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Intra-Industry Trade and Job Turnover

by

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Abstract

In this study we examine the widely held view that intra-industry trade (IIT) entails relatively low adjustment costs. We construct a panel of industry level trade, production and employment data for Ireland. IIT is calculated using the conventional Grubel-Lloyd index as well as alternative measures of marginal IIT. Our measure of labour adjustment is defined as the share of intra-industry job turnover in an industry's total job turnover. *Ceteris paribus*, we find no relationship between the Grubel-Lloyd index of IIT and our measure of labour adjustment; however, we find that marginal IIT has a small positive effect on the reallocation of labour within an industry. These results support the "smooth adjustment hypothesis" if IIT is understood in the sense of marginal IIT, and if labour reallocation is less costly within than between industries.

Outline

1 Introduction

Intra-industry trade (IIT), the international two-way exchange of goods with similar input requirements, has been the focus of countless theoretical and empirical studies since the early 1960s. There are two principal reasons for this interest. First, the observation of substantial IIT flows runs counter to the predictions of neo-classical trade theory. The IIT phenomenon therefore motivated the development of the "new trade theory", which can account for such trade patterns (for a survey, see Helpman and Krugman, 1985). Second, and crucial to this paper, simple intuition suggests that IIT expansion is concomitant with greater factor reallocation within rather than between industries. To redeploy workers or machinery in another plant within the same sector is likely to be easier than to adapt them for production in an entirely different industry. The liberalisation of trade between countries with high or growing IIT is therefore believed to entail relatively low adjustment costs. This has become known as the "smooth-adjustment hypothesis" (SAH) and has found widespread acceptance among economists (for a survey, see Brülhart, 1998). Yet, there exists no formal theoretical underpinning for this assumption. Perhaps even more importantly, empirical tests of the SAH have been crude and rather indirect, and the results have been inconclusive.

In this paper we carry out more direct and better specified empirical tests of the SAH using a panel data set of Irish manufacturing industries. Ireland is one of the most trade-oriented industrialised countries and thus serves as an ideal case study for examining the link between IIT and labour adjustment. There are four ways in which we improve on previous research on the SAH. First, we use a superior proxy for the dependent variable, labour reallocation, which is constructed from plant-level employment data. This allows us to calculate measures of intra- as well as inter-industry job reallocation. Second, we go beyond simple bivariate relationships and attempt to isolate the effect of IIT on labour adjustment, controlling for other relevant variables. Third, we use marginal measures of IIT (MIIT) as well as the more traditional static measures. Recent research suggests that the former are more appropriate when considering adjustment issues. Fourth, we consider more general dynamic models.

Our panel data set suggests a positive, albeit small, statistically significant and robust relationship between a measure of MIIT and the share of intra-industry job turnover, a result which is consistent with the SAH. However, the results provide no support for the SAH when IIT is measured using the Grubel-Lloyd index. In addition, we find that low concentration ratios and high trade exposure tend to increase the share of intra-industry job reallocation, and that trade changes precede changes in job turnover.

The paper is organised as follows. Section 2 consists of an overview of the theoretical and empirical literature relating to the SAH. In Section 3, we clarify the SAH and relate it to the empirical measures of IIT and labour adjustment used in this paper. We also describe the other variables in our data set. The econometric model and results are described in Section 4. Our results are summarised in Section 5.

2 The Hypothesis of IIT and Smooth Adjustment in the Literature

The supposition that IIT entails lower costs of factor-market adjustment than inter-industry trade was first made by Balassa (1966, p. 472), who wrote that “the difficulties of adjustment to freer trade have been generally overestimated”, because “it is apparent that the increased (intra-industry) exchange of consumer goods is compatible with unchanged production in every country”. Over the following three decades the SAH has become firmly established as part of conventional wisdom. The degree of acceptance is well captured by Grant *et al*. (1993, p. 32f.): “A (...) purported characteristic of intra-industry trade is its allegedly low adjustment costs in the face of trade liberalisation. It has become an article of faith that the European Community’s early liberalisation succeeded because of intra-industry trade”.

In Table 1, we have compiled a list of relevant studies published since 1987 to illustrate the pervasive use economists have made of the SAH. Even though this list is certainly not exhaustive, and not all the included studies accept the hypothesis uncritically, it becomes clear that the SAH has been invoked in the academic analysis of most recent episodes of trade liberalisation.

Sceptics of the SAH on theoretical grounds have been rare. The monopolistic-competition model of IIT is generally invoked as the main underpinning of the SAH. Krugman’s (1981, p. 970) model, for instance, yields the hypothesis that IIT “poses fewer adjustment problems than inter-industry trade”. However, use of the term “adjustment” in the interpretation of such a model is misleading. The welfare effects Krugman alluded to did not relate to transition costs but to end-state utility distributions before and after trade liberalisation. This result, valid in its own right but not to be confused with the SAH, has been expressed succinctly by Rodrik (1994, p. 7): “intra-industry trade will make everyone better off: it will increase the number of varieties available for consumption without reducing anyone’s real income". The mainstream models of IIT in horizontally differentiated goods assume the products of an industry to be perfectly homogenous in terms of quantitative and qualitative factor requirements and thus eliminate transitional costs by assumption. Oligopoly models of "reciprocal dumping" can also account for IIT. Intuitively, adjustment seems likely to be more disruptive in homogenous industries with concentrated market structures than in sectors with differentiated products and large firm numbers. However, these issues have been formally explored neither in terms of their implications for real factor rewards nor in terms of transitional adjustment costs. The main theories capable of explaining IIT, therefore, do not provide a coherent underpinning for the SAH.

There exists a rich literature on trade-induced adjustment (see e.g. Neary, 1982). However, these analyses are firmly rooted in neo-classical trade models with perfect competition in two homogeneous goods. IIT does not feature in these models. Hence, while we have some models that are able to generate IIT, and some theories linking trade and factor-market adjustment, we have no integrated theory of IIT and adjustment.

