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Has the EU's Single Market Programme Fostered Competition? Testing for a Decrease in Markup Ratios in EU Industries

Harald Badinger

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Editorial

On the occasion of the 65th birthday of Governor Klaus Liebscher and in recognition of his commitment to Austria's participation in European monetary union and to the cause of European integration, the Oesterreichische Nationalbank (OeNB) established a "Klaus Liebscher Award". It will be offered annually as of 2005 for up to two excellent scientific papers on European monetary union and European integration issues. The authors must be less than 35 years old and be citizens from EU member or EU candidate countries. The "Klaus Liebscher Award" is worth EUR 10,000 each. The winners of the third Award 2007 were Harald Badinger and Gert Peersman. Harald Badinger's winning paper is presented in this Working Paper, while Gert Peersman's contribution is contained in Working Paper 136.

In this paper Harald Badinger uses a panel approach to test whether the EU's Single Market Programme has led to a reduction in firms' markups over marginal costs. The analysis covers 10 EU Member States over the period 1981 to 1999, for each of three major industry groups (manufacturing, construction, and services) and 18 more detailed industries. The paper addresses explicitly the uncertainty with respect to the timing of the changeover and allows for a possibly continuous regime shift in a smooth transition analysis. Where regime shifts can be found, the velocity of transition is extremely high, making the linear model a justifiable approximation. The author also tests for discrete structural breaks in the time window from 1988 to 1996, taking up endogeneity concerns in a GMM framework. Markup reductions are found for aggregate manufacturing (though it is also suggested that mark-ups increased in some manufacturing industries in the pre-completion period at the end of the 1980s) and – less robustly – for construction. In contrast, markups have gone up in most service industries since the early 90s, which confirms the weak state of the Single Market for services and suggests that anti-competitive defence strategies have emerged in the 1990s in service industries.

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Has the EU's Single Market Programme fostered competition?
Testing for a decrease in markup ratios in EU industries*

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Abstract: We use a panel approach, covering 10 EU Member States over the period 1981 to 1999, for each of three major industry groups (manufacturing, construction, and services) and 18 more detailed industries to test whether the EU's Single Market Programme has led to a reduction in firms' markups over marginal costs. We address explicitly the uncertainty with respect to the timing of the changeover and allow for a possibly continuous regime shift in a smooth transition analysis. Where regime shifts can be found, the velocity of transition is extremely high, making the linear model a justifiable approximation. We also test for discrete structural breaks in the time window from 1988 to 1996, taking up endogeneity concerns in a GMM framework. Markup reductions are found for aggregate manufacturing (though it is also suggested that markups increased in some manufacturing industries in the pre-completion period at the end of the 1980s) and – less robustly – for construction. In contrast, markups have gone up in most service industries since the early 90s, which confirms the weak state of the Single Market for services and suggests that anti-competitive defense strategies have emerged in the 1990s in service industries.

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I. Introduction

The Single Market Programme of the European Union (EU) has been in effect for more than 10 years now. Since 1 January 1993, goods, services and production factors move freely across intra-EU borders. Notwithstanding remaining problems in coverage and implementation, the removal of barriers to trade and factor flows together with the introduction of flanking policies, most notably the EU's common competition policy, marks another milestone in the economic integration of the European economies and completes the vision of a Common European market, dating back to the treaties of Rome in 1957.

The Single Market (SM) Program, launched by the European Commission (1985) as a remedy against the 'Eurosclerosis', was designed to increase competition in the European Markets and to improve the EU's international competitiveness, and thus ultimately to enhance efficiency and welfare. Smith and Venables (1988) simulated the potential welfare gains from two consequences of the removal of barriers: the reduction of trade costs and full market integration. Since the latter channel turns out more important in their analysis, the investigation of the Single Market's pro-competitive effect should be at the heart of an evaluation of the track record after 10 years.

Unfortunately, however, evidence on SM effects is still very limited. The Commission's review of the Single Market of 1996 provides an analysis up to 1992; this was clearly too early to give a conclusive ex-post assessment. Only a few further studies were carried out since then. Allen et al. (1998), building on their work in the Commission's review, use data up to 1994; they derive the SM's effect on price-cost margins from the estimation of price and demand functions for 15 'selected' industries (assumed to be particularly sensitive to the SM according to Buiges et al. (1990)) of the four largest EU countries (DE, FR, IT, and the UK). Their analysis provides valuable insights by combining these estimates with a welfare analysis in a CGE framework. Nevertheless, in light of the fact that their sectors make up only one third of total manufacturing output, and the time period considered, their conclusion that the SM "has indeed had a substantial pro-competitive effect in European markets, and has led to significant reductions in price-cost margins" (Allen et al., 1998, p. 467) has to be interpreted with caution (see the comment by Flam (1998) for a criticism).

At the country level Bottasso and Sembenelli (2001) use a similar industry classification and a large sample of Italian firms to test for a structural break due to the Single Market, using the markup-estimation method suggested by Hall (1988). Again, significant reductions in markups (and increases in productivity) are only found for the group of "most sensitive firms". The most recent study is Sauner-Leroy (2003), which covers 9 European countries and the period from 1987 to 2000. It uses data from firms' financial statements of the Commission's BACH database, aggregated at the manufacturing level. These data enable him to directly calculate price-cost margins and to test for the impact of the Single Market in a simple regression framework with further control variables. Though country-specific results differ somewhat, the analysis suggests that markups decreased in the period from 1987 to 1992, along with a decrease in prices; in the post-completion period from 1993 to 2000, however, markups recovered in line with the realization of efficiency gains. Summing up, the overall evidence on the Single Market's achievements is mixed at best and far from comprehensive.

This paper adds some further results on the pro-competitive effect of the Single Market, supplementing the previous studies both in its scope and methodology. We use a sample of 10 European countries and 18 sectors (including 5 service sectors on which no evidence exists so far), which covers the period from 1981 to 1999. Thus our level of disaggregation is between the studies by Allen et al. and Sauner-Leroy. As to the method, we employ the Roeger (1995) approach for markup estimation, which none of the previous studies on the consequences of the SM has adopted so far. Using a panel approach, we test for a possibly continuous regime shift using smooth transition analysis (see Granger and Teräsvirta, 1997) as well as for a discrete regime shift and check the robustness with respect to endogeneity concerns, which are likely to be present in the Roeger approach as well as shown by Hylleberg and Jorgenson (1998).

In sum, evidence for a pro-competitive effect is mixed. We find significant reductions in markups in manufacturing and to some extent in construction. As far as services are concerned, results are less encouraging: markups appear to have gone up in the service sectors since the mid-90s, reflecting the weak state of implementation in the SM in services. The remainder of the paper is organized as follows. Section II briefly describes the methods of markup estimation used. Section III contains a description of the data. Section IV discusses the pertinent issues in identifying effects of the Single Market. Section V sets up the empirical models and presents the estimation results. Section VI summarizes the main conclusions.

