HOME-BIASED DEMAND AND INTERNATIONAL SPECIALISATION: A TEST OF TRADE THEORIES

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ABSTRACT

We develop and apply a discriminating criterion to distinguish the two principal paradigms of international trade theory: constant-returns perfectly competitive models on the one hand, and increasing-returns monopolistically competitive models on the other. Our criterion rests on the existence of home-biased demand. It predicts a positive relationship between countries' relative output and their relative home bias in increasing-returns sectors, and no relationship in constant-returns sectors. In implementing the test on data for OECD countries we find that industries accounting for up to two thirds of manufacturing output conform to the increasing-returns monopolistically competitive model.

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I. INTRODUCTION

International trade theory is dominated by two major paradigms. One paradigm belongs to the neo-classical world with constant returns to scale in production (CRS) and perfectly competitive product markets (PC). The other paradigm rests on the assumption of increasing returns to scale (IRS) and, in its most prominent formulation, monopolistically competitive markets (MC). While other important models exist which combine features of both paradigms, the theoretical and empirical debate has concentrated on these two powerful benchmark cases.

To distinguish between these two paradigms is of more than academic interest. Trade policies, market integration, migration, and other economic changes may have very different positive and welfare consequences depending on the underlying paradigm. It is therefore important to find a way of distinguishing the two paradigms in the data, and to quantify their respective importance in shaping industrial specialisation patterns. This is the purpose of our study.

In the theoretical part, we develop a discriminating criterion suitable for empirical estimation. The discriminating criterion relies on the assumption that demand is home biased. It posits that the home bias influences the pattern of international specialisation in sectors that are characterised by increasing returns and monopolistic competition while such bias is inconsequential for sectors characterised by constant returns and perfect competition. We test this hypothesis in 29 industries, covering 22 OECD countries for 1970-85. Our results suggest that 17 industries, accounting for about two thirds of industrial output, can be associated with the IRS-MC paradigm, while 11 industries can be associated with the CRS-PC paradigm. For one industry the results are inconclusive.

The paper is structured as follows. In Section II we provide a selective review of the relevant literature. Section III sets out our theoretical model and derives the discriminatory hypothesis. We operationalise this test empirically in Section IV. Section V concludes.

II. RELATED LITERATURE

II.1 Searching for Evidence of Increasing Returns in Industrial Specialisation Patterns

Several studies have attempted to gauge the relative explanatory power of the two main paradigms in trade theory directly or indirectly. A first group of studies pertains to the empirical industrial organisation literature (for reviews see Scherer and Ross, 1990; and Tybout, 1993). These studies estimate the incidence of plant-level increasing returns directly and generally do not find them to be pervasive. Estimates of industries' susceptibility to scale economies were analysed in conjunction with measures of industrial specialisation across the US states (Kim, 1995) and across EU countries and regions (Brülhart, 1998), and it was diagnosed that "scale-driven" geographical specialisation was mainly a phenomenon of the past.

A second group of studies focused on intra-industry trade as evidence of the importance of the IRS-MC paradigm (see Greenaway and Milner, 1986; and, for a critical appraisal, Leamer and Levinsohn, 1995). Since intra-industry trade was generally associated with IRS-MC models, the observed large and increasing shares of intra-industry trade were interpreted as evidence of the growing relevance of non-neoclassical trade models. This evidence became less persuasive when some studies, like Falvey and Kierzkowski (1987) and Davis (1995), demonstrated that intra-industry trade could also be generated in suitably amended versions of the CRS-PC framework.

A third approach was to enlist the excellent empirical performance of the gravity equation in support of the IRS-MC paradigm, since the gravity equation has a straightforward theoretical counterpart in the IRS-MC model (Helpman, 1987). This view was challenged by studies that showed that the gravity equation can also arise in a variety of other models (Davis and Weinstein, 1998b; Deardorff, 1998; Evenett and Keller, 1998; Feenstra, Markusen and Rose, 1999; Haveman and Hummels, 1997). On the empirical grounds, the challenge came from the evidence that the gravity equation fits excellently also on the set of non-OECD countries, a piece of evidence at odds with the assumptions of the IRS-MC paradigm (Hummels and Levinsohn, 1995).

II.2 The Magnification Effect

The scope of the relevant empirical literature has until recently been limited by the lack of a testable discriminating hypothesis that could serve to distinguish among theoretical paradigms in statistically rigorous fashion. A major breakthrough in this direction came in a series of papers by Davis and Weinstein (1996, 1998a, 1999). They developed a separation criterion based on the feature of IRS-MC models that demand idiosyncrasies are reflected in the pattern of specialisation more than one for one, thus giving rise to a magnification effect (pointed out originally in Krugman, 1980). Conversely, in a CRS-PC model, there is no magnification effect. Hence, with IRS-MC, a country will tend to export those goods on which it spends a larger share of its income than the world as a whole, and import those on which it spends a relatively smaller income share; and the reverse with CRS-PC. This feature can serve as the basis for empirical investigation. Sectors that exhibit the magnification effect are associated with the IRS-MC paradigm, while sectors that do not exhibit the magnification effect are associated with CRS-PC. Davis and Weinstein have estimated the magnification effect empirically in data for OECD countries (1996, 1998a) and for Japanese regions (1999), which allowed them to attribute industrial sectors to one of the two paradigms. In their first paper with international data (1996) they have found little evidence of the magnification effect and therefore scant support for increasing returns. However, in their later studies, which took account of cross-border demand dissipation (1998a) and of region-level specialisation patterns (1999), they produced evidence of pervasive magnification effects, suggesting that over half of industrial output was in industries that conform with the IRS-MC paradigm.

The Davis-Weinstein test is intuitively compelling and marks a significant step in bringing rigour to the empirical analysis of the new trade theory. Its major advantage is that it provides a reduced form that is common to all model variants within the IRS-MC paradigm, and absent from all models in the CRS-PC vein. Thus, it makes possible to distinguish between paradigms without having to examine and compare each of the multidimensional features of each of the paradigms (such as product differentiation or homogeneity, scale economies or constant returns, perfect competition or imperfect competition, etc). Of course, the test should not be interpreted literally, since no model in the CRS-PC or in the IRS-MC family is a precise description of any real-world economic sector. Yet, we may think it reasonable to believe that each real-world industry will resemble one of the two classes of models more than the other.

The magnification effect is only valid as a separation criterion if it is truly general to all (or to "all reasonable") models of one class and to none of the other. It is therefore important that this discriminating criterion be subjected to theoretical as well as empirical "sensitivity tests", so as to explore its robustness and degree of generality. Some papers have pointed out limits to the applicability of the magnification effect as a discriminatory criterion. Feenstra, Markusen and Rose (1999) find that the magnification effect may be generated also in CRS models with reciprocal dumping. Instead of the magnification effect they use a discriminating criterion according to which, in a gravity equation, the income elasticity of exports should be higher for differentiated goods than for homogeneous commodities. Head and Ries (1999) also show that the magnification effect can arise in settings that do not necessarily conform with the IRS-MC paradigm.

Another aspect of the test based on the magnification effect is that it appears to be sensitive to the modelling of trade costs. As outlined above, the empirical results of the two Davis-Weinstein (1996, 1998a) studies on OECD data differ substantially as a result of different assumptions about the effect of national borders on demand. In the purely theoretical context, Davis (1998) has pointed out that the magnification effect hinges on the relative size of trade costs in the CRS and IRS sectors, and that it vanishes if equal trade costs arise in both types of sectors. Head and Ries (1999) have turned this sensitivity to trade costs into a useful feature. They find that the size of the magnification effect increases with trade costs in CRS sectors and decreases with trade costs in IRS sectors, and they use this feature as a discriminating criterion. Our discriminating criterion does not depend on the presence of trade costs, and it remains valid even if trade costs are zero. This is an attractive feature given the fragility of the magnification effect with respect to assumptions about trade costs.