In the absence of rigorous theoretical foundations for the SAH, it is interesting to see whether empirical work has produced supporting evidence. Unfortunately empirical evidence is also scant. The studies devoted to this topic all explored the issue in an indirect fashion. The main approach was to examine whether factor intensities are less heterogeneous within than between industries. Considerable heterogeneity has been found within industries, but differentials between industries tend also to be significant.[1] The persuasiveness of such factor-ratio analyses, however, is constrained by the crude measures of production factors. These measures are unable to distinguish, for instance, between industry-specific and transferable capital. An alternative empirical approach to the SAH is via political-economy considerations. Lundberg and Hansson (1986) and Marvel and Ray (1987) conjectured that the fast trade liberalisation in sectors subject to high initial IIT levels resulted from a lower demand for protection in these industries, which in turn suggests that IIT has relatively benign welfare effects. In a study of Australian trade liberalisation, however, Ratnayake and Jayasuriya (1991) argued that previous single-

1 See Lundberg and Hansson (1986). It would also be appropriate to consider the empirical literature on "vertical IIT" as part of the effort to gauge the heterogeneity of industries (see Greenaway et al., 1995). Quality differentiation within industries is likely to be accompanied by substantial intra-industry variance of factor requirements.

equation estimations had suffered from simultaneity bias, and they detected no effect of IIT on tariff reduction when estimated through a simultaneous-equation model. Finally, a number of recent studies have reported simple correlations between various IIT measures and industry-level employment changes (see Brülhart and Hine, 1998). The results of these exercises provide some support for the SAH with MIIT. However, due to their methodological limitations, such bivariate analyses have to be interpreted with caution. The available empirical evidence on the SAH therefore appears to be as inconclusive as the relevant theoretical work.

3 The Smooth-Adjustment Hypothesis: From Theory to Empirics

In the absence of formal scrutiny of the SAH, the precise meaning of the SAH has never been explicitly spelt out. Moreover, the two core concepts, trade-induced adjustment costs and IIT, have been implicitly interpreted in a number of different ways.

3.1 Trade as an Exogenous Variable?

There are two principal conceptions of trade as a source of adjustment. In partial-equilibrium, small open economy (SOE) models, adjustment is traditionally analysed by departing from a change in world market prices. Such price changes are exogenous to the SOE and can originate from many different sources, such as changes in demand, factor endowments or trade policies of trading partners. These changes can be labelled “trade-induced”, since they would not affect the SOE in autarky. The second concept of trade as a source of adjustment centres on changes in trade costs in multi-country general-equilibrium models, holding everything else constant. Under that definition, “trade-induced” means sparked by a change in the level of barriers to international trade. In a nutshell, domestic adjustment is trade-induced either if caused by a reduction in trade barriers, holding everything else constant, or if caused by relevant and independent changes in foreign markets, holding trade costs constant.[2]

3.2 Adjustment Costs

Adjustment costs can also be divided into two categories. First, they can arise in perfectly competitive markets with flexible prices. If factors are subject to any degree of

[2] In reality, of course, demand and production structures change continuously. Therefore, integration occurs simultaneously with other changes, and the two types of trade-induced adjustment, while separable in theory, are difficult to disentangle empirically.

heterogeneity and product specificity, then trade-induced reallocation will inevitably divert resources to make the transition possible. Hence, production will occur inside the long-run production possibility frontier for the duration of adjustment, as resources are used to re-train, move and match labour, and to adapt the capital stock. Temporary factor-price disparities are needed to incite resource use on such "adjustment services". When arising from a fall in the relative price of importables (e.g. through integration), adjustment costs of this nature do not lead to an aggregate welfare loss, and their impact is purely distributional.[3] In theory, lump-sum transfers can be designed so as to compensate all individuals for transitional income losses.[4] In practice, however, transitional wage and income disparities often go uncompensated, thus producing net losers and feeding protectionist pressures.

The second class of adjustment costs arises in the presence of market imperfections. The most commonly analysed imperfection is that of downwardly rigid nominal wages. In these circumstances, adjustment costs might outweigh the gains from trade, hence trade liberalisation might be Pareto inferior.[5] The cost-benefit balance depends on the magnitude of adjustment costs and trade gains as well as on the social discount rate.

3.3 Measuring Intra-Industry Trade

In the context of the SAH different commentators have implicitly held different conceptions of IIT. The standard IIT measure is the Grubel-Lloyd (GL) index:

$$GL_t = \left(1 - \frac{|M_t - X_t|}{(M_t + X_t)}\right) \tag{1a}$$

where *M* stands for imports in a particular industry, *X* represents corresponding exports, and *t* is the reference year.[6] The GL index leaves room for at least two interpretations of "IIT" in

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[3] See Baldwin *et al*. (1980, p. 408).

[4] See Feenstra and Lewis (1994, p. 202). Dixit and Norman (1986) have proposed an incentive-compatible taxation scheme which ensures Pareto gains.

[5] See Baldwin *et al*. (1980, p. 408ff.). Brecher and Choudhri (1994) have formalised this proposition in an efficiency-wage model.

[6] A survey of the statistical properties and suggested adjustments of this index can be found in Greenaway and Milner (1986).

the adjustment context. IIT could refer to either the GL index at the start or end of the relevant period (GL_t), or to the growth of the GL index over that period (ΔGL_t):

$$\Delta GL_t = GL_t - GL_{t-n} \tag{1b}$$

where Δ stands for the difference between years *t* and *t-n*.

The GL index is a static measure, in the sense that it captures IIT for one particular year. However, adjustment is a dynamic phenomenon. By suggesting the concept of *marginal* IIT (MIIT), Hamilton and Kniest (1991) have addressed the problem of the mismatch between the static measures of IIT and the dynamic nature of the adjustment process in the SAH. They argued that the observation of a high proportion of IIT in one particular time period does not justify *a priori* any prediction of the likely pattern of *change* in trade flows. Even an observed increase in static IIT levels between two periods (positive ΔGL) could "hide" a very uneven change in trade flows, concomitant with *inter*- rather than *intra*-industry adjustment.