II. Markup estimation: methodological background

Our approach to estimating the markups factors relies on the paper by Roeger (1995). This approach is an extension and variant of the seminal paper by Hall (1988), who showed that the Solow residual under market power is

$$\Delta \ln q_t - \alpha_t \Delta \ln l_t = (\mu_t - 1)\alpha_t \Delta \ln l_t + \Delta \ln e_t\,, \qquad (1)$$

where q_t is the output/capital ratio (Q_t/K_t), l_t is the labour/capital ratio (L_t/K_t), e_t stands for the level of Hicks-neutral technological progress, α_t is the factor share earned by labour (i.e. the ratio of labour compensation L_tW_t to total revenue $Y_t = P_tQ_t$), and μ_t is the markup ratio P_t/MC_t (MC denoting marginal costs). Assuming a constant markup ratio, μ can be estimated from (1). The problem, however, is the endogeneity of the right hand side variable; thus instruments, i.e. variables correlated with output which are neither the cause nor the consequence of technological change, are required for a consistent estimation and valid inference. Hall, in his empirical analysis of US industries, uses military expenditures, the political party of the president and the oil price. Obviously, it is hard if not impossible to find good instruments that are exogenous under all views of macroeconomic fluctuations.

Roeger (1995) develops an approach that avoids some of these problems, by first recognizing that the primal technology residual given by (1) (which is calculated from the production function), can also be written in extensive form as

$$(\Delta\ln Q_t - \Delta\ln K_t) - \alpha_t(\Delta\ln L_t - \Delta\ln K_t) = \beta(\Delta\ln Y_t - \Delta\ln K_t) + (1-\beta)\Delta\ln e_t\,, \tag{2}$$

where the parameter β corresponds to the Lerner index which is directly related to the markup ratio via $\mu = 1/(1-\beta)$. He then derives the price based Solow residual (calculated from the dual cost function), which is given by

$$\alpha_t\,\Delta\ln W_t + (1-\alpha_t)\Delta\ln R_t - \Delta\ln P_t = -\beta(\Delta\ln P_t - \Delta\ln R_t) + (1-\beta)\Delta\ln e_t\,, \tag{3}$$

where R_t denotes the user costs of capital.

Under perfect competition ($\mu = 1$ or $\beta = 0$), both the primal and the dual Solow residual are an exact measure of technological progress (leaving measurement problems aside). Under imperfect competition, prices depart from marginal costs and the technology residual can be decomposed into a technical innovation term and i) the rate of change in the capital productivity, multiplied by β (primal residual, see (2)), or ii) the rate of change in output prices minus the rate of change in capital costs, also multiplied by β (dual residual, see (3)).

Substituting the expression for $\Delta\ln e_t$ implied by (3) into (2), Roeger derives the following expression suitable for the estimation of β:

$$\begin{aligned}(\Delta\ln Q_t + \Delta\ln P_t) - \alpha_t(\Delta\ln L_t + \Delta\ln W_t) - (1-\alpha_t)(\Delta\ln K_t + \Delta\ln R_t)\\ = \beta[(\Delta\ln Q_t + \Delta\ln P_t) - (\Delta\ln K_t + \Delta\ln R_t)] + u_t\,,\end{aligned} \tag{4}$$

where u_t is a standard error term. Equation (4) gives the difference between the primal and the dual residual; under perfect competition it should equal zero. To simplify notation, we rewrite (4) as

$$z = \beta x + u_t \,, \tag{5}$$

where z may be interpreted as nominal Solow residual, and x is the growth rate of the nominal output/capital ratio; u_t is an error term reflecting the difference of the measurement errors from the two productivity terms. The attractive feature of this apprach, at least at a first glance, is that the productivity term vanishes and that no instruments are needed for the estimation of β. Additionally, by focussing on nominal variables it allows to overcome some problems in data availability.

It should be noted that both (1) and (5) are derived under the assumption of constant returns to scale; there is, however, good reason to believe that in many cases, market power exists as a results of economies of scale. Martins et al. (1996) and Hylleberg and Jorgensen (1998), show that under increasing returns, (5) becomes

$$z = [\lambda(\beta - 1) + 1]x + u_t \,. \tag{6}$$

where λ is an index of returns to scale, defined as ratio of average to marginal costs. It follows that the estimates of β and μ are downward biased in the presence of increasing returns. Similarly, the markup over marginal costs is underestimated in the presence of sunk costs, downward rigidities of the capital stock or labour hoarding; thus it has been suggested to interpret the markup implied by the estimate of β from (5) as lower bound (Martins et al., 1996).

III. Data

The most comprehensive database providing consistent data for European countries at the industry level is the STAN database of the OECD. Still, the available data are not perfectly suitable for our investigation: in particular, sufficient data is available only for employment instead of hours worked so that our estimates may be biased as a result of labour hoarding. Also, data for gross value added is mainly available only at basic prices (rather than at factor costs) and thus includes some indirect taxes (other than the value added tax), inducing an additional upward bias. This has to be borne in mind when interpreting the results.

As to the country coverage, Ireland is not in the database; four further countries (Luxembourg, Greece, Denmark, and Portugal) had to be excluded due to numerous missing values, mainly with respect to fixed capital formation. Thus we had to reduce our sample to 10 countries (Austria, Belgium, Finland, France, Western Germany, Italy, Netherlands, Spain, Sweden, and United Kingdom); for some sectors, further countries (mainly Sweden, the United Kingdom and Spain) had to be excluded. As far as the time dimension is concerned, we restrict our estimation period to the time from 1981 to 1999 in order to avoid the use of unjustifiably imbalanced samples.

Our goal to set up a sample comprising most of the EU countries prevented us from moving down to very detailed industry level, because this would have required a further reduction in the cross country dimension; the level of aggregation used here is in between the Allen et al. (1998) and the Sauner-Leroy (2003) studies. Table 1 gives an overview of our sample, which covers 10 countries, 18 industries (and three aggregates) and the period from 1981 to 1999.

< Table 1 here >

Three further issues are worth mentioning: First, as opposed to previous studies, we do not restrict our attention to so called ‘sensitive industries’. After all, the goal of the Single Market was to make European industries more competitive, not only selected samples of firms; thus, a substantial change in market structures should also show up at the level of aggregation used here, if evidence is more than episodic. Table 1 also gives an overview of the quantitative importance of the respective industries in terms of their shares in total manufacturing and total value added respectively.

Second, as a first study on the Single Market we also include service sectors on which no evidence exists so far. However, this is a delicate choice, bearing in mind that the available data and standard capital stock measures may be less reliable there. Thus evidence on service sectors must be treated with caution.

Finally, like Hall and Roeger we use value added rather than output; equation (4) is specified in terms of output, however. As pointed out by Hall, this induces an upward bias in the estimated markup ratio, but under the assumption of a roughly constant value added/gross output ratio (V), the 'true' mark up ratio over marginal costs can be computed by adjusting the estimate of β with the factor (1-V). Table 1 shows the averages (over country and time) of the gross value added to output ratio for the different industries.

IV. Modelling the changeover: issues in identifying the Single Market effects

We go on to estimate a panel data model for each of the three major industry groups and each of the 18 more detailed industries. The empirical model corresponding to (5) is given by

$$z_{i,t} = \alpha_i + \beta_1 x_{i,t} + u_{i,t}\ , \tag{7}$$

where $u_{i,t}$ is a standard error term. The cross-section dimension (i) ranges from 7 to 10 countries (depending on the data availability in the respective sector) and t denotes time (t = 1981, . . . , 1999).[1] Restricting the coefficient to be equal across countries ($\beta_{1i} = \beta_1$), which is equivalent to assuming the same markup for all countries, is clearly a somewhat unrealistic assumption. However, if markups do not differ dramatically accorss countries, which is a reasonable assumption, focussing on the pattern of the average value may be justified in light of the purpose of our study. The same argument is used by Botasso and Sembenelli for their

[1] Since we estimate one separate panel for each sector we do not introduce another subscript to denote the industry dimension.

study using Italian firms.[2] In general terms the hypothesis of a structural change in markups can be written as

$$z_{i,t} = \alpha_i + \beta_1 x_{i,t} + \beta_2 \mathrm{F}(t) x_{i,t} + u_{i,t}\ , \tag{8}$$

where F(t) is an transition function, which remains to specified.