Turning to our contribution in this paper, we allow for the widely documented reality that goods from different countries are *ipso facto* considered imperfect substitutes (the Armington assumption), and that buyers are for a variety of reasons biased in favour of either home- or foreign-produced goods.[1] In such a model, a different type of home-market effect emerges, one that arises from the relative degree of home bias in expenditure. Our theoretical result is as

[1] Feenstra *et al*. (1999) and Head and Ries (1999) have used the Armington assumption in a similar context. Note that we parametrise the home bias in the utility function instead of defining it in terms of the income elasticity of imports, as in, e.g., Anderson and Marcouiller (1999).

follows. In an IRS-MC setting, relatively strong home bias in a country's aggregate expenditure on a good will make the country relatively specialised in the production of that good; whereas in a CRS-PC framework, relative home biases have no impact on the location of production. This result forms the basis for our empirical separation criterion.

Our model hinges on the existence of home-biased demand. We argue that this is a sensible claim, given the strong empirical evidence in its support. For example, Winters (1984) argued that, while demand for imports is not completely separable from demand for domestic goods, substitution elasticities between home and foreign goods are nevertheless finite. Davis and Weinstein (1998b) and Trefler (1995) found that by allowing for home-biased demand the predictive power of the HOV model could be improved very significantly. Head and Mayer (1998) identify home bias in expenditure as one of the most potent sources of market fragmentation in Europe. McCallum (1995), Wei (1996) and Helliwell (1996, 1997) find that trade volumes among regions within countries are generally a multiple of trade volumes among different countries even after controlling for geographical distance and other barriers. Finally, Brülhart and Trionfetti (1998) find evidence of home bias in public procurement. The assumption of home bias therefore seems to rest on solid empirical grounds.

III. THEORY: DERIVATION OF A DISCRIMINATING CRITERION

A model suitable for our analysis needs to accommodate both the CRS-PC and the IRS-MC paradigms. For this purpose, we use a framework close to that of Helpman and Krugman (1985, part III).

The world is composed of two countries indexed by i (i=1,2) and two homogeneous factors of production indexed by V (V=L,K). Each country is endowed with a fixed and exogenous quantity L_i and K_i of the factors. L and K are used to produce three commodities indexed by S (S =Y, X and Z). As will become clear below, the presence of trade costs results in the loss of one degree of freedom. To assure factor price equalisation we therefore use a three-goods-two-factors framework.

III.1 Technologies and Factor Markets

It is assumed that commodities Y and Z are produced by use of a CRS technology and traded in perfectly competitive markets without transport costs. Commodity X is assumed to be subject to an IRS technology and to trade costs. These trade costs are of the conventional “iceberg” type, where for each unit shipped only a fraction $t \in [0,1]$ arrives at its destination. The average and marginal cost function associated with a CRS sector is $c_S(w,r)$, where w and r are the rewards to L and K respectively. Production of X entails a fixed cost $f(w,r)$ and a constant marginal cost $m(w,r)$. As usual, it is assumed that technologies are identical across countries. In order to make factor intensities independent of plant scale, it is assumed that the functions $m(w,r)$ and $f(w,r)$ use factors in the same relative proportion. Thus, factor proportions depend only on relative factor prices. It is also assumed that there are no factor intensity reversals. The average cost function in the X sector is $c_X(w,r,x) = m(w,r)+f(w,r)/x$, where x is output. The number of varieties of X produced in the world, denoted by N, and the number of varieties produced in country i, denoted by n_i, are determined endogenously. The industry-level demand functions for L and K obtain from the cost functions through Shephard’s lemma and are denoted by $l_S(w,r)$ and $k_S(w,r)$.

The efficiency conditions and factor-market clearing conditions are:

$$p_S = c_S(w,r), \qquad S = Y, Z \qquad (1a)$$

$$p_X(1-1/s) = m(w,r), \qquad (1b)$$

$$p_X = c_X(w,r,x), \qquad (2)$$

$$l_Y(w,r)Y + l_X(w,r)xn_i \quad l\ (w,r)Z_i = L_i \qquad i = 1,2 \qquad (3a)$$

$$k_Y(w,r)Y_i + k_X(w,r)xn_i + k_Z(w,r)Z_i = K_i \qquad i = 1,2 \qquad (3b)$$

where s is the elasticity of substitution among varieties of *X*. Equations (1a) and (1b) state the usual conditions that marginal revenue equals marginal cost in all sectors and countries. Equation (2) states the zero-profit condition in sector *X* in all countries. Equations (3a) and (3b) state the market-clearing conditions for factors in all countries. These equations describe the supply side of the model. Free trade assures commodity price equalisation for goods *Y* and *Z*. The f.o.b. price of *X* is the same across countries and the c.i.f. price is simply p_X / t .

III.2 Demand

Households' preferences feature love for variety, represented by the traditional nested CES-Cobb-Douglas utility function. We extend the basic model by assuming that household demand is home biased. For simplicity, we model the home bias parametrically at the Cobb-Douglas level of the utility function, and represent it by the parameter h_i Î [0,1]. This way of parametrising the home bias is similar to Trionfetti (1999), but other structures are conceivable. One alternative representation would be by inserting a parameter as a weight in the CES aggregate. In Appendix 1 we show that the salient results also hold if the home bias is modelled in this alternative form.

When h_i=0, the household is not home biased. As h_i increases the household becomes increasingly home biased, and when h_i=1 the household purchases only domestically produced commodities. Thus, the representative utility function for the consumer in country i is as follows:

$$U_i = X^{(1-h_{Xi})a_{Xi}} X_i^{h_{Xi}a_{Xi}} Y^{(1-h_{Yi})a_{Yi}} Y_i^{h_{Yi}a_{Yi}} Z^{(1-h_{Zi})a_{Zi}} Z_i^{h_{Zi}a_{Zi}}, \qquad \text{with } \sum_S a_{Si} = 1,$$

and with CES sub-utility

$$X = \left(\sum_{k \in n_1} c_k^{(s-1)/s} + \sum_{k \in n_2} c_k^{(s-1)/s} \right)^{s/(1-s)}.$$

Denoting with E_{Si} the aggregate expenditure of households of country i on commodity S, we have $E_{Si} = a_{Si} I_i$, where I_i is aggregate income of households in country i (households have claims on capital). Two-stage utility maximisation and aggregation over individuals yields the aggregate expenditure of households of country i on each domestic variety of the commodity X. The ratio $E_{Si} / \sum_i E_{Si}$ is country i's share of world demand for good S. This ratio represents idiosyncratic demand and is the basis for the calculation of magnification effects.

III.3 Equilibrium in Product Markets

The equilibrium conditions in the product market require that demand equals supply for each commodity and each variety. The market-clearing conditions are:

$$p_X x = \frac{p_X^{1-s}}{P_{X1}^{1-s}}(1-h_{X1})E_{Xi} + \frac{q\, p_X^{1-s}}{P_{X2}^{1-s}}(1-h_{X2})E_{X2} + \frac{1}{n_1} h_{X1} E_{X1} \tag{4}$$

$$p_X x = \frac{q\, p_X^{1-s}}{P_{X1}^{1-s}}(1-h_{X1})E_{Xi} + \frac{p_X^{1-s}}{P_{X2}^{1-s}}(1-h_{X2})E_{X2} + \frac{1}{n_2} h_{X2} E_{X2} \tag{5}$$

$$E_{Y1} + E_{Y2} = p_Y \left(Y_1 + Y_2 \right) \tag{6}$$

where $q = t^{s-1}$, and P_{Xi} is the usual CES price index. Equation (4) states the equilibrium condition for the varieties of IRS good produced in country 1, expression (5) states the equilibrium condition for the varieties produced in country 2, and expression (6) states the equilibrium condition in the market for CRS good Y. By Walras' law, the equilibrium condition for the other CRS good Z is redundant. The model so far is standard except for the home bias. The system (1)-(6) is composed of eleven independent equations and twelve unknowns (p_X, p_Y, p_Z, x, n_1, n_2, Y_1, Y_2, Z_1, Z_2, w, r). Taking p_Z as the numéraire, the system is perfectly determined. In the absence of trade costs, equations (4) and (5) are not independent. Hence, a two-by-two model would guarantee full dimensionality of the factor-price-equalisation set (eight independent equations and nine unknowns). The presence of trade costs segments the market for the differentiated commodity and, therefore, requires two equations for that market. Thus, in the presence of trade costs, a two-by-two model would have too many equations for the factor-price-equalisation set to be of full dimensionality. To restore full dimensionality we need one more commodity. This is why we use a three-by-two model. While it is analytically convenient to work within the factor-price-equalisation space, our results do not depend on its existence. Under appropriate conditions, the dimensionality of the model could be extended to any V, S, and i, with $S > V$.

