

A TEST OF TRADE THEORIES WHEN EXPENDITURE IS HOME BIASED^{*}

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Abstract

We develop a criterion to distinguish two dominant 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 (the "home-bias effect"), while no such relationship exists in constant-returns sectors. This discriminating criterion turns out to be robust to a number of generalisations of the baseline model. Our empirical results suggest that the increasing-returns model fits particularly well for the mechanical and electrical engineering industries, which account for close to half of manufacturing output.

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1. Introduction

International trade theory is dominated by two major paradigms. One paradigm belongs to the neoclassical 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 frequently employed formulation, monopolistically competitive markets (MC). While other important models exist which combine features of both paradigms, much of the theoretical and empirical literature 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 model. It is therefore worthwhile to look for ways of separating the two theoretical approaches in the data, and to attempt to quantify their respective relevance for observed industrial specialisation patterns. This is the purpose of our study.

In the theoretical part, we develop a discriminating criterion that is amenable to empirical estimation. The criterion rests 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 the location of sectors characterised by constant returns and perfect competition (CRS-PC). We find this discriminating criterion to be robust to a number of generalisations of the baseline model, including imperfectly elastic sectoral factor supplies and multiple non-equidistant countries.

We test the discriminating hypothesis across 17 industries, based on a cross-country dataset for 1997. By combining production data with trade data, we can compute internal trade volumes and thereby estimate country-sector level home biases via a generalised gravity specification. By

matching trade and production data with input-output tables, we can compute final expenditure values, which the theory prescribes as another ingredient to the testing equation. Our results suggest that the IRS-MC model fits particularly well for the engineering industries (fabricated metal products, non-electrical machinery, electrical machinery and precision engineering, and transport equipment), which account for close to half of manufacturing output in our sample.

The paper is structured as follows. In Section 2 we review the relevant literature. Section 3 sets out our theoretical model and derives the discriminatory criterion. Section 4 discusses the robustness of that criterion. We operationalise the theoretical criterion empirically in Section 5. Section 6 concludes.

2. Related Literature

Numerous studies have directly or indirectly attempted to gauge the relative explanatory power of the 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. The theoretical relevance of this evidence became uncertain when some studies, such as 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. It has indeed been shown that the gravity equation has a straightforward theoretical counterpart in the IRS-MC model (Helpman, 1987). However, gravity-type predictions have also been derived from a variety of other models (Davis and Weinstein, 2001; Deardorff, 1998; Eaton and Kortum, 2002; Evenett and Keller, 2002; 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 Hummels and Levinsohn (1995) plausibly interpret as being at odds with IRS-MC paradigm.

A third approach was to derive a testable discriminating hypothesis from the theory that can serve to distinguish among theoretical paradigms through statistical inference. Work along this line started with Davis and Weinstein (1996, 1999, 2003). 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 “home-market effect” (HME, first identified by Krugman, 1980). Since the HME does not appear in a CRS-PC model, this feature can serve as the basis for discriminating empirically between paradigms. Davis and Weinstein have estimated the HME in data for Japanese regions (1999) and for OECD countries (1996, 2003), which allowed them to associate industrial sectors with one of the two paradigms.

The work of Davis and Weinstein has stimulated a lively research programme. Head and Ries (2001) have exploited the sensitivity of the HME to trade costs for an alternative discriminating hypothesis: in CRS sectors (with product differentiation by country of origin) the HME is amplified by trade costs, whilst in IRS sectors it decreases with trade costs. They estimated this

prediction in a panel of 3-digit Canadian and U.S. industry data covering the period 1990-1995. Alternatively using cross-sectional and time series variation in the data, they computed the slope of the line relating a country's share of output in an industry to its share of expenditure in that industry. Their sample period included a tariff reduction (NAFTA) that allowed them to relate the slope to the changes in trade costs (after controlling for other factors). They found evidence in support of both models depending on whether parameter identification comes from the cross section or from the time series, but the CRS model with product differentiation by country of origin seems to be supported more strongly.

Some researchers have classified sectors according to extraneous information on their characteristics, and tested whether those classifications map into different structural relationships predicted by the theory. Feenstra, Markusen and Rose (2001) have estimated countries' income elasticity of exports in a gravity model. According to the HME (which they derived in a variety of theoretical settings), this elasticity should be larger for differentiated goods than for homogeneous goods. Their results strongly support this hypothesis. Hanson and Xiang (2004) have employed a difference-in-difference gravity specification in order to allay concerns about endogeneity bias or specification bias. Their version of the HME is that larger countries tend to export relatively more of high-transport-cost, strong-scale-economies goods and relatively less of low-transport-cost, weak-scale-economies goods. They tested this prediction on country pairs' exports to third markets and found evidence of HMEs in high transport-cost, strong-scale-economies industries, as predicted by the theory. Weder (2003) has formulated the HME in terms of relative exports: a country tends to export more of the goods for which it has a larger home market, and the strength of this relationship increases in the importance of scale economies. His empirical findings, based on US-UK trade, support the theoretical predictions:

HMEs become stronger the larger are an industry's economies of scale, measured by average firm size.

Recent work has shown that the association between HMEs and the imperfectly competitive model with differentiated goods is neither necessary nor exclusive once one departs from the benchmark variant of the model. Three issues have been identified that limit the generality of the HME as a discriminatory criterion. First, as demonstrated by Davis (1998), the existence of HMEs relies on trade costs in the CRS-PC sector being sufficiently smaller than those of the IRS-MC sector. Second, Head and Mayer (2004) have shown that the HME may fail as a discriminating criterion when the elasticity of factor supplies across sectors and/or countries is finite. In fact, the model that led to the derivation of the HME (Krugman, 1980) implies that countries' intersectoral transformation curves are linear - once one allows for sufficiently imperfect factor substitutability across sectors, the HME will vanish even if the world otherwise conforms with the IRS-MC model. Third, Behrens, Lamorgese, Ottaviano and Tabuchi (2004) have taken the study of HMEs from the standard two-country model to a setting with multiple non-equidistant countries. They showed that, depending on the distribution of expenditure among neighbouring countries, IRS-MC sectors may or may not exhibit HMEs.

These issues notwithstanding, the role of expenditure as a determinant of international specialisation, and in particular the stark relationship implied by the HME, of course remains an important dimension of trade theory and empirics, as well as of policy-related research. In view of the significant challenges to the HME *as a discriminatory criterion*, however, we seek a robust and empirically implementable alternative feature of the theory that can serve to discriminate among alternative trade models.

Our approach, building on Trionfetti (2001a), is based on the widely documented reality that buyers consider imports and domestically produced goods *ipso facto* as imperfect substitutes, and that they 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. Specifically, in an IRS-MC setting, relatively strong home bias in a country's aggregate expenditure will make that country relatively specialised in the production of the good concerned (the "home-bias effect"), 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 test.

Our whole study hinges on the pervasive existence of home-biased demand. We believe that this is a sensible premise, given the strong empirical evidence in its support. For example, Winters (1984) has 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 (2001) and Trefler (1995) find that by allowing for home-biased demand the predictive power of the HOV model can be improved very significantly. Head and Mayer (2000) identify home bias in expenditure as one of the most potent sources of market fragmentation in Europe. Anderson and van Wincoop (2003), McCallum (1995) and Wei (1996) find that trade volumes among regions within countries significantly exceed trade volumes among different countries even after controlling for geographical distance and other barriers. The assumption of home bias therefore rests on solid empirical ground.

¹ This paper extends the work of Trionfetti (2001a) in five principal ways. We formally spell out the implications of home bias on the HME as well as on international specialisation, we show that the discriminatory criterion based on the home-bias effect is robust to a number of critical extensions of the benchmark (Helpman-Krugman) model, we estimate sector-country level home biases directly using a generalised gravity specification, we take account of the N -country model in the empirics, and we base estimation on a world-wide cross-country data set.

Finally, it might be useful at this point to clarify the principal difference between our approach and that chosen by Head and Ries (2002). Their analysis pits a model of product differentiation by firm and firm-level IRS (implying MC) against a model of product differentiation by country of origin with firm-level CRS (implying PC). The latter is referred to as the “national product differentiation” model. We instead assume that differentiation by country of origin (the home bias) is present in *all* sectors (to varying extents), and we test a model of product differentiation by firm (IRS-MC) against a model of no product differentiation by firm (CRS-PC). We consider the support found by Head and Ries (2002) for the differentiation-by-country-of-origin model as further confirmation that our consideration of home bias as a pervasive phenomenon is empirically well founded.

