

A TEST OF TRADE THEORIES WHEN EXPENDITURE IS HOME BIASED

November 2001

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ABSTRACT

We develop and apply a criterion to distinguish two paradigms of international trade theory: constant-returns perfectly competitive models, and increasing-returns monopolistically competitive models. Our analysis makes use of the pervasive presence of home-biased expenditure. It predicts that countries' relative output and their relative home biases are positively correlated in increasing-returns sectors, while no such relationship exists in constant-returns sectors. We estimate country-level sectoral home biases through a gravity equation for international and intranational trade, and we use those estimates to implement our test on input-output data for six European Union economies.

JEL classification: F1, R3

Keywords: international specialisation, new trade theory, home-market effects, border effects

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Acknowledgements:

We thank David Greenaway, Thierry Mayer and Alan Winters for useful suggestions, and we gratefully acknowledge financial support from the Swiss National Science Foundation and from the British Economic and Social Research Council ("Evolving Macroeconomy" programme, ESRC grant L138251002).

1. 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, much of the theoretical and empirical debate has concentrated on these two benchmark cases.

To distinguish between 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 worthwhile to look for 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 criterion relies on the assumption that demand is home biased. It posits that the home bias influences international specialisation in sectors that are characterised by increasing returns and monopolistic competition (IRS-MC), while such bias is inconsequential for sectors characterised by constant returns and perfect competition (CRS-PC). We test this hypothesis in 18 industries, based on data for the six major EU economies for 1970-85. Our results suggest that five industries, accounting for about a quarter of industrial output, can be associated with the IRS-MC paradigm.

The paper is structured as follows. In Section II we briefly review 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.

2. Related Literature

Numerous studies have directly or indirectly attempted to gauge the relative explanatory power of the two main paradigms in trade theory.

A first 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 second 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 could also arise in a variety of other models (Davis and Weinstein, 2001a; Deardorff, 1998; Evenett and Keller, 2001; Feenstra, Markusen and Rose, 2001; Haveman and Hummels, 1997). Furthermore, it was found that the gravity equation is an excellent predictor of trade volumes among non-OECD economies, a piece of evidence that is

prima facie at odds with the assumptions of the IRS-MC paradigm (Hummels and Levinsohn, 1995).

A third approach was to derive a testable discriminating hypothesis from the models that could serve to distinguish among theoretical paradigms through rigorous statistical inference. Work along this line started with Davis and Weinstein (1999, 2001b). 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). The magnification effect does not appear in a CRS-PC model. 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 Japanese regions (1999) and for OECD countries (2001b), which allowed them to attribute industrial sectors to one of the two paradigms. This work has stimulated several further developments. Feenstra, Markusen and Rose (2001) found that the magnification effect may be generated also in CRS models with reciprocal dumping. Instead of the magnification effect they used 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 (2001) and Head, Mayer and Ries (2001) also showed that the magnification effect can arise in settings that do not necessarily conform with the IRS-MC paradigm. It has furthermore been found that the magnification effect can be sensitive to the modelling of trade costs (Davis, 1998). Head and Ries (2001) recognised this sensitivity to trade costs as a useful feature. They found that in CRS sectors the size of the magnification effect increases with trade costs whilst in IRS sectors it decreases with trade costs, and they used this difference as a discriminating criterion.

Our study follows this third approach in extracting a discriminating criterion from the theory to distinguish between the main trade models. We make use of 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 magnitude of home bias in expenditure. Our theoretical result is as 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 (2001a) 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 (2000) identified home bias in expenditure as one of the most potent sources of market fragmentation in Europe. Helliwell (1996, 1997), McCallum (1995) and Wei (1996) found that trade volumes among regions within countries are generally a multiple of trade volumes among

¹ Feenstra *et al.* (2001) and Head and Ries (2001) 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 (2000).

different countries even after controlling for geographical distance and other barriers. The assumption of home bias therefore seems to rest on solid empirical grounds.

Finally, our discriminating criterion remains valid even if trade costs are zero. This is an attractive feature, considering how difficult it is to quantify trade costs empirically.

3. 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).²

3.1 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.³ Good X is assumed to be subject to an

² 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.

³ The discriminating criterion that we develop does not hinge on this assumption; it is also valid if we assume positive trade costs in the CRS-PC sector. See Appendix 1.

IRS technology and to trade costs. These trade costs are of the conventional “iceberg” type, where for each unit shipped only a fraction $\tau \in [0,1]$ 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 $m(w,r)$. 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:

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

$$p_X(1 - 1/\sigma) = m(w,r), \quad (1b)$$

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

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

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

where σ is the elasticity of substitution among varieties of X ($\sigma > 1$). 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 , p_X , is the same across countries and the c.i.f. price is simply p_X / τ .

3.2 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 $h_i \in [0,1]$. This is a common way of parametrising the home bias (see, e.g. Miyagiwa, 1991; Trionfetti, 2001), but other structures are conceivable. One alternative representation would be through a parameter that is inserted inside the CES aggregate, as in Head and Ries (2001). 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:

$$U_i = X^{(1-h_{xi})\alpha_{xi}} X_i^{h_{xi}\alpha_{xi}} Y^{(1-h_{yi})\alpha_{yi}} Y_i^{h_{yi}\alpha_{yi}} Z^{(1-h_{zi})\alpha_{zi}} Z_i^{h_{zi}\alpha_{zi}}, \quad \text{with } \sum_S \alpha_{Si} = 1,$$

and with CES sub-utility

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

Denoting with E_{Si} the aggregate expenditure of households of country i on commodity S , we have $E_{Si} = \alpha_{Si} I_i$, 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 will serve as the basis for the calculation of magnification effects.