III.4 A Discriminating Criterion

There is a difference between the two CRS-PC sectors and the IRS-MC sector that can be immediately found by inspection of equations (4)-(6). The difference is that the parameter representing the home bias cancels out of equation (6), while it does not cancel out in equations (4) and (5). Hence, the home bias does not affect international specialisation in the CRS-PC sectors but it affects international specialisation in the IRS-MC sectors. This is the essence of our discriminating criterion. We associate sectors with the IRS-MC paradigm if the home bias is significant in explaining international specialisation in the sector in question. Conversely, if the home bias is not significant, we associate the sector with CRS-PC.

It is useful at this point to rewrite equations (4)-(5) in terms of the ratios $h \equiv n_1 / n_2$ and $e \equiv E_{X1} / E_{X2}$. This gives:

$$
\begin{aligned}
& qh_2 h^3 + \left[(1-q)(1-h_2) + \left(1+q^2\right) h_2 - q(1-q)e(1-h_1) - qeh_1 \right] h^2 + \\
& + \left[q(1-q)(1-h_2) - (1-q)e(1-h_1) + qh_2 - \left(1+q^2\right) eh_1 \right] h + qeh_1 = 0
\end{aligned}
\tag{7}
$$

It is well known that the roots of a third degree polynomial, especially when there are so many parametrical coefficients, are unwieldy expressions. Fortunately, we can glean the fundamental features of the model by simple inspection of (7), without having to use the explicit solutions. To examine the relevant issues, we consider an exogenous idiosyncratic preference shock $da \equiv da_{X1} = -da_{X2}$ that generates the idiosyncratic expenditure shock $de \equiv dE_{X1} = -dE_{X2}$.

1. *The benchmark case*. It is immediate that if demand is unbiased in both countries (i.e., if h_1=h_2=0) then the left-hand side of (7) reduces to a first-degree polynomial whose solution is: $h = (e-q)/(1-qe)$. This is the benchmark case of the Davis-Weinstein studies, and it has the property that $dh/de > 1$, i.e. it gives rise to a magnification effect. Since the magnification effect is a feature of IRS-MC sectors and not of CRS-PC sectors, it has been employed as a discriminating criterion. If trade costs in the IRS-MC sector are zero, the discriminating criterion based on the magnification effect is no longer valid, because at t=1 the derivative is zero for all paradigms.
2. *No magnification effect in IRS-MC sectors*. If demand is home biased, the derivative in question is not necessarily larger than one for the IRS-MC sector. To show this, we should differentiate (7) around each of the roots in the admissible range of h and then evaluate the size of the derivative at each admissible root. This is a complex procedure that we relegate to Appendix 1. Here, we show a simple case that can be understood by inspection of (7). Suppose that households in country 1 are not home biased while those in country 2 are completely home biased (i.e., h_1 = *0*; and h_2 = 1). Then, equation (7) reduces to a second-degree polynomial whose only positive root is $h = (1-q)e - q$. It is immediate that

$dh/de < 1$ in this case. In general, the effect of idiosyncratic demand can be offset by idiosyncratic home bias. The fact that the derivative in question may be smaller than one invalidates the generality of the magnification effect as a discriminating criterion.

3. *Zero trade costs*. A test based on home-biased demand, unlike one based on the magnification effect, gives a discriminating criterion that is valid even in the absence of trade costs. To see this it suffices to set $q = 1$ in equation (7). This gives the unique solution $h = e(h_1/h_2)$. Clearly, in this extreme case, the home bias drives the pattern of specialisation. Note that, if neither country is home biased (h_1=h_2=0), and there are no trade costs, the solution is undetermined (0/0) and the derivative is zero in all paradigms. The independence of our discriminating criterion from trade costs is a welcome feature. This is not because trade costs are unimportant in reality, but because we have reduced the sensitivity of the discriminating criterion to the way trade costs are modelled or measured.

4. *The discriminating criterion.* It is straightforward to show that $dh/dh\big|_{dh\equiv dh_1=-dh_2>0} > 0$ for any parameter value. Consider a symmetric change dh_1=-dh_2>0. As a consequence of the change, the term on the right-hand side of (4) increases by $(qn_2/n_1)(n_1+qn_2)^{-1}E_{X1} + q(qn_1+n_2)^{-1}E_{X2} > 0$. Since the left-hand side of (4) is constant, the increase in the right-hand side requires an increase in n_1 in order to satisfy (4). The same holds, *mutatis mutandis*, for (5). This is true for any values of parameters and for any associated solution of the system. The explicit solutions for most configurations are extremely complex; but, as a useful example, the case of identical countries can be solved in a manageable way. Setting $e = 1$ and h_1=h_2=h, then solving for h and differentiating around the solution we have:

$$dh/dh\big|_{dh\equiv dh_1=-dh_2>0} = 4q(1+q)\Big/\left[(1-q)^2 + 4qh\right] > 0,$$

which confirms a positive response of production shares to idiosyncrasies in the degree of home bias.

We can use these findings to devise a discriminating criterion based on home biased demand. Differentiating (7) around any of its solutions with respect to an idiosyncratic change $de \equiv dE_{X1} = -dE_{X2}$ and to a change $dh \equiv dh_1 = -dh_2$ gives:

$$dh = c_1 de + c_2 dh \,, \tag{8}$$

where the coefficients c_1 and c_2 are the partial derivatives. Based on the fact that home biased demand matters only for IRS-MC sectors we can derive the following discriminating criterion. Sector S is IRS-MC if the estimated c_2 is larger than zero, and CRS-PC if the estimated c_2 equals zero. This discriminating criterion and its empirical implementation are the focus of this paper. In addition, as discussed above, if trade costs are zero idiosyncratic demand does not affect international specialisation. Accordingly, an estimated value of c_1 larger than zero reveals the importance of trade costs.

Association of Parameter Values with Paradigms in the New Test

c_1	**Trade costs**	c_2	**Paradigm**
+	$q < 1$	+	IRS-MC
+	$q < 1$	0	CRS-PC
0	$q = 1$	+	IRS-MC
0	$q = 1$	0	CRS-PC

IV. EMPIRICAL IMPLEMENTATION

In operationalising our discriminating criterion, we need to proceed in two stages. First, we estimate home biases across industries and countries. Those bias estimates can then be used as an ingredient to the estimation of our testing equation (8).

IV.1 Estimating Home Bias

We estimated home biases across countries and industries using an augmented form of the standard gravity equation. Thanks to the general compatibility of this approach with the major theoretical paradigms, using of the gravity equation at the first stage of our exercise should not prejudice the testing of theories in stage two.[2] We estimated variants of the following regression equation:

$$\begin{aligned} &LOGTRADE_{ij,t} = a + b_1 HOMEDUM_{ij} + b_2 LOGGDP_{ij,t} + b_3 LOGCGDP_{ij,t} + \\ &b_4 LOGDIST_{ij} + b_5 LOGREMOTE_{ij,t} + b_6 BORDUM_{ij} + b_7 LANGDUM_{ij} + \\ &b_8 PTADUM_{ij,t} + b_9 CLOSEDDUM_{ij} + u_{ij,t} \end{aligned} \quad , \qquad (9)$$

where the acronyms have the following meanings (for details on the construction of these variables, see Appendix 2):

LOGTRADE$_{ij,t}$ = log of (imports + exports) between countries *i* and *j* in year *t*,

HOMEDUM = dummy which is 1 if *i*=*j*, and zero otherwise,

LOGGDP = log of (GDP_i*GDP_j),

LOGCGDP= log of the product of per-capita GDP in *i* and *j*,

LOGDIST = log of geographical distance between the two countries,

LOGREMOTE = a measure of remoteness of *i* and *j* from the other sample countries,

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[2] On the generality of the gravity equation, see Davis and Weinstein, 1998b; Deardorff, 1998; Evenett and Keller, 1998; Feenstra *et al*., 1999; Haveman and Hummels, 1997. The gravity model has been shown to be successful even at the level of individual industries *inter alia* by Bergstrand, 1990; Davis and Weinstein, 1998b; Feenstra *et al*., 1999; Head and Mayer, 1999.

