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## **A TEST OF TRADE THEORIES WHEN EXPENDITURE IS HOME BIASED**

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and  
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**DT-GREQAM**

# A TEST OF TRADE THEORIES WHEN EXPENDITURE IS HOME BIASED<sup>\*</sup>

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## Abstract

We develop a criterion to distinguish two dominant paradigms of international trade theory: homogeneous-goods perfectly competitive models, and differentiated-goods 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 differentiated-goods sectors (the "home-bias effect"), while no such relationship exists in homogeneous-goods sectors. This discriminating criterion turns out to be robust to a number of generalisations of the baseline model. Our empirical results, based on a world-wide cross-country data set, suggest that the differentiated-goods model fits particularly well for the machinery, precision engineering and transport equipment industries, which account for some 40 percent of sample manufacturing output.

**JEL classification:** F1, R3

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

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## 1. Introduction

International trade theory is dominated by two major paradigms. One paradigm belongs to the neoclassical world with homogeneous goods and perfectly competitive product markets (PC). The second paradigm, frequently referred to as the “new trade theory”, rests on the assumption of differentiated goods and 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.

We develop a discriminating criterion that is amenable to empirical estimation. The criterion rests on the assumption that demand is home biased, in the sense that some buyers perceive home-produced goods *ipso facto* to be different from imports. It posits that the home bias influences international specialisation in sectors that are characterised by product differentiation associated with monopolistic competition and increasing returns (MC), while such bias is inconsequential for the location of sectors characterised by homogeneous goods associated with perfect competition and constant returns (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.

In a second step, 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 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. One prominent approach has enlisted the excellent empirical performance of the gravity equation in support of the MC paradigm.<sup>1</sup> It has indeed been shown that the gravity equation has a straightforward theoretical counterpart in the 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 the MC paradigm.

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<sup>1</sup> An earlier literature had focused on intra-industry trade as evidence of the importance of the MC paradigm (see Greenaway and Milner, 1986; and, for a critical appraisal, Leamer and Levinsohn, 1995). Since intra-industry trade was generally associated with 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 PC framework.

Another, more direct, 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 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 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 PC sectors (with product differentiation by country of origin) the HME is amplified by trade costs, whilst in MC 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 PC 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. 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.

Related 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 PC sector being sufficiently smaller than those of the 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 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, 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 distinguish among alternative trade models.

Our study, building on Trionfetti (2001), 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 aggregate 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. Evidence that home-bias relates to the location of production and not to the nationality of firms is found by Evans (2001). The assumption of bias in favour of goods that are “made in” the home country therefore rests on solid empirical ground.



Finally, it might be useful at this point to clarify the principal difference between our approach and that chosen by Head and Ries (2001). Their analysis pits a model of product differentiation by firm and firm-level increasing returns (implying MC) against a model of product differentiation by country of origin with constant returns at the firm level (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 (MC) against a model of no product differentiation by firm (PC). The support found by Head and Ries (2001) for the differentiation-by-country-of-origin model offers confirmation that home bias is a pervasive phenomenon.

### 3. Theory: Derivation of a Discriminating Criterion

A suitable model for our analysis needs to accommodate both the PC and the 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  $S \in \{X, Y\}$ . In the next four sub-sections, we first describe the demand side, the supply side and the equilibrium of the model, and then we derive our discriminating criterion.

#### 3.1 Demand

Preferences feature love for variety, represented by the traditional nested CES-Cobb-Douglas utility function. We extend the standard model by assuming that demand is home biased, and 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.<sup>2</sup> Furthermore, we assume that there are two types  $t \in \{u, b\}$  of buyers (but one could easily extend the model to a continuum): the “unbiased” type ( $u$ ) and the “home-biased” type ( $b$ ), with  $\mu^i$  and  $(1 - \mu^i)$  denoting, respectively, the population share of type- $b$  and type- $u$  buyers in country  $i$ . 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.<sup>3</sup>

We allow home biased consumers to have different degrees of home bias for the two goods. We model the degree of 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 from country  $i$  are not home biased in sector  $S$ . As  $\delta_S^{it}$  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 degree of home bias, but other ways are perfectly conceivable.<sup>4</sup> For simplicity, we assume that biased and unbiased consumers have the same expenditure shares.