Some major issues have to be addressed in identifying potential effects of the SM on markups. First, how can we be ‘sure’ that a structural break occurring around 1993 can be attributed to the SM and is not due to other changes occurring simultaneously? In econometric terms: which controls should be added? The most obvious and important issue to be controlled for is the state of the economy. There is a long debate in the literature on the relation between the business cycle and markup ratios. Rotemberg and Woodford (1991), as a prominent example, find evidence for a counter-cyclical markup ratio, but the issue is unsettled both theoretically and empirically.[3] We include the economy wide output gap (*GAP*) as a control variable, to ensure that our results are not drive by business cycle effects.

A further indicator one might want to include is concentration. This variable, however, had to be omitted mainly for reasons of data availability; there are no time series of concentration measures (such as C4 or the Herfindahl index) for all EU countries, not even at our fairly crude level of aggregation. But there is a further reason to leave concentration out: The study by the Commission (1996) suggests that the evolution (i.e. the increase) of firm concentration since the mid-80s was itself strongly influenced by the SM programme and its

[2] At his point, assuming parameter heterogeneity would only reduce the degrees of freedom. Later on, when we adopt a GMM approach for estimation, the model would become practically inestimable, if parameter heterogeneity were allowed for.

[3] Domowitz et al. (1988), for example, find mild evidence for a pro-cyclical markup. A new interpretation is given by Banerjee and Russel (2004): using the dataset of Rotemberg and Woodford, they argue that markups are best modelled as integrated process and cannot be related to (stationary) business cycles measures in the long-run.

announcement in the Commisson's White paper. Thus, with a view to our goal of estimating the Single Market's bottom line effect, the omission of concentration may also be justified from an economic perspective. Nevertheless, it would be clearly desirable to disentangle the SM effects more in detail in future research.[4]

Finally, the panel approach allows us to include time specific effects, which helps us to control for common shocks (and common components of the business cycle), which – if related to markups – may also overlap and dilute (or enforce) the estimated effects of the Single Market. Thus, our final empirical model is

$$z_{i,t} = \alpha_i + \beta_1 x_{i,t} + \beta_2 \mathrm{F}(t) x_{i,t} + \beta_3 GAP_{i,t} + \eta_t + u_{i,t} \,. \tag{9}$$

But even if we assume that all crucial controls have been included, there is another important issue to be addressed in the estimation. Previous studies simply assume a discrete break, i.e. an instantaneous change in the markup in a certain year. The most obvious choice of the break year is 1993 (as in Sauner-Leroy (2003)); Allen et al. (1998) use the year 1992; Botasso and Sembenelli (2001), in their investigation for Italian firms, consider data only up to 1993 and assume the break to have occurred in 1988. These choices may appear somewhat arbitrary (and are to some extent motivated by data availability at the time of the study), but reflect a more fundamental uncertainty: the SM was announced in 1985 by the Commission's White Paper on the completion of the internal market and is thus likely to have been anticipated by rational agents (note the wave of mergers and acquisitions that followed the announcement in the late 1980s). Thus (part of) its effect may have set in even before 1993; on the other hand

[4] Import penetration is a further candidate; again we omit this variable since the SM is likely to have affected intra-EU trade. One could use extra EU import measures, but they are of comparably little importance and varied little in out period of investigation, relative to intra-EU measures. It is thus very unlikely that extra EU imports played a relevant role in shaping the development of the markups in the 90s (as it is also confirmed in several trial regressions). We thus omit import measures here as well from the beginning since they would add little at the cost of introducing (further) endogeneity problems.

implementation did not take place immediately (see, e.g., European Commission 2002a), suggesting that part of the effects materialized with some delay. At the same time, it is unclear a priori, how quickly the structural break (i.e. change in markups) was realized.

To deal with this problem, the most general approach is to specify a smooth transition model, where the change in markups is estimated simultaneously with the transition midpoint and the velocity of transition. Whether the requirement to use a non-linear regression model is a drawback or a virtue clearly depends on the nature of the transition process. If the actual transition process is highly non-linear, the regime shift will hardly be detected using a linear specification. On the other hand, if the changeover can be well approximated using a discrete change (and if we even have some reliable a priori information when the transition took place), the test for a regime shift using a linear model will have more power. In our particular case, it is reasonable to assume (and as will be confirmed by our estimates later on) that the regime shift occurred in the time window from 1988 to 1996. But we have no reliable information on the velocity of transition. We thus started with a smooth transition model to check for pronounced non-linearities in the transition and the see whether a discrete change model is a reasonable approximation.

Finally, endogeneity is also very likely to be an issue here. By eliminating the technology term the Roeger approach (5) was argued to eliminate (at least alleviate) endogeneity problems of the Hall approach (1). However, measurement error (in variables such as capital stock, user costs of capital) may be pronounced, particularly in service industries. But even if problems with measurement and omitted variables are left aside, Hylleberg and Jorgensen (1998) show that slightly relaxing the assumption of a constant markup makes the Roeger approach vulnerable for similar lines of criticism, i.e. the endogeneity of x. Hence it will be important to check the sensitivity of the results with respect to the exogeneity assumption.

V. Empirical models and estimation results

1. Smooth transition analysis

We start with the most general approach using smooth transition analysis. What remains in order to make (9) estimable is to specify the transition function F(t).We opt for a simple form and use a symmetric logistic function

$$\mathrm{F}(t) = \frac{1}{1+e^{[-\gamma(t-\tau)]}}\ , \tag{10}$$

which maps t onto the interval (0,1). Applied to our model (9) this allows for a smooth transition between the initial state ($t \to -\infty$)

$$z_{i,t} = \alpha_i + \beta_1 x_{i,t} + \beta_3 GAP_{i,t} + \eta_t + u_{i,t}\ , \tag{11}$$

and the final state ($t \to +\infty$)

$$z_{i,t} = \alpha_i + (\beta_1 + \beta_2)x_{i,t} + \beta_3 GAP_{i,t} + \eta_t + u_{i,t}\ . \tag{12}$$

The parameter γ determines the speed of transition, while τ is associated with the transition mid-point, since F(t) = 0.5 for $t = \tau$. For $\gamma \to \infty$, (9) collapses into a linear model, with an instantaneous structural break at $t = \tau$. Hence the nonlinear model (9) is more general, nesting the linear discrete break model as a special case. Of course, more general forms of the transition function F(t) are conceivable, using higher order polynomials in t and including the dependent and/or the exogenous variables. However, for our purposes, a transition process described by such a logistic smooth transition model (LST, see Granger and Teräsvirta, 1997, chapter 4) appears to be a reasonable choice and allows us to address our two main concerns: to allow for a gradual change and to endogenize the timing (i.e. the transition mid-point). Thus we have

$$z_{i,t} = \alpha_i + \beta_1 x_{i,t} + \beta_2 \frac{1}{1+e^{[-\gamma(t-\tau)]}}\, x_{i,t} + \beta_3 GAP_{i,t} + \eta_t + u_{i,t}\ . \tag{13}$$

The problem in testing the hypothesis of the constancy of the regression parameter β (i.e. H_0: $\gamma = 0$)[5] against the alternative of a continuous structural change is that τ remains unidentified under the null. Lin and Teräsvirta (1994) suggest to approximate F(t) using a Taylor series around $\gamma = 0$, which allows the reparameterization of (13) in terms of identified parameters. The null hypothesis $\gamma = 0$ can then be tested using a Lagrange multiplier (LM) test of excluding restrictions applied to this reparameterized model. We use a third order Taylor series approximation, which implies the use of interaction terms between x and t up to the third order. The second column of Table 2 reports the results of the χ^2-tests of the excluding restrictions for the various industries.[6] Since we find evidence for a break in most cases we go on to estimate the LST model for the full sample of all industry groups and industries.