BORDUM = dummy which is 1 if *i* and *j* share a common border, and zero otherwise,

LANGDUM = dummy which is 1 if *i* and *j* share a common language, and zero otherwise,

PTADUM = dummy which is 1 if *i* and *j* are fellow members of a preferential trade agreement,

CLOSEDUM = dummy which is 1 if at least one of the two countries was described as "closed" by Sachs and Warner (1995).

The object of our interest is b_1, the coefficient on trade within countries. A positive (negative) coefficient is interpreted as positive (negative) home bias. By including variables for distance, adjacency, language, PTAs and institutional obstacles to trade we aim to control for physical transport costs, tariff and non-tariff barriers as well as for informational and marketing costs in accessing foreign markets.[3] To the extent that we manage to control for supply-side-driven cost differentials between domestic and foreign suppliers through inclusion of these variables, *HOMEDUM* will pick up the effect of home-biased demand.

The difficulties in implementing this approach are twofold. First, one has to find a measure of "trade within countries", and, second, the distance variable also has to be defined for intra-country trade. Following Wei (1996) and Helliwell (1997), we define trade within countries as output minus exports.[4] The validity of this measure rests on the assumption that all output recorded in the statistics is sold in a different location from its place of production, i.e. neither consumed *in situ* nor used as an intermediate input in the original plant. The official definition

[3] See Rauch (1999) on informational trade costs. Anderson and Marcouiller (1999) have shown that corruption and imperfect contract enforcement can act as additional impediments to international trade. Data limitations do not allow us to control for such factors. It seems plausible, however, that these effects will be less significant in our OECD-dominated country sample than in their data set, which included a large share of developing countries. Note also that we could invest greater confidence in our estimates if better measures of tariff and non-tariff barriers were available, ideally at sector and year level. Unfortunately, such data do not exist for the time period analysed in this paper.

[4] Hence, $LOGTRADE_{ii,t} = \log(2*[Output\text{-}Exports]_{ii,t})$.

of the “production boundary” in national accounts statistics supports us in making this assumption: “goods and services produced as outputs must be such that they can be sold on markets or at least be capable of being provided by one unit to another (…). The System [of national accounts] includes within the production boundary all production actually destined for the market” (OECD, 1999).

For estimates of “intra-country distances” we used the approach of Keeble, Offord and Walker (1986) and Leamer (1997), who defined them as equivalent to a fraction of the radius of a circle with the same area as the country in question.[5] This method may appear crude, but Head and Mayer (1999) found it to produce strikingly similar results to a more sophisticated approach that could draw on regional data for the EU. The main weakness of this approach is its sensitivity to the choice of divisor *x*, which is arbitrary.

Having constructed the intra-country variables, we estimated equation (9) on data for 29 industrial sectors, 22 OECD countries and 16 years (1970-85) drawing mainly on the OECD’s COMTAP database as made available by Feenstra, Lipsey and Bowen (1997). This yielded a data set with over 212,000 year- and industry-level bilateral observations. A full description of variables and data sources can be found in Appendix 2.

We began by running several variants of equation (9) on the entire data set. These results are given in Table 1. First, we simply pooled the data and estimated our gravity equation using OLS (column (1) of Table 1). With the exception of the border dummy, all our coefficients have the expected signs and magnitudes and are statistically significant. A coefficient of 1.31 on *HOMEDUM* suggests that *ceteris paribus* a country’s trade with itself is on average 2.7

(=$e^{1.31}$-1) times as large as trade with another country. This estimate may seem implausibly high. However, it falls within the range of results found by Wei (1996) for aggregate trade among a smaller OECD sample, which lie between 1.3 and 8.7, and it is identical to the coefficient estimated by Helliwell (1997, Table 2 vii) in his most comparable specification.[6] The magnitude of the estimated home bias is, however, sensitive to the way we proxy intra-national distances. We have therefore experimented with different definitions of $LOGDIST_{ii}$. *LOGDIST1* (column (2) of Table 1) assumes intra-national distances to be equivalent to the full radius of a circle with the country's surface area, that is *LOGDIST1* assumes larger average intra-national distances than the default *LOGDIST*. As a consequence, the estimated home bias increases to a factor 12.7 (=$e^{2.62}$-1). The converse is true in the case of *LOGDIST2*, which assumes smaller average intra-national distances than *LOGDIST*, using a divisor x of 6 (column (3) of Table 1). With *LOGDIST2*, the estimated home-bias factor falls to 0.6 (=$e^{0.48}$-1). Obviously, we need to be careful in interpreting the absolute magnitude of the home-bias estimates. Fortunately, this is not a problem in the context of our paper, since what we need in order to apply our discriminating criterion is an estimate of *relative* home biases, and these are unaffected by the definition of *LOGDIST*.

We have explored the robustness of our findings by applying different estimators. We ran a Tobit model to take account of the censoring in our data set due to the omission of observations with zero trade (column (5) of Table 1). This resulted in a slightly higher point estimate on *HOMEDUM* than OLS (1.34 vs. 1.31), which should not surprise, given that intra-country trade

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[5] Hence, $LOGDIST_{ii} = {}^{1}\!/_{x}\sqrt{\dfrac{Area_i}{p}}$.

[6] This figure might also seem at first sight to be incompatible with openness ratios (trade/GDP), which, in our sample period, range between 0.14 (US) and 0.90 (Belgium). However, it must be borne in mind that this is a bilateral and conditional exercise. A coefficient on HOMEDUM of 1.31 suggests that, on average, a representative agent facing two potential trade partners, one domestic and one foreign, will be 2.7 times more likely to trade with

is defined in all industries, whilst zero observations only appear for inter-country observations. In view of the small proportion of zero observations in our data set (0.6% of all observations) and of the resulting marginal impact on coefficient estimates, we proceeded nevertheless to run the disaggregated gravity regression using OLS.

We also ran fixed-effects and random-effects panel models with year dummies to relax the restriction of identical intercepts across years, countries and sectors (columns (6) to (9) of Table 1). In this instance, we found that the size of the estimated coefficient on *HOMEDUM* was substantially larger than in the pooled OLS run (1.44-1.56 vs. 1.31). Imposing identical coefficients across the three dimensions of our panel is clearly restrictive. This result supports our approach of estimating the gravity model at industry level. Hence, our next step was to run equation (9) separately for each of the 638 country-industry observations (22 countries*29 industries), so as to get individual home-bias estimates. We used fixed-effects panel regressions, allowing for variation in year intercepts. The resulting coefficients on *HOMEDUM* are reported in Table 2. These numbers are crucial to our subsequent analysis. They are empirical proxies for h_i in our theoretical model, and therefore represent the key ingredient to the empirical test of competing paradigms.

It is worth taking a look at the results in Table 2 in order to assess their plausibility. Average home biases across countries are shown in the rightmost column. The highest averages appear for ISIC 313 (beverages), 311 (food products) and 312 (food n.e.s.). Given the strong home bias implied in most OECD countries' agricultural policies, this seems plausible. The lowest home-bias factors are found for ISIC 354 (petroleum and coal products), 372 (basic metal products) and 353 (petroleum refining). The figures in the bottom row of Table 2 report average home

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the domestic partner, even after controlling for distance and other cost factors, but it *does not* mean that the biases across industries. The highest average home bias are those of Japan, Turkey and Greece, and the lowest those of New Zealand, Australia and the Netherlands. However, the estimated average coefficients for New Zealand (-23.7) and for Australia (-8.5) seem excessively low, even though those countries are known to have particularly strong trade links with faraway fellow OECD countries. Indeed, the magnitude of those estimates is extremely sensitive to the construction of the remoteness measure (*LOGREMOTE*) in the gravity equation, which itself is subject to a substantial degree of arbitrariness. We therefore felt compelled to drop observations relating to these two countries from our data set for the subsequent estimations.

