Intra-Industry Trade and Job Turnover

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Abstract:

Trade expansion is widely believed to entail comparatively low labour-market adjustment costs if it takes the form of intra-industry trade (IIT). We examine this hypothesis using a panel of trade, production and employment data for Ireland. The share of intra-industry job turnover in an industry's total job turnover is used as a proxy for the degree of low-cost labour-market adjustment. IIT is calculated applying the conventional static index as well as alternative measures of marginal IIT. *Ceteris paribus*, we find no relationship between the static IIT index and our measure of labour adjustment. However, marginal IIT has a small positive effect on the reallocation of labour within an industry. These results are consistent with the "smooth adjustment hypothesis" in the sense of marginal IIT, provided that labour reallocation is less costly within than between industries.

Keywords: intra-industry trade, adjustment costs, job turnover
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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 will induce factor reallocation within rather than between industries; and the redeployment of workers or machinery in another plant within the same sector has been shown 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 "smooth-adjustment hypothesis" (SAH) has found widespread acceptance among economists. Yet, there exists relatively little formal theoretical or empirical analysis of the SAH which would support this assumption.

In this paper we carry out empirical tests of a central element of the SAH, the link between IIT and labour reallocation. We use a panel data set for Irish manufacturing industries. Ireland is one of the most trade-oriented industrialised countries and thus serves as a useful case study for examining the link between IIT and labour-market adjustment. Our paper presents four innovations relative to previous research. First, our dependent variable, intra-industry job turnover, has been shown in the labour economics literature to be a good proxy for labour-market adjustment costs, which are rather difficult to measure. Second, we move beyond simple bivariate relationships and aim to isolate the effect of IIT on job turnover, controlling for other relevant variables. Third, we use measures of marginal IIT (MIIT) as well as the more traditional static measures. Recent research suggests that the former may be more appropriate when considering adjustment issues. Fourth, we consider a range of general dynamic econometric 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, our findings provide no support for the SAH when IIT is measured using the static measures. 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 rates.

The paper is organised as follows. Section 2 motivates the paper by way of an overview of the theoretical and empirical literature relating to the SAH. In Section 3, we present and discuss our measures of IIT and job turnover. The econometric model and our empirical results are described in Section 4. Section 5 concludes.

2 Literature Background

The Smooth Adjustment Hypothesis

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, Papadakis and Richardson (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".[1]

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 may be misleading. The welfare effects that 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

[1] The SAH has been invoked in the scientific analysis of most recent episodes of trade liberalisation, including NAFTA (e.g. Caves, 1991; Globerman, 1992; Little, 1996; Shelburne, 1993) and European integration (e.g. Cadot *et al.*, 1995; Greenaway and Hine, 1991; Neven, 1995; Sapir, 1992).

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. Turning to the link between trade and factor markets, there exists a rich literature on trade-induced adjustment (see e.g. Neary, 1982; Furusawa and Lai, 1999). However, these analyses are firmly rooted in neo-classical trade theory with perfect competition in two homogeneous goods. IIT does not feature in these models.

A useful advance in the theoretical literature has been made by Lovely and Nelson (2000). Building on the general-equilibrium framework of Ethier (1982), they have modelled a reduction in trade costs for a monopolistically competitive sector where all trade is intra-industry. Falling trade costs trigger a symmetric change in the volume of imports and exports among identical countries, i.e. perfect marginal IIT[2]; but they also entail a rise in the sector's productivity through scale economies, which can reduce the factor demands from this sector in each country and thus stimulate *inter*-industry adjustments. The Lovely and Nelson (2000) analysis reveals that the link between IIT and intra-industry adjustment is not necessarily positive. Most importantly for our study, their demonstration that trade-induced productivity effects can reverse the conventionally assumed relationship between IIT and adjustment supports a multivariate empirical approach, where adjustment is related to IIT as well as to other determinants, notably absorption.

A number of empirical studies have also been devoted to the SAH. One approach has been to examine whether factor intensities are less heterogeneous within than between industries. Considerable heterogeneity has been found within industries, but differentials

[2] The concept of marginal IIT is explained in Section 3.

between industries tend also to be significant.[3] 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-equation estimations had suffered from simultaneity bias, and they detected no effect of IIT on tariff reduction when estimated through a simultaneous-equation model.[4] Finally, a number of studies have reported correlations between various IIT measures and industry-level employment changes (see Brülhart and Hine, 1999). 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.