For most industries we ran into convergence problems or obtained implausible coefficients when all parameters (α_i, β_1, β_2, γ, τ) were estimated at the same time. We thus pursue a grid search strategy, imposing the velocity of transition (γ) and estimating the other coefficients using nonlinear least squares; γ was varied from 0.2 to 10 with a step size of 0.01. As can be seen from Figure 1, which shows the corresponding transition functions for τ = 1993, these values cover a broad spectrum, ranging from a very slow transition process (γ = 0.2) to the case of a de facto instantaneous change (γ = 10). Since there is strong reason to assume a break a priori which is also enforced by the LM tests, the optimization criterion used was to minimize the p-value of the estimate of β_2. This approach may be tilted towards finding a break; in practice, minimizing the sum of squared residuals produces almost the

[5]As Lin and Teräsvirta (1994) show, F(t) can be transformed to F*(t) = F(t) - 0.5 without any loss of generality; in this case F*(t,0) = 0 for $\gamma = 0$, making $\gamma = 0$ the natural hypothesis for parameter constancy in (13).

[6] To be more specific, column (2) of Table 3 reports, for each industry, the LM-test of the joint hypothesis that $\delta_1 = \delta_2 = \delta_3 = 0$ in the test regression $z_{i,t} = \alpha_i + \beta_1 x_{i,t} + \beta_3 GAP_{i,t} + \delta_1 t x_{i,t} + \delta_2 t^2 x_{i,t} + \delta_3 t^3 x_{i,t} + \eta_t + u_{i,t}$. The F-test variant for the excluding restrictions (recommended by Lin and Teräsvirta for small samples) produces de facto identical results.

same results. Using this grid search strategy, quick convergence was achieved for all industries.[7] The results of the nonlinear least squares estimates for model (12) for the various industries are given as of column three in Table 2.

< Figure 1 here >

< Table 2 here >

We do not go into the details at this point but emphasize three important issues: First, the estimates of the average markups (β_1) all turn out significant at the one per cent level: not very surprisingly, there is strong evidence against perfect competition in European industries, suggesting that monopolistic and oligopolistic competition prevails. The (adjusted) markup ratios (P/MC) implied by the estimates of the Lerner index β_1 vary between 1.21 and 4.16. These figures are roughly in line with the results by Martin et al. (1996) who consider a sample of OECD countries in the pre-completion period 1970-1992, using a more detailed industry classification, however.

Second, we find significant changes in markups although with different transition points and different signs: in manufacturing markups went down from 1.38 to 1.29 since 1993 (a relative reduction by 28 per cent) and from 1.56 to 1.32 in construction since 1986 (by 43 per cent in relative terms). In contrast, markups have increase in services industries as of 1997 by 21 per cent (from 1.31 to 1.39).

Third, the coefficient of the output gap is mainly insignificant; most significant coefficients are negative, indicating a countercyclical markup. Overall, the results do not speak strongly on the relation between markups and cycles; since the output gap acts only as a control and is not our primary concern here, we do not pursue this issue further here.

Finally, and this is the most important message from our estimates so far, the velocity of transition is extremely high in all models. Where regime shifts occurred, they materialized

[7] Starting values were taken from the estimates of the linear break model (see equation (14)); τ was set to 1993.

extremely quickly. This is somewhat surprising, given the long (and still incomplete) way from the announcement of the SM to its implementation. In light of this result, however, it may be argued that nonlinearities are very weak and that the hypothesis of an instantaneous regime shift is a reasonable, justifiable approximation. This has the advantage of allowing us to sharpen our alternative hypothesis in testing for a structural break, and hence to improve the power of the test. Moreover, it brings us back to linearity which simplifies the estimation considerably. It also provides some ex-post justification for the previous studies that employed a linear model from the beginning.

2. Analysis assuming an instantaneous break

As mentioned above the instantaneous break model is obtained by letting t approach infinity in F(t). Practically, F(t) is replaced by a level dummy D^T, which takes values of 1 for $t \geq T$ and zero otherwise. Thus we have

$$z_{i,t} = \alpha_i + \beta_1 x_{i,t} + \beta_2 D_t^T x_{i,t} + \beta_3 GAP_{i,t} + \eta_t + u_{i,t} \,. \tag{14}$$

While a fast velocity of transition appears to be a reasonable assumption, there is still uncertainty with respect to the time of the break. Again we determine the breakpoint endogenously, estimating the model for alternative breakpoints ranging from 1988 to 1996 and choosing that which maximizes the p-value of the estimate of β_2.[8]

a) Results for three major industry groups

For the moment we focus on three major industry groups. The standard least square dummy variable (LSDV) estimates of the linear panel data model (14) are given in Table 3 (first row of panel with respective industry group). The results are comparable with that of our results from the LST models: again we find significant changes, which are de facto identical to that

[8] Again, the results using the sum of squared residuals criterion produces very similar results.

given in Table 2; for construction and services the breakpoints implied by our search within the time range 1988 to 1996 differ somewhat from the LST estimates; for manufacturing the implied breakpoints are de fact identical.[9]

< Table 3 here >

The linear framework offers a convenient way to address endogeneity problems discussed in section IV. Hylleberg and Jorgensen (1998), who show that the Roeger approach is likely to be vulnerable to similar endogeneity problems as the Hall approach, step back from an instrumental variable approach in their time series setting, given the absence of proper instruments. In our panel with a much larger number of observations, however, there are numerous moment restrictions that can be exploited in a GMM framework. We check the sensitivity of the results, starting with a GMM estimation in first-differences (GMM-FD) of model (14). Thereby the individual effects[10] are eliminated by first differencing. As instruments for the variable x in first differences ($\Delta x_{i,t}$, as well as $\Delta(D^T x_{i,t})$), all available lags of its level dated t-2 and earlier may be used ($x_{i,t-s}$, $s \geq 2$) Though originally introduced in the context of dynamic panels (Arellano and Bond, 1991), the extension of this GMM approach to problems like measurement error and endogeneity of the right hand side variables is straightforward (Bond et al. 2001). The validity of instruments requires absence of second order serial correlation in the residuals and can also be checked by means of a Sargan test.

In some cases, lagged levels may be only weak instruments for first differences, leading to a poor performance of the GMM-FD estimator. Blundell and Bond (1998) suggested

[9] Estimating all models assuming the breakpoint to coincide with the year 1993, when the SM came into force, we obtained only one significant results (GMM-SYS for 15-37). As outlined before, however, we argue that confining the test of SM effects only to the case of T = 1993 is to narrow.