IV.2 An Empirical Test of the Discriminating Criterion

Having computed empirical estimates of countries' home bias for each industry in our data set, we could proceed to an implementation of the testing equation (8). We estimated the following variant of this equation for each of the 29 industries *s* across all 22 countries *i*,*j* and 16 years *t*:

$$Output_{it}^{s} = a^{s} + g_{1}^{s} SHARE_{it}^{s} + g_{2}^{s} IDIODEM_{it}^{s} + g_{3}^{s} IDIOBIAS_{it}^{s} + d_{t}^{s} + u_{it}^{s}, \qquad (10)$$

where subscripts denote countries and years, superscripts denote industries, and:

$$SHARE_{it}^{s} = \frac{\sum_{i} Output_{it}^{s}}{\sum_{i}\sum_{s} Output_{it}^{s}} * \sum_{s} Output_{it}^{s},$$

$$IDIODEM_{it}^{s} = \left(\frac{Expenditure_{it}^{s}}{\sum_{s} Expenditure_{it}^{s}} - \frac{\sum_{i} Expenditure_{it}^{s}}{\sum_{i}\sum_{s} Expenditure_{it}^{s}}\right) * \sum_{s} Output_{it}^{s}, \text{ and}$$

$$IDIOBIAS_{it}^{s} = (Bias_{i}^{s} - [(\sum_{j \neq i} Bias_{j})/21]) * \sum_{s} Output_{it}^{s},$$

21 being the number of countries in our sample, minus one.

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average openness ratio Ω should satisfy Ω=(1-Ω)/2.7=0.27.

According to our hypothesis, industries with estimated g_3 of zero conform with the CRS-PC paradigm, whereas industries with positive estimated g_3 conform with the IRS-MC scenario. The other variables are constructed equivalently to the Davis-Weinstein testing equation. In our data set, *IDIOBIAS* and *IDIODEM* are not correlated (see Appendix 2).

Three issues warrant discussion. First, there is the question of sectoral disaggregation. Davis and Weinstein have nested the IRS-MC model in an encompassing Heckscher-Ohlin framework *à la* Helpman (1981). They assumed in most of their empirical specifications that factor endowments determine specialisation at the "industry" level (2- or 3-digit statistical headings) whilst the distinction between IRS-MC and CRS-PC is only relevant at the sub-industry level of "goods" (3- or 4-digit headings). However, neither theory nor existing empirical work gives us any strong priors as to the correct nesting of the theory. In our model, there is no hierarchy between "industries" and "goods". As a consequence, factor endowments do not appear in the reduced-form testing equation (8). This is a result of the realistic assumption in our model that there are more goods than factors, and it is a convenient feature in view of empirical implementation.[7]

Second, u_{it}^{s} is likely to be heteroskedastic, as the variance of errors may well be positively correlated with the size of countries.[8] Our significance tests are therefore based on heteroskedasticity-corrected standard errors, using the residual variance estimator of MacKinnon and White (1985). We make this conservative adjustment in order to minimise the

[7] Note that Davis and Weinstein (1998a) found *SHARE* to be highly collinear with their endowment variables, and therefore dropped it from their testing specification. Since we are doing the reverse, omitted-variable bias is likely to be small. Another reason for working at a higher level of aggregation is that running the testing equation at the level of 4-digit industries severely curtails degrees of freedom and therefore produces few robust results (see Davis and Weinstein, 1996).

[8] A Cook-Weisberg heteroskedasticity test on the pooled model strongly rejects the null of constant error variance.

risk of wrongly attributing sectors to the IRS-MC paradigm due to underestimation of the standard error of g_3.

Third, we cannot rule out simultaneity of expenditure and output, and therefore bias in the estimates of g_1 and g_2. The assumption underlying our test is that expenditure shares are an exogenous determinant of output location. Some support for the assumption that demand determines production and not the other way around is found in Davis and Weinstein (1998a, p. 29ff.) and in Trionfetti (1999).

IV.3 Results

We first ran our model on the pooled sample. This yielded the following point estimates:

$$\text{Output} = a + \underset{[t = 21.5]}{(1.2*10^8)*SHARE} + \underset{[t = 4.51]}{1.4 * IDIODEM} + \underset{[t = 3.51]}{0.0005 * IDIOBIAS} + d_t,$$

with 8920 observations and $R^2 = 0.46$.

This gives us a preliminary indication about the importance of increasing returns. The parameter estimate on *IDIOBIAS* is positive and significant. The coefficient on *IDIODEM*, being larger than one, suggests presence of the magnification effect. However, we cannot reject the hypothesis that $g_2 \leq 1$, hence the magnification effect is not statistically significant on aggregate.[9]

We know *a priori* that pooled runs impose too much structure, since our fundamental expectation is to find different parameter estimates for individual industries. The results of our industry-by-industry regressions with year fixed-effects are given in Table 3. The equation generally performs well, yielding R^2s in the range 0.78 to 0.96. Coefficient estimates on

[9] This is true also if we drop *IDIOBIAS* from the estimating equation.

IDIOBIAS are in the expected positive or insignificant range for all industries bar one (ISIC 332: furniture).

Again, we apply a conservative criterion for accepting the IRS hypothesis. We only reject the null that g_3=0 if the heteroskedasticity-adjusted *t* test is significant at the 99% level. Using this stringent criterion, we find that of the 29 industries, 17 conform with the IRS-MC paradigm, 11 conform with CRS-PC, and 1 produces a meaningless result (i.e. a significantly negative estimated g_3). The allocation of sectors looks broadly plausible. For instance, the four "canonical examples" of increasing-returns industries according to Davis and Weinstein (1998, 1999) all belong to the IRS-MC category according to our exercise: textiles, iron and steel, transport equipment, and instrument engineering.

It is interesting to measure the relative importance of the two paradigms in terms of percentage of industrial output. IRS-MC sectors appear to account for nearly two thirds of industrial production, the rest being accounted for by CRS-PC sectors. Precisely, the combined output share of the 17 IRS sectors in our data set was 60.5% in 1970 and 63.9% in 1985, whilst the output share of the 11 CRS industries amounted to 38.0% in 1970, and to 35.0% in 1985.

Some caution is of course warranted in the interpretation of our results. The sector with a negative estimated g_3 sounds a warning. Furthermore, some of our results are at odds with informed intuition. For instance, it does not seem very plausible that ISIC 382 (non-electrical machinery) is characterised by constant returns while, for instance, ISIC 323 (leather goods) is driven by increasing returns. Overall, however, the association of sectors with paradigms resulting from the empirical investigation seems quite plausible.

A final comment is in order. Our estimates of the parameter g_2 are consistent with those of Davis and Weinstein (1996), namely, we do not find evidence of the magnification effect. Our interpretation is different however. We do not interpret the absence of magnification effects as validation of the CRS-PC paradigm, because, as discussed in the theory section, g_2 may be smaller than one in both CRS-PC and IRS-MC sectors when demand is home biased.

V. SUMMARY AND CONCLUSIONS

We have developed and applied an empirical test which serves to separate the two principal paradigms of international trade: models with constant returns and perfect competition (CRS-PC), and models with increasing returns an monopolistic competition (IRS-MC). The test rests on the assumption that demand is home biased, an assumption that is strongly supported by empirical evidence. We show theoretically that specialisation patterns are affected by intercountry differences in home bias if an industry conforms to the IRS-MC paradigm, but not if it is characterised by CRS-PC. This finding provides us with a discriminating criterion that we implement empirically. In the empirical part we estimate industry- and country-level home biases through a gravity approach, and use these estimates to apply our test to a data set with 29 industries, covering 22 OECD countries in 1970-85. The results suggest that 17 industries, accounting for almost two thirds of industrial output, can be associated with IRS-MC; 11 industries, accounting for little over one third of industrial output, can be associated with CRS-PC; and for one industry the results are inconclusive.

A particularly fruitful generalisation of the theoretical framework would be to consider Ricardian technological differences across countries. This undoubtedly important determinant of international trade and specialisation patterns deserves explicit consideration. On the empirical side, two types of data would be useful in refining the work reported in this study. First, availability of country- and industry-level measures of trade costs would allow greater confidence that the coefficient on home dummies in the gravity equation picks up the residual effect due purely to demand bias. Second, more sectorally disaggregated data would likely make for better approximations of what constitutes a “good” in competing theoretical paradigms.