Brülhart (1994) has suggested the following index to measure MIIT:[7]

$$A_t = 1 - \frac{|\Delta X_t - \Delta M_t|}{|\Delta X_t| + |\Delta M_t|} \tag{2}$$

This index, like the GL coefficient, varies between 0 and 1, where 0 indicates that marginal trade in an industry is exclusively inter-industry and 1 indicates that it is exclusively intra-industry. The *A* index shares most of the statistical properties of the GL index.[8] One such property is that the *A* index of an industry is independent of the size of that industry. However, it has been argued that such measures of trade composition should be related to gross trade or production (Greenaway *et al*., 1994). Therefore, an alternative measure, based on absolute values of MIIT, was suggested in Brülhart (1994):

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[7] Hamilton and Kniest (1991), Greenaway *et al*. (1994), Menon and Dixon (1997) and Thom and McDowell (1998) have proposed alternative measures of MIIT. For a survey, see Brülhart (1998).

[8] Oliveras and Terra (1997) have shown that the statistical properties of the *A* index differ from those of the GL index in two respects. First, the *A* index is not subject to a growing downward bias as the level of statistical disaggregation is increased. Second, there is no functional relationship between the *A* index for a given period and the *A* indices of constituent sub-periods.

$$C_t = \left(|\Delta X_t| + |\Delta M_t|\right) - |\Delta X_t - \Delta M_t| \tag{3}$$

which can be scaled even at the disaggregated industry level:

$$CW_t = \frac{C_t}{W_t} \tag{4}$$

where *W* is some relevant scaling variable, such as output or employment.

3.4 Measuring Intra-Industry Labour Adjustment

In order to capture labour adjustment, we construct a measure of intra-sectoral job reallocation in the spirit of Davis and Haltiwanger (1992). Numerous studies of gross job flows have confirmed the existence of job turnover well beyond that necessary to accommodate employment reallocation across sectors (see, e.g., Davis *et al*., 1996). Strobl (1996) has shown that this stylised fact also applies to Irish industry: the degree of job reallocation within sectors is considerably higher than both net aggregate employment changes and employment reallocation across sectors, even with high sectoral disaggregation. To date, research in this area has focused on documenting and explaining the intertemporal pattern of job turnover, mainly with respect to its synchronisation with the business cycle. However, there has been little corresponding analysis of the cross-sectional characteristics of plant-level job turnover.[9]

Our proxy for labour adjustment is derived as follows. We aggregate gross changes in plant-level job numbers within a particular industry separately for plants whose employment has expanded (*POS*) and those whose employment has contracted (*NEG*) over the period bounded by *t-n* and *t*:

$$POS_t = \sum_i E_{it} - E_{it-1} \text{ for } E_{it} - E_{it-1} > 0 \tag{5a}$$

$$NEG_t = \sum_i |E_{it} - E_{it-1}| \text{ for } E_{it} - E_{it-1} < 0 \tag{5b}$$

where *E* stands for the number of employees and *i* denotes plants. From this, we derive an industry-level measure of excess job reallocation:

9 For a rudimentary characterisation of job flows across sectors, see Davis *et al*. (1996).

$$INTRA_t = (POS_t + NEG_t - |POS_t - NEG_t|) / (POS_t + NEG_t) \qquad (6)$$

where $INTRA_t$ is the share of total plant-level job reallocation that is due to job reallocation in excess of net aggregate employment change of the particular industry. The values of *INTRA* fall within the interval [0,1]. The left endpoint corresponds to instances where all plants within the sector experience either net job creation or job destruction; and the right endpoint corresponds to instances where the net change in job numbers of the sector is zero, and hence every job lost is offset by a job created simultaneously in the same sector.

How does *INTRA* relate to the adjustment concept of the IIT literature? This is best explained by stating what *INTRA* does *not* capture, i.e. what we have to posit by assumption. First, *INTRA* is not a direct measure of labour adjustment *costs*, as it contains no information about flows into and out of unemployment, nor about relative wage changes and "adjustment services". Therefore, our measure does not tell us whether or not intra-industry job reallocation is less costly than reallocation from one industry to another. The first assumption we have to make is that labour moves more easily within than between industries. While we cannot subject this assumption to an empirical test with the available data, Shin (1997) has pointed out that a number of recent studies for the US provide evidence of relatively higher movement costs across industries due to industry-specific human capital.[10] Second, plant-level data carry no information on labour reallocation within establishments. However, the adjustment concept underlying the SAH encompasses reallocation both within and between plants. We therefore have to resort to a second assumption, stating that plant-level turnover correlates positively and significantly with total job turnover. Third, our data do not track individual workers as they move between jobs. Some redundant workers may well move to a different industry, even if vacancies are available in plants belonging to their original industry and filling the vacancies in the workers' original industry would be less costly in terms of retraining. For *INTRA* to capture the share of job-switches which occur within industries, we must also assume that excess job turnover correlates positively and significantly with the share of workers who move jobs

[10] See also Fallick (1993), Kletzer (1996) and Neal (1995). Indirectly, relatively lower intra-sectoral reallocation costs are reflected in the finding of Davis and Haltiwanger (1992) that job turnover within US manufacturing sectors significantly outstrips turnover between sectors. Similarly, using Irish data, Strobl *et al*. (1998) have found that even if one disaggregates the manufacturing sectors into 208 industries, intra-industry job reallocation accounts for more that half of total job turnover.

within an industry.[11] Given these three fairly plausible assumptions, *INTRA* serves as a valid inverse proxy for labour adjustment costs.

3.5 The Panel Data Set

We constructed an industry-level panel of job turnover, trade and other potentially relevant variables for the Irish manufacturing sector using three sources: a plant-level employment data set provided by the Irish Agency for Enterprise and Technology (Forfás), the Census of Industrial Production (CIP) published by the Irish Central Statistical Office, and a trade data set provided by Eurostat. Further details of the data are set out in the Data Appendix. The panel used for estimation consists of observations on 64 industries over a twelve year period, 1979 to 1990. Our choice of relevant variables, other than our proxies for IIT, MIIT, and labour adjustment, was restricted both by the lack of strong theoretical priors and by data availability. In our model we included the following set of variables which may be important: *TECH*, a proxy for the technology-intensity of an industry; *WAGE*, the average real wage[12]; *CONC*, the four-plant concentration ratio; *FOREIGN*, the share of employment accounted for by foreign-owned plants; Δ*CONS*, the change in apparent demand; and *TRADE*, the degree of trade exposure.