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

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

and with CES sub-utilities

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<sup>2</sup> See, e.g., Baldwin (1984), Miyagiwa (1991), Trefler (1995), Hummels (2001), and Head and Ries (2001).

<sup>3</sup> 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 preferences whereby they reserve some share of expenditure for home-produced goods, while the remaining expenditure is allocated without country bias.

<sup>4</sup> One alternative representation would be through a parameter inserted inside the CES aggregator, as in Head and Ries (2001), Hummels (2001) and Combes, Lafourcade and Mayer (2005). We discuss this alternative in Section 4 and show that the salient results of our model remain unchanged.

$$X = \left( \int_{k \in \Omega^A} c_k^{(\sigma-1)/\sigma} dk + \int_{k \in \Omega^B} c_k^{(\sigma-1)/\sigma} dk \right)^{\sigma/(\sigma-1)}, \quad X_i = \left( \int_{k \in \Omega^i} c_k^{(\sigma-1)/\sigma} dk \right)^{\sigma/(\sigma-1)},$$

where  $X$  is the CES aggregate of all varieties,  $X_i$  is the CES aggregate of varieties produced in country  $i$ ,  $Y_i$  refers to output of  $Y$  produced in country  $i$ ,  $c_k$  is consumption of variety  $k$ ,  $\Omega^i$  represents the set of varieties produced in country  $i$ , and  $\sigma > 1$  is the elasticity of substitution between any two varieties of  $X$ . 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  $\delta$  parameter.

Utility maximisation under the budget constraint and aggregation over individuals yields the demand functions for individual domestic and foreign varieties. Let  $x_{ii}^d$  denote the former and  $x_{ji}^d$  the latter, where the first subscript stands for the country of production and the second subscript stands for the destination country. Denoting aggregate expenditure by buyers of type  $t$  in country  $i$  on sector  $S$  with  $E_S^{it}$ , we have that  $E_S^{ib} = \alpha_S^i \mu^i w L^i$  and  $E_S^{iu} = \alpha_S^i (1 - \mu^i) w L^i$ . Furthermore, denoting aggregate expenditure by buyers of all types in country  $i$  on sector  $S$  with  $E_S^i$ , we have that  $E_S^i = E_S^{iu} + E_S^{ib} = \alpha_S^i w L^i$ . Finally, the parameters  $\mu^i$  and  $\delta_S^i$  always appear multiplicatively in the demand functions. We can thus define the parameter  $h_S^i \equiv \mu^i \delta_S^i$ , aggregating the two components of home bias. We refer to this parameter simply as “the home bias” of country  $i$  in sector  $S$ . The demand functions resulting from profit maximisation can then be written as:

$$x_{ii}^d = \frac{P_X^{-\sigma}}{(P_X^i)^{1-\sigma}} (1 - h_X^i) E_X^i + \frac{P_X^{-\sigma}}{(P_X^{hi})^{1-\sigma}} h_X^i E_X^i, \quad i = A, B.$$

$$x_{ji}^d = \frac{\phi p_X^{1-\sigma}}{(P_X^i)^{1-\sigma}} (1 - h_X^i) E_X^i, \quad i = A, B. \quad j \neq i.$$

Demand for any domestic variety ( $x_{ii}^d$ ) is composed of two elements: the first summand represents demand from unbiased buyers, and the second summand represents demand from home-biased buyers. Demand for any foreign ( $x_{ji}^d$ ) variety has only one component since only the unbiased buyers demand foreign varieties. In the demand functions,  $P_X^i$  denotes the perfect price index associated with unbiased demand of residents of country  $i$  and is given by the following CES:

$$P_X^i = \left[ \int_{k \in n_i} (p_X(k))^{1-\sigma} dk + \int_{k \in n_j} \left( \frac{p_X(k)}{\tau} \right)^{1-\sigma} dk \right]^{\frac{1}{1-\sigma}} = p_X [n^i + \phi n^j]^{\frac{1}{1-\sigma}},$$

where  $\phi \equiv \tau^{\sigma-1}$  is the “phi-ness” representation of trade openness (Baldwin *et al.*, 2003), and  $n^i$  represents the number of varieties produced in country  $i$ . The second equality makes use of the fact that, in equilibrium, all varieties have the same factory-gate price as will become clear below. The term  $P_X^{hi}$  denotes the perfect price index associated with home-biased demand. By definition, the demand of home-biased buyers falls only on domestic varieties; therefore  $P_X^{hi}$  is an aggregate of prices of domestic varieties only, and it takes the following CES form:

$$P_X^{hi} = \left[ \int_{k \in \Omega_i} (p_X(k))^{1-\sigma} dk \right]^{\frac{1}{1-\sigma}} = p_X [n^i]^{\frac{1}{1-\sigma}}.$$