3.3 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:

$$p_X x = \frac{p_X^{1-\sigma}}{P_{X1}^{1-\sigma}} (1-h_{X1}) E_{X1} + \frac{\theta p_X^{1-\sigma}}{P_{X2}^{1-\sigma}} (1-h_{X2}) E_{X2} + \frac{1}{m} h_{X1} E_{X1} \quad (4)$$

$$p_X x = \frac{\theta p_X^{1-\sigma}}{P_{X1}^{1-\sigma}} (1-h_{X1}) E_{X1} + \frac{p_X^{1-\sigma}}{P_{X2}^{1-\sigma}} (1-h_{X2}) E_{X2} + \frac{1}{n_2} h_{X2} E_{X2} \quad (5)$$

$$E_{Y1} + E_{Y2} = p_Y (Y_1 + Y_2) \quad (6)$$

where $\theta = \tau^{\sigma-1}$, 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 . According to 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.

Note that for analytical convenience we have built a tree-by-two model instead of the more usual two-by-two structure. This allows us to work within the factor-price equalisation set.⁴

⁴ A two-by-two model would guarantee full dimensionality of the factor-price-equalisation set in the absence of trade costs (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

3.4 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. In Appendix 1 we show that this criterion remains valid even if we assumed trade costs in the homogeneous goods. 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 $\eta \equiv n_1/n_2$ and $\varepsilon \equiv E_{x_1}/E_{x_2}$. This gives:

$$\begin{aligned} & \theta h_2 \eta^3 + \left[(1-\theta)(1-h_2) + (1+\theta^2)h_2 - \theta(1-\theta)\varepsilon(1-h_1) - \theta\varepsilon h_1 \right] \eta^2 + \\ & + \left[\theta(1-\theta)(1-h_2) - (1-\theta)\varepsilon(1-h_1) + \theta h_2 - (1+\theta^2)\varepsilon h_1 \right] \eta + \theta\varepsilon h_1 = 0 \end{aligned} \quad (7)$$

It is well known that the roots of a third degree polynomial, especially when there are so many parametric 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 $d\alpha \equiv d\alpha_{x_1} = -d\alpha_{x_2}$ that generates the idiosyncratic expenditure shock $d\varepsilon \equiv dE_{x_1} = -dE_{x_2}$.

costs, it is necessary that there is one good more than there are factors in order to have full dimensionality of the factor price equalisation set. While working within the factor-price equalisation set is convenient, the results also hold outside it.

1. *The magnification effect when expenditure is unbiased.* 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: $\eta = (\varepsilon - \theta)/(1 - \theta\varepsilon)$. This is the benchmark case. It has the property that $d\eta/d\varepsilon > 1$, 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.
2. *The magnification effect when expenditure is home biased.* If demand is home biased, the derivative $d\eta/d\varepsilon$ is not necessarily larger than one for the IRS-MC sector (see Appendix 1). That implies that the magnification effect may fail to constitute a discrimination criterion. However, it is possible in this framework to derive an alternative discriminating criterion.
3. *The discriminating criterion when expenditure is home biased.* It is straightforward to show that $d\eta/dh|_{dh=dh_1=-dh_2>0} > 0$ 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 $(\theta n_2 / n_1)(n_1 + \theta n_2)^{-1} E_{X1} + \theta(\theta n_1 + n_2)^{-1} E_{X2} > 0$.

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), and 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 $\varepsilon = 1$ and $h_1=h_2=h$, then solving for η and differentiating around the solution we have:

$$d\eta/dh|_{dh=dh_1=-dh_2>0} = 4\theta(1 + \theta)/[(1 - \theta)^2 + 4\theta h] > 0,$$

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

4. *Robustness to zero trade costs.* The test based on home-biased demand gives a discriminating criterion that is valid even in the absence of trade costs. To see this it suffices to set $\theta=1$ in equation (7). This yields the unique solution $\eta = \varepsilon(h_1/h_2)$. In this extreme case, the home bias fully determines 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 the discriminating criterion from trade costs is a welcome feature. This is not because we think trade costs are unimportant in reality, but because it makes the discriminating criterion more robust.

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 $d\varepsilon \equiv dE_{x_1} = -dE_{x_2}$ and to a change $dh \equiv dh_1 = -dh_2$ gives:

$$d\eta = c_1 d\varepsilon + c_2 dh, \tag{8}$$

where the coefficients c_1 and c_2 are the partial derivatives. The interpretations of parameters c_1 and c_2 are summarised in Table 1.

Based on the fact that home-biased demand matters only for IRS-MC sectors, we can derive the following discriminating criterion. Sector S is associated with IRS-MC if the estimated c_2 is larger than zero, and with CRS-PC if the estimated c_2 equals zero. That is:

- if $c_2 > 0$ for sector S , then S is associated with IRC-MC, and
- if $c_2 = 0$ for sector S , then S is associated with CRS-PC.

This discriminating criterion and its empirical implementation are the focus of this paper. In addition, our model implies that idiosyncratic demand does not affect international specialisation if trade costs are zero. Accordingly, an estimated value of c_1 larger than zero reveals the importance of trade costs.