4 Econometric Model and Results

4.1 The Econometric Model

In terms of our *a priori* expectations of what determines *INTRA*, three variables stand out. First, one may expect highly concentrated industries to experience relatively low intra-sectoral employment reallocation, *ceteris paribus*, the stronger the competitive pressures within an industry, the higher will be the share of intra-industry labour turnover. The expected sign on a the four-plant concentration ratio *CONC*, therefore, is negative. Second, based on similar reasoning, there is likely to be a positive relationship between *INTRA* and trade exposure *TRADE*. Third, and most importantly, we have strong priors about the

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[11] There is empirical evidence in support of this assumption. For instance, Fallick (1993) found that improvements in the job prospects of displaced US workers in their previous industry reduces their search intensity in other industries.

[12] Appendix Table 1 shows that there are some implausible values for the wage variable. We chose not to make any adjustment to the published data. However, we ran all relevant regressions with as well as without this variable, and did not detect any significant impact.

coefficients on *IIT*. If the SAH, as used in the MIIT literature, is valid, the GL index and *INTRA* should be unrelated, *ceteris paribus*. However, one would expect to find a significant positive relationship with a measure of MIIT. In addition, we consider four variables which are known to be important in shaping industrial employment patterns in Ireland, but for which there are no clear-cut priors on expected coefficient signs. These variables are sectoral wages, exposure to foreign ownership, technology intensity, and changes in apparent demand.[13]

We started with the following fixed-effects panel data model:

$$\begin{aligned} INTRA_{it} = {} & \alpha_i + \beta_1 TECH_{it} + \beta_2 WAGE_{it} + \beta_3 CONC_{it} + \beta_4 FOREIGN_{it} \\ & + \beta_5 \Delta CONS_{it} + \beta_6 TRADE_{it} + \beta_7 IIT_{it} + \lambda_t + \varepsilon_{it} \end{aligned} \tag{7}$$

where α is a fixed effect, λ is a time dummy, and ε is an iid random error term. The subscripts i and t refer to industries and years respectively. We chose to estimate fixed-effects rather than random-effects panel data models, because our data set consists of essentially the population of all manufacturing industries. The panel is balanced, since we have a full set of observations on our 64 industries for the twelve years 1979 to 1990. When estimating our models we always included time dummies, although we do not report the estimated coefficients in the Tables. All first-differenced variables are calculated for one-year intervals.

4.2 Panel Results

Our main results are set out in Tables 2 and 3. The first column of results in Table 2 uses the GL index to measure IIT. Apart from the concentration *CONC* and trade exposure *TRADE* variables, all the other explanatory variables, including the current and lagged values of the GL index, are statistically insignificant. The concentration and trade exposure variables are correctly signed according to our priors.[14] We experimented with various lags,

[13] On foreign ownership and technology intensity in Irish industry, see Foley and McAleese (1991).

[14] We used lagged concentration rather than current concentration, and lagged *TRADE* rather than current trade exposure, as our explanatory variables, since it is most plausible that *CONC* and *TRADE* predetermine *INTRA*. However, we obtained similar results using current concentration and trade exposure.

leads and moving averages of the GL index, as well as with ΔGL, but obtained similar results - the index was always insignificant and often “incorrectly” signed.

Columns (2) to (4) of Table 2 all use the *A* index as the IIT variable. The one-year lagged *A* index consistently gave the best results. This suggests that trade changes precede labour changes. In column (4), we report results based on the *A* index calculated over two-year intervals, but this variable is insignificant and incorrectly signed. We also experimented with the *C* measure (not reported in the Table), but found no significant and robust relationships.[15]

Column (3) of Table 2 sets out a parsimonious model, which is our preferred specification on both *a priori* and statistical grounds. In addition to the time dummies, the explanatory variables are lagged concentration, lagged trade exposure and the lagged one-year *A* index. These three variables are all statistically significant and correctly signed according to our priors. In addition, the fit of the model is reasonable, and the RESET test suggests no misspecification.[16]

In Table 3, we have explored the possibility that labour turnover adjusts to changes in the explanatory variables over a number of years, rather than in only one year. We did this by including the lagged dependent variable as an explanatory variable. In order to obtain consistent estimates, we first-differenced the data and used instrumental variables, employing a generalised method of moments (GMM) procedure. Since differencing may be viewed as a test of specification, the results in Table 3 amount to an informal test of the robustness of the Table 2 results.

When the data are first-differenced and the lagged dependent variable is included as an explanatory variable, one has to be careful about the choice of instruments and the dating of the instruments, since first-differencing induces a first-order moving average, or MA(1), error term. We used predetermined variables as instruments and tested for the validity of the instruments/over-identifying restrictions using the standard minimised GMM criterion

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[15] Specifically, we explored the *C* measure and indices proposed by Hamilton and Kniest (1991) and by Greenaway *et al*. (1994). Results are available from the authors.

[16] Given the bounded nature of our dependent variable, we re-ran all the regressions using a logit transformation of *INTRA*, setting values of 0 and 1 to 0.001 and 0.999 respectively, and using trimmed values of *INTRA*. None of the main findings were altered.

function to form the test statistic. In all cases we failed to reject the validity of the instruments. In addition, we tested for a MA(2) error term, which would render our GMM results inconsistent. In all cases, we failed to reject a zero MA(2) error term.

The results in columns (1) and (2) of Table 3 are presented for comparison with the results in column (2) of Table 2. The data are first-differenced, but none of the variables is instrumented. The lagged dependent variable appears to be highly significant and, implausibly, negative in column (1). However, it is reassuring to note that the column (1) results are inconsistent, since the lagged dependent variable is not instrumented. Where it is instrumented, i.e. in columns (3), (4) and (6), it is always statistically insignificant. We explored the insignificance of the lagged dependent variable further, and we conclude that it is a robust finding. The results in columns (4) and (5) suggest that there is no need to instrument the other current dated variables, which happen to be statistically insignificant. The final column sets out our preferred parsimonious model in first differences. The coefficient estimates are similar to those in column (3) of Table 2, although concentration is now statistically insignificant. In conclusion, we believe that we have estimated a robust, parsimonious panel data model of labour turnover and found a significant, positive relationship between the lagged *A* index and *INTRA*.