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APPENDIX 1: THEORETICAL EXTENSIONS

1. No Magnification Effect in IRS-MC Sectors When Demand is Home Biased

We can show that the derivative dh/de is not necessarily larger than one when demand is home biased. Let us define the left-hand side of (7) as a $P(h,e,h_1,h_2,q)$. Unfortunately, in general, the roots of this polynomial are long expression that are not manageable (not even with math packages) except in the following two cases.

Case A: Identical Countries

Setting $e=1$ and $h_1=h_2=h$ the only positive root of the polynomial is $h=1$. Differentiating totally around this root gives:

$$\frac{dh}{de}(h,q)=-\frac{\partial P(.)/\partial e}{\partial P(.)/\partial h}=\frac{\left(1-q^2\right)(1-h)+(1+q)^2 h}{(1-q)^2(1-h)+(1+q)^2 h}>1$$

This derivative, although it is larger than one, i.e., it exhibit the magnification effect, it is everywhere smaller than the derivative under the benchmark case, which is:

$$\frac{dh}{de}(q)=-\frac{\partial P(.)/\partial e}{\partial P(.)/\partial h}=\frac{(1+q)}{(1-q)}>1$$

Hence, home biased demand attenuates the size of the magnification effect.

Case B: Non-Identical Countries

In this example we drop the assumption of identical countries by assuming that only country 2 is home biased (i.e., $h_1=0$) so that the polynomial $P(h,e,h_1,h_2,q)$ reduces to a second-degree polynomial. Differentiating around its only positive root gives:

$$\frac{dh}{de}(q,e,h_2)=\frac{(1-q)\left[(e+h_2)^2q^2+2(2eh_2+h_2-e)q+1\right]^{\frac{1}{2}}+(e+h_2)q^2+(h_2-e-1)q+1-2h_2}{2h_2\left[(e+h_2)^2q^2+2(2eh_2+h_2-e)q+1\right]^{\frac{1}{2}}}$$

This expression is not particularly informative. Yet, after assigning specific values to one of the three parameters, the expression can be plotted in a three-dimensional space. An example is in Figure A1, where the expression for $\frac{dh}{de}(q,e,h_2)$ is compared with the constant 1. The figure (computed at $e=1$) shows a large domain in which $\frac{dh}{de}(q,1,h_2)<1$, i.e., no magnification effect exists. Many other combinations of parameter values also yield no magnification effect.

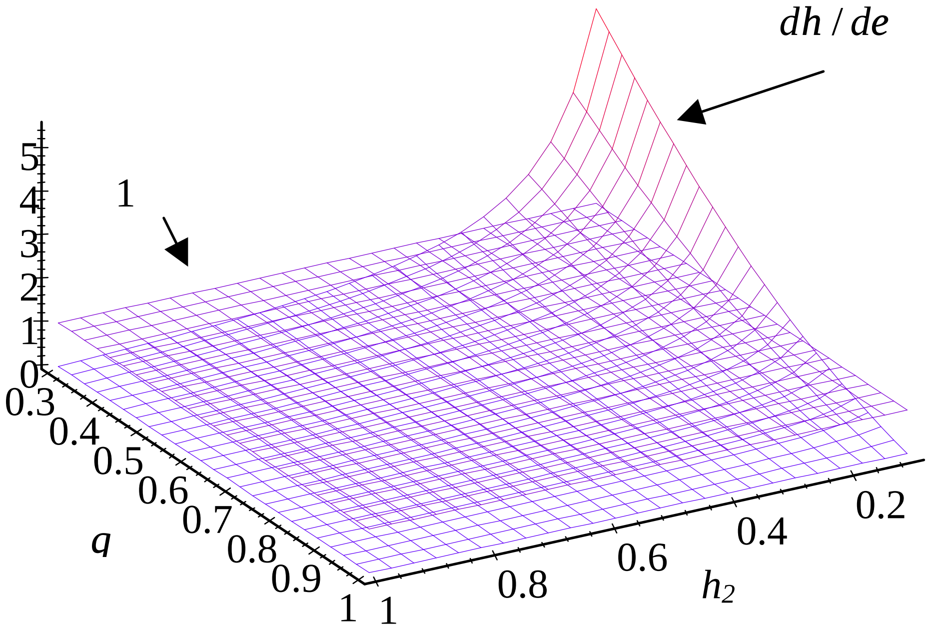

Figure A.1: An Example of the Domain with No Magnification Effect.

2. Home Bias in the CES Aggregate

The home bias may be inserted in the utility function also at the CES sub-utility level, for instance in the following way.

$$u_i = \left[\sum_{k \in n_i} \left(a_i c_{iik} \right)^{\frac{s-1}{s}} + \sum_{k \in n_j} \left(b_i c_{ij} \right)^{\frac{s-1}{s}} \right]^{\frac{s}{s-1}},$$

where the weights represent the bias. Then, defining $h_i \equiv a_i / b_i$, the market equilibrium equations (compacted into one) for the IRS-MC product becomes

$$\frac{h_1 - q}{h_1 n_1 + q n_2} E_{X1} + \frac{q - h_2}{n_1 q + h_2 n_2} E_{X2} = 0 \ ,$$

and the derivative is

$$\frac{dh}{de}(q, e, h_1, h_2) = (h_1 - q) \frac{h_2^2 h_1 - h_2 h_1 q - q^2 h_2 + q^3}{\left(q^2 e + h_2 h_1 - h_1 q e - h_1 q \right)^2},$$

which is not necessarily larger than one, even at $e = 1$. Hence, the magnification effect cannot serve as a robust discrimination criterion here either.

APPENDIX 2: DATA DESCRIPTION

Bilateral trade data are taken from the World Trade Database, and production data are from the OECD Comtap database, both circulated by Feenstra *et al.* (1997). Since data recorded by the importing countries were retained, all flows are c.i.f. Hence, our estimates of trade within country can be considered conservative. Observations for which estimated intra-country trade was, implausibly, negative (i.e. Output-Exports<0) were dropped from the sample. 366 observations, accounting for 0.48 % of total output and 3.73% of total trade in the full sample, were thus omitted.

The following 22 countries are contained in our sample: Australia, Austria, Belgium-Luxembourg, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Turkey, UK, USA, Yugoslavia. A list of the 29 ISIC industries in our data set can be found in Table 3. Distance data measure Great Circle distances between capital cities and are taken from Jon Haveman, Purdue University (http://www.eiit.org/Trade.Resources/TradeData.html).

Inclusion of a remoteness variable in the gravity model has been advocated persuasively by Polak (1996). *LOGREMOTE* is defined as follows (in the spirit of Helliwell, 1997):

$$LOGREMOTE_{ij,t} = \sum\nolimits_{k \neq i,j} [LOGDIST_{ik} + LOGDIST_{jk} - 2LOGGDP_{k,t}] .$$

The linguistic groupings underlying *LANGDUM* are defined as follows:
English: Australia, Canada, Ireland, New Zealand, UK, USA; *French*: Belgium, Canada, France; *German*: Austria, Germany; *Dutch*: Belgium, Netherlands; *Scandinavian*: Denmark, Sweden, Norway.

The preferential trade areas underlying *PTADUM* are defined as follows:
EU: Belgium, Denmark (1973-), Greece (1981-), France, Germany, Ireland (1973-), Italy, UK (1973-).
EFTA: Austria, Denmark (-1972), Norway, Portugal, Spain, Sweden, UK (-1972).
Australia-New Zealand Free Trade Agreement: Australia, New Zealand.

The dummy for trade “closedness” is from Sachs and Warner (1995). The dummy is set to zero if a country satisfies four tests: (1) average tariff rates below 40 percent; (2) average quota and licensing coverage of imports of less than 40 percent; (3) a black market exchange rate premium that averaged less than 20 percent during the 1970s and 1980s; and (4) no extreme controls (taxes, quotas, state monopolies) on exports. It turns out that this measure is time invariant in our data set. According to the Sachs-Warner criteria, New Zealand, Turkey and Yugoslavia were “closed” for the whole time interval 1970-85, whilst the remaining 19 countries were “open” throughout this period.