4. Empirical Implementation

In operationalising our discriminating criterion, we 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).

4.1 Estimating Home Bias

We estimated home separately for each country-industry pair, 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 our inference in stage two.⁵ We estimated variants of the following regression equation separately for six importing countries and 18 manufacturing sectors:

$$\begin{aligned} LOGIM_{ij,t} = & \alpha + \beta_1 HOMEDUM_{ij} + \beta_2 LOGGDPH_{i,t} + \beta_3 LOGGDPF_{j,t} + \\ & \beta_4 LOGPOPH_{i,t} + \beta_5 LOGPOPF_{j,t} + \beta_6 LOGDIST_{ij} + \beta_7 LOGREMOTE_{ij,t} + \\ & \beta_8 BORDUM_{ij} + \beta_9 LANGDUM_{ij} + \beta_{10} PTADUM_{ij,t} + \beta_{11} CLOSEDUM_{ij} + \\ & \beta_{12} NTB_{ij} + u_{ij,t} \end{aligned} \quad (9)$$

⁵ On the generality of the gravity equation, see Davis and Weinstein, 2001; Deardorff, 1998; Evenett and Keller, 2001; Feenstra *et al.*, 2001; Haveman and Hummels, 1997. The gravity model has been shown to be successful even

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

$LOGIM_{ij,t}$ = log of imports of country i from country j in year t ,

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

$LOGGDPH(F)$ = log of home-country (foreign-country) GDP,

$LOGPOPH(F)$ = log of home-country (foreign-country) population,

$LOGDIST$ = log of geographical distance between the two countries,

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

$BORDUM$ = 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,

$CLOSEDDUM$ = dummy which is 1 if at least one of the two countries was described as “closed” by Sachs and Warner (1995),

NTB = Harrigan (1993) measure of non-tariff import barriers by importer, exporter and industry in 1983.

The object of our interest is β_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

at the level of individual industries *inter alia* by Bergstrand, 1990; Davis and Weinstein, 2001; Feenstra *et al.*, 2000; Head and Mayer, 2000.

accessing foreign markets.⁶ 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.

Another potentially important issue concerns the degree of substitutability of goods contained within an industry. As argued, among others, by Deardorff (1998) and demonstrated by Evans (2000), border effects depend not only on home biases and trade costs, but also on the elasticity of substitution among an industry's products: border effects are higher if imports and domestic products are close substitutes in terms of their objective attributes, *ceteris paribus*. This issue is important for inter-industry comparisons. The main purpose of our home-bias estimates, however, is to allow for comparison across countries and over time, industry by industry, and hence our final exercise is unlikely to be affected by this concern.

The practical 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.⁷ 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 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

⁶ See Rauch (1999) on informational trade costs. Anderson and Marcouiller (2000) have shown that corruption and imperfect contract enforcement can act as additional impediments to international trade. It seems plausible 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.

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:

$$LOGDIST_{ii} = \log\left(\frac{1}{x} \sqrt{\frac{Area_i}{\pi}}\right). \quad (10)$$

This method may appear crude, but Head and Mayer (2000) 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. For most of our analysis, we set $x=3$. As we will discuss below, however, this arbitrariness is of no consequence for our study, since our model requires a measure of relative, not absolute, home biases.

Having constructed the intra-country variables, we estimated equation (9) on data for 18 industrial sectors, six importing countries (Belgium, France, Germany, Italy, Netherlands, UK), 22 exporting countries (OECD members) 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 close to 38,000 year- and industry-level bilateral observations. A full description of variables and data sources can be found in Appendix 2.

⁷ Hence, $LOGIM_{ii,t} = \log(Output-Exports)_{ii,t}$.

We began by running several variants of equation (9) on the entire data set. These results are given in Table 2. First, we pooled the data and estimated the gravity equation using OLS, excluding the *NTB* variable (column (1) of Table 2). All coefficients have the expected signs and magnitudes and are statistically significant. A coefficient of 0.64 on *HOMEDUM* suggests that *ceteris paribus* a country's trade with itself is on average 1.9 ($=e^{0.64}$) times as large as trade with another country. This estimate may seem high. However, it fits at the lower end of the range found by Wei (1996) for aggregate trade among a smaller OECD sample, which lie between 1.3 and 8.7, and it is smaller than the coefficient estimated by Helliwell (1997, Table 3) in his most comparable specifications.⁸

In a second step, we re-ran the full equation (9) including the *NTB* variable. In principle, this should be regarded as a valuable control variable if the coefficient on *HOMEDUM* is to pick up home bias rather than cost differences between imports and domestic produce. However, due to incomplete coverage of the *NTB* data, inclusion of this regressor reduced the sample size by more than 40 percent.⁹ With the exception of foreign GDP, all coefficients remained statistically significant and correctly signed. Somewhat surprisingly, the coefficient on *HOMEDUM* increased to 1.54. This result, however, is due entirely to the sample selection forced through the inclusion of *NTB*; when we estimated specification (1) on those observations for which *NTB* was

⁸ This figure might also seem at first sight to be incompatible with openness ratios (trade/GDP), which, in our data sample, 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 0.64 suggests that, on average, a representative agent facing two potential trade partners, one domestic and one foreign, will be 1.9 times more likely to trade with the domestic partner, even after controlling for distance and other cost factors, but it *does not* mean that the average openness ratio Ω should satisfy $\Omega=(1-\Omega)/1.9=0.34$.