4.3 Panel Unit Root Tests

It is important to verify whether the non-stationarity of some variables could invalidate our findings. We therefore examined the orders of integration of the data using the "*t*-bar" panel unit root test proposed by Im, Pesaran and Shin (1997). Further details are given in the Data Appendix. The results suggest that one can reject the null hypothesis that *INTRA*, *CONS* and all measures of MIIT have a unit root. The unit root hypothesis cannot be rejected for the GL index. More importantly, *CONC* and *TRADE* may also be subject to unit roots. However, Dickey-Fuller tests along the lines of Kao (1997) suggest that these two variables, which appear as explanatory variables in our preferred models, are cointegrated. This means that the left and right hand sides of our preferred equations are I(0) and so are balanced. Hence, our main panel results do not seem to be affected by unit root problems.

4.3 OLS Cross Section Results

On prior econometric grounds there are good reasons for preferring fixed-effect panel regressions to OLS cross-section results. Fixed-effect panel data models are more general and robust than cross-section models. In cross sections, the intercept term is restricted to be the same in every industry, whereas in fixed-effect panel data models it is allowed to vary by industry. Fixed effects models may be capturing the effects of omitted variables. Consistent estimates of the slope coefficients are obtained even if the fixed effects are correlated with some of the right-hand side variables.

Some researchers, however, suggest that cross-section regression results may be more informative than fixed-effects panel regressions. This point has been made, e.g., by Durlauf and Quah (1998) in their review of panel data analysis of economic growth. In particular Durlauf and Quah suggest that by using panel data techniques which condition out or remove fixed effects, the researcher "winds up analysing a left-hand side (...) variable purged of its long-run variation across countries. Such a method, therefore, leaves unexplained exactly the long-run cross-country growth variation originally motivating this empirical research" (p.53). The relevant question is: over what time horizon is the model supposed to apply?

Our results in Tables 2 and 3 suggest that we are not dealing with a dynamic equation so, conditional on the exogenous and predetermined variables, the short and long run versions of the equations are the same. Moreover, economic theory has little to say about the appropriate time horizon to use in our context. The *INTRA* measure is constructed using annual data and is not a long run measure. In the long run, theory suggests that this variable should either be zero or constant. If this is the case, Durlauf and Quah's objection to the use of fixed effect panel data models is not relevant in the present context. Of course, *INTRA* and the *A* index are quite variable from year to year. However, when we used three-period moving averages of the these indices, we obtained similar results to those reported in Tables 2 and 3.

Nevertheless, we have calculated some time-averaged cross-section regressions using OLS. The OLS results are given in Table 4. Industry concentration *CONC* and trade exposure *TRADE* are again the most significant variables, both with the expected signs. GL indices of IIT, calculated in various ways, are never statistically significant. Time-averaged year-on-year *A* indices of MIIT also are not significant, but coefficients on some "long-term" MIIT measures, calculated over the entire 14-year interval, have the expected sign and border on being statistically significant. It must be noted, however, that the latter result is not robust to variations in the span of data used to construct the indices.

Economic Significance

Our analysis so far has concentrated on the signs and statistical significance of estimated regression coefficients. However, it is conceivable that a precisely estimated coefficient with the anticipated sign nevertheless has little economic importance. This is the case if the size of estimated regression coefficients is so small that movements in the independent variable will have a negligible impact on the dependent variable (McCloskey and Ziliak, 1996).

In order to evaluate the economic significance of our regression results, we have calculated "beta coefficients", as suggested, for instance, by Leamer (1984). The beta coefficients measure the change in *INTRA* (expressed in standard deviation units) for unit changes in each of the explanatory variables (in standard deviation units) holding other variables constant.[17] Table 5 reports the results. It is difficult to attribute meaning to the absolute size of the beta coefficients. We can, however, draw inferences from the relative coefficient sizes. If the ranking of explanatory variables in terms of their beta coefficients were to differ substantially from a ranking based on *t* statistics, then we would have to cast doubt over the economic significance of our results. It turns out that we find a close overlap between economic and statistical significance. The concentration ratio *CONC* and trade exposure *TRADE* are the most significant explanatory variables both in the economic sense and statistically. We also find that MIIT variables score better than IIT measures in terms of beta coefficients.[18]

5 Conclusion

In this paper, we used a panel of Irish data for the 1980s to examine the "smooth-adjustment hypothesis" (SAH), which states that IIT is associated with lower factor-reallocation costs than inter-industry trade. For the first time in empirical work on IIT, we used the share of intra-industry job turnover as a proxy for labour-adjustment. In addition, we modelled the share of intra-industry job turnover by incorporating a number of potentially relevant explanatory variables rather than just conducting a bivariate analysis.

[17] Note that there is no relationship between the beta coefficients and the simple or partial correlation coefficients when there is more than one explanatory variable in a model.

[18] The only noticeable discrepancy between results based on economic significance and those based on statistical significance appears for the *FOREIGN* variable, which gives substantially stronger results in terms of beta coefficients.

We also supplemented traditional, static measures of IIT with more recently developed, dynamic measures of marginal IIT (MIIT).

The econometric results provide no support for the SAH when IIT is understood in the static sense of the Grubel-Lloyd index. However, one measure of MIIT, the *A* index, showed a positive, statistically significant and robust relationship with the rate of intra-industry job turnover. This lends empirical support to the SAH as defined in the MIIT literature.

Our study reveals considerable scope for future work. It would be interesting to conduct similar analyses for other countries and time periods, in order to establish the robustness of our results. In addition refinements of the adjustment cost measure could be examined. Finally, our study has revealed a need for further theoretical work on the link between intra-industry trade and labour turnover.