3. Theory: Derivation of a Discriminating Criterion

A suitable model 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), where the world is composed of two countries, labelled with superscripts $i \in \{A, B\}$, and each country is endowed with an exogenous quantity L_i of labour which is employed to produce two commodities indexed by $S \in \{X, Y\}$.

3.1 *Technologies and Factor Markets*

We assume that the homogeneous good Y is produced by use of a CRS technology and traded costlessly in perfectly competitive markets.² Given that technologies are identical across countries and that there is perfect inter-sectoral mobility of labour, free trade in Y yields factor

² The discriminating criterion that we develop remains valid if we assume positive trade costs in the CRS-PC sector (see Section 4.1).

price equalisation.³ Varieties of good X are subject to IRS 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 the CRS sector is $w\ell_Y$, where w is the wage and ℓ_Y is the input requirement per unit of output. Production of X entails a fixed cost wF and a constant marginal cost wm , where F is the fixed labour input and m is labour input per unit of output. Therefore, average cost in the X sector is $wm + wF/x$, where x is output per firm. Firms are identical and face identical demand functions, and hence, the optimal price and output levels are identical for all firms.

The conditions for efficiency and factor-market clearing are:

$$p_Y = w\ell_Y, \quad (1a)$$

$$p_X(1 - 1/\sigma) = wm, \quad (1b)$$

$$m(1 - 1/\sigma)^{-1} = m + F/x, \quad (2)$$

$$\ell_Y Y^A + (F + mx)n^A = L^A, \quad (3)$$

$$\ell_Y Y^B + (F + mx)n^B = L^B, \quad (4)$$

where $\sigma > 1$ is the elasticity of substitution among varieties of X . Equations (1a) and (1b) state the usual conditions that marginal revenue equals marginal cost in both sectors and countries. Equation (1a) also represents the zero-profit condition in sector Y . The zero-profit condition in sector X in all countries is in turn represented by equation (2). Since the mark-up is constant, this equation determines the level of output per firm independently of prices and wages. Equations (3) and (4) state the market-clearing conditions for L in both countries, where n^i and Y^i are,

³ Another necessary condition for factor price equalisation is the “no-corner solution” assumption we introduce in Section 3.3.

respectively, the number of varieties of X and the output of Y produced in country i . These equations describe the supply side of the model. Free trade assures commodity price equalisation in the Y sector. In the Dixit-Stiglitz framework, trade costs are borne entirely by buyers. Therefore, the equilibrium price of imported varieties of X is simply p_X/τ .

3.2 Demand

Preferences feature love for variety, represented by the traditional nested CES-Cobb-Douglas utility function. We extend the basic model by assuming that demand is home biased. We follow the related trade literature in assuming that the home bias is exogenous, because we too are interested in studying the consequences of home bias and not its causes.⁴ Furthermore, we assume that there are two types $t \in \{u, b\}$ of buyers (but one could easily extend the model to a continuum of them): the “unbiased” type (u) and the “home-biased” type (b). We could for instance think of these two types as private-sector and public-sector purchasers, knowing that public procurement is typically characterised by particularly strong home bias (see, e.g., Trionfetti, 2001b).⁵

We model the home bias parametrically at the Cobb-Douglas level of the utility function, and represent it by the parameter $\delta_S^{it} \in [0, 1]$, with $\delta_S^{iu} < \delta_S^{ib}$. When $\delta_S^{it} = 0$, buyers are not home biased in sector S . As δ_S^{it} increases, buyers become increasingly home biased, and when $\delta_S^{it} = 1$ type- t buyers purchase sector- S goods solely from domestic producers. Parameterisation at the Cobb-Douglas level is a common way of introducing the home bias, but other ways are perfectly conceivable. One alternative representation would be through a parameter inserted inside the

⁴ See, e.g., Baldwin (1984), Miyagiwa (1991), Trefler (1995), Hummels (2001), and Head and Ries (2001).

⁵ None of our findings hinge on this segmentation of the pool of buyers. All of our results would go through if, for example, we assumed instead that all buyers share identical “schizophrenic” preferences whereby they reserve some share of expenditure for home-produced goods, while the remaining expenditure is allocated without country bias.

CES aggregator, as in Head and Ries (2001), Hummels (2001) and Combes, Lafourcade and Mayer (2005). We use this alternative in Appendix 1 and show that the salient results of our model remain unchanged.

With these assumptions, the utility function of buyer type t in country i is:

$$U^{it} = X^{(1-\delta_X^t)\alpha_X^i} X_i^{\delta_X^t\alpha_X^i} Y^{(1-\delta_Y^t)\alpha_Y^i} Y_i^{\delta_Y^t\alpha_Y^i},$$

$$\text{with CES sub-utility } X = \left(\int_{k \in n^i} c_k^{(\sigma-1)/\sigma} dk + \int_{k \in n^j} c_k^{(\sigma-1)/\sigma} dk \right)^{\sigma/(\sigma-1)},$$

and with $\alpha_X^i + \alpha_Y^i = 1$. For simplicity, we assume that $\alpha_S^{iu} = \alpha_S^{ib}$ for all sectors and countries, so that the t superscript does not appear in the expenditure shares. To simplify notation further, we assume that $\delta_S^{iu} = 0$ and $\delta_S^{ib} \in (0,1]$ for all sectors and countries. This allows us to suppress the t superscript of the δ parameter.

In each country and sector, a certain share μ_S^i of the population is of type b , while the remaining share $(1 - \mu_S^i)$ is of type u . Denoting aggregate expenditure by buyers of type t in country i on sector S with E_S^{it} , we have that $E_S^{ib} = \alpha_S^i \mu_S^i wL^i$ and $E_S^{iu} = \alpha_S^i (1 - \mu_S^i) wL^i$. Furthermore, denoting aggregate expenditure by buyers of all types in country i on sector S with E_S^i , we have that

$$E_S^i = E_S^{iu} + E_S^{ib} = \alpha_S^i wL^i.$$

3.3 Equilibrium in Product Markets

Product-market equilibrium requires that demand equals supply for each sector and each variety.

In writing the market equilibrium equations, it turns out that, after simplifying through for

$E_S^{iu} + (1 - \delta_S^i)E_S^{ib}$, μ_S^i and δ_S^i always appear multiplicatively. We can thus define the convenient

parameter $h_S^i \equiv \mu_S^i \delta_S^i$, aggregating the two components of home bias. We refer to this parameter simply as “the home bias” of country i in sector S . Using this notation, the world market-clearing conditions are:

$$p_X x = \frac{P_X^{1-\sigma}}{(P_X^A)^{1-\sigma}} (1 - h_X^A) E_X^A + \frac{\phi P_X^{1-\sigma}}{(P_X^B)^{1-\sigma}} (1 - h_X^B) E_X^B + \frac{1}{n^A} h_X^A E_X^A, \quad \text{and} \quad (5)$$

$$p_X x = \frac{\phi P_X^{1-\sigma}}{(P_X^A)^{1-\sigma}} (1 - h_X^A) E_X^A + \frac{P_X^{1-\sigma}}{(P_X^B)^{1-\sigma}} (1 - h_X^B) E_X^B + \frac{1}{n^B} h_X^B E_X^B, \quad (6)$$

where $\phi \equiv \tau^{\sigma-1}$ is the “phi-ness” representation of trade openness (Baldwin *et al.*, 2003), and P_X^i is the usual CES price index. Equation (5) states the equilibrium condition for any IRS-MC variety produced in country A , and equation (6) states the equilibrium condition for any variety produced in country B .