Trade and Job Turnover

In the empirical literature job turnover is conventionally defined as the gross change in plant-level payroll numbers, due either to changes in plant scale or to plant births and deaths. Our focus is on job turnover within sectors, i.e. simultaneous creation and destruction of jobs by different plants of the same industry. Numerous studies of gross job flows have confirmed the existence of such "excess" job turnover (see, e.g., Davis, Haltiwanger and Schuh, 1996). There is substantial evidence in the labour literature that job moves within industries are less costly than flows between industries. Studies that

[3] 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.

[4] This reversed direction of causation, from trade liberalisation to IIT, also appears in the model of Lovely and Nelson (2000).

confirm this point include Fallick (1993), Kletzer (1996), Neal (1995) and Shin (1997), who all used US data; and Greenaway, Upward and Wright (2000) and Haynes, Upward and Wright (2000), who used data for the United Kingdom.

Some authors have also investigated the link between job turnover and trade. Using a panel of Canadian industry and plant data, Baldwin and Caves (1997) have shown that job turnover significantly increases in trade exposure. Andersson, Gustafsson and Lundberg (2000), drawing on Swedish data, found that job turnover increased in the degree of import penetration. Davidson and Matusz (2000) looked at the relationship but reversed the direction of causality. They found, using US data for 1973-1986, that job destruction rates had a negative impact on sectoral net exports. Finally, Klein, Schuh and Triest (2003), working with US data for 1973-1993, found that long-term changes in the real exchange rate significantly affect gross job turnover of traded manufacturing sectors, even though they have no significant impact on net employment. These studies did not investigate the link between IIT and job turnover.

3 The Smooth-Adjustment Hypothesis: From Theory to Empirics

The SAH has implicitly been subject to varying interpretations. Two concepts in particular need to be defined carefully: adjustment costs and IIT.

Adjustment Costs

Adjustment costs can arise in perfectly competitive markets with flexible prices. If factors are heterogeneous and product specific, then trade-induced reallocation will inevitably divert resources to make the transition possible.[5] 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 discounted welfare loss, and their impact is purely distributional.[6] In theory, lump-sum transfers can be designed so as to compensate all individuals intertemporally for transitional income losses.[7] In practice, however, transitional wage and income disparities often go uncompensated, thus producing net losers and feeding protectionist pressure.

A 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.[8] The cost-benefit balance depends on the magnitude of adjustment costs and trade gains as well as on the social discount rate.

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:

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[5] "Trade induced" can mean either triggered by a change in trade costs, or caused by a change in relative prices that originates outside the home country.

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

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

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

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

where M stands for imports in a particular industry, X represents corresponding exports, and t is the reference year.[9] The GL index leaves room for at least two interpretations of "IIT" in the adjustment context. IIT could refer to either the GL index at the start or end of the relevant period (GL_{t-n}, GL_t), or to the growth of the GL index over that period (ΔGL_t):

$$\Delta GL_t = GL_t - GL_{t-n} \qquad \textbf{(1b)}$$

where Δ is the first-difference operator.

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) addressed the problem of the mismatch between static measures of IIT and the dynamic nature of the adjustment process in the SAH. They argued that observing 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.

Following up on the point raised by Hamilton and Kniest (1991), Brülhart (1994) developed the following index to measure MIIT:

$$A_t = 1 - \frac{|\Delta X_t - \Delta M_t|}{|\Delta X_t| + |\Delta M_t|} \qquad \textbf{(2)}$$

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-

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

industry. The A index shares most of the statistical properties of the GL index.[10] 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, Hine, Milner and Elliott, 1994). Therefore, an alternative measure, based on absolute values of MIIT, was developed in Brülhart (1994):

$$C_t = (|\Delta X_t| + |\Delta M_t|) - |\Delta X_t - \Delta M_t| \qquad \textbf{(3)}$$

which can be scaled even at the disaggregated industry level:

$$CW_t = \frac{C_t}{W_t} \qquad \textbf{(4)}$$

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

Finally, Menon and Dixon (1997) pointed out that the impact of trade on *inter*-industry factor reallocation is captured most directly through a measure of "unmatched changes in trade" (UMCIT):

$$UMCIT_t = |\Delta X_t - \Delta M_t|.^{11} \qquad \textbf{(5)}$$

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 *et al.* (1996). Our proxy for labour adjustment is defined as follows. We aggregate gross changes in plant-level payroll numbers within a particular

[10] 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.

[11] Hamilton and Kniest (1991), Greenaway; Hine, Milner and Elliott (1994) and Thom and McDowell (1998) have proposed alternative measures of MIIT. For a survey of MIIT measurement literature, see Brülhart (2002).