Correlations among the variables of our testing equation (underlying Table 3, no. obs. = 9816):

	Output	*SHARE*	*IDIODEM*	*IDIOBIAS*
Output	1.000			
SHARE	0.624	1.000		
IDIODEM	0.219	-0.002	1.000	
IDIOBIAS	0.287	0.364	-0.049	1.000

TABLE 1: Gravity Equations: Full Sample

(22 countries, 29 sectors, 1970-85: 212,236 observations; dependent variable = log of imports+exports; beta values in brackets)

	Pooled OLS				***Tobit***	***Panel, fixed effects***		***Panel, random effects***	
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**	**(9)**
HOMEDUM	1.31 (0.09)	2.62 (0.18)	0.48 (0.03)	1.85 (0.12)	1.34 (0.09)	1.56 (0.10)	1.44 (0.10)	1.53 (0.10)	1.44 (0.10)
LOGGDP	0.71 (0.46)	0.71 (0.46)	0.71 (0.46)	0.68 (0.44)	0.70 (0.46)	0.77 (0.50)	0.79 (0.51)	0.77 (050)	0.79 (0.51)
LOGCGDP	0.07 (0.18)	0.07 (0.18)	0.07 (0.18)	0.11 (0.30)	0.06 (0.18)	0.12 (0.33)	0.13 (0.36)	0.13 (0.34)	0.13 (0.36)
LOGDIST	-1.20 (-0.51)			-1.50 (-0.64)	-1.17 (-0.50)	-0.91 (-0.39)	-0.93 (-0.39)	-0.90 (-0.39)	-0.93 (-0.39)
LOGDIST1		-1.20 (-0.47)							
LOGDIST2			-1.20 (-0.54)						
LOGREMOTE	0.02 (0.22)	0.02 (0.22)	0.02 (0.22)	0.04 (0.32)	0.02 (0.21)	-0.0002# (-0.01)	-0.001 (-0.05)	-0.0005# (-0.01)	-0.001 (-0.05)
BORDUM	-0.002# (-0.0002)	-0.002# (-0.0002)	-0.002# (-0.0002)	-0.02# (0.002)	0.03# (0.003)	0.23 (0.03)	0.22 (0.03)	0.26 (0.03)	0.22 (0.03)
LANGDUM	0.91 (0.10)	0.91 (0.10)	0.91 (0.10)		0.90 (0.10)	1.03 (0.12)	1.05 (0.12)	0.99 (0.12)	1.05 (0.12)
PTADUM	0.29 (0.04)	0.29 (0.04)	0.29 (0.04)		0.30 (0.04)	0.24 (0.03)	0.34 (0.05)	0.34 (0.04	0.34 (0.05)
CLOSEDDUM	-0.95 (-0.12)	-0.95 (-0.12)	-0.95 (-0.12)		-0.91 (-0.11)	-0.66 (-0.09)	-0.64 (-0.09)	-0.54 (-0.08)	-0.64 (-0.09)
Year dummies	No	No	No	No	No	Yes	Yes	Yes	Yes
Country dummies	No	No	No	No	No	Yes	No	(random ef.)	No
Sector dummies	No	No	No	No	No	No	Yes	No	(random ef.)
R^2	0.60	0.60	0.60	0.58	n.a.	0.61	0.61	0.61	0.61

Note: All coefficients pass the *t* test at the 0.01% level, except those marked by #.

TABLE 2: Estimated Home-Country Biases by Country and ISIC Industry, 1970-85

ISIC	AUS	AUT	BEL	CAN	DEN	FIN	FRA	GER	GRE	IRE	ITA	JAP	NOR	NTH	NZL	POR	SPA	SWE	TUR	UK	USA	YUG	*Avg*
311	0.83	2.61	1.84	6.47	2.65	5.37	4.94	2.11	9.11	1.76	1.19	17.39	3.80	1.55	-1.50	4.83	4.72	3.03	9.09	5.21	4.86	5.73	*4.44*
312	-8.97	n.a.	0.84	5.83	n.a.	3.58	3.37	2.13	n.a.	n.a.	0.04	21.48	3.38	2.45	n.a.	n.a.	n.a.	2.63	n.a.	2.85	3.69	n.a.	*3.33*
313	-8.40	3.54	3.33	6.24	6.23	3.47	4.79	1.86	9.85	-1.08	2.39	16.74	5.38	2.40	10.17	2.54	4.67	1.66	17.43	4.10	2.33	8.31	*4.91*
314	-27.60	3.52	-1.22	6.60	1.28	-0.09	2.96	-1.75	19.12	2.34	2.51	14.98	2.41	2.61	-78.78	6.35	7.61	2.69	n.a.	8.81	7.60	7.00	*-0.53*
321	-7.86	0.62	-0.20	5.48	0.43	2.15	2.73	1.19	8.33	-0.33	1.78	11.25	0.81	-2.10	-21.15	4.41	4.80	0.89	6.12	3.66	3.36	4.28	*1.39*
322	-4.24	-1.67	-3.19	6.58	-2.72	-0.84	1.68	-1.39	3.55	-0.95	-0.91	4.00	-0.50	-4.44	2.91	3.34	5.01	-0.87	7.31	0.83	3.87	5.58	*1.04*
323	-8.77	-2.52	-1.58	3.93	-1.91	0.59	2.40	-0.89	5.92	3.68	-0.23	19.78	0.12	-1.44	-0.52	9.05	4.48	0.16	7.59	0.62	1.19	5.01	*2.12*
324	-2.12	-0.02	-2.77	4.34	-0.60	0.68	1.50	-1.13	10.33	2.15	-0.02	-1.68	0.14	-3.89	21.62	6.69	6.25	-1.66	9.61	1.44	3.20	6.11	*2.74*
331	2.84	-3.41	0.96	5.33	-0.38	1.66	1.35	0.03	2.69	-0.05	-0.51	31.08	1.68	-0.51	-12.40	5.47	0.83	0.38	2.88	2.74	2.44	1.31	*2.11*
332	-0.60	-2.21	-1.15	4.55	-0.16	1.71	-0.07	-1.37	7.66	4.81	0.04	-5.04	1.57	-2.35	-8.39	6.64	3.30	-0.20	11.60	0.03	1.60	8.24	*1.37*
341	6.57	-1.86	-1.12	2.60	-0.28	3.86	1.20	0.41	3.00	-1.87	1.53	1.83	1.20	-0.75	-32.70	1.21	0.73	1.35	6.76	1.64	0.77	4.45	*0.02*
342	-10.32	1.77	0.88	4.44	1.61	3.64	3.52	0.60	8.29	4.90	2.93	14.00	2.72	2.77	-21.98	8.52	4.52	2.54	9.19	4.91	3.20	7.50	*2.73*
351	-9.97	-0.91	-2.00	2.75	-1.43	1.92	1.17	1.17	2.95	0.17	0.20	14.63	0.24	-0.54	-19.31	2.07	1.22	0.96	2.20	2.49	1.29	0.75	*0.09*
352	-12.56	1.84	0.89	4.88	-0.67	1.83	3.26	2.36	6.74	n.a.	1.28	20.62	1.76	0.02	-21.35	2.95	2.43	1.36	5.12	2.94	3.31	4.27	*1.58*
353	n.a.	4.53	-1.84	5.25	-1.94	3.61	1.07	0.15	4.39	-1.48	-1.72	27.92	0.89	-1.92	-65.81	3.45	-0.91	-1.79	4.65	0.21	2.81	-0.78	*-0.92*
354	-4.02	1.07	0.87	4.98	-1.29	0.38	-5.26	0.04	3.78	n.a.	-2.14	-11.13	1.48	-1.72	-50.45	n.a.	1.54	1.14	9.37	2.60	1.99	4.16	*-2.13*
355	-12.14	1.23	0.09	6.54	-0.89	0.61	1.12	0.16	4.58	2.80	0.79	-9.18	-0.65	-1.04	-19.28	5.36	3.22	0.32	6.31	2.11	1.48	3.32	*-0.14*
356	-5.45	0.73	1.00	5.07	-0.87	1.37	2.39	-0.40	7.07	2.24	1.43	11.67	1.42	0.54	-18.24	4.72	4.56	0.19	8.03	2.06	2.76	5.22	*1.71*
361	-17.27	0.26	2.17	7.87	0.59	1.90	2.19	-0.52	n.a.	4.27	0.75	1.53	1.93	-4.12	-18.90	7.14	5.61	0.05	6.11	2.79	3.10	0.88	*0.40*
362	-6.13	2.51	-0.51	4.65	-0.91	1.41	2.38	0.94	2.57	3.44	0.11	14.24	0.59	-0.89	-24.15	5.35	3.08	0.77	6.02	2.69	1.83	4.11	*1.10*
369	-11.44	-0.11	0.80	5.88	1.72	2.95	2.13	0.94	5.66	5.70	0.30	12.21	2.65	0.77	-24.26	5.36	3.75	1.43	5.65	3.65	3.41	0.89	*1.37*
371	-10.55	-1.49	-0.65	5.72	-5.05	0.57	0.04	0.10	7.20	0.50	0.27	7.10	-0.27	-2.41	-39.77	6.41	2.06	0.41	4.73	2.70	2.05	1.61	*-0.85*
372	-8.40	-0.60	n.a.	3.82	-2.62	2.11	0.92	0.73	6.45	-1.47	-0.14	6.56	-0.09	-1.70	-49.29	2.96	2.04	-0.01	6.99	1.26	0.15	3.29	*-1.29*
381	-5.75	0.55	0.40	5.23	-0.24	1.44	1.51	0.98	8.27	1.96	0.80	-0.56	1.06	1.48	-12.51	4.81	4.14	1.73	7.17	2.78	3.05	3.90	*1.46*
382	-7.51	-0.55	0.59	2.50	-0.46	1.31	1.23	0.05	4.30	1.19	0.53	-1.12	0.59	-0.40	-17.93	2.75	0.55	0.92	8.21	1.24	1.22	2.69	*0.09*
383	-10.91	-0.41	-1.23	4.51	-2.03	0.36	1.97	0.58	5.65	3.43	1.19	1.71	1.06	0.45	-30.70	5.02	1.42	0.92	6.90	2.88	2.04	3.91	*-0.06*
384	-15.20	-1.42	-2.38	2.80	-1.37	-0.26	1.21	0.34	6.04	4.57	0.94	4.62	-0.52	-0.72	-31.72	5.79	1.88	1.43	5.94	2.94	0.11	2.43	*-0.57*
385	-11.71	0.20	-0.83	3.26	-3.71	0.25	1.42	0.54	3.97	1.47	-0.50	6.12	-0.44	n.a.	-31.19	0.89	0.98	0.37	5.09	1.69	1.51	2.60	*-0.86*
390	-9.94	n.a.	n.a.	4.54	0.54	1.22	2.56	0.31	3.96	1.95	0.53	12.29	1.73	n.a.	-23.72	n.a.	2.90	0.47	7.28	2.05	1.80	4.30	*0.82*
Avg	*-8.46*	*0.20*	*-0.30*	*4.86*	*-0.63*	*1.55*	*1.67*	*0.29*	*6.24*	*1.77*	*0.51*	*8.84*	*1.16*	*-0.67*	*-23.70*	*4.77*	*3.06*	*0.72*	*7.09*	*2.53*	*2.40*	*3.90*	