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Table 1: Recent Empirical Studies with Reference to the Smooth-Adjustment Hypothesis

Region	Episode	Studies
World	Global integration	Fischer and Serra (1996)
	Regional integration	Hoekman and Kostecki (1995)
European Union	Iberian enlargement	Greenaway (1987)
		Krugman (1987)
		Hine (1989)
	Single Market	Balasubramanyam and Greenaway (1993)
		CEC (1996)
		CEPII (1988)
		Fontagné, Freudenberg and Péridy (1997)
		Greenaway and Hine (1991)
		Sapir (1992)
	EU Eastward enlargement	Cadot, Faini and de Melo (1995)
		Drábek and Smith (1995)
		Gatsios and Dimelis (1994)
		Hoekman and Djankov (1996)
		Lemoine (1995)
		Neven (1995)
		Thom and McDowell (1998)
	Swedish trade with EU	Lundberg (1992)
North America	NAFTA	Globerman (1992)
		Gonzalez and Velez (1993)
		Little (1996)
		Shelburne (1993)
Latin America	Regional integration	Guell and Richards (1998)
		Primo Braga, Safadi and Yeats (1994)
Asia	Trade expansion by Japan and NICs	Grant, Papadakis and Richardson (1993)
		Lincoln (1990)
		Noland (1990)
		OECD (1994)
		Rajan (1996)
	Regional integration	Drysdale and Garnaut (1993)
		Khalifah (1996)
	Arab countries	Havrylyshyn and Kunzel (1997)
Australia	Regional integration	Hamilton and Kniest (1991);
		Menon and Dixon (1995)
South Africa	Regional integration	Parr (1994)

Table 2: Labour Turnover and Intra-Industry Trade: Fixed-Effects Panel Data Model Estimates

Dependent variable = *INTRA* measure of intra-industry labour turnover

Explanatory Variables	**(1) GL index**		**(2) *A* index**		**(3) *A* index lagged, parsimonious model**		**(4) Two-period *A* index, parsimonious model**	
	Coefficient	*t* statistic	Coefficient	*t* statistic	Coefficient	*t* statistic	Coefficient	*t* statistic
TECH	-0.227	-1.165	-0.222	1.154	-	-	-	-
WAGE	-0.141	-1.155	-0.147	-1.186	-	-	-	-
CONC lagged	-0.420	-2.031	-0.410	-1.200	-0.396	-2.028	-0.412	-2.101
FOREIGN	-0.137	-0.871	-0.117	-0.753	-	-	-	-
Δ*CONS*	-0.019	-0.165	-0.019	-0.167	-	-	-	-
Δ*TRADE*	-0.663	-1.616	-0.637	-1.641	-	-	-	-
TRADE lagged	0.340	3.081	0.341	3.216	0.311	3.433	0.283	3.032
GL	0.056	0.543	-	-	-	-	-	-
GL lagged	-0.037	-0.336	-	-	-	-	-	-
A	-	-	-0.007	-0.265	-	-	-	-
A lagged	-	-	0.054	1.835	0.055	1.921	-	-
A calculated over two-year intervals	-	-	-	-	-	-	-0.035	-1.227
No. of explanatory variables	20		20		14		14	
Residual sum of squares	36.59		36.41		36.76		36.88	
Equation standard error	0.2313		0.2307		0.2308		0.2312	
RESET test (P value)	65.9%		58.0%		57.7%		49.0%	
Adjusted R^2	0.2362		0.2399		0.2394		0.2368	

Notes to Table 2: The sample size is 768, since the panel is balanced, consisting of observations on 64 industries for 12 years (1979 to 1990). The standard deviation of the dependent variable INTRA is 0.265. Eleven year-dummies are included in all of the models although the coefficient estimates and associated t statistics are not reported. The t statistics shown are based on heteroscedasticity-consistent standard errors.

Table 3: Labour Turnover and Intra-Industry Trade: GLS/GMM Estimates of First-Differenced Models with and without Lagged Dependent Variable (LDV)

Dependent variable = *INTRA* measure of intra-industry labour turnover

Explanatory Variables	**(1) GMM with LDV**		**(2) GMM, no LDV**		**(3) GMM with LDV**		**(4) GMM with LDV**		**(5) GMM, no LDV**		**(6) GMM with LDV**		**(7) GMM, no LDV**	
	Coeff.	*t* stat.	Coeff.	*t* stat.	Coeff.	*t* stat.	Coeff.	*t* stat.	Coeff.	*t* stat.	Coeff.	*t* stat.	Coeff.	*t* stat.
INTRA lagged	-0.51	-20.9	-	-	-0.05	-0.83	0.01	0.15	-	-	-0.06	-0.92	-	-
TECH	-0.33	-1.24	-0.50	-2.04	-0.44	-1.82	-0.46	-0.73	-0.50	-0.80	-	-	-	-
WAGE	-0.32	-1.95	-0.52	-2.02	-0.47	-1.98	-0.43	-0.72	-0.86	-1.31	-	-	-	-
CONC lagged	-0.56	-1.71	-0.38	-1.13	-0.48	-1.48	-0.25	-0.74	-0.20	-0.58	-0.43	-1.31	-0.33	-0.96
FOREIGN	-0.36	-1.22	-0.64	-1.55	-0.44	-1.18	-3.06	-1.60	-3.28	-1.66	-	-	-	-
Δ*CONS*	0.004	0.06	-0.018	-0.14	-0.018	-0.35	-0.016	-0.14	-0.027	-0.25	-	-	-	-
TRADE lagged	0.42	3.15	0.31	2.27	0.36	2.88	0.43	2.67	0.42	2.72	0.40	3.25	0.37	2.43
A lagged	0.04	1.51	0.07	2.16	0.07	2.23	0.09	2.92	0.08	2.58	0.07	2.31	0.07	2.29
No. of explanatory vars	18		17		18		18		17		14		13	
No. of instruments	0		0		19		23		21		15		0	
Variables instrumented	-		-		*INTRA lagged*		*INTRA lagged, TECH, WAGE, FOREIGN,* Δ*CONS*		*TECH, WAGE, FOREIGN,* Δ*CONS*		*INTRA lagged*		-	
Minimised GMM criterion	-		-		2.386		10.476		7.769		2.158			
No. of overidentifying restrictions	-		-		1		5		4		1			
Test of restrictions (P value)	-		-		12.2%		6.28%		10.0%		14.2%			
Test of MA(2) error (P value)	90.0%		97.2%		80.2%		79.4%		92.2%		77.0%		97.3%	

Notes to Table 3: (i) The sample size is 704, since the panel is balanced consisting of observations on 64 industries for 11 years (1980 to 1990). (ii) Ten year-dummies are included in all of the models, although the coefficient estimates and associated *t* statistics are not reported. (iii) The GMM coefficient estimates and *t* statistics shown above are all based on the optimal GMM variance-covariance matrix. (iv) The instruments for the (first differenced) lagged dependent variable (LDV) are the first difference and level of the dependent variable lagged two periods. Current dated (first differenced) variables are instrumented by their lagged levels and first differences. (v) The test statistics for the validity of the instruments/over-identifying restrictions and MA(2) errors are both distributed as chi-squares under the null. The degrees of freedom of the former test are equal to the difference between the number of instruments and the number of explanatory variables. The MA(2) error test has one degree of freedom.