Comparing the two price indices, we see that  $P_X^{hi} > P_X^i$ , since the latter contains a larger number of varieties than the former. We shall use this inequality in our discussion of the discriminating criterion.

### 3.2 Supply

We assume that the homogeneous good  $Y$  is produced by use of a constant-returns technology and traded costlessly in perfectly competitive markets.<sup>5</sup> Given that technologies are identical across countries and that there is perfect inter-sectoral mobility of labour, free trade in  $Y$  yields factor price equalisation.<sup>6</sup> Varieties of good  $X$  are subject to increasing returns at firm level 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 constant-returns sector is  $w\ell_Y$ , where  $w$  is the wage and  $\ell_Y$  is the input requirement per unit of output. Perfect intersectoral labour mobility ensures that the wage is the same in both sectors. Profit maximisation in the perfectly competitive  $Y$  sector implies that price equals marginal cost:

$$p_Y = w\ell_Y. \tag{1a}$$

Free trade assures commodity price equalisation in the  $Y$  sector. Therefore, equation (1a) applies to both countries.

Production of  $X$  entails a fixed cost  $wF$  and a constant marginal cost  $w m$ , where  $F$  is the fixed labour input and  $m$  is labour input per unit of output. Therefore, average cost in the  $X$  sector is

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<sup>5</sup> The discriminating criterion that we develop remains valid if we assume some positive trade costs in the PC sector (see Section 4.1).

<sup>6</sup> Another necessary condition for factor price equalisation is the “no-corner solution” assumption we introduce in Section 3.3.

$wF/x + wm$ , where  $x$  is firm output. A country- $i$  firm's profit is given by the difference between total revenue and total cost:  $p_X x_{ii}^d + p_X^* x_{ij}^d - wm(x_{ii}^d + x_{ij}^d / \tau) - wF$ , where  $p_X$  is the factory-gate price, and  $p_X^*$  is the price that any firm quotes on the foreign market.<sup>7</sup> The first-order conditions for profit maximisation under monopolistic competition yield the following optimal pricing rules, which apply to both countries:

$$p_X = \frac{\sigma}{\sigma-1} wm, \quad (1b)$$

$$p_X^* = \frac{\sigma}{\sigma-1} \frac{1}{\tau} wm. \quad (1c)$$

Free entry in both sectors implies that equilibrium profits are zero everywhere. In the perfectly competitive sector, given the linear technology and perfectly elastic demand, the zero profit condition (average cost equal average revenue) coincides with profit maximisation condition (marginal cost equal marginal revenue). Thus, equation (1a) represents the zero-profit condition in sector  $Y$ . The zero-profit condition in sector  $X$  in all countries can be written as follows:

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

The left-hand side of this equation represents average revenue and the right-hand side is average cost. This equation determines the output per firm independently of wages, since mark-ups are constant in this model. Solving (2) for  $x$  yields output per firm:  $\bar{x} = F(\sigma - 1)/m$ .

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<sup>7</sup> Firms have identical cost functions and face identical demand functions. Consequently, the optimal price and output levels are identical for all firms. There is no need, therefore, to use firm-specific subscripts.

Labour is in fixed supply in each country and demanded by both sectors. Let  $Y^i$  denote the quantity of  $Y$  produced in country  $i$ . The labour-market equilibrium conditions are then:

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

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

The first element on each left-hand side is the demand for labour from the  $Y$  sector, and the second element is the demand for labour from the  $X$  sector.