[10] The correlation between $x_{i,t}$ and the individual effects (α_i) is not of primary concern here, but it is controlled for by the GMM approach as well.

extending the GMM-estimator to a system approach, where the equations in first differences are supplemented by the original equations in levels, where now lagged first differences are used as instruments for the endogenous variable. Other approaches, exploiting additional moment restrictions have also been suggested (see Baltagi, 2001, for an overview), none of which can claim to be superior under general conditions. We will focus on the GMM-FD and GMM-SYS approach here; even in this case we are not able to exploit all moment conditions, since the time dimension of our panel is fairly large, such that the instruments matrix grows quickly to dimensions that are hard to handle (i.e. make it invertible). Hence we restricted the maximum number of lags to be included to a number of 10. The results of the GMM estimation in first differences (GMM-FD) and in combined first differences and levels (GMM-SYS) are also given in Table 3. Appendix B shows more explicitly how the two GMM approaches are applied to model (14). Since the weighted two step estimator turned out rather sensitive against small variations in the lag length of the instruments for many industries we opted for the unweighted one step approach using robust standard errors. Unfortunately, the Sargan test cannot be used in this case since it rests on the assumption of homogeneity; for checking the validity of instruments we thus have to rely on the serial correlation tests (columns eight and nine of Table 3).[11] Only for manufacturing is the p-value of the test for second order serial correlation close to the 10 per cent level; we thus started our instruments from the third lag and check for third order serial correlation (which turns out insignificant); anyway, the change in the results is negligible.

As far as manufacturing and services are concerned, controlling for endogeneity buttresses the estimated change in markups. Both coefficients remain significant at the one

[11] Results for the Sargan test are ambiguous. The null of valid instruments tends to be rejected using the one step results. However, since this test is based on the homoscedasticity assumption, these results could be driven by heterogeneity; this is likely to be the case, since the Sargan tests of the two step estimates are never significant with p-values close to 1.

per cent level, suggesting that our result is not driven by misspecification. Results for construction, however, are weakened: with the GMM-FD approach the significance deteriorates and the coefficient becomes insignificant when the GMM-SYS approach is used. Note that also the breakpoint, i.e. the year with the minimum p-value for the estimate of β_2, changes from 1988 to 1992. Intuitively, one would have expected the requirement for EU wide tenders in the area of public projects to have a nonnegligible impact on the markups in the construction industry. The results provide some support for this presumption but are not completely robust against different estimation methods. There is no convincing explanation for the differences between the GMM-FD and the GMM-SYS estimator. Thus, the markup reduction in construction obtained in all estimations except the GMM-SYS approach should not be completely rejected but interpreted with due caution.

b) Results for detailed industries

We now step down to a more detailed industry classification to get some idea, which industries have been most affected. We show only the results from the GMM-SYS approach here, but for most industries the GMM-FD and LSDV estimates are rather close.

< Table 4 here >

For manufacturing, we find significant changes in 8 of the 13 industries. The fact that only half of the significant changes are negative can be reconciled with the aggregate effect obtained in Table 3, bearing in mind that i) the share of the sectors with a negative effect makes up more than one third, and ii) that the increases in markups are found mainly for the end of the 80s. Industries where a decrease in markups could be identified are textiles etc., rubber and plastic products, basic metals and fabricated metal products as well as part of machinery and equipment. With the exception of textiles, all regime shifts with markup reductions took place in the period 1994 to 1995, i.e. after the SM had come effectively into force.

At least with respect to the timing this result is in contrast to Sauner-Leroy (2003), who finds that markups fell in the period from 1987 to 1992, and recovered as of 1993 back to their old levels. Our estimates suggest the opposite: markups increased in some industries at the end of the 80s, and went down after the introduction of the SM. This is not implausible. In their evaluation of the Single Market the Commission (1996) documents significant increases in concentration and firm size after the announcement of the Single Market in 1985. Firms appear to have used their larger market power to increase their margins before the SM eventually came into force. Given the bird eye's perspective of our study it is difficult to compare our results with the Allen et al. study, who focus only on selected (particularly sensitive) sectors and at a much lower level of aggregation. Our findings suggest that their conclusion that the SM "has indeed had a substantial pro-competitive effect in European markets" (Allen et al. 1998, p. 467) cannot be generalized without qualification. Their results for selected manufacturing industries hold up on average at the aggregate manufacturing level and (less robustly) for construction as well, but evidence for services is different.

In none of the service industries, a decrease in markups could be identified. In contrast: markups have increased since the early 90s, particularly in wholesale and retail trade, hotels and restaurants and financial intermediation. This result confirms the Commissions assessment in its report on the state of the internal market in services that there is "a huge gap between the vision of an integrated EU economy and the reality as experienced by European citizens and European service providers." (European Commission, 2002b, p. 70). This may explain the absence of pro-competitive effects, but not why markups have even gone up. A tentative and somewhat speculative explanation for this puzzling result could be as follows: During the 1990s firms appear to have developed strategies, partly in response to the SM (or to its non-functioning) and globalisation, which may have actually reduced competition. Two such practices are observed rather frequently in service industries according to the European Commission (2002b, pp. 55-69): i) Anti-competitive defensive strategies of service firms; *existing constraints on* cross-border green-field investments together with the risk of becoming acquisition targets create incentives for service firms to grow through acquisition or agreements at national level to defend their markets. The merger between Promodes-Carrefour in response to the acquisitions of Wal-Mart in Europe in 1998-1999 is cited as the most impressive example for many other cases. ii) Arrangement strategies arising from legal uncertainty and the lack of enforcement. Obstacles to market entry that are hard to remove often force service providers to negotiate an arrangement with national authorities or to form a partnership with local firms to circumvent the resistance of certain authorities to grant access to their national market. Once such an arrangement is obtained, interest in eliminating barriers (by the involved parties) disappears and the arrangements act as a further entry barrier to competitors, buttressing the anti-competition effects. Such practices are particularly harmful in service industries, which are intricately intertwined, by triggering knock-on effects from one service industry to another and multiplying through the chain of value added. These explanations are suggestive, but more research is needed for a fully convincing explanation for the puzzling increase in markups in European service industries despite apparent progress in European and international integration throughout the 1990s.

VI. Summary and conclusions

This paper investigates the pro-competitive effects of the EU's Single Market in terms of its effect on firms' market power as measured by the Lerner index. Using a panel of 10 EU Member States covering the period 1981 to 1999 for each of three industry groups (manufacturing, construction, services) and 18 more detailed industries we test for structural breaks in the framework of the markup estimation method suggested by Roeger (1995). We explicitly address the uncertainty with respect to the timing of the changeover, allowing for a possibly continuous regime shift and an endogenous breakpoint in a smooth transition analysis. Where regime shifts can be found, however, the velocity of transition turns out extremely high, making the linear model a justifiable approximation. We thus also test for discrete structural breaks between 1988 and 1996, the time window where SM effects are likely to have occurred. Endogeneity concerns are addressed in a GMM framework as well.

What previous studies have found for selected manufacturing industries and some countries, is confirmed more generally by our results for aggregate manufacturing of 10 EU Member States: markups have been reduced since the early 1990s, down from 1.41 to 1.28 (a relative reduction by 32 per cent). At a more detailed level, it is also found that in some industries markups increased in the pre-completion period around the end of the 1980s, which may be explained by an increase in concentration and average firm size in European industries in this period. Firms appear to have used their increase in market power to raise their markups; after the SM came into force the increased competition has decreased markups again. This reflects two fundamentally different effects of economic integration: on the one hand it spurs competition; on the other hand it may also result in a concentration of firms with potentially opposite effects. This paper has only considered the bottom line effect, but disentangling the Single Market effects more in detail is a challenging line for future research: in light of the data limitations, however, these questions may be better addressed in single country studies, using a more detailed level of aggregation or firm data as well as detailed case studies.

For construction, we also find reductions in markups, down from 1.37 to 1.32 since 1992, although this result is not completely robust against alternative estimation methods. The most puzzling results, however, is that markups have increase in services industries since the mid-1990s, up from 1.31 to 1.39 (by 21 per cent). This reflects the weak state of the Single Market for services, but the anti-competitive effect deserves further explanation. It is likely that the (non-functioning of the) SM and globalisation have induced anti-competitive strategies by firms such as acquisitions at the national level to defend the market and ‘arrangements’ with local partners and authorities to circumvent existing obstacles. These strategies, which are quite prevalent in service industries according to the European Commission (2002b), may even have reduced competition in certain industries.

The economy-wide importance of services – they account for some 70 per cent of GDP and employment in most EU Member States – makes this a particularly alarming result. The European Commission (2004) has launched a new proposal, which is designed to abolish the many remaining barriers in the service industries. Its success in enabling firms to realize the full potential for cross-border services in the Single Market will be of considerable importance for the EU’s future economic development, given that the improvement of the functioning of the Single Market has been identified as one of the EU’s chief requirements to enhance its growth performance (see Sapir et al., 2004).

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Appendix A: Data sources and definition of variables

$Q_{i,t}$ = real gross value added, from OECD: *Industry Structural Analysis* (STAN).

$P_{i,t}$ = deflator of gross value added, calculated from ratio of nominal to real gross value added; taken from OECD: STAN.

$K_{i,t}$ = real capital stock in millions of national currency, calculated using a perpetual inventory method: $K_t = K_{t-1}(1-\delta) + I_{t-1}$. The (industry specific) depreciation rate (δ) was calculated from data on average service life in the respective sector from the International Sectoral Database (ISDB) of the OECD. Initial value of capital stock was calculated according to $K_{1975} = \bar{I}/(g_{I,70-00} + \delta)$, where $\bar{I}$ is average investment over the period 1970 to 1980, $g_{I,70-00}$ is growth of investment over the period 1970-2000 (some time periods had to be adjusted for some countries/sectors due to data availability) (see Grilliches 1980, Coe and Helpman 1995). $I_{i,t}$ is real gross fixed capital formation calculated from nominal gross fixed capital formation and deflated with the respective investment deflator OECD: STAN.

$R_{i,t}$ = user costs of capital, approximated by $R_{i,t} = ((i-\pi^e)+\delta)\,P^*_{i,t}$ as in Martins et al. (1996); i is the country-specific long-term nominal interest rate, π^e is the expected rate of inflation, proxied by the HP-filtered component of the GDP deflator (λ = 500), and $P^*_{i,t}$ is the economy wide deflator for business investment (Source: OECD).

$L_{i,t}$ = total employment in million persons, taken from OECD: STAN.

$W_{i,t}$ = average nominal wage in sector i, given by $LC_{i,t}/L_{i,t}$, where LC is labour compensation in millions of national currency, taken from OECD: STAN.

$\alpha_{i,t}$ = share of labour compensation in nominal gross value added ($LC_{i.t}/\ Q_{i.t}P_{i.t}$).

Notes: DE refers to West Germany; data as of 1993 constructed using growth rates of the reunified Germany. Data were taken from the SourceOECD database.

Appendix B: GMM estimation of model (14)

a) GMM estimation in first differences (Arellano and Bond, 1991)

Applied to model (14) in section V we obtain

$$\Delta z_{i,t} = \beta_1 \Delta x_{i,t} + \beta_2 \Delta(D_t^T x_{i,t}) + \beta_3 \Delta GAP_{i,t} + \Delta \eta_t + \Delta u_{i,t} \tag{B1}$$

for $t = 1983, \ldots, 1999$ and $i = 1, \ldots, N$, where $x_{i,t-2}$ and all previous lags are used as instruments for $\Delta x_{i,t}$ and $\Delta(D_t^T x_{i,t})$ assuming that $E[u_{i,t} u_{i,s}] = 0$ for $i = 1, \ldots, N$ and $s \neq t$ and exploiting the moment conditions that $E[x_{i,t-s} u_{i,t}] = 0$ for $t = 1983, \ldots, 1999$ and $s \geq 2$. Of course, differencing cancels out the fixed effect ($\Delta \alpha_i = 0$).

b) GMM system estimation (Blundell and Bond, 1998)

The 17 equations in firsts differences given by (B1) are supplemented by the following 18 equations in levels

$$z_{i,t} = \beta_1 x_{i,t} + \beta_2 D_t^T x_{i,t} + \beta_3 GAP_{i,t} + \alpha_i + \eta_t + u_{i,t} \tag{B2}$$

for $t = 1982, \ldots, 1999$ and $i = 1, \ldots, N$, where lagged first differences ($\Delta x_{i,t-1}$) are used as instruments[12] for $x_{i,t}$ and $D_t^T x_{i,t}$ in the level equations, based on the assumption that $E[\alpha_i \Delta x_{i,2}] = 0$ for $i = 1, \ldots, N$, which (together with the standard assumptions for (B1)) yields the additional moment conditions $E[\upsilon_{i,t} \Delta x_{i,t-1}] = 0$ for $i = 1, \ldots, N$ and $t = 1983, \ldots, 1999$, where $\upsilon_{i,t} = \alpha_i + u_{i,t}$.[13] Using Monte Carlo studies, Blundell and Bond (1998) and Blundell et al. (2000) showed that the finite sample bias of the GMM estimator in first differences can be reduced substantially with the system GMM estimator.

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[12] Note that there are no instruments for the first observation $x_{i,2}$ available.

[13] This requires the first moment of $x_{i,t}$ to be stationary (which is fulfilled here).

Table 1. Overview of countries and industries used in the estimation

industry	code	share in TM in per cent[1)]	share in TVA in per cent[1)]	GVA/PROD in per cent	Excl. countries	N	T	obs.
Total manufacturing (TM)	**15-37**	**100.00**	**21.60**	**32.7**	**--**	**10**	**19**	**190**
Food products, beverages and tobacco	15-16	11.84	2.56	24.9	SE	9	19	171
Textiles, textile products, leather and footwear	17-19	5.70	1.23	35.1	SE	9	19	171
Wood and products of wood and cork	20	2.69	0.58	34.1	FR, SE	8	19	152
Pulp, paper, paper products, printing and publishing	21-22	10.04	2.17	38.4	SE	9	19	171
Coke, refined petroleum products and nuclear fuel	23	1.75	0.38	17.05	ES, SE	8	19	152
Chemicals and chemical products	24	9.39	2.03	32.9	FR, ES, SE	7	19	133
Rubber and plastics products	25	4.02	0.87	38.4	FR, ES, SE	7	19	133
Other nonmetallic mineral products	26	4.41	0.95	42.5	SE	9	19	171
Basic metals and fabricated metal products	27-28	12.93	2.79	35.2	SE	9	19	171
Machinery and equipment, n.e.c.	29	11.39	2.46	37.9	BE, ES, SE	7	19	133
Electrical and optical equipment	30-33	11.97	2.59	40.2	BE, ES, SE	7	19	133
Transport equipment	34-35	10.00	2.16	28.9	SE	9	19	171
Manufacturing n.e.c.	36-37	3.87	0.83	39.4	SE	9	19	171
Construction	**45**		**5.59**	**43.07**	**NL, ES**	**8**	**19**	**152**
Services	**50-74[2)]**		**44.46**	**62.4**	**NL, ES, SE, UK**	**6**	**19**	**114**
Wholesale and retail trade; repairs	50-52		11.51	62.7	NL, ES, UK	7	19	133
Hotels and restaurants	55		2.28	46.8	NL, ES, SE, UK	6	19	114
Transport and storage and communication	60-64		7.01	55.9	ES, SE,	8	19	152
Financial intermediation	65-67		5.23	67.8	NL, ES, UK	7	19	133
Real estate, renting, and business activities	70-74		18.43	67.2	NL, ES, UK	7	19	133