To these two equations we have to add the following “no-corner-solution” conditions: $p_X \tilde{n}^i x > h_X^i E_X^i$ and $p_Y \tilde{Y}^i > h_Y^i E_Y^i$, where the variables with a tilde represent counterfactual equilibrium values in the absence of home bias. These conditions imply that the size of home-biased expenditure on goods from sector S is smaller than the equilibrium output of S that would satisfy (5)-(6) in the absence of the home bias; i.e. that buyers’ home-biased expenditure does not exceed the hypothetical undistorted (by home bias) free-trade level of domestic production.⁶

⁶ This could be a strong assumption at the level of certain narrowly defined industries (for example in the defence sector) and on a small spatial scale (think of “buy local” practices of certain municipal governments). However, given that our empirical analysis is based on broad manufacturing sectors at the country level, this assumption is unlikely to be constraining.

By Walras' law, the equilibrium condition for Y is redundant. We write it out nonetheless, as it will be useful in the discussion below:

$$p_Y(Y^A + Y^B) = (1 - h_Y^A)E_Y^A + h_Y^A E_Y^A + (1 - h_Y^B)E_Y^B + h_Y^B E_Y^B. \quad (7)$$

The model so far is standard except for the home bias. The system (1)-(6) is composed of seven independent equations and eight unknowns ($p_X, p_Y, x, n^A, n^B, Y^A, Y^B, w$). Taking p_Y as the numéraire, the system is perfectly determined.⁷

3.4 A Discriminating Criterion

There is a difference between the CRS-PC sector and the IRS-MC sector that can be identified by simple inspection of equations (5)-(7): the parameter representing the home bias cancels out of equation (7), while it does not cancel out of equations (5) and (6). 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.

Consider a shock to the home bias $dh_X^A = -dh_X^B > 0$. As a consequence of this shock, the right-hand side of (5) increases by $(\phi n^B / n^A)(n^A + \phi n^B)^{-1} E_X^A + \phi(\phi n^A + n^B)^{-1} E_X^B > 0$. Since the left-hand side of (5) remains unchanged, the shock on the right-hand side requires an increase in n^A and a decrease in n^B in order to satisfy (5). The same applies, *mutatis mutandis*, to (6). These results hold for any set of parameter values and their associated solutions of the system. An

⁷ Note that, although buyers perceive domestic and foreign-produced Y as different, there is only one price for Y . This is because we assume that some share of expenditure in each country remains unbiased, which, combined with the no-corner-solution condition for Y , implies that the price of Y is equalised internationally.

increase in the own home bias, *ceteris paribus*, increases the own share of X -sector output. We call this the “home-bias effect” (HBE).⁸

Consider now a shock $dh_Y^A = -dh_Y^B > 0$. Such a shock to the home bias has no effect on the right-hand side of (7). Therefore, the CRS-PC sector, unlike the IRS-MC sector, is unaffected by countries’ relative home biases. This result is akin to Baldwin’s (1984) neutrality proposition, whereby home-biased government expenditure has no effect on international specialisation in a (CRS-PC) Heckscher-Ohlin model. The intuition is straightforward: provided the no-corner-solution condition holds, any increase in the home bias of home-biased buyers is compensated by an increase in the import share of unbiased buyers, since, for the latter, domestic and foreign-produced Y are perfect substitutes. This explains why the right-hand side of (7) remains unchanged (as does the sum of the first two terms and the sum of the second two terms in (7)).

It will be useful to express our main variables in terms of the share of country A : $\eta_X \equiv n^A / (n^A + n^B)$, $\eta_Y \equiv Y^A / (Y^A + Y^B)$, $\lambda \equiv L^A / (L^A + L^B)$, and $\varepsilon_S \equiv E_S^A / (E_S^A + E_S^B)$.

Differentiation of the system (1)-(6) yields the estimable equation and the formal derivation of our discriminating criterion:

$$d\eta_S = c_{1S}dh_S + c_{2S}d\varepsilon_S, \quad \text{for } S \in \{X, Y\}, \quad (8)$$

where $dh_S \equiv dh_S^A - dh_S^B > 0$ represents the change to relative home biases and $d\varepsilon_S > 0$ represent the idiosyncratic change in the size of expenditure.⁹ The coefficients of (8) – given below – is the basis of the discriminating criterion:

⁸ This is equivalent to what we have termed the “pull effect” in the context of public procurement (Brühlhart and Trionfetti, 2004).

$$c_{1X} \equiv \frac{\partial \eta_X}{\partial h_X} = \frac{\phi(1+\phi)}{(1-\phi)^2 + 4\phi h_X} > 0,$$

$$c_{1Y} \equiv \frac{\partial \eta_Y}{\partial h_Y} = 0,$$

$$c_{2X} \equiv \frac{\partial \eta_X}{\partial \lambda} = \frac{(1+\phi)(1-\phi + 2\phi h_X)}{(1-\phi)^2 + 4\phi h_X} > 1,$$

$$c_{2Y} \equiv \frac{\partial \eta_Y}{\partial \lambda} = -\frac{1}{(1-\alpha)} \frac{\alpha(1+\phi)(1-\phi + 2\phi h_X) - (1-\phi)^2 - 4\phi h_X}{(1-\phi)^2 + 4\phi h_X} < 1.$$

Hence:

- if $\hat{c}_{1S} > 0$ for sector S , then S is associated with IRC-MC, and
- if $\hat{c}_{1S} = 0$ for sector S , then S is associated with CRS-PC.

This discriminating criterion and its empirical implementation are the focus of our paper.¹⁰

⁹ Formally, system (1)-(6) is differentiated with respect to $d\lambda > 0$ and $dh_S^A = -dh_S^B$. The differentiation point is taken where countries are identical in all parameters, including the home bias ($h_S^A = h_S^B \equiv h_S$). The discrimination criterion does not depend on shocking the home bias symmetrically, i.e., $dh_S^A = -dh_S^B$. The HBE obtains also from independent shocks, such as $dh_S^A > 0$ and $dh_S^B = 0$. This will become important when we discuss robustness, in section 4.3. Regarding the source of the expenditure shock, $d\lambda > 0$ implies $dE_S^A = -dE_S^B > 0$ thus resulting in $d\varepsilon_S > 0$. Alternatively, a shock to preferences such as $d\alpha_X^A = -d\alpha_X^B > 0$, resulting in $dE_X^A = -dE_X^B > 0$ and $dE_Y^A = -dE_Y^B < 0$, would have equivalent implications. Although the specific algebraic value of the c_{2S} in equation (8) depends on the source of the expenditure shock, qualitative results and the discriminating criterion are unaffected. The expressions stated for c_{2S} below equation (8) are based on $d\lambda$ as the source of idiosyncratic expenditure.

¹⁰ Note that in the benchmark model underlying equation (8), c_{2X} exhibits the HME, and could thus serve as a discriminating feature as well. However, when we depart from this benchmark case, c_{2X} may be smaller than 1, and therefore no longer offer a valid discriminating criterion. This occurs under the generalisations discussed in Section 4, and/or if the home bias is modelled differently (see Appendix 1). Note also that, in the benchmark case, the coefficient c_{2Y} could be negative. The reason is that the larger is α_X , the larger is the X sector compared to the Y sector. If X is large enough, its expansion may require so much labour moving into X that Y may shrink despite the increase in λ . Finally, under autarky ($\phi = 0$), $c_{1X} = c_{1Y} = 0$, and $c_{2X} = c_{2Y} = 1$; and in the absence of home bias (i.e., if $h_X^A = h_X^B = 0$) the system (1)-(6) yields the familiar HME solution $\eta_X = 1/2 + [(1+\phi)/(1-\phi)](\varepsilon_X - 1/2)$.

The HME and the HBE criteria are similar in one respect: they both test theories by looking at the relationship between the geographical distribution of demand and the geographical distribution of output. The crucial difference between them is in that the HBE, unlike the HME, uses a component of the geographical distribution of expenditure (the home bias) which is independent of country size and of trade costs. This feature makes it robust to a number of generalisations of the benchmark model to which the HME has turned out to be sensitive.