### 3.3 Equilibrium in Product Markets

Product-market equilibrium requires that demand equals supply for each sector and each variety. Using the demand functions, and recalling from equations (1) and (2) that equilibrium prices and firm output are equal across the two countries, the equilibrium equations are:

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

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

Equation (5) states the equilibrium condition for any  $X$  variety produced in country  $A$ , and equation (6) states the equilibrium condition for any  $X$  variety produced in country  $B$ . The left-hand side of each equation represents the value of supply of any variety and the right-hand side represents the corresponding value of demand. Recalling the demand functions, it is immediately clear that the first summand on the right-hand side of equation (5) represents demand from unbiased buyers of country  $A$  for any domestic variety, the second summand represents demand

from unbiased buyers of country  $B$  for any variety produced in  $A$ , and the last summand represents demand from biased buyers for any variety produced in  $A$ . The interpretation of the three terms on the right-hand side of equation (6) is analogous, *mutatis mutandis*.

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.<sup>8</sup> 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 home bias; i.e. that buyers’ home-biased expenditure does not exceed the hypothetical undistorted (by home bias) free-trade level of domestic production.<sup>9</sup>

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 left-hand and right-hand sides represent world supply and world demand for  $Y$ , respectively.

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<sup>8</sup> These conditions can be expressed in terms of exogenous variables as follows:  $h_X^i < \frac{\alpha_X (L^i + \phi L^j)}{(1 - \phi) L^i}$ , and

$h_Y^i < 1 - h_X^i$ .

<sup>9</sup> 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.



The model so far is standard except for the home bias. The system (1)-(6) is composed of eight independent equations and nine unknowns  $(p_X, p_X^*, p_Y, x, n^A, n^B, Y^A, Y^B, w)$ . Taking  $p_Y$  as the numéraire, the system is perfectly determined.<sup>10</sup>

### 3.4 *A Discriminating Criterion*

There is a difference between the PC sector and the 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 PC sectors but it affects international specialisation in the MC sectors. This is the essence of our discriminating criterion.

Consider, for instance, a shock to the home bias of country  $A$ ,  $dh_X^A > 0$ . This shock causes a change to the value of the right-hand side of (5) equal to  $\left( -\frac{p_X^{1-\sigma}}{(P_X^A)^{1-\sigma}} E_X^A + \frac{p_X^{1-\sigma}}{(P_X^{hA})^{1-\sigma}} E_X^A \right) dh_X^A > 0$ . The positive sign of this change is ensured by the fact that, as discussed above,  $P_X^{hi} > P_X^i$ . The right-hand side of (5) represents demand for any of the varieties produced in country  $A$ . The increase in the right-hand side is therefore an increase in the demand for any variety produced in country  $A$ . Since the output per firm and the price (i.e., the left-hand side of (5)) remain unchanged, this increase in demand must be absorbed via the entry of new firms in country  $A$ . The same applies, *mutatis mutandis*, to (6): the shock causes a change in the demand for any of the varieties

produced in  $B$  equal to  $-\left( \frac{\phi p_X^{1-\sigma}}{(P_X^A)^{1-\sigma}} E_X^A \right) dh_X^A < 0$ . This decline in demand for any variety

produced in country  $B$  will cause exit of some firms in country  $B$ . In the  $Y$  sector, instead, a

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<sup>10</sup> Note that, although home-biased consumers 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.

shock  $dh_Y^A > 0$  will have no effect on output. This neutrality is apparent by inspection of the right-hand side of (7) where the home bias  $h_Y^A$  cancels out from the equation. Therefore, the PC sector, unlike the MC sector, is unaffected by countries' relative home biases.

The economic mechanism can be explained as follows. Consider the  $X$  sector first. A euro of unbiased demand is spent on all varieties, whereas a euro of home-biased demand is spent only on domestic varieties. Therefore, an increase in the share of home-biased demand induces an increase in the expenditure on each domestic variety and a decline in the expenditure on each foreign variety. Consequently supply must increase in the domestic country and must decline in the foreign country. In conclusion, in order to recover the demand-supply equilibrium in all markets after a shock  $dh_X^i > 0$ , it is necessary that  $n^A$  increase and that  $n^B$  decrease. Thus, an increase in the own home bias, *ceteris paribus*, increases the own share of  $X$ -sector output. We call this the “home-bias effect” (HBE).<sup>11</sup>

Consider now the  $Y$  sector. The intuition for the neutrality of the home bias in this sector is simple: 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)). This result is akin to Baldwin's (1984) neutrality proposition, whereby home-biased government expenditure has no effect on international specialisation in a (PC) Heckscher-Ohