)	0.008 (0.75)
CONC	-0.0007 (-0.32)	0.0006 (0.31)
$\mathbf{R^2}$		0.998
$\mathbf{X^2}$ (p-value) Hausman test on Hausman-Taylor-RE versus Within Estimation of all time-varying variables	4.17 (.999)	-

Notes: t-values in brackets, n = 1630, 179 firms, time-dummies not reported

5 Conclusion

We have explored the impact of two kinds of board level, worker codetermination on the productivity of firms. The move to almost parity after 1976 seems to increase productivity slightly in the affected firms. One-third codetermination also has a positive coefficient in one specification. There is certainly no suggestion of the negative effects predicted by opponents in both industry and academia.

As emphasized initially, our short panels both include major recessions, and are thus likely to exaggerate the short-term costs of co-determination in the form of more labour hoarding. This could not be explicitly tested due to lack of data on working time. The long term benefits from greater job security and consequent worker commitment and investment in firm specific skills noted by Gorton and Schmid(2001), Freeman and Lazear (1995) and many earlier writers would probably require longer data series to be identified. Although co-determination had a statistically significant positive effect only after 1976, the fact that we find no evidence of any deterioration in productivity or growth following the move to almost parity in 1976 does reject the critical view that this (political) move was primarily redistributional. Theoretical arguments suggest there is scope for real efficiency gains in terms of both productivity and job satisfaction. The latter is notoriously difficult to measure, and the former is probably most evident in a long run context that goes beyond our static model and limited data. More research is clearly needed to explore the ramifications of employee involvement.

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Appendix 1: The Hausman-Taylor instrumental variable estimator

Although this approach is very useful in many circumstances, it is rarely used in practice. This is most likely the case, because the application is very complicated. Another reason might be that the original article and the presentation in textbooks is not very user-friendly (Baltagi 1995) or too short for practical use (Verbeek 2000, Wooldridge 2002). Therefore we present a simple summary of the model.

The crucial difference between the random effects model and the fixed effects model are the assumptions about the correlation between the individual specific effects and the explanatory variables. If no time-invariant variables are included, both models can be estimated and a Hausman test on the differences between the coefficients can be applied concerning the question whether a significant correlation is present or not.

If one or more time-invariant variables are used, the fixed effects model cannot be estimated and a comparison with the random effects estimator is also impossible. A solution for this frequent problem is the Hausman-Taylor estimator. In their model, some variables are related to individual specific effects, while others are not. Hausman and Taylor consider four types of variables:

X_1 time variant variables, which are not correlated with the individual fixed effects

X_2 time variant variables, which are correlated with the individual fixed effects

Z_1 time invariant variables, which are not correlated with the individual fixed effects

Z_2 time invariant variables, which are correlated with the individual fixed effects

A fixed effects estimator will result in consistent estimates for the time-variant variables, however the time invariant covariates would vanish. Despite of this, in the first place a within estimation is carried out, based on the variables Y, X_1, X_2 (Y is the dependent variable) in deviations from their mean ($\tilde{X}_{1it} = X_{1it} - \bar{X}_{1i}$), leading to

$$\tilde{y} = \tilde{X}_{1it}\beta_1 + \tilde{X}_{2it}\beta_2 + \tilde{\nu}_{it}$$

The residuals are used to calculate the variance of the idiosyncratic error component:

$$\sigma_v^2 = \frac{\sum \tilde{\nu}_{it}}{N(T-1)}$$

with N the number of units and T the number of period is. K stands for the number of variables. In the next step, group means of the residuals are then used as the dependent variable in an instrumental variable regression on Z_1 and Z_2 where Z_1 and $\bar{X}_1$ are the instruments. These estimation is used in order to obtain an estimate of the variance of the random effect:

$$\hat{e}_i = (\bar{y}_i - \bar{X}_{1i}\beta_{1w} - \bar{X}_{2i}\beta_{2w}) - Z_{1i}\gamma_{1IV} - Z_{2i}\gamma_{2IV}$$

which produces

$$s^2 = \frac{1}{N}\sum_{i=1}^{N}\left(\frac{1}{T_i}\sum_{i=1}^{T_i}\hat{e}_{it}\right)^2$$

The total variance is accordingly (for the case of an unbalanced panel)

$$s^2 = \bar{T}\sigma_\mu^2 + \sigma_v^2 \text{ with } \bar{T} = \frac{\sum_{i=1}^{n} 1/T_i}{n}$$

Based on this estimator it is possible to calculate

$$\hat{\sigma}_\mu^2 = \frac{\left(s^2 - \hat{\sigma}_v^2\right)}{\bar{T}}$$

We modify this approach in the that we calculate the observation specific $\hat{\sigma}_{\mu_i}^2$ (Greene 2000, 580) and therefore in our case the random effects weight reads:

$$\hat{\theta}_i = 1 - \left[\frac{\hat{\sigma}_v^2}{\hat{\sigma}_v^2 + T_i\hat{\sigma}_{\mu_i}^2}\right]^{1/2}$$

The final step is the random effect instrumental variable estimation with the following variables:

$$w_{it}' = (Y_i, X_{1it}', X_{2it}', Z_{1it}', Z_{2it}')$$

These variables are transformed accordingly

$$w_{it}^{*'} = w_{it}' - \hat{\theta}_i\bar{w}$$

The instrumental variables are in turn:

$$v_{it}^{'} = \left[(X_{1it} - \bar{X}_{1i})^{'}, (X_{2it} - \bar{X}_{2i})^{'}, Z_{1i}^{'}, \bar{X}_{1i} \right]$$

The instrumental variables for X_1, X_2 are their deviations from the means. (It doesn't matter whether the deviations form means or the levels of the X_1 variables are used.) The instrumental variables for Z_1 are the variables themselves and finally the instrumental variables for the critical variables Z_2 are the means of X_1, as these variables are independent of the individual fixed effects. Hence the great advantage of the Hausman-Taylor instrumental variable model is, that all variables are identified in the estimation provided that the number of variables X_1 is at least as large as Z_2. No additional instrumental variables are needed for identification, as the X_1 variables identify themselves via deviations from means and identify the time-invariant variables Z_2 by their means. However if additional instrumental variables are available, they can be used in order to improve efficiency. (Verbeek 2000, Wooldridge 2002). We use the industry dummies in addition to the means of X_1.

Subsequently Amemiya and MaCurdy (1986) as well as Breusch, Mizon and Schmidt (1989) proposed different sets of instrumental variables, but these are based on rather restrictive exogeneity assumptions. Therefore we do not consider these models.

We modify this approach in the following way: As our time-invariant variables codetermination is a dummy-variable, we use probit instead of a least square procedure to instrument it. In the standard Hausman-Taylor model codetermination would be predicted by use of least squares, which implies non-normality and heteroscedasticity of the residuals and, most importantly, does not exclude the possibility, that the predicted variable is outside of the zero-one interval. Predictions outside of the unit interval would be very awkward in our case. Therefore we prefer a standard method for binary variables, namely probit at this stage. Tests with the estimation of a heteroscedastic probit model were unsuccessful, as the model did not converge.

Now the question arises, from where the researcher knows, what variables are affected by the individual effects (X_2) and which ones are not (X_1)? The Hausman-Taylor approach can be varied with respect to the variables classified into the two groups, then two

estimations are carried out, one based on a within and one based on a random effects models and Hausman tests on differences between coefficient vectors show, whether the coefficients are affected by the modification of the estimation method.