4. Robustness

For the HBE-based test to be worth taking to data, we first need to ascertain that it is robust to three important generalisations of the model: trade costs in both sectors or in none, inelastic sectoral factor supplies, and real-world geography with multiple non-equidistant countries.¹¹

4.1 Trade Costs

Davis (1998) has highlighted the importance of exploring the implications of trade costs in the CRS-PC sector. The irrelevance of the home bias for the location of output in the CRS-PC sector, even if there are trade costs in that sector, can be ascertained by inspection of the suitably amended equilibrium equations as follows. Assume that for each unit of Y shipped only a fraction $\vartheta \in (0,1)$ arrives at its destination. Introducing trade costs in the Y sector results in the non-equalisation of goods and factor prices. Thus, prices of Y and wages become country specific. This implies that three endogenous variables (p_Y^i , p_X^i , and w^i) and three independent equations must now be added to system (1)-(6). The first two additional equations are equivalent to equations (1a,b), which now become country specific. The third equation is a new version of

¹¹ A fourth issue concerns the precise modelling of home bias. We show in Appendix 1 that the HBE criterion is robust to some alternative ways of introducing home bias in the utility function.

(7), which is no longer redundant. For clarity of exposition and without loss of generality we consider the case where country A is a net exporter of Y . Then, equation (7) becomes

$$p_Y^A [(1 - h_Y^A) E_Y^A + h_Y^A E_Y^A] + \mathcal{G} p_Y^B [(1 - h_Y^B) E_Y^B + h_Y^B E_Y^B] = p_Y^A Y^A + \mathcal{G} p_Y^B Y^B. \quad (7')$$

Once again, h_Y^i cancels out. This means that here too the home bias is irrelevant for international specialisation in the CRS-PC sector. The reason is that the increase in expenditure on the domestic good due to the increase in the home bias of the biased buyers crowds out expenditure on the domestic good by the unbiased buyers, who switch to imports. This crowding-out effect is independent of the presence of trade costs, since (a) as long as there is trade, the consumer price of domestic goods is identical to that of corresponding imports, and (b), as far as unbiased expenditure on Y goods is concerned, domestically produced goods and imports are perfect substitutes. Since equations (5) and (6) remain unchanged, it is straightforward that home bias still matters for the IRS-MC sector. Therefore, the HBE criterion remains valid.

A second feature of the HBE criterion concerning trade costs is that, unlike the HME, it is valid even in the absence of trade costs. To see this, it suffices to set $\phi = 1$ in equations (5)-(6). This yields the unique solution $\eta_X = h_X^A \lambda / [h_X^A \lambda + h_X^B (1 - \lambda)]$. It is easily verified that $\partial \eta_X / \partial \lambda > 0$ in this case as well. By contrast, if neither country is home biased ($h_X^A = h_X^B = 0$) and there are no trade costs, the HME cannot provide a discriminating criterion. The reason is that, in this case, the solution is indeterminate (0/0), and the derivative $\partial \eta_S / \partial \varepsilon_S$ is zero in all sectors.

4.2 *Imperfectly Elastic Labour Supply*

Head and Mayer (2004, Section 6.4) have demonstrated that the HME need not occur in IRS-MC sectors when sectoral factor supplies are imperfectly elastic. We can show that the criterion based on the HBE is robust to the presence of imperfectly elastic factor supplies.

Assume that, in addition to L , sector Y uses a production factor K , with a Cobb-Douglas technology. The resulting structure is similar to the one adopted by Puga (1999) and implied by Head and Mayer (2004): if the IRS-MC sector demands more labour, wages will rise because of the decreasing marginal productivity of labour in the CRS-PC sector.¹² Let γ denote the labour share of total costs in the production of Y . Given the production technology in the CRS-PC sector, labour demand from sector Y is $(\gamma/w^i)^{\gamma/(1-\gamma)}\bar{K}$, and factor rewards are $w^i = \gamma(k_Y^i/l_Y^i)^{1-\gamma}$ and $r^i = (1-\gamma)(\gamma/w^i)^{\gamma/(1-\gamma)}$. After some normalisations, the labour market-clearing conditions can be written as:

$$n^i = \bar{L} - (\gamma/w^i)^{\gamma/(1-\gamma)}\bar{K}, \quad i = A, B.$$

The market-clearing conditions for goods are given by equations (5) and (6), taking account of the wage and rental rate equations given above.

Totally differentiating the resulting system around its symmetric equilibrium, we again obtain our testing equation (8), where the coefficients now are as follows:¹³

¹² Obviously, in this modified structure, factor-price equalisation no longer holds. All factors are assumed to be internationally immobile for simplicity. Factor K earns a Ricardian surplus rk_Y , which is maximised subject to $Y = (l_Y)^\gamma (k_Y)^{1-\gamma}$, taking wages as given. For the sake of symmetry and simplicity, we assume that $K^A = K^B = \bar{K}$ and $L^A = L^B = \bar{L}$.

¹³ As before, the system is differentiated around identical parameter values for the two countries. The symmetric equilibrium yields solutions $n^A = n^B = 2\alpha_X\bar{L}/(1+\alpha_X)$, from which the symmetric-equilibrium values of all other endogenous variables can be recovered. The resulting expressions for c_{1X} and c_{2X} are involved and not

$$c_{1X} = \frac{-8\phi(1+\phi)}{-3 + \phi(8-14h_X) - \phi^2(5-2h_X) - 8\sigma\phi(1-h_X)} > 0, \quad (9)$$

$$c_{2X} = \frac{-8(1+\phi)(1+2h_X\phi-\phi)}{-3 + \phi(8-14h_X) - \phi^2(5-2h_X) - 8\sigma\phi(1-h_X)} > 0. \quad (10)$$

It is clear by inspection that $c_{1X} > 0$, since the denominator is negative for any combination of parameter values. This means that the country that is relatively more home biased will, *ceteris paribus*, tend to specialise in the production of X , which confirms the validity of the HBE criterion even in the presence of imperfectly elastic labour supply (c_{1Y} , of course, remains zero in the current context). This result is in fact quite intuitive. A shock $dh_X \equiv dh_X^A - dh_X^B > 0$ causes an increase in demand for the varieties produced in country A . This increased demand can be absorbed in three possible ways: (1) if labour supply is perfectly elastic, the demand increase is entirely absorbed by an increase in output; (2) if labour supply is imperfectly inelastic, the demand increase is absorbed partially by wages and partially by an increase in output; and (3) if labour supply is perfectly inelastic, the demand increase is entirely absorbed by an increase in wages. The first case is considered in Section 3.4. The third case applies when there is neither intersectoral nor international labour mobility. In this extreme case, which our model does not accommodate, $\frac{\partial \eta_X}{\partial \varepsilon_X} = \frac{\partial \eta_X}{\partial h_X} = 0$, simply because there are no additional inputs available to

expand production of X .

As for the parameter c_{2X} , inspection of (10) reveals that it is positive but not necessarily larger than 1. Note that c_{2X} is not necessarily larger than one even if we eliminate the home bias from

particularly informative. They can be substantially simplified if we set $\gamma = 1/2$ and $\alpha_X = 1/2$, which is what we do to obtain (9) and (10). The qualitative results do not depend on this simplification.

the model. This can easily be seen by setting $h_X = 0$ in equation (10). The same result has been found numerically by Head and Mayer (2004) using exogenous values for the elasticity of sectoral labour supply. Here we provide an explicit analytic expression corresponding to their simulations.

4.3 *Multiple Non-Equidistant Countries*

Results derived in the two-country model extend to the many-country case if we assume that the world's M countries are equidistant from each other. This is of course not realistic. As pointed out by Head and Mayer (2004), this issue restricts the validity of the test based on the HME. Davis and Weinstein (2003) address this empirically, and a theoretical analysis is provided by Behrens *et al.* (2004).

We can show that our proposed discriminating criterion remains valid in an asymmetric multi-country world. The multi-country model complicates the analysis of HMEs because the expenditure shocks applied to the system are subject to the constraint that $\sum_{i=1}^M dE_X^i = 0$.

Therefore, a shock to country i 's own expenditure $dE_X^i > 0$ must imply a change in expenditure to at least one other country. Without information on the distribution of the expenditure changes across the $M-1$ other countries it is impossible to know the total effect of $dE_X^i > 0$. This problem does not appear in the two-country case because, by construction, $dE_X^B = -dE_X^A$. In the case of more than two countries, we could have that $dE_X^i > 0$ is accompanied by $-dE_X^j > dE_X^i > 0$ while

still satisfying the constraint that $\sum_{i=1}^M dE_X^i = 0$. If i - j bilateral trade costs are sufficiently low, the output response to an increase in E_X^i may be less than proportional (or indeed negative).¹⁴

This problem happens not to afflict the test based on the HBE, because there is no constraint on the sum of all home biases. So, we can indeed consider a shock $dh^i > 0$, *ceteris paribus*. It is then quite straightforward from inspection of equations (5) and (6) that a shock $dh_X^A > 0$ increases aggregate demand for country A 's varieties regardless of the number of countries and of the structure of trade costs. Given that the left-hand side of the equations is constant, an increase in η_X is necessary to absorb the excess demand.