⁹ The variable *CLOSEDDUM* had to be dropped in this specification, since the *NTB* variable is available for none of the countries for which *CLOSEDDUM*=1.

defined, the estimated coefficient on *HOMEDUM* was virtually identical. In other words, inclusion of the *NTB* variable limited the size of our data set but did not significantly affect the size of the estimated coefficient of interest. We therefore proceeded without including *NTB*.

Our next step was to replace OLS by the Tobit estimator, to take account of the censoring in our data set due observations with zero trade (column (3) of Table 2). *LOGIM* is recorded as zero in 1,574 observations (4%), whereof 86 are intra-country.¹⁰ This resulted in a slightly lower point estimate on *HOMEDUM* than OLS (0.54 vs. 0.64), which should not surprise, since zero observations account for a somewhat larger share of intra-country observations than of inter-country observations.

Wei (1996) has shown that the magnitude of the estimated home bias is sensitive to the way we proxy intra-national distances. We have therefore experimented with different definitions of $LOGDIST_{ii}$. *LOGDIST1* (column (4) of Table 2) 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*, for which we set $x=3$. As a consequence, the estimated home bias increases to a factor 6.1 ($=e^{1.81}$). 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 (5) of Table 2). With *LOGDIST2*, the estimated home-bias factor turns negative, to -1.3 ($=e^{-0.25}$). Obviously, one must 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*.

¹⁰ We set intra-country trade to zero where exports exceeded production.

Imposing identical coefficients across the three dimensions of our panel is restrictive. Our paper builds on the presumption, that home biases differ across countries and sectors.¹¹ Hence, our next step was to run equation (9) separately for each of the 108 country-industry observations (6 importing countries times 18 industries), so as to obtain individual home-bias estimates. We used the same specification as that of model (3) in Table 2. The resulting coefficients on *HOMEDUM*, averaged by country, are reported in Table 3. On average, we find the strongest home biases for France and Germany, while the mean home bias for Italy is negative. As Table 3 shows, however, the mean home bias estimates hide large variances. This is borne out by Table 4, where we report the same estimation results, but averaged by industry. The highest average home biases are found in the sectors tobacco, meat, and other food; whilst the lowest average home biases appear in the sectors motor vehicles, timber and furniture, and paper and printing. These results appear to be quite plausible. They are the key ingredient to our testing equation, which we estimate in the second part of our empirical exercise.

4.2 An Empirical Test of the Discriminating Criterion

We estimated the following variant of equation (8) for each of the 18 industries s across the six importing countries i, j and four years t (1970, 75, 80, 85):

$$Output_{it}^s = \alpha^s + \gamma_1^s SHARE_{it}^s + \gamma_2^s IDIODEM_{it}^s + \gamma_3^s IDIOBIAS_{it}^s + u_{it}^s, \quad (11)$$

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

¹¹ 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 and found that the panel effects were significant and had non-negligible effect on the estimated coefficient on *HOMEDUM*. These results are available from the authors on request.

$$SHARE_{it}^s = \frac{\sum_i Output_{it}^s}{\sum_i \sum_s Output_{it}^s} * \sum_s Output_{it}^s,$$

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

$$IDIOBIAS_{it}^s = (Bias_i^s - [\text{median}(Bias_i^s)]) * \sum_s Output_{it}^s.$$

According to our discriminating criterion, industries with estimated γ_3 of zero conform with the CRS-PC model, whereas industries with positive estimated γ_3 conform with the IRS-MC model. The other variables are constructed equivalently to the Davis-Weinstein testing equation. In our data set, *IDIOBIAS* and *IDIODEM* are not significantly correlated (see Appendix 2).

Four issues warrant discussion. First, there is the question of sectoral disaggregation. Neither theory nor existing empirical work give us any strong priors as to the correct definition of an “industry” and constituent “goods” in the data (see Maskus, 1991). 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, as it obviates the need to draw a dividing line between the two levels of sectoral aggregation.¹²

¹² Davis and Weinstein (2001b) 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 unimportant.

Second, there may exist potential for simultaneity of expenditure and output, and therefore bias in the estimates of γ_2 . The use of input-output tables allows us to obviate to this problem. One source of simultaneity could arise from the use of the same data source in the construction of expenditure and output variables. If expenditure is constructed by subtracting net exports from output, measurement error will simultaneously affect dependent and independent variables, and thus bias estimated coefficients. This is why we draw our expenditure and output data from input-output tables. As pointed out by Trionfetti (2001), input-output tables should be immune from the simultaneity problem through measurement error, because output and expenditure data (horizontal and vertical entries) are collected from independent sources of raw data. In addition, simultaneity bias could arise if some sectoral expenditure represents demand for intermediate inputs that are classified under the same sector heading. Our testing equation (8) implies the assumption that expenditure shares are an exogenous determinant of output location, but this assumption is rarely satisfied in the data. It is for this reason that we have chosen to compute net expenditure per sector as well as gross expenditure, which is possible in input-output tables. The definition of “net expenditure” should include expenditure from those sources that use the industry’s output for final consumption, and exclude expenditure from those sources that use the output as intermediate inputs. Net expenditure is therefore computed as expenditure on the industry’s output by private households and by the public sector, excluding expenditure by the industry itself and by all other manufacturing industries.¹³ We call *IDIODEM1* the variable computed from gross expenditure values and *IDIODEM2* the variable computed from net expenditure values.