Notes: Full sample contains 10 countries (AT, BE, FI, FR, DE, IT, NL, ES, SE, UK), time period ranges from 1981 to 1999 for all industries. – TVA ... total value added. – TM ... total manufacturing. – GVA/PROD ... ratio of gross value added to production. – [1)] shares in per cent of value added of total manufacturing and total value added respectively (averages over countries and years). – [2)] group of services comprises 50-52, 55, 60-67, 70-74; public services and private non-market services are not included.

Table 2. Estimation results for logistic smooth transition (LST) model (13)

	χ^2	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_3$	$\hat{\tau}$	γ	$\overline{R}^2$	SEE	NT	$\hat{\mu}_{t<T}$ [1)]	$\hat{\mu}_{t\geq T}$ [1)]
15-37	1.21	0.409***	-0.075*	-0.205**	1994.7**	9.89	0.822	0.018	190	1.38	1.29
		(20.08)	(-1.78)	(-1.98)	(2.31)						
15-16	6.48*	0.510***	0.048	0.035	1989.6***	7.97	0.878	0.019	171	1.62	-
		(16.39)	(1.10)	(0.32)	(2.62)						
17-19	8.52**	0.418***	-0.073	-0.142	1988.0***	10.00	0.703	0.026	171	1.37	-
		(9.06)	(-1.28)	(-0.97)	(21.36)						
20	5.29	0.540***	0.163**	-0.014	1992.1***	1.67	0.763	0.038	152	1.55	1.86
		(13.43)	(2.10)	(-0.06)	(8.22)						
21-22	20.31***	0.388***	0.218***	-0.273*	1990.0***	3.71	0.781	0.026	171	1.31	1.56
		(13.14)	(4.04)	(-1.66)	(46.55)						
23	4.20	0.916***	0.094**	0.413**	1993.5**	9.97	0.961	0.042	152	4.16	6.18
		(27.33)	(2.21)	(2.51)	(1.88)						
24	25.11***	0.375***	0.233***	-0.045	1983.9***	7.41	0.841	0.022	133	1.34	1.69
		(10.97)	(5.83)	(-0.31)	(23.12)						
25	5.19	0.444***	0.158**	-0.009	1989.6**	8.51	0.760	0.028	133	1.38	1.59
		(14.96)	(2.47)	(-0.05)	(2.11)						
26	0.95	0.518***	0.037	-0.080	1993.5	9.72	0.835	0.025	171	1.42	-
		(19.43)	(0.70)	(-0.61)	(0.41)						
27-28	1.11	0.447***	-0.126**	-0.226	1994.9***	6.16	0.690	0.029	171	1.41	1.26
		(14.69)	(-2.00)	(-1.27)	(13.68)						
29	3.40	0.490***	-0.137***	-0.058	1994.6	9.89	0.794	0.024	133	1.44	1.28
		(16.70)	(-2.72)	(-0.43)	(0.90)						
30-33	7.75*	0.463***	0.144**	0.042	1987.7**	0.68	0.862	0.024	133	1.38	1.570
		(16.50)	(2.56)	(0.22)	(2.43)						
34-35	11.25**	0.449***	0.164	0.281	1996.6**	9.90	0.562	0.045	171	1.47	-
		(10.37)	(1.36)	(1.20)	(1.79)						
36-37	1.06	0.466***	-0.040	-0.192*	1992.5	9.95	0.784	0.025	171	1.39	-
		(13.40)	(-0.76)	(-1.84)	(0.39)						
45	13.37***	0.628***	-0.200***	0.059	1986.4**	7.84	0.805	0.024	152	1.56	1.32
		(18.41)	(-4.16)	(0.47)	(2.49)						
50-74	10.86**	0.630***	0.121**	-0.070	1997.1***	10.00	0.972	0.010	114	1.31	1.39
		(28.10)	(2.03)	(-0.92)	(45.56)						
50-52	20.23***	0.640***	-0.120	-0.042	1990.9***	5.53	0.801	0.028	133	1.31	-
		(9.60)	(-1.28)	(-0.29)	(8.53)						
55	5.68	0.442***	0.114	0.201	1994.0***	10.00	0.752	0.023	114	1.31	-
		(9.57)	(1.59)	(1.24)	(32.94)						
60-64	12.64***	0.389***	0.144***	-0.193***	1984.6***	5.36	0.864	0.017	152	1.21	1.31
		(14.19)	(5.24)	(-2.64)	(6.73)						
65-67	10.09**	0.616***	0.222***	0.061	1990.6***	9.80	0.876	0.029	133	1.25	1.37
		(13.08)	(2.95)	(0.34)	(2.90)						
70-74	34.78***	0.882***	-0.129***	-0.089	1984.8***	3.21	0.982	0.010	133	1.41	1.33
		(61.50)	(-5.92)	(-1.56)	(8.83)						

Notes: Country-specific and time specific fixed effects included in all models. [1)] Price/marginal cost ratios implied by the estimates of β_1 and β_2, adjusted using the average value added to gross output ratios. ***, **, * indicate significance at the 10, 5, and 1 per cent level, respectively.

Table 3. Estimation results for discrete break model with endogenous breakpoint T (14), major industry groups

	T	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_3$	SEE	$\overline{R}^2$	m_1	m_2	NT	$\hat{\mu}_{t<T}$ 1)	$\hat{\mu}_{t\geq T}$ 1)
					manufacturing (15-37)						
LSDV	1994	0.409*** (13.86)	-0.075*** (-3.10)	-0.205*** (-3.06)	0.018	0.823			190	1.38	1.29
GMM-FD	1995	0.430*** (15.60)	-0.094*** (-3.81)	-0.113 (-1.63)	0.023	-	-2.81***	-1.58	170	1.41	1.29
GMM-SYS	1995	0.430*** (14.40)	-0.103*** (-3.07)	-0.148*** (-2.68)	0.018	-	-2.78***	-1.57	350	1.41	1.28
					construction (45)						
LSDV	1992	0.515*** (8.20)	-0.083*** (-3.37)	-0.017 (-0.32)	0.025	0.786			152	1.42	1.33
GMM-FD	1992	0.470*** (7.10)	-0.046** (-1.99)	0.017 (0.29)	0.033	-	-2.46**	0.39	136	1.37	1.32
GMM-SYS	1992	0.441*** (5.57)	-0.011 (-0.31)	-0.061 (-0.93)	0.026	-	-2.49**	0.38	280	1.34	-
					services (50-74)						
LSDV	1996	0.630*** (14.35)	0.093*** (2.92)	-0.068 (-1.30)	0.010	0.972			114	1.31	1.37
GMM-FD	1996	0.615*** (13.40)	0.107*** (3.17)	-0.078 (-1.35)	0.011	-	-1.89*	-0.38	102	1.30	1.37
GMM-SYS	1996	0.609*** (12.20)	0.121*** (2.93)	-0.096 (-1.58)	0.011	-	-1.80*	-0.26	210	1.30	1.38