In sum, we find that the HBE-based discriminating criterion is robust (a) to various ways of modelling trade costs, (b) to imperfectly elastic intersectoral factor supply, and (c) to the assumption of $M > 2$ asymmetrically spaced countries.

5. Empirical Implementation

We operationalise our discriminating criterion in two stages. First, we estimate home biases across industries and countries. Those bias estimates are then used as an ingredient to the estimation of our testing equation (8).

¹⁴ Behrens *et al.* (2004) find that, while the two-country HME does not generalise to an M -country setting, a one-to-one relationship between output shares and *spatially filtered* expenditure shares characterises IRS-MC sectors in such a general model. We consider that relationship as a complement to the HBE test in our empirical analysis.

5.1 Estimating Home Bias

We estimate home bias separately for each country-industry pair, using a gravity equation that substitutes fixed effects for country-specific variables. Thanks to the general compatibility of this approach with the major theoretical paradigms, using the gravity equation at the first stage of our exercise should not prejudice our inference in stage two.¹⁵ Specifically, we estimate the following regression equation:

$$\begin{aligned}
 LOGIM_S^{ij} = & \\
 \alpha + \beta_1 HOMEDUM^{ij} + \beta_2 LOGDIST^{ij} + \beta_3 BORDUM^{ij} + \beta_4 LOGTARIFF_S^{ij} & \\
 + \beta_5 NTB_S^{ij} + \beta_6 COLONYDUM^{ij} + \beta_7 SAMECTRYDUM^{ij} + \beta_8 OFFLANGDUM^{ij} & \\
 + \beta_9 SPKLANGDUM^{ij} + \delta' \mathbf{M}^i + \theta' \mathbf{X}^j + \kappa' \mathbf{S} + \varepsilon_S^{ij} & \quad (11)
 \end{aligned}$$

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

$LOGIM_S^{ij}$	=	log of sector S imports of country i from country j ,
$HOMEDUM$	=	dummy equal to one if $i = j$, and zero otherwise,
$LOGDIST$	=	log of geographical distance between the two countries,
$BORDUM$	=	dummy equal to one if i and j are different countries that share a common border, and zero otherwise,
$LOGTARIFF$	=	log of applied tariff rate,
NTB	=	frequency measure of non-tariff barriers,
$COLONYDUM$	=	dummy equal to one if i and j are different countries that have or have had a colonial link,

¹⁵ The gravity model has been shown to be successful even at the level of individual industries *inter alia* by Bergstrand (1990), Chen (2004), Davis and Weinstein (2001), Feenstra *et al.* (2001) and Head and Mayer (2000). See Feenstra (2004) for a discussion of the advantages of the approach based on country fixed effects.

SAMECTRYDUM = dummy equal to one if i and j are different countries that have been part of the same nation at some time in modern history,

OFFLANGDUM = dummy equal to one if i and j are different countries that share a common official language, and zero otherwise,

SPKLANGDUM = dummy equal to one if i and j are different countries that share a common spoken language, and zero otherwise,

\mathbf{M}^i = vector of importer-specific fixed effects,

\mathbf{X}^j = vector of exporter-specific fixed effects,

\mathbf{S} = vector of industry fixed effects, and

ε = a potentially heteroskedastic stochastic error term.

The object of our interest is β_1 , the coefficient on imports within countries, often referred to as the “border effect”. A positive (negative) coefficient is interpreted as positive (negative) home bias. By including variables for distance, adjacency, tariffs, NTBs, common colonial and national heritage, and common language, we aim to control for physical and policy-induced trade costs as well as for informational and marketing costs in accessing foreign markets. 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.

A 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 (2003), border effects depend not only on home biases and trade costs, but also on the elasticity of substitution among an industry’s products: measured 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 between-industry comparisons. The purpose of our home-bias estimates, however, is to allow for comparison across countries, industry by industry, and hence our final exercise is unlikely to be affected significantly by this concern.

Two practical difficulties remain. First, one has to find a measure of “trade within countries”, and, second, the distance variable has to be defined for intra-country trade. Following Wei (1996), we define trade within countries as output minus exports.¹⁶ 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 is consistent with 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 draw on the estimates by CEPII, which establish consistency between international and internal distance measures, as they base the latter on size-weighted distances among the main cities inside each country.¹⁷ For comparison, we also consider a frequently employed approximation initially suggested by Leamer (1997), according to which internal distance is defined as two thirds of the radius of a circle with the same area as the country in question ($DIST_DISC^{ii} = \ln(0.67\sqrt{Area^i / \pi})$).

Having constructed the intra-country variables and drawing on the World Bank’s Trade and Production Database, our data cover 17 industrial sectors, 60 importing countries, up to 164

¹⁶ Hence, $LOGIM^i = \log(Output^i - \sum_j Exports^{ij})$.

¹⁷ See www.cepii.fr/anglaisgraph/bdd/distances.htm.

exporting countries in 1997. This provides a data set with 112,010 industry-level bilateral observations. A full description of variables and data sources is given in Appendix 2.

We begin by running equation (11) on the entire data set. These results are shown in Table 1. The first data column reports OLS results for a model that retains only observed trade flows (which make up 52 percent of the observations for which we know the values of the right-hand-side variables). As usual, the gravity model provides a good fit to the data, and all estimated coefficients have the expected signs. The pooled OLS regression suggests very strong home bias: on average, purchases from national sources are predicted to be 41 ($=e^{3.71}$) times larger than purchases from sources the same distance away but located in a different country speaking a different language. Even by the standards of the literature on border effects, which has been struggling to explain the surprisingly large coefficients found in numerous settings, this number is implausibly high. It is therefore reassuring that a regression that considers also the large number of zero trade flows, using the Tobit estimator, yields an estimated mean home bias of 1.9 ($=e^{0.64}$), which may still seem large but fits at the lower end of the range of estimates obtained elsewhere (e.g. Chen 2004; Head and Mayer, 2002). If we replace the cities-based measure of internal distances with the Leamer approximation (model 3 of Table 1), the mean estimated home bias rises to 2.1 ($=e^{0.73}$). Given the unavoidably imprecise nature of distance measures, one must be careful in interpreting the absolute magnitude of these home-bias estimates. Fortunately, this is but a minor problem in the context of our paper, since what we need for our discriminating criterion is an estimate of *relative* home biases across countries, and these are not significantly affected by the method used to measure internal distances.

Imposing identical coefficients across the two dimensions of our panel is restrictive. Our paper builds on the presumption that home biases differ across countries and sectors. Sector-country

estimates of *IDIOBIAS* are obtained via sector-by-sector pooled regressions with a separate dummy for each intra-country observation and full sets of importer and exporter fixed effects. Thus, we obtain individual home-bias estimates per country-industry pair, which we call *IDIOBIAS*. We use the same specification as that of model 2 in Table 1, minus industry and importer fixed effects. *IDIOBIAS*, the estimated matrix of country-industry specific Tobit coefficient on *HOMEDUM*, provides the key ingredient to our testing equation.¹⁸

5.2 An Empirical Test of the Discriminating Criterion

We begin the estimation of the discriminatory criterion by taking equation (8) literally and estimating the following equation for each of the 17 industries S across the importing countries i :

$$OUTPUTSHARE_S^i = c_{0S} + c_{1S}IDIOBIAS_S^i + c_{2S}EXPENDISHARE_S^i + \nu_S^i, \quad (12)$$

where superscripts denote countries, subscripts denote industries, and:

$$OUTPUTSHARE_S^i = \frac{Output_S^i}{\sum_i Output_S^i},$$

$IDIOBIAS_S^i$ = estimated coefficient on $HOMEDUM_S^i$ from disaggregated Tobit estimation of equation (11), and

$$EXPENDISHARE_S^i = \frac{NetExpenditure_S^i}{\sum_i NetExpenditure_S^i}.$$

¹⁸ Note that, by focusing on unadjusted Tobit coefficients, we retain the effects on a latent variable that might be called “propensity to trade”, where positive trade flows obtain only if this propensity exceeds a certain threshold. Implied coefficients for strictly positive trade flows (estimated for mean values of the explanatory variables) are given in brackets in Table 1. These coefficients on observed trade flows are consistently smaller in absolute value than the Tobit coefficients, which suggests, as expected, that the probability that trade is observed relates to the regressors qualitatively in the same way as the volume of trade conditional on trade being observed.