¹³ Specifically, net expenditure is calculated as the sum of the following four expenditure headings in the input-output tables (NACE-Clio R44 codes in brackets): “final consumption of households on the economic territory” (01), “general public services” (810), non-market services of education and research (850), and non-market services of health (890).

Third, u_{it}^S is likely to be heteroskedastic, as the variance of errors may well be positively correlated with the size of countries.¹⁴ Our significance tests are therefore based on White-corrected standard errors. We make this conservative adjustment in order to minimise the risk of wrongly attributing sectors to the IRS-MC paradigm due to underestimation of the standard error of γ_3 .

Fourth, *IDIODEM* is a generated regressor, which could lead to bias in the coefficient estimates on it and on all other explanatory variables (Pagan, 1980). No unbiased or consistent estimator has as yet been derived analytically for the situation where an estimated coefficient of one equation enters as an explanatory variable in another. We therefore resort to bootstrap techniques. Resampling the data 5,000 times with replication, we re-estimated the coefficient vectors and standard errors for each model. The difference between the original regression coefficients and their bootstrap equivalents is a measure of estimation bias. We followed Efron's (1982) rule that bias is only a serious concern when the estimated bias is larger than 25 percent of the standard error. It turned out that the estimated biases were significantly below that threshold in all of the specifications that we estimated. Hence, we report the original coefficient estimates together with original as well as bootstrap estimated standard errors.¹⁵

¹⁴ A Cook-Weisberg heteroskedasticity test on the pooled model strongly rejects the null of constant error variance.

¹⁵ One might think that the average estimated bootstrap coefficient is superior to the original regression estimate. However, the bootstrap coefficient estimates have an indeterminate amount of random error and may thus have greater mean square error than the (potentially biased) original regression estimates (Mooney and Duval, 1993). We therefore report the original regression estimates.

4.3 Results

We first ran our model on the full data sample. The results are given in Table 5. In specification (1), we use OLS to estimate the specification based on gross expenditure values (*IDIODEM1*). This yields a coefficient on *IDIOBIAS* that is positive but not significantly different from zero, hence we would infer that on the whole the CRS-PC model dominates the IRS-MC model. In specification (2) we re-ran the same equation with *IDIODEM2*, which is based on net expenditure values. We find that the coefficient on *IDIODEM2* is significantly smaller than *IDIODEM1*, the two coefficients being 0.60 and 1.17 respectively. This confirms that the coefficient on demand idiosyncrasies is biased upward if demand is measured in gross terms. We also find that the estimated coefficient on *IDIOBIAS* is larger in the second specification and has become statistically significantly different from zero. This suggests that, on average, the IRS-MC model has stronger explanatory power than the CRS-PC model. Specifications (3) and (4), where we account for observed heteroskedasticity across sectors and countries by estimating a panel GLS model, confirm this result: the coefficient on *IDIOBIAS*, corresponding to c_2 of our testing equation (8) is always statistically significantly positive.

We know *a priori* that pooled runs impose too much structure, since our motivating hypothesis is to find different parameter estimates for individual industries. The results of industry-by-industry regressions are given in Table 6. The equation generally performs well, yielding R^2 s in the range 0.56 to 0.99. Coefficient estimates on *IDIOBIAS* are in the expected positive or insignificant range for all industries.

At the 95-percent confidence level we find that, of the 18 industries, five conform with the IRS-MC paradigm. The allocation of sectors looks broadly plausible.

It is interesting to measure the relative importance of the two paradigms in terms of their share of industrial output (Table 6, columns 5/6). According to our results based on the six largest EU economies, IRS-MC sectors account for around one quarter of industrial production. This share was increasing over our sample period. The combined output share of the five IRS sectors in our data set was 24.7 percent in 1970 and 26.7 percent in 1985.

A final comment is in order. Our estimates of the parameter γ_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, γ_2 may be smaller than one in both CRS-PC and IRS-MC sectors when demand is home biased. According to our model, γ_2 will take positive values, and will reflect important trade costs if it is significantly larger than zero. Our results are consistent with this prediction. The estimated coefficients on *IDIODEM* are never significantly negative, but they are significantly positive in ten of our 18 sectors. This is another indication that the empirical testing equation, which is tied closely to the underlying theoretical model, performs well in our data set.

5. Conclusions

We have developed and applied an empirical test to separate two paradigms of international trade theory: a model with constant returns and perfect competition (CRS-PC), and a model with increasing returns a monopolistic competition (IRS-MC). The discriminating criterion makes use of the assumption that demand is home biased, an assumption that is strongly supported in the empirical literature. We show theoretically that specialisation patterns are affected by inter-country differences in the degree of home bias if an industry conforms to the IRS-MC paradigm, but not if it is characterised by CRS-PC. This result provides us with a discriminating criterion that we implement empirically. In the empirical part we estimate industry- and country-level home biases through disaggregated gravity regressions, and use these estimates to apply our test to a data set with 18 industries for the six major EU countries in 1970-85. The results suggest that five industries, accounting for around a quarter of industrial output, can be associated with IRS-MC.