Table 4: Labour Turnover and Intra-Industry Trade: OLS Cross-Section Results

Dependent variable = *INTRA* measure of intra-industry labour turnover; 64 observations

	(1) Average GL Index		**(2) Average ΔGL Index**		**(3) ΔGL Index; 1977-79 base,1988-90 end**		**(4) Average *A* Index**		**(5) *A* Index; 1977 base,1990 end**		**(6) *A* Index; 1977-79 base, 1988-90 end**		**(7) Average Weighted *C* Index**	
Explanatory variables	Coeff.	*t* stat.	Coeff.	*t* stat.	Coeff.	*t* stat.	Coeff.	*t* stat.	Coeff.	*t* stat.	Coeff.	*t* stat.	Coeff.	*t* stat.
TECH	0.12	0.78	0.12	0.76	0.13	0.79	0.11	0.71	0.11	0.75	0.11	0.72	0.11	0.69
WAGE	-0.12	-0.87	-0.12	-0.91	-0.11	-0.94	-0.12	-0.91	-0.16	-1.33	-0.15	-1.22	-0.13	-0.87
CONC	-0.40	-6.23	-0.40	-6.30	-0.39	-6.27	-0.40	-6.13	-0.38	-6.41	-0.39	-6.69	-0.40	-6.29
FOREIGN	-0.03	-0.69	-0.04	-0.76	-0.04	-0.76	-0.04	-0.78	-0.06	-1.39	-0.06	-1.43	-0.04	-0.84
ΔCONS	-0.04	-1.08	-0.03	-1.00	-0.03	-0.97	-0.03	-1.00	-0.03	-0.89	-0.03	-0.94	-0.02	-0.69
TRADE	0.01	1.59	0.01	1.61	0.01	1.64	0.01	1.55	0.01	1.77	0.01	1.78	-0.01	1.18
Average GL	-0.03	-0.55	-	-	-	-	-	-	-	-	-	-	-	-
Average ΔGL	-	-	0.21	0.41	-	-	-	-	-	-	-	-	-	-
ΔGL (1977-79 base, 1988-90 end)	-	-	-	-	0.03	0.58	-	-	-	-	-	-	-	-
Average A	-	-	-	-	-	-	-0.003	-0.03	-	-	-	-	-	-
A (1977 base, 1990 end)	-	-	-	-	-	-	-	-	0.05	1.59	-	-	-	-
A (1977-79 base, 1988-90 end)	-	-	-	-	-	-	-	-	-	-	0.05	1.67	-	-
Average Weighted C	-	-	-	-	-	-	-	-	-	-	-	-	0.09	2.13
Adjusted R^2	0.589		0.588		0.590		0.587		0.604		0.604		0.592	
F (P value)	13.90 (0.0%)		13.86 (0.0%)		13.94 (0.0%)		13.80 (0.0%)		14.70 (0.0%)		14.71 (0.0%)		14.08 (0.0%)	
RESET test (P value)	0.09 (76.5%)		0.04 (84.5%)		0.02 (88.6%)		0.05 (82.8%)		0.46 (50.1%)		0.70 (40.6%)		0.05 (82.2%)	

Notes to Table 4: All non-IIT data, as well as GL, are averaged across the 14 years 1977-90 for each of the 64 industries. ΔCONS, • GL and A are averaged across 13 year-intervals 1978-90. The t statistics are based on heteroscedasticity-consistent standard errors.

Table 5: Measures of "Economic Significance" - Beta Coefficients

Explanatory variables	**(1)** **Fixed-effect panel results**	**(2)** **OLS cross section results**
TECH	-0.071	0.061
WAGE	-0.084	-0.110
CONC	-	-0.709
CONC lagged	-0.380	-
FOREIGN	-0.131	-0.139
ΔCONS	-0.005	-0.089
TRADE	-	0.181
TRADE lagged	0.218	-
ΔTRADE	-0.095	-
GL / Average GL	0.053	-0.044
GL lagged	-0.035	-
Average ΔGL	-	0.033
ΔGL (1977-79 base, 1988-90 end)	-	0.050
A /Average A	-0.009	-0.003
A lagged	0.067	-
A calculated over two-year intervals	-0.043	-
A (1977 base, 1990 end)	-	0.129
A (1977-79 base, 1988-90 end)	-	0.133
Average Weighted C	-	0.075

Notes to Table 5: The beta coefficient can be interpreted as the number of standard error changes in INTRA resulting from a standard error change in the relevant explanatory variable. The fixed effects panel data results in column (1) are based on the estimates taken from model (2) of Table 3; except for GL and GL lagged (model 1) and A calculated over two-year intervals (model 4). Standard deviations are calculated over pooled data. The OLS results in column (2) are based on the estimates set out in Table 4.

Data Appendix

Data Sources

We constructed an industry-level panel of job turnover, trade and other potentially relevant variables for the Irish manufacturing sector using three sources: an employment data set provided by the Irish Agency for Enterprise and Technology (Forfás), the Census of Industrial Production (CIP) published by the Irish Central Statistical Office, and a trade data set provided by Eurostat. The Forfás data are compiled from an annual employment survey that has been carried out since 1973 and covers all known plants in the Irish manufacturing sector. A detailed description of the Forfás employment survey is given in Strobl (1996). The overall response rate to this survey has been high, covering on average over 99 per cent of the relevant population. The unit of observation is the individual plant, for which the number of permanent full-time employees is reported. Plants are identified by a 4-5 digit NACE sector and nationality of ownership. Plants are classified as foreign if at least 50 per cent of shareholdings are owned by non-Irish nationals.