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*:

$$POS_t = \sum_i E_{it} - E_{i,t-n} \ \text{ for } E_{it} - E_{i,t-n} > 0, \qquad \textbf{(6a)}$$

$$NEG_t = \sum_i | E_{it} - E_{i,t-n} | \ \text{ for } E_{it} - E_{i,t-n} < 0, \qquad \textbf{(6b)}$$

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

$$INTRA_t = (POS_t + NEG_t - |POS_t - NEG_t|) / (POS_t + NEG_t), \qquad \textbf{(7)}$$

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. We cannot subject this assumption to an empirical test with the available data; but, as mentioned in the previous Section, there is strong support for this proposition in the empirical labour literature. Second, plant-level employment data carry no information on labour reallocation within establishments. However, the adjustment concept underlying the SAH encompasses reallocation both within and between plants. We assume that plant-level turnover correlates positively and significantly with total job turnover. Baldwin, Beckstead and Caves (2001), for example, found that the Canada-US Free Trade Agreement led to significant commodity specialisation of individual Canadian plants. Trade expansion does, therefore, appear to lead to intra-plant as well as inter-plant adjustment. However, it would be difficult to argue against the assumption that intra-plant job moves are generally less costly than inter-plant inter-industry job moves. 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 within an industry. 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. Given these three plausible assumptions, *INTRA* serves as a good proxy for low labour adjustment costs.

4 Econometric Model and Results

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 Appendix (in particular, Appendix Table 1). The panel used for estimation consists of observations on 64 industries over a twelve year period, 1979 to 1990.

As shown in Appendix Table 1, intra-industry job turnover is a pervasive phenomenon. On average, 63 percent of jobs created have been matched by a contemporaneous job destroyed in the same industry in our sample period. The high mean level of intra-industry job turnover in Ireland is very similar to results found for other industrialised countries (Strobl, Walsh and Barry, 1998). On the trade side, we find that the average GL index in our sample is 0.66, and the average year-on-year A index is 0.31. These shares of IIT and MIIT are also in line with patterns observed for other comparable economies (Brülhart and Hine, 1999). In addition to job turnover and IIT, we use the following set of variables: *CONC*, the four-plant concentration ratio; *TRADE*, the degree of trade exposure; *ΔCONS*, the change in apparent demand; *TECH*, a proxy for the technology-intensity of an industry; *WAGE*, the average real wage[12]; and *FOREIGN*, the share of employment accounted for by foreign-owned plants.

The Econometric Model

Although we cannot draw on a unified theoretical framework for the specification of an empirical model, we do have some clear priors to guide model specification. The theoretical reasoning and existing empirical results that we have described in the previous two Sections of this paper suggest four main explanatory variables in an empirical model of *INTRA*.

First, one may anticipate 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 the four-plant concentration ratio *CONC*, therefore, is negative.[13]

Second, based on similar reasoning, there is likely to be a positive relationship between *INTRA* and trade exposure *TRADE*.[14] The positive relationship between trade exposure and intra-industry job turnover has been formalised in a dynamic IIT model with heterogeneous plants by Melitz (2003). Empirically, both plant concentration and trade exposure have previously been found to be significant determinants intra-industry job turnover by Baldwin and Caves (1997) in Canadian data, and by Andersson, Gustafsson and Lundberg (2000) in a data set for Sweden.

Third, inclusion of *ΔCONS*, which measures sector-level changes in demand, is also warranted *a priori*. The rationale goes as follows. It is implied in the SAH that sectoral shares in global output and expenditure remain constant. However, if the weight of a

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[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.

[13] It has been suggested that the number of plants or average plant size (both of which are trending variables) might be better at explaining *INTRA* (a stationary ratio) than *CONC*. We did not find this to be the case.

[14] The positive relationship between trade exposure and intra-industry job turnover has been formalised in a dynamic IIT model with heterogeneous plants by Melitz (1999). Empirically, both plant concentration and trade exposure have previously been found to be significant determinants intra-industry job turnover by

certain sector in world output and expenditure increases (decreases) while productivity and trade propensities remain unchanged, then there will be net inter-industry factor reallocation into (out of the) sector, even if trade changes are pure MIIT.[15] Fourth, and most importantly, we have strong priors about the coefficients on *IIT*. If the SAH, as used in the MIIT literature, is valid, the GL index (contemporaneous, lagged or first-differenced) and *INTRA* should be unrelated, *ceteris paribus*. However, one would expect to find a significant positive relationship with a measure of MIIT.

Finally, we also introduced three other variables - sectoral wages, exposure to foreign ownership, and technology intensity - which are known to be important in shaping industrial employment patterns in Ireland.[16]

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

$$INTRA_{it} = \alpha_i + \beta_1 CONC_{it} + \beta_2 TRADE_{it} + \beta_3 \Delta CONS_{it} + \beta_4 TECH_{it} + \beta_5 WAGE_{it} + \beta_6 FOREIGN_{it} + \beta_7 IIT_{it} + \lambda_t + \varepsilon_{it}, \qquad \textbf{(8)}$$

where α is an industry 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.

Baldwin and Caves (1997) in Canadian data, and by Andersson, Gustafsson and Lundberg (2000) in a data set for Sweden.

[15] This point has been developed formally by Lovely and Nelson (2002) in a one-country specific-factors model. We would thus anticipate a negative relationship between sectoral changes in apparent demand and INTRA, ceteris paribus.

[16] See e.g. Barry (1999) on the increasing importance of high tech, foreign-owned multinational companies and the absolute decline in more traditional Irish-owned manufacturing.

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.

Panel Results

Our main results are set out in Tables 2 and 3. The first column of results in Table 1 reports coefficient estimates of equation (8), using the GL index for IIT. The estimated coefficients on the variables that represent concentration, trade exposure and changes in demand are correctly signed according to our priors, and the first two are statistically significant.[17] The coefficients on the GL index, however, are statistically insignificant; and the coefficient on the lagged GL index does not even have the expected sign. We have experimented with various lags, leads, averages and interactions 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 1 all use the *A* index as the IIT variable. The one-year lagged *A* index consistently gave the best results.[18] This suggests that MIIT matters, and 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. Interacting the *A* index with trade exposure never yielded significant results (not reported). We also experimented extensively with the *C* and *UMCIT* measures but found no significant and robust relationships. The problem in our context with the *C* and *UMCIT* measures is that they are highly collinear with the trade exposure variable. Whilst measures

[17] 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, similar results were obtained using current concentration and trade exposure.

of total matched/unmatched trade changes such as *C* and *UMCIT* perform well in bivariate correlations with indicators of factor-market adjustment, their effect is difficult to isolate in a specification where trade exposure is separately controlled for.[19]

Column (3) of Table 1 sets out a parsimonious model, which is our statistically preferred specification.[20] 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 estimation bias from model misspecification.[21]

In Table 2, 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

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[18] Sensitivity tests on different lag lengths and interval sizes (n) show that the once-lagged one-year *A* index statistically dominates all other dynamic MIIT specifications (see Brülhart, 2000).

[19] We also explored indices proposed by Hamilton and Kniest (1991) and by Greenaway *et al*. (1994), as well as interaction terms of IIT measures and other explanatory variables, without finding significant results. In addition, we estimated the model with absolute values of *INTRA* as regressand instead of *INTRA*, and found the results to be qualitatively unaffected. Detailed tables are available from the authors.

[20] We used the automated econometric model selection and testing procedures in Pc-Gets (Hendry and Krolzig, 2001) to check on the validity of reducing the general model in column (2) to the parsimonious one in column (3). We also used PcGets to examine some more general specifications involving, *inter alia*, combinations of the various IIT indices and additional lags for all of the explanatory variables.

[21] 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.

viewed as a test of specification, the results in Table 2 amount to an informal test of the robustness of the Table 1 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 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 2 are presented for comparison with the results in column (2) of Table 1. 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, 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 1, 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*.

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 (2003). Further details are given in the Appendix (Appendix Table 2). The results suggest that one can reject the null hypothesis that *INTRA*, *ΔCONS* 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 (1999) 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.

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 (1999) 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 3. Industry concentration *CONC* and trade exposure *TRADE* are again the most significant variables, both with the expected signs. We also find as before that coefficients on $\Delta CONS$ have the expected sign but are not statistically significant. 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.[22] Table 4 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

[22] 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.

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.[23]

5 Conclusions

In this paper, we have drawn on a panel of Irish data for the 1980s to examine an important element of the "smooth-adjustment hypothesis" (SAH), which states that IIT is associated with lower factor-reallocation costs than inter-industry trade. We used the share of intra-industry job turnover as a proxy for labour-market adjustment. The choice of proxy is motivated by supportive results obtained in the labour economics literature. In addition, when modelling the share of intra-industry job turnover, we allowed for other potential influences by incorporating additional explanatory variables. We also supplemented traditional, static measures of IIT with more recently developed, dynamic measures of marginal IIT (MIIT).

The econometric results reject the existence of a significant relationship between intra-industry job turnover and static IIT in the 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 some empirical support to the SAH, as defined in the MIIT literature. We also found that the share of intra-industry

[23] 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.

job turnover increases in the degree of an industry's openness to trade and decreases in the four-plant concentration ratio.