In addition to the assumption that demand is home biased, our study distinguishes itself from the literature on home-market effects in two principal ways. One feature is that the discriminating criterion holds regardless of whether trade costs exist, and regardless of whether they exist only in the IRS-MC sector or in both. This is a welcome feature especially in the light of the sensitivity, highlighted in the recent literature, of the magnification effect to the way trade costs appear in the model. Another innovation is that, in the empirical part, we draw on input-output data. This allows us to relate production to final expenditure, and hence to avoid simultaneity problems that might arise from regressing production on total expenditure where there are intra-sectoral input-output linkages.

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TABLE 1: Interpretation of Parameter Values in the Testing Equation

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

TABLE 2: Gravity Equations: Full Sample

(6 importing countries, 22 exporting countries, 18 sectors, 1970-85: dependent variable = log of imports; beta coefficients in brackets)

	<i>OLS</i>		<i>Tobit</i>		
	(1)	(2)	(3)	(4)	(5)
<i>HOMEDUM</i>	0.64 (0.04)	1.54 (0.13)	0.54 (0.04)	1.81 (0.12)	-0.25 (-0.02)
<i>LOGGDPH</i>	1.33 (0.16)	2.78 (0.35)	1.33 (0.18)	1.33 (0.18)	1.33 (0.18)
<i>LOGGDPF</i>	1.63 (0.26)	-0.09 [#] (-0.01)	1.63 (0.30)	1.63 (0.30)	1.63 (0.30)
<i>LOGPOPH</i>	0.68 (0.14)	0.72 (0.17)	0.70 (0.17)	0.70 (0.17)	0.70 (0.17)
<i>LOGPOPF</i>	0.56 (0.17)	0.28 (0.10)	0.57 (0.20)	0.57 (0.20)	0.57 (0.20)
<i>LOGDIST</i>	-1.14 (-0.43)	-0.69 (-0.33)	-1.15 (-0.49)		
<i>LOGDIST1</i>				-1.15 (-0.45)	
<i>LOGDIST2</i>					-1.15 (-0.52)
<i>LOGREMOTE</i>	0.04 (0.25)	0.04 (0.27)	0.04 (0.27)	0.04 (0.27)	0.04 (0.27)
<i>BORDUM</i>	0.32 (0.04)	0.38 (0.05)	0.31 (0.04)	0.31 (0.04)	0.31 (0.04)
<i>LANGDUM</i>	0.93 (0.09)	0.62 (0.07)	0.97 (0.10)	0.97 (0.10)	0.97 (0.10)
<i>PTADUM</i>	0.72 (0.10)	0.99 (0.15)	0.72 (0.11)	0.72 (0.11)	0.72 (0.11)
<i>CLOSEDDUM</i>	-1.29 (-0.12)		-1.39 (-0.15)	-1.39 (-0.15)	-1.39 (-0.15)
<i>NTB</i>		-2.53 (-0.21)			
No. of observations	37,968	22,176	37,968	37,968	37,968
No. of censored obs.	n.a.	n.a.	1,526	1,526	1,526
R ²	0.44	0.42	n.a.	n.a.	n.a.

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

TABLE 3: Estimated Home-Country Biases by Country, 1970-85

Country	Mean coefficient on <i>HOMEDUM</i>	Standard deviation of coefficients on <i>HOMEDUM</i>
Belgium-Luxembourg	0.50	2.24
France	4.34	2.82
West Germany	3.29	1.45
Italy	-0.79	1.16
Netherlands	1.11	2.90
United Kingdom	2.84	1.22

Note: Based on Tobit regression (specification (3) of Table 2) estimated separately for each of the 6 importing countries and 18 sample industries.

TABLE 4: Estimated Home-Country Biases by NACE Industry, 1970-85

NACE code: Industry	Mean coefficient on <i>HOMEDUM</i>	Standard deviation of coefficients on <i>HOMEDUM</i>
1170: Chemicals	2.21	2.22
1190: Metal goods	1.15	1.39
1210: Machinery	1.69	1.36
1230: Office machines	0.71	4.22
1250: Electrical goods	1.90	2.14
1270: Motor vehicles	-0.21	3.59
1290: Other transp. eq.	1.94	1.54
1310: Meat products	3.50	3.35
1330: Dairy products	2.70	2.43
1350: Other food	3.45	1.92
1370: Beverages	2.26	2.97
1390: Tobacco products	4.76	4.70
1410: Textiles, clothing	2.18	3.15
1430: Leather, footwear	1.84	2.64
1450: Timber, furniture	0.39	1.60
1470: Pulp, paper, printing	0.41	1.25
1490: Rubber, plastic	1.63	1.99
1510: Instrum. engineering and other manuf.	2.29	2.28

Note: Based on Tobit regression (specification (3) of Table 2) estimated separately for each of the 6 importing countries and 18 sample industries.

TABLE 5: Pooled Estimation of the Discriminating Criterion

(Dependent variable = *Output*; 400 observations)

	<i>OLS</i>		<i>GLS (panel heteroskedasticity)</i>	
	(1)	(2)	(3)	(4)
<i>SHARE</i>	1.0066 (0.015, 0.00) [0.015, 0.00]	1.0086 (0.025, 0.00) [0.026, 0.00]	1.0087 (0.011, 0.00) [0.021, 0.00]	0.9921 (0.011, 0.00) [0.016, 0.00]
<i>IDIODEMI</i>	1.1745 (0.045, 0.00) [0.049, 0.00]			
<i>IDIODEM2</i>		0.604 (0.058, 0.00) [0.058, 0.00]	0.543 (0.024, 0.00) [0.038, 0.00]	0.512 (0.036, 0.00) [0.039, 0.00]
<i>IDIODIAS</i>	0.0002 (0.00017, 0.19) [0.00017, 0.19]	0.0005 (0.00020, 0.01) [0.00021, 0.01]	0.0005 (0.00011, 0.00) [0.00010, 0.00]	0.0006 (0.00024, 0.01) [0.00017, 0.00]
No. of observations	400	400	400	400
R ²	0.98	0.94	n.a.	n.a.
Panels (no.)	None	None	Sectors (18)	Countries (6)

Notes: *Round brackets:* estimated standard errors (White corrected for OLS models) and P values. *Square brackets:* Bootstrapped standard errors (5000 replications with replacement) and P values. *Constant term* included in all regressions but not reported (never statistically significantly different from zero).