The CIP provides a range of other relevant data series on Irish manufacturing sectors, derived from the annual survey of all industrial establishments and enterprises employing three or more workers and aggregated to the 2-3 digit NACE level. The response rate of the CIP is around 92 per cent on average. The variables of particular interest to the purpose here are total expenditure on wages and salaries, the number of employees, and gross output per industry. Finally, import and export series were available for 3-4 digit NACE sectors from Eurostat. Initially, the data set contained 68 industries. Because of incomplete coverage, we excluded four industries: mineral oil refining, extraction of gas, water supply, and railway rolling stock. These industries accounted for less that 3 per cent of Irish industrial employment in all sample years. The combination of the three data sets yielded an integrated data set with 64 2-3 digit NACE sectors covering the entire Irish manufacturing sector. Variable definitions and summary statistics are set out in Appendix Table 1.

Panel Unit Root Tests

We examined the orders of integration of the variables using the "*t*-bar" panel unit root test proposed by Im, Pesaran and Shin (1997), since this test is more general than other panel unit root tests. The Im *et al*. (1997) test statistic is the sample average of the *t* statistics on the lagged level of the dependent variable in the Dickey-Fuller or augmented Dickey-Fuller regressions, calculated for each of the 64 industries in the panel. The null hypothesis is that of a unit root. Critical values are smaller in absolute size than the standard unit root critical values. Im *et al.* (1997) point out that panel unit root tests are more powerful than standard unit root tests. However, the power of these tests is still low when there are not many time periods in the panel. In addition, the Im *et al*. (1997) critical values are based on the assumption that the data generation processes are independent across industries. Finally, there appear to be structural breaks in some of the data series, in which case the null hypothesis of a unit root is less likely to be rejected even when it is false.

Appendix Table 2 reports our panel unit root results. Dickey-Fuller and augmented Dickey-Fuller regressions with one and two lags, and with and without time trends, were estimated for each of the variables. *A priori*, the test statistics obtained from the regressions with a time trend are the preferred ones, since Dickey-Fuller regressions which include a time trend have a useful property. As noted by DeJong *et al*. (1992) and Hamilton (1994), the *t* statistic on the lagged level of the dependent variable is invariant to whether the true coefficient on the time trend is zero or not. The results suggest that one can reject the null hypothesis that *INTRA*, Δ*CONS* and all measures of MIIT have a unit root. The unit root hypothesis cannot be rejected for the GL index. More importantly, *CONC* and *TRADE* may also be subject to unit roots. However, Dickey-Fuller tests along the lines of Kao (1997) suggest that these two variables are cointegrated. Hence, our main panel results do not seem to be affected by unit root problems.

Appendix Table 1: Variable Descriptions and Summary Statistics

Variable	Description	Source	Mean	Std. Dev.	Max.	Min.	Correl. with *INTRA*
INTRA	Intra-industry plant-level job reallocation as a share of the industry's gross job reallocation (see equation 6)	Forfás	0.63	0.26	1.00	0.00	1.00
TECH	Share of industrial workers in total (inverse proxy for technology intensity)	CIP	0.70	0.09	0.93	0.42	0.01
WAGE	Annual wages and salaries per employee (1976 ECU '00,000)	CIP	0.05	0.12	1.48	0.01	-0.12
CONC	Share of employment accounted for by the four biggest plants in the industry	Forfás	0.47	0.25	1.00	0.08	-0.42
FOREIGN	Share of employment accounted for by plants under majority non-Irish ownership	Forfás	0.43	0.30	0.98	0.00	-0.17
ΔCONS	Year-on-year change in apparent consumption (output + imports – exports, 1976 ECU mn)	CIP, Eurostat	0.02	0.07	0.79	-0.78	0.03
TRADE	Imports plus exports divided by output	CIP, Eurostat	0.12	0.18	1.70	0.001	0.16
IIT	• Grubel-Lloyd index of IIT (see equation 1) • A index of MIIT (see equation 2) Weighted C index of MIIT (see equation 3, 1976 ECU '000, employment used as weight)	Eurostat Eurostat Eurostat	0.66 0.31 6.65	0.25 0.33 28.43	0.9997 0.9994 494.25	0.04 0.00 0.00	-0.04 -0.01 0.02

Appendix Table 2; Panel Unit Root Tests Using the Im-Pesaran-Shin t-Bar Statistic

Variables		DF/ADF regressions with no time trend			DF/ADF regressions with time trend		
		0 Lags	**1 Lag**	**2 Lags**	**0 Lag**	**1 Lag**	**2 Lags**
INTRA		-3.52**	-2.46**	-2.10**	-3.75**	-2.68**	-2.35
TECH		-2.07	-1.82	-1.44	-2.65	-2.37	-1.78
WAGE		-1.37	-1.51	-1.23	-1.87	-1.95	-1.75
CONC		-1.51	-1.63	-1.47	-1.85	-2.06	-2.09
FOREIGN		-1.24	-1.33	-1.23	-1.96	-2.11	-1.95
ΔCONS		-3.33**	-2.76**	-1.95**	-3.42**	-2.84**	-2.08
TRADE		-1.36	-1.35	-1.21	-2.07	-2.23	-1.79
IIT	GL	-1.64	-1.61	-1.45	-2.31	-2.26	-2.03
	A	-3.43**	-2.52**	-2.05**	-3.63**	-2.75**	-2.16
	CW	-3.91**	-2.97**	-2.19**	-4.13**	-3.15**	-2.35
Approximate 5% critical value		-1.93			-2.55		
Approximate 10% critical value		-1.84			-2.46		

Notes to Appendix Table 2: The t-bar statistic is the sample average of the t statistics on the lagged level of the dependent variable in the Dickey-Fuller (DF) or augmented Dickey-Fuller (ADF) regressions for each of the 64 industries in the panel. The samples size in the DF regressions are either 13, or 12 in the case of ΔCONS and the A and weighted C variables. Statistically significant t bar statistics are indicated with an asterisk. The critical values are interpolated from those given in Table 4 of Im et al. (1997).