According to our discriminating criterion, industries with estimated c_{1S} of zero conform with the CRS-PC model, whereas industries with positive estimated c_{1S} conform with the IRS-MC model.

Five issues warrant discussion. First, there is the question of sectoral disaggregation. Neither theory nor existing empirical work give us strong priors as to the correct definition of an “industry” and constituent “goods” in the data (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 (we think realistic) assumption that there are more goods than factors, and it is a convenient feature in view of empirical implementation, as it does away with the need to draw a dividing line between the two levels of sectoral aggregation.¹⁹

Second, we must suspect potential for simultaneity of expenditure and output, and therefore bias in the parameter estimates. Our testing equation (8) implies the assumption that expenditure shares are an exogenous determinant of output location, but this assumption is unlikely to be satisfied in the data. The use of input-output tables allows us to attenuate this problem. The main source of potential simultaneity bias is sectoral expenditure representing demand for intermediate inputs that are classified under the same sector heading (see, e.g., Hillberry and Hummels, 2002). It is for this reason that we compute net (i.e. final) expenditure per sector. The definition of *NetExpenditure* includes expenditure from sources that use the industry’s output for

¹⁹ Davis and Weinstein (2003) found *EXPENDIHARE* to be highly collinear with their endowment variables, and therefore dropped it from their testing specification. Since we are doing the reverse (including *EXPENDISHARE* but not the endowment variables), omitted-variable bias is unlikely to be important.

final consumption, and exclude expenditure from those sources that use the output as intermediate inputs.²⁰

Third, ν_s^i is likely to be heteroskedastic, as the variance of errors may well be positively correlated with the size of countries.²¹ Our significance tests are therefore based on heteroskedasticity-consistent 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 c_{1S} .

Fourth, *IDIOBIAS* 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-estimate 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 follow 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 turns out that the estimated biases are significantly below that threshold

²⁰ See Appendix 2 for details on the computation of *NetExpenditure*. Another form of simultaneity could potentially arise if an unobserved country-specific factor drove both the left-hand side and the right-hand side of equation (12). For instance, expenditure shares might be affected by historical specialisation patterns and adjust slowly to changes in specialisation. While probably less likely, such a link might in principle also exist with respect to relative home biases. This issue could be addressed if we ran the two steps of our estimation procedure in panel data. This would allow the differencing-out of time-invariant country effects (at the considerable cost, however, of losing the between-country variance as a source of identification for equation 12), but data constraints (particularly regarding the time profile of sectoral bilateral trade barriers) put such an exercise beyond the scope of this paper. These considerations, together with the result found by Head and Ries (2002) that home-market effects estimated "within" are significantly smaller than those estimated "between", lead us to view our empirical results as upper-bound estimates of c_2 , and, possibly, of c_1 .

²¹ A Breusch-Pagan test on the pooled model strongly rejects the null of constant error variance.

in all of the specification we estimate. Hence, we report OLS coefficient but base hypothesis tests on bootstrap error distributions.²²

Fifth, the theory underlying our testing equation is couched in a two-country setting. While we show that the HBE generalises to a world of $M > 2$ asymmetrically spaced countries, Behrens *et al.* (2004) demonstrate that an equivalent generalisation to the relationship between expenditure shares and output shares would be erroneous. Instead, they show that, in an asymmetric M -country IRS-MC model, there will be a one-to-one relationship between sectoral output shares and *spatially filtered* expenditure shares. Specifically, they show that (in our notation)

$$\eta_X = \mathbf{\Omega} \varepsilon_X, \quad \text{where } \mathbf{\Omega} = [\text{diag}(\mathbf{\Phi}^{-1} \mathbf{1}) \mathbf{\Phi}]^{-1}, \quad (13)$$

and $\mathbf{\Phi}$ is the $M \times M$ matrix of ϕ s (Behrens *et al.*, 2004, eq. 17). Head and Mayer (2004) have shown that an IRS-MC model suggests the following formula for estimating ϕ :

$$\hat{\phi}^{ij} = \sqrt{M^{ij} M^{ji} / M^{ii} M^{jj}}. \text{ We apply this formula to compute } \mathbf{\Omega}, \text{ and estimate the testing equation}$$

(12) in a version that replaces *EXPENDISHARE* with its filtered counterpart that we name *EXPENDISHARE_BLOT*. Note that it is not clear, *a priori*, which of the two expenditure measures should be preferred, since application of the Behrens *et al.* (2004) filter presupposes that IRS-MC provides the appropriate model. We therefore estimate both versions of the testing equation.

²² 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).

Furthermore, given that Behrens *et al.* (2004) show that (13) holds in the IRS-MC sector, we can use a test on the null hypothesis that the coefficient on spatially filtered expenditure shares ($\Omega_{\mathcal{E}X}$) equals one as a complementary strategy to identify the IRS-MC paradigm in the data.

5.3 Results: Pooled Estimates

We first run our model on the full data sample. The results are given in Table 2. Columns (1) and (2) report estimates of equation (12), without and with country fixed effects.²³ In both cases, the estimated coefficients on *IDIOBIAS* are statistically significantly larger than zero, which, according to equation (8), suggests that on average the data support the IRS-MC model. The same patterns are found in regressions (3) and (4), where we use spatially filtered expenditure shares: we again find statistically significantly positive coefficients on *IDIOBIAS*, consistent with the IRS-MC model.

As a third exercise, we estimate the testing equation using gross expenditure (columns 5 and 6). As expected, this tends to increase the estimated coefficient on the expenditure share. It also biases the estimated coefficient on *IDIOBIAS* downwards. Purging expenditure measures of intermediate expenditure in order to avoid simultaneity bias is therefore confirmed as an important component of our estimation strategy.

The results of Table 1 appear inconsistent in an important respect. Bootstrap tests on the hypothesis that the coefficients on *EXPENDISHARE_BLOT* are equal to one reject that hypothesis strongly. Since this is inconsistent with equation (13), our estimates on *EXPENDISHARE_BLOT* reject the IRS-MC model while our estimates on *IDIOBIAS* support that model. Although the practicalities of constructing *EXPENDISHARE_BLOT* inevitably

²³ Industry fixed effects are redundant, because *IDIOBIAS* represents deviations from industry means, and the other variables represent industry shares.

increase the scope for measurement error, this inconsistency could cast doubt on the robustness of our findings. It should be noted, however, that by imposing equal coefficients across sectors the pooled estimations of Table 2 violate the basic premise of our research that sectors differ in their relevant characteristics. These regressions therefore impose too much structure, which is why the main focus of our empirical exercise should be on sector-level estimation.

5.4 Results: Industry-by-Industry Estimates

The regression results of the baseline specification derived from the two-country model are given in Table 3 for each of our 17 sample industries. The equation generally performs well, yielding R^2 s between 0.66 and 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 seven of the 17 sectors conform with the IRS-MC paradigm. The allocation of sectors looks plausible, as it comprises all the machinery and engineering sectors (ISIC 38) plus textiles, non-ferrous metals and “other manufactures” (jewellery, music instruments, sports equipment and non-classifiable items).

Taking these results at face value, we can measure the relative importance of the two paradigms in terms of their share of industrial output (Table 3, last column). The seven sectors that, applying the 95-percent confidence criterion, conform with the IRS-MC prediction account for exactly 50 percent of sample output.