TABLE 6: Industry-by-Industry Estimation of the Discriminating Criterion(OLS, dependent variable = *Output*; bootstrapped standard errors brackets)

	(1)	(2)	(3)		(4)	(5)	(6)
NACE	<i>SHARE</i>	<i>IDIODEM</i>	<i>IDIODIAS</i>	R ²	Para- digm	Output share in 1970 (%)	Output share in 1985 (%)
170	0.95 (0.06)*	-0.15 (0.30)	0.0019 (0.0017)	0.98	CRS	10.0	12.9
190	1.18 (0.06)*	1.29 (0.63)	-0.0021 (0.0062)	0.98	CRS	9.8	8.4
210	1.29 (0.08)*	1.63 (1.03)	-0.0021 (0.0062)	0.95	CRS	9.7	9.5
230	1.05 (0.06)*	0.61 (0.09)*	0.0005 (0.0002)*	0.99	IRS	2.0	2.7
250	1.16 (0.10)*	0.74 (0.53)	0.0005 (0.0034)	0.97	CRS	8.4	8.8
270	1.01 (0.08)*	1.03 (0.22)*	0.0037 (0.0013)*	0.98	IRS	7.8	8.4
290	0.80 (0.15)*	0.68 (0.22)*	0.0015 (0.0035)	0.82	CRS	2.5	3.1
310	0.71 (0.05)*	0.42 (0.04)*	0.0014 (0.0005)*	0.98	IRS	5.3	4.9
330	0.80 (0.05)*	0.42 (0.07)*	0.0021 (0.0004)*	0.97	IRS	2.8	3.0
350	0.79 (0.04)*	0.34 (0.26)	-0.0024 (0.0017)	0.97	CRS	9.4	8.9
370	0.97 (0.07)*	0.35 (0.07)*	0.00003 (0.00006)	0.94	CRS	3.3	2.4
390	0.97 (0.04)*	0.39 (0.04)*	0.0001 (0.0002)	0.98	CRS	2.4	1.2
410	0.63 (0.19)*	1.21 (0.56)	0.0066 (0.0063)	0.82	CRS	9.8	7.4
430	0.74 (0.20)*	1.75 (0.37)*	0.0002 (0.0010)	0.85	CRS	1.6	1.7
450	1.06 (0.08)*	0.35 (0.19)	-0.0043 (0.0018)	0.95	CRS	4.3	3.7
470	0.84 (0.06)*	0.27 (0.21)	0.0070 (0.0028)*	0.98	IRS	6.8	7.7
490	1.12 (0.03)*	0.58 (0.23)*	-0.0001 (0.0003)	0.99	CRS	3.3	3.9
510	0.56 (0.26)*	0.39 (0.15)*	-0.0004 (0.0014)	0.56	CRS	1.0	1.4

Notes: * denotes statistical significance at the 95% confidence level.

Appendix 1: Theoretical Extensions

A1.1 No Magnification Effect in IRS-MC Sectors When Demand is Home Biased

We can show that the derivative $d\eta/d\varepsilon$ is not necessarily larger than one when demand is home biased. Let us define the left-hand side of (7) as a $P(\eta, \varepsilon, h_1, h_2, \theta)$. Unfortunately, in general, the roots of this polynomial are long expression that are not manageable (even for the computer) except in the following two cases.

Case A: Identical Countries

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

$$\frac{d\eta}{d\varepsilon}(h, \theta) = -\frac{\partial P(\cdot)/\partial \varepsilon}{\partial P(\cdot)/\partial \eta} = \frac{(1-\theta^2)(1-h) + (1+\theta)^2 h}{(1-\theta)^2(1-h) + (1+\theta)^2 h} > 1$$

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

$$\frac{d\eta}{d\varepsilon}(\theta) = -\frac{\partial P(\cdot)/\partial \varepsilon}{\partial P(\cdot)/\partial \eta} = \frac{(1+\theta)}{(1-\theta)} > 1$$

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 $P(\eta, \varepsilon, h_1, h_2, \theta)$ reduces to a second-degree polynomial. Differentiating around its only positive root gives:

$$\frac{d\eta}{d\varepsilon}(\theta, \varepsilon, h_2) = \frac{(1-\theta)[(\varepsilon+h_2)^2\theta^2 + 2(2\varepsilon h_2 + h_2 - \varepsilon)\theta + 1]^{\frac{1}{2}} + (\varepsilon+h_2)\theta^2 + (h_2 - \varepsilon - 1)\theta + 1 - 2h_2}{2h_2[(\varepsilon+h_2)^2\theta^2 + 2(2\varepsilon h_2 + h_2 - \varepsilon)\theta + 1]^{\frac{1}{2}}}$$

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 $\frac{d\eta}{d\varepsilon}(\theta, \varepsilon, h_2)$ is compared with the constant 1. The figure

(computed at $\varepsilon = 1$) shows a large domain in which $\frac{d\eta}{d\varepsilon}(\theta, 1, h_2) < 1$, i.e., no magnification effect exists. Many other combinations of parameter values also yield no magnification effect.