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 highlights the need for further theoretical work on the link between (intra-industry) trade dynamics and labour-market adjustment.

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Table 1

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
CONC lagged	-0.420	-2.031	-0.410	-1.200	-0.396	-2.028	-0.412	-2.101
TRADE lagged	0.340	3.081	0.341	3.216	0.311	3.433	0.283	3.032
Δ*TRADE*	-0.663	-1.616	-0.637	-1.641	-	-	-	-
ΔCONS	-0.019	-0.165	-0.019	-0.167	-	-	-	-
TECH	-0.227	-1.165	-0.222	1.154	-	-	-	-
WAGE	-0.141	-1.155	-0.147	-1.186	-	-	-	-
FOREIGN	-0.137	-0.871	-0.117	-0.753	-	-	-	-
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 heteroskedasticity-consistent standard errors.

Table 2

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	-	-
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
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
ΔCONS	0.004	0.06	-0.018	-0.14	-0.018	-0.35	-0.016	-0.14	-0.027	-0.25	-	-	-	-
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	-	-	-	-
FOREIGN	-0.36	-1.22	-0.64	-1.55	-0.44	-1.18	-3.06	-1.60	-3.28	-1.66	-	-	-	-
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, ΔCONS*		*TECH, WAGE, FOREIGN, ΔCONS*		*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 2: (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 3

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.
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
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
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
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
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
Average GL	-0.03	-0.55	-	-	-	-	-	-	-	-	-	-	-	-
Average ΔGL	-	-	0.21	0.41	-	-	-	-	-	-	-	-	-	-
Δ*GL (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 3: All non-IIT data, as well as *GL*, are averaged across the 14 years 1977-90 for each of the 64 industries. *ΔCONS*, Δ*GL* and *A* are averaged across 13 year-intervals 1978-90. The *t* statistics are based on heteroskedasticity-consistent standard errors.

Table 4

Measures of "Economic Significance" - Beta Coefficients

Explanatory variables	(1) Fixed-effect panel results	(2) OLS cross section results
CONC	-	-0.709
CONC lagged	-0.380	-
TRADE	-	0.181
TRADE lagged	0.218	-
ΔTRADE	-0.095	-
ΔCONS	-0.005	-0.089
TECH	-0.071	0.061
WAGE	-0.084	-0.110
FOREIGN	-0.131	-0.139
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	-
Weighted C	0.016	
Weighted C lagged	0.050	
UMCIT	0.002	
UMCIT lagged	-0.036	
A (1977 base, 1990 end)	-	0.129
A (1977-79 base, 1988-90 end)	-	0.133
Average Weighted C	-	0.075

Notes to Table 4: 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 1; except for *GL* and *GL* lagged (model 1), *A* calculated over two-year intervals (model 4), and the *C* and *UMCIT* measures (not reported in Table 1). Standard deviations are calculated over pooled data. The OLS results in column (2) are based on the estimates set out in Table 4.

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. 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 (2003), since this test is more general than other panel unit root tests. The Im *et al*. (2003) 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.* (2003) 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*. (2003) 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*, *ΔCONS* 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 (1999) 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 7)	Forfás	0.63	0.26	1.00	0.00	1.00
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
TRADE	Imports plus exports divided by output	CIP, Eurostat	0.12	0.18	1.70	0.001	0.16
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
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
FOREIGN	Share of employment accounted for by plants under majority non-Irish ownership	Forfás	0.43	0.30	0.98	0.00	-0.17
IIT	• Grubel-Lloyd index of IIT (see equation 1)	Eurostat	0.66	0.25	0.9997	0.04	-0.04
	• *A* index of MIIT (see equation 2)	Eurostat	0.31	0.33	0.9994	0.00	-0.01
	• Weighted *C* index of MIIT (see equations 3 and 4, 1976 ECU '000, employment used as weight)	Eurostat	6.65	28.43	494.25	0.00	0.02
	• *UMCIT* measure of marginal inter-industry trade (see equation 5, 1976 ECU '000)	Eurostat	10,755	20,022	176,656	8.93	0.08

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
CONC		-1.51	-1.63	-1.47	-1.85	-2.06	-2.09
TRADE		-1.36	-1.35	-1.21	-2.07	-2.23	-1.79
ΔCONS		-3.33**	-2.76**	-1.95**	-3.42**	-2.84**	-2.08
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
FOREIGN		-1.24	-1.33	-1.23	-1.96	-2.11	-1.95
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.* (2003).