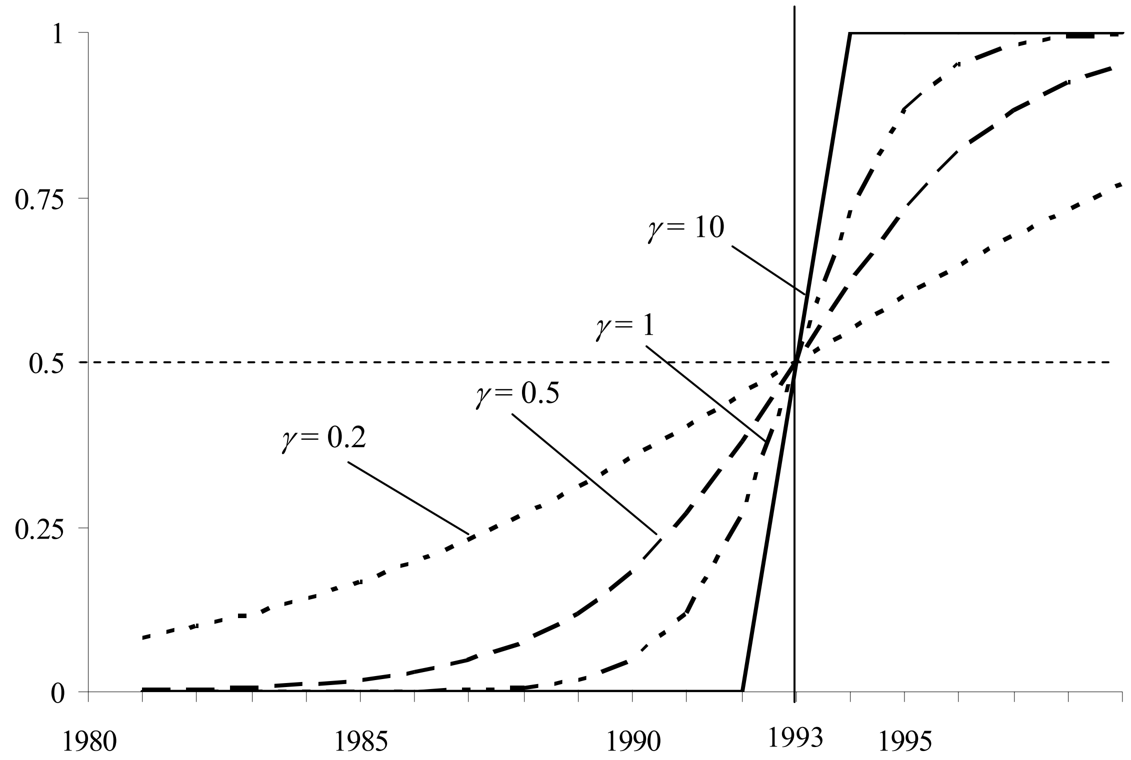
GMM estimation was carried out using the DPD package for Ox (Doornik et al. 2002). m_1 (m_2) … first (second) order serial correlation test. See also Table 2.

Table 4. GMM-system estimates of discrete break model with endogenous breakpoint T (14), detailed industries

	T	$\hat{\beta}_1$	$\hat{\beta}_2$	$\hat{\beta}_3$	SEE	m_1	m_2	NT	$\hat{\mu}_{t<T}$ [1)]	$\hat{\mu}_{t\geq T}$ [1)]
15-37	1995	0.430***	-0.103***	-0.148***	0.018	-2.78***	-1.57	350	1.41	1.28
		(14.40)	(-3.07)	(-2.68)						
15-16	1989	0.460***	0.083***	-0.007	0.019	-2.44**	0.80	315	1.53	1.69
		(13.80)	(2.60)	(-0.10)						
17-19	1990	0.380***	-0.060***	-0.091	0.026	-2.42**	-1.19	315	1.33	1.26
		(8.19)	(-3.17)	(-0.63)						
20	1991	0.493***	0.180**	-0.058	0.039	-1.83*	-0.82	280	1.48	1.80
		(18.60)	(2.01)	(-0.68)						
21-22	1988	0.386***	0.151***	-0.214**	0.026	-2.48**	-1.24	315	1.31	1.49
		(12.00)	(3.07)	(-2.51)						
23	1993	0.933***	0.090	0.456***	0.045	-2.19**	-0.74	280	4.42	-
		(18.50)	(1.42)	(7.35)						
24	1995	0.587***	-0.048	0.057	0.025	-2.28**	-1.01	245	1.65	-
		(16.60)	(-0.64)	(0.58)						
25	1994	0.535***	-0.107***	0.105	0.030	-2.12**	-1.39	245	1.49	1.36
		(10.30)	(-3.53)	(1.04)						
26	1990	0.481***	0.025	-0.109	0.025	-2.55**	-0.95	315	1.38	-
		(11.60)	(0.55)	(-1.13)						
27-28	1995	0.442***	-0.138***	-0.219*	0.028	-2.60***	-1.46	315	1.40	1.25
		(6.83)	(-4.13)	(-1.72)						
29	1994	0.481***	-0.128***	-0.045	0.024	-2.29**	-1.74*	245	1.43	1.28
		(19.00)	(-3.45)	(-0.96)						
30-33	1994	0.599***	-0.017	0.075	0.026	-2.19**	-1.74*	245	1.56	-
		(9.09)	(-0.32)	(0.87)						
34-35	1996	0.414***	0.116	0.180	0.047	-2.12**	-0.70	315	1.42	-
		(18.40)	(1.46)	(1.30)						
36-37	1988	0.395***	0.059*	-0.176**	0.025	-2.68***	-0.42	315	1.31	1.38
		(8.43)	(1.90)	(-2.11)						
45	1992	0.441***	-0.011	-0.061	0.026	-2.49**	0.38	280	1.33	-
		(5.57)	(-0.31)	(-0.93)						
50-74	1996	0.609***	0.121***	-0.096	0.011	-1.80*	-0.26	210	1.30	1.38
		(12.20)	(2.93)	(-1.58)						
50-52	1992	0.529***	0.060***	-0.149	0.029	-1.74*	-1.11	245	1.25	1.28
		(5.40)	(3.46)	(-1.63)						
55	1989	0.300***	0.224***	-0.017	0.024	-1.96*	-1.50	210	1.19	1.39
		(4.92)	(6.25)	(-0.31)						
60-64	1988	0.488***	0.049	-0.119	0.018	-1.95*	-0.26	280	1.27	1.31
		(10.60)	(1.14)	(-1.53)						
65-67	1990	0.597***	0.248***	-0.057	0.030	-2.17**	-0.57	245	1.24	1.37
		(12.40)	(4.37)	(-0.61)						
70-74	1988	0.790***	-0.027	-0.095*	0.011	-1.79*	0.69	245	1.35	1.33
		(23.40)	(-1.00)	(-1.71)						

Estimation was carried out using the DPD package for Ox (Doornik et al. 2002). See also Table 3.

Figure 1. Transition function F(t) for $\tau = 1993$ and different values of γ



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