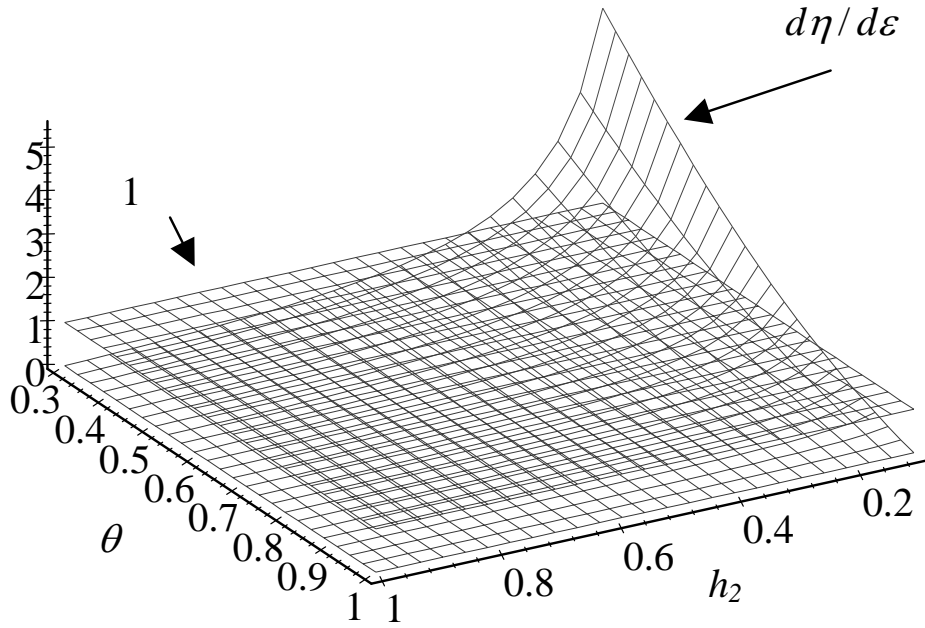


Figure A1.1: An Example of the Domain with No Magnification Effect

A1.2 Home Bias in the CES Aggregate

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

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

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

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

and the derivative is

$$\frac{d\eta}{d\varepsilon}(\theta, \varepsilon, h_1, h_2) = (h_1 - \theta) \frac{h_2^2 h_1 - h_2 h_1 \theta - \theta^2 h_2 + \theta^3}{(\theta^2 \varepsilon + h_2 h_1 - h_1 \theta \varepsilon - h_1 \theta)^2},$$

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

A1.3. Trade Costs in the CRS-PC Sector

In this section we show that home-biased demand is irrelevant for international specialisation in the CRS-PC sector even if these goods are traded at a transport cost. Introducing trade costs in the homogeneous goods results in the non-equalisation of factor prices. The major consequence for our paper is that we cannot derive equation (7) and the expression for $d\eta/dh$ as neatly as we do under factor price equalisation. This notwithstanding, we can show the irrelevance of the home bias in CRS-PC sectors by inspection of the market equilibrium equations. For clarity of exposition we consider the situation where, say, country 1 is a net exporter of Y . Naturally, the choice of the country is irrelevant. Assume that for each unit of Y shipped, only a fraction $\vartheta \in (0,1)$ arrives at destination. Then, equation (6) becomes

$$p_{Y1}[(1-h_{Y1})E_{Y1} + h_{Y1}E_{Y1}] + \vartheta p_{Y2}[(1-h_{Y2})E_{Y2} + h_{Y2}E_{Y2}] = p_{Y1}Y_1 + \vartheta p_{Y2}Y_2$$

Once again, h_{Yi} cancels out of this expression. This means that the home bias is irrelevant for international specialisation in the CRS-PC sector and, therefore, that the assumption of free trade in Y and Z made in Section 3 is purely a matter of analytical convenience.

Appendix 2: Data Description

Output data are taken from input-output tables in Eurostat’s “National Accounts ESA” series. Bilateral trade data are taken from the World Trade Database, available through Feenstra *et al.* (1997). We retained trade data recorded by the importing countries. These data, original classified under SITC headings, were concorded with the NACE input-output data used for the output values. Hence, all trade flows are c.i.f., and our estimates of trade within countries can be considered conservative. Observations for which estimated intra-country trade was, implausibly, negative (i.e. Output-Exports<0) were set to zero.

The following 22 exporting 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 18 NACE industries in our data set can be found in Table 6. Distance data measure great circle distances between the countries’ capital cities. They are from Jon Haveman (www.macalester.edu/research/economics/PAGE/HAVEMAN/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):

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

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 4, 9816 obs.):

	<i>Output</i>	<i>SHARE</i>	<i>IDIODEM</i>	<i>IDIODIAS</i>
<i>Output</i>	1.000			
<i>SHARE</i>	0.624	1.000		
<i>IDIODEM</i>	0.219	-0.002	1.000	
<i>IDIODIAS</i>	0.287	0.364	-0.049	1.000