Finally, Table 4 reports the corresponding results for the specification that includes spatially filtered expenditure shares, *EXPENDISHARE_BLOT*. This equation too performs well, with R^2 s in the ~~somewhat lower~~ range 0.45 to 0.96. Coefficients on *IDIOBIAS* are still predominantly

positive or insignificantly different from zero, as predicted by the theory. However, two sectors, food products and wood products, exhibit statistically significantly negative coefficients (at the 90-percent confidence level), which runs plainly against the theory. If we retain the 95-percent criterion for our test based on *IDIOLIAS*, only two sectors conform with the IRS-MC prediction, “fabricated metal products” and “other manufactures”. Given the additional scope for measurement error implied in these estimations, they probably should be considered to be approximative. It is, however, interesting to note that the one-to-one relationship predicted to hold between *OUTPUTSHARE* and *EXPENDISHARE_BLOT* by the IRS-MC model of Behrens *et al.* (2004) seems to fit best for the engineering sectors (ISIC 38) and for “other manufactures”, which also have a high incidence of statistically positive coefficients on *IDIOLIAS*. Together with our findings of Table 3, we therefore conclude that these sectors, accounting for 45 percent of sample output, fit the predictions of the IRS-MC model best.

6. Conclusions

We develop and apply an empirical test to distinguish two paradigms of international trade theory: a model with constant returns and perfect competition (CRS-PC), and a model with increasing returns and monopolistic competition (IRS-MC). The discriminating criterion makes use of the assumption that demand is home biased, an assumption that is well 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 show to be robust to a number of theoretical generalisations, including imperfectly elastic sectoral factor supply and multiple non-equidistant countries.

Our discriminating criterion can be taken to data. In the empirical part we estimate industry- and country-level home biases through disaggregated gravity regressions in a cross section of up to 60 importing countries, and use these estimates to apply our test separately for 17 manufacturing industries. The results suggest that the engineering industries (fabricated metal products, non-electrical machinery, electrical machinery and precision engineering, and transport equipment), plus “other manufacturing”, which together account for close to half of manufacturing output value in our sample, conform with the predictions of the IRS-MC model.

Our paper opens some potentially fruitful avenues for future research. In terms of theory, one could broaden the focus beyond the two benchmark models that we study here, to map the incidence of home bias on specialisation in a variety of setups with different combinations of assumptions on market structure and production technologies - similar to Head, Mayer and Ries’s (2002) exploration of the home-market effect across different trade models. Empirically, it would be particularly interesting to estimate our model in a panel data set. The data requirements would be formidable, but such an analysis could in principle allow the differencing-out of time-invariant features that might simultaneously affect countries’ production patterns and relative home biases in spite of our best efforts at eliminating such factors through the use of input-output data.

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TABLE 1: Gravity Equations: Full Sample¹

dependent var. = $\ln(M)$	Tobit ²		
	OLS	(2)	(3)
	(1)	(2)	(3)
<i>HOMEDUM</i>	3.714** (0.103)	0.640** (0.152) [0.429]	0.726** (0.153) [0.489]
<i>DIST</i>	-1.337** (0.017)	-2.186** (0.020) [-1.391]	
<i>DIST_DISC</i>			-2.080** (0.020) [-1.319]
<i>ADJACENCYDUM</i>	0.568** (0.063)	0.907** (0.087) [0.621]	0.881** (0.087) [0.600]
<i>TARIFF</i>	-0.044** (0.009)	-0.055** (0.010) [-0.035]	-0.051** (0.010) [-0.033]
<i>NTB</i>	-0.432** (0.188)	-0.466** (0.177) [-0.296]	-0.502** (0.178) [-0.319]
<i>COLONYDUM</i>	0.602** (0.055)	0.997** (0.076) [0.686]	0.973** (0.076) [0.666]
<i>SAMECOUNTRYDUM</i>	0.637** (0.091)	0.653** (0.127) [0.439]	0.667** (0.126) [0.447]
<i>OFFLANGDUM</i>	0.077 (0.057)	0.819** (0.067) [0.548]	0.799** (0.067) [0.532]
<i>SPOKLANGDUM</i>	0.448** (0.055)	0.299** (0.066) [0.194]	0.388** (0.066) [0.252]
Importer fixed effects	yes	yes	yes
Exporter fixed effects	yes	yes	yes
Industry fixed effects	yes	yes	yes
Year fixed effects	yes	yes	yes
Observations	57,948	112,010	112,010
R ²	0.580		
Pseudo R ²		0.236	0.235

¹ Robust standard errors in parentheses. Distance and tariff variables in natural logs (see text for precise variable definitions). ** (*) denotes rejection of H₀: coeff. = 0 at 95% (90%) confidence level, based on heteroskedasticity-consistent standard errors.

² Tobit coefficients on latent dependent variable. Estimated elasticities of observed dependent variable at sample means in brackets.

TABLE 2: Pooled Estimation of the Discriminating Criterion ¹
(dependent variable = *OUTPUTSHARE*)

	(1)	(2)	(3)	(4)	(5)	(6)
<i>IDIOBIAS</i>	0.126** (0.040)	0.107** (0.053)	0.119** (0.060)	0.087** (0.065)	0.018 (0.018)	-0.015 (0.025)
<i>EXPENDISHARE</i>	0.974** (0.049)	0.458** (0.117)				
<i>EXPENDISHARE_BLOT</i>			0.774** / ## (0.058)	0.287** / ## (0.135)		
<i>EXPENDISHARE_GROSS</i>					0.958** (0.001)	1.166** (0.002)
Country fixed effects	no	yes	no	yes	no	yes
<i>N</i>	652	652	588	588	649	649
<i>R</i> ²	0.862	0.923	0.657	0.868	0.969	0.976

¹ Constant term included in all regressions but not reported. Coefficients and standard errors reported with respect to (*IDIOBIAS* * 1000). Bootstrap standard errors (5,000 iterations with replacement). ** (*) denotes rejection of H0: coeff. = 0 at 95% (90%) confidence level, based on bias-corrected bootstrap confidence intervals (two-tail test for *EXPENDISHARE*, one-tail test for *IDIOBIAS*). ## (#) denotes rejection of H0: coeff. = 1 at 95% (90%) confidence level (two-tail test), based on bias-corrected bootstrap confidence intervals

TABLE 3: Industry-by-Industry Estimation of the Discriminating Criterion, Two-Country Model
(bootstrapped OLS; dependent variable = *OUTPUTSHARE*)

ISIC (Rev. 2)	Description	<i>IDIOBIAS</i> ¹		<i>EXPENDISHARE</i>		R ²	N	Size share ²
		<i>coefficient</i> ³	<i>std error</i> ⁴	<i>coefficient</i> ³	<i>std. error</i> ⁴			
311	Food products	-0.157	0.139	0.967**	0.078	0.985	38	0.109
313/4	Beverages, tobacco	0.028	0.186	0.928**	0.037	0.996	39	0.032
321	Textiles	0.188**	0.060	1.090**	0.101	0.941	41	0.033
322	Clothing	0.106	0.145	0.764**	0.154	0.950	32	0.014
323/4	Leather, footwear	0.015	0.391	0.585**	0.406	0.494	37	0.007
331/2	Wood products	0.129	0.300	0.998**	0.125	0.971	34	0.030
341/2	Paper products, publishing	0.096	0.376	1.022**	0.297	0.919	40	0.073
351/2/5/6	Chemicals	0.353*	0.464	1.231**	0.223	0.929	40	0.129
353/4	Petroleum products	-0.071	0.164	1.142**	0.127	0.976	38	0.034
361/2/9	Non-metall. mineral prods	0.183	0.251	0.972**	0.245	0.718	40	0.033
371	Iron, steel	-0.151	0.363	1.218**	0.359	0.664	39	0.037
372	Non-ferrous metals	0.216**	0.135	1.264**	0.291	0.859	39	0.019
381	Fabricated metal products	0.708**	0.414	1.213**	0.366	0.772	39	0.054
382	Non-electrical machinery	0.136**	0.082	1.089**	0.224	0.944	40	0.115
383/5	Electrical machinery, precision engineering	0.131**	0.101	1.099**	0.284	0.918	39	0.137
384	Transport equipment	0.170**	0.111	0.917**	0.217	0.941	39	0.130
390	Other manufactures	0.313**	0.241	0.769**	0.424	0.865	35	0.012

¹ Coefficients and standard errors reported with respect to (*IDIOBIAS* * 1000)

² Share in total output of sample countries

³ ** (*) denotes rejection of H0: coeff. = 0 at 95% (90%) confidence level (one-tail test), based on bias-corrected bootstrap confidence intervals.