TABLE 3: Industry-by-Industry Estimation of the Discriminating Criterion

(OLS with year fixed-effects, dependent variable = *OUTSHARE*; normalised (beta) coefficients, *t* values in brackets)

		(1)	(2)	(3)			(4)	(5)	(6)
ISIC	**Description**	***SHARE***	***IDIODEM***	***IDIOBIAS***	R^2	**No. of obs.**	**Para-digm**	**Output share in 1970 (%)**	**Output share in 1985 (%)**
311	Food products	0.88 (11.34)*	-0.16 (-1.77)	**-0.15 (-2.38)**	0.90	320	**CRS**	10.9	10.1
312	Food n.e.s.	0.82 (10.39)*	-0.06 (-0.79)	**0.22 (3.62)***	0.89	192	**IRS**	2.1	2.2
313	Beverages	0.99 (11.92)*	0.21 (6.41)*	**0.03 (0.52)**	0.86	317	**CRS**	2.2	1.9
314	Tobacco	0.56 (5.31)*	0.30 (5.71)*	**0.49 (5.36)***	0.93	288	**IRS**	1.2	0.9
321	Textiles	0.91 (10.37)*	0.10 (2.01)	**0.18 (5.13)***	0.93	320	**IRS**	5.3	3.3
322	Clothing	0.61 (9.03)*	0.17 (7.35)*	**0.34 (5.91)***	0.96	306	**IRS**	2.6	1.7
323	Leather goods	1.13 (9.35)*	0.43 (4.41)*	**0.15 (4.58)***	0.84	320	**IRS**	0.4	0.3
324	Leather footwear	0.95 (17.87)*	0.19 (4.67)*	**-0.03 (-0.84)**	0.83	311	**CRS**	0.6	0.4
331	Wood products	0.92 (10.19)	0.14 (1.90)	**0.10 (1.45)**	0.88	308	**CRS**	2.1	1.9
332	Furniture etc.	0.88 (7.77)*	0.02 (0.78)	**-0.12 (-3.68)***	0.87	320	**?**	1.3	1.2
341	Paper products	0.36 (5.35)*	0.41 (10.28)*	**0.41 (8.51)***	0.96	319	**IRS**	3.7	3.6
342	Printing and publishing	0.81 (9.51)*	0.23 (4.98)*	**0.10 (3.14)***	0.87	320	**IRS**	3.7	3.8
351	Industrial chemicals	0.84 (7.92)*	0.02 (0.66)	**0.12 (2.91)***	0.83	320	**IRS**	4.5	5.6
352	Other chemicals	0.83 (8.16)*	0.15 (7.14)*	**0.10 (2.53)**	0.87	304	**CRS**	3.5	3.6
353	Petroleum refining	0.40 (4.40)*	0.55 (5.86)*	**0.35 (5.01)***	0.88	320	**IRS**	3.4	9.6
354	Petroleum and coal products	0.89 (9.71)*	0.30 (9.73)*	**0.05 (0.77)**	0.82	288	**CRS**	0.4	0.6
355	Rubber products	0.93 (13.36)*	0.04 (1.01)	**-0.07 (-1.30)**	0.90	320	**CRS**	1.2	1.0
356	Plastic products	0.55 (4.42)*	-0.38 (-2.33)	**0.78 (4.05)***	0.88	320	**IRS**	1.5	2.2
361	Pottery, china etc.	0.82 (4.77)*	0.91 (43.53)*	**0.33 (2.60)**	0.78	304	**CRS**	0.2	0.2
362	Glass products	0.90 (10.60)*	0.04 (0.88)	**0.07 (2.02)**	0.89	320	**CRS**	0.7	0.7
369	Other non-metallic mineral pr.	0.46 (4.78)*	-0.16 (-2.43)	**0.58 (7.83)***	0.94	320	**IRS**	2.3	2.1
371	Iron and steel	0.60 (6.61)*	0.18 (5.02)*	**0.47 (7.15)***	0.95	320	**IRS**	6.4	4.9
372	Non-ferrous metals	0.82 (8.75)*	0.15 (3.36)*	**0.12 (2.79)***	0.90	304	**IRS**	2.8	2.3
381	Fabricated metal products	0.79 (10.63)*	0.24 (5.99)*	**0.09 (1.69)**	0.91	320	**CRS**	6.7	5.7
382	Non-electrical machinery	0.86 (9.93)*	0.09 (2.30)	**0.02 (0.28)**	0.86	320	**CRS**	9.5	8.9
383	Electrical machinery	0.20 (3.20)*	0.29 (15.86)*	**0.65 (10.93)***	0.96	320	**IRS**	7.3	7.4
384	Transport equipment	0.49 (5.10)*	0.37 (5.50)*	**0.38 (6.54)***	0.93	320	**IRS**	10.8	11.2
385	Instrument engineering	0.46 (6.21)*	0.37 (5.79)*	**0.21 (3.35)***	0.93	304	**IRS**	1.4	1.8
390	Misc. manufactures	0.73 (9.81)*	0.16 (2.82)*	**0.16 (5.19)***	0.92	255	**IRS**	1.1	1.0

Notes: * denotes statistical significance at the 99% confidence level. *t* statistics are calculated on the basis of MacKinnon-White (1985) heteroskedasticity consistent standard errors.