⁴ Bootstrap standard errors (5,000 iterations with replacement).

TABLE 4: Industry-by-Industry Estimation of the Discriminating Criterion , *M*-Country Model
(bootstrapped OLS; dependent variable = *OUTPUTSHARE*)

ISIC (Rev. 2)	Description	<i>IDIOTBIAS</i> ²		<i>EXPENDISHARE_BLOT</i> ¹		R ²	N
		<i>coefficient</i> ⁵	<i>std error</i> ⁴	<i>coefficient</i> ³	<i>std. error</i> ⁴		
311	Food products	-0.470**	0.225	0.630 ^{##}	0.068	0.955	35
313/4	Beverages, tobacco	-0.117	0.300	0.681 ^{##}	0.093	0.936	37
321	Textiles	0.068	0.210	0.693	0.179	0.694	38
322	Clothing	0.079	0.144	0.605 [#]	0.179	0.766	30
323/4	Leather, footwear	-0.175	0.332	0.821	0.264	0.648	36
331/2	Wood products	-0.481*	0.361	0.657	0.198	0.771	33
341/2	Paper products, publishing	0.766	0.908	0.738	0.377	0.517	37
351/2/5/6	Chemicals	0.472*	0.792	0.766	0.297	0.633	37
353/4	Petroleum products	-0.073	0.167	0.711	0.168	0.843	35
361/2/9	Non-metall. mineral prods	-0.146	0.538	0.697	0.319	0.438	37
371	Iron, steel	-0.014	0.328	0.785	0.460	0.413	36
372	Non-ferrous metals	0.018	0.107	0.572 [#]	0.191	0.646	36
381	Fabricated metal products	0.825**	0.756	1.016	0.485	0.453	36
382	Non-electrical machinery	0.049	0.162	1.035	0.308	0.826	37
383/5	Electrical machinery, precision engineering	0.163*	0.229	1.031	0.383	0.703	36
384	Transport equipment	0.229	0.359	0.913	0.274	0.814	36
390	Other manufactures	0.738**	0.525	0.947	0.312	0.793	33

¹ Expenditure share weighted by spatial trade-freeness matrix *à la* Behrens *et al.* (2004); see text for details

² Coefficients and standard errors reported with respect to (*IDIOTBIAS* * 1000)

^{3##} (#) denotes rejection of H0: coeff. = 1 at 95% (90%) confidence level (two-tail test), based on bias-corrected bootstrap confidence intervals

⁴ Bootstrap standard errors (5,000 iterations with replacement)

^{5**} (*) denotes rejection of H0: coeff. = 0 at 95% (90%) confidence level (one-tail test), based on bias-corrected bootstrap confidence intervals

Appendix 1: The HBE When Home Bias Features in the CES Sub-Utility Function

We can show that the HBE remains valid when the home bias is inserted in the CES sub-utility, for instance in the following way:

$$u^i = \left(\psi^i \int_{k \in n^i} c_k^{\frac{\sigma-1}{\sigma}} dk + \omega^i \int_{k \in n^j} c_k^{\frac{\sigma-1}{\sigma}} dk \right)^{\frac{\sigma}{\sigma-1}},$$

where ψ^i and ω^i are weights, and $h_X^i \equiv (\psi^i / \omega^i)^\sigma \in (0, \infty)$ represents the home bias. This is the modelling approach chosen by Head and Ries (2001), Hummels (2001) and Combes *et al.* (2005). Expressing the equilibrium equations for the IRS-MC sector in terms of η_X and λ , we have:

$$\frac{h_X^A - \phi}{h_X^A \eta_X + \phi(1 - \eta_X)} \lambda + \frac{\phi - h_X^B}{\eta_X \phi + h_X^B(1 - \eta_X)} (1 - \lambda) = 0,$$

which has solutions $\eta_X = \frac{h_X^A h_X^B \lambda + \phi^2 (1 - \lambda) - \phi h_X^B}{(h_X^A - \phi)(h_X^B - \phi)}$ and derivatives:

$$c_{1X} \equiv \frac{\partial \eta_X}{\partial h_X} = \frac{\phi(1 + \lambda)}{(h_X^A - \phi)^2} + \frac{\phi \lambda}{(h_X^B - \phi)^2} > 0,$$

$$c_{2X} \equiv \frac{\partial \eta_X}{\partial \lambda} = \frac{h_X^A h_X^B - \phi^2}{(h_X^A - \phi)(h_X^B - \phi)}.$$

The coefficient c_{1X} is positive, which confirms the validity of the discriminating criterion based on the HBE.²⁴

²⁴ The coefficient c_{2X} is larger than unity if both countries are home biased but it may be positive and less than one if both countries are “foreign biased” (i.e. $h_X^i < 1$). The exact condition for $c_{2X} \in (0, 1)$ is $\sqrt{h_X^A h_X^B} > \phi > (h_X^A + h_X^B)/2$.

Appendix 2: Data Description

Sectoral trade and output data as well as input-output tables are taken from the World Bank database (Nicita and Olarreaga, 2001). The trade and output data, originally classified into 28 ISIC industries, were aggregated up to 17 industries, so as to be compatible with the sector classification of input-output tables. We retained trade data recorded by the importing countries. Hence, all trade flows are c.i.f., and our estimates of within-country trade can be considered conservative. Observations for which estimated intra-country trade was, implausibly, negative (i.e. output - exports < 0) were set to zero.

Based on the variables available from the input-output tables, net (final) expenditure was computed as follows (sector subscripts omitted for simplicity):

$$\begin{aligned} \text{NetExpenditure} = & (\text{Output} - \text{Exports})(1 - \text{ShareIntermSales}) \\ & + \text{Imports} - (\text{Output} * \text{ShareIntermImp} * \text{ShareIntermOut}) \end{aligned}$$

where *ShareIntermSales* is the share of output sold to other sectors as intermediate inputs, *ShareIntermImp* is the share of intermediates used that is imported, and *ShareIntermOut* is the share of own-sector intermediates needed to produce one unit of output value. The first summand represents home-produced domestic final consumption, while the remaining terms represent foreign-produced domestic final consumption.

The reference year for our analysis is 1997. Where observations for trade and/or production in 1997 were missing, we used analogous data for the nearest available years. 85% of observations are for 1997, 6% are for 1998, 4% are for 1996, 3% are for 1999, and 2% are for 1995. This explains why we included time dummies in the gravity regressions even though, in an intertemporal sense, the estimations are strictly cross-section.

The sample countries underlying Tables 4 and 5 are: Austria, Bolivia, Canada, Chile, China, Cameroon, Colombia, Costa Rica, Denmark, Ecuador, Egypt, Spain, Ethiopia, Finland, France, United Kingdom, Greece, Guatemala, Honduras, Hungary, Indonesia, India, Ireland, Italy, Japan, Korea, Sri Lanka, Latvia, Morocco, Moldova, Mexico, Malaysia, Netherlands, Norway, Nepal, New Zealand, Philippines, Poland, Portugal, Sweden, Turkey, United States, Venezuela and South Africa.

For definitions of *LOGDIST*, *COLONYDUM*, *SAMECTRYDUM*, *OFFLANGDUM*, and *SPKLANGDUM*, see www.cepii.fr/anglaisgraph/bdd/distances.htm.

Tariff data are from CEPII's MAcMap database (see www.cepii.fr/anglaisgraph/bdd/macmap.htm). Simple averages of the bilateral *ad valorem* tariffs at HS6 level were used to aggregate up to our 17 sectors. Data on bilateral non-tariff barriers are from CEPII's trade and production database (see www.cepii.fr/anglaisgraph/bdd/TradeProd.htm). They are computed as frequencies of basic HS6 tariff lines that are affected by various classes of non-tariff barriers identified in Unctad's TRAINS data base (categorised as "threat", "price", "quantity" and "quality" NTBs). These frequencies are aggregated up to our 17 sectors from 28 ISIC industries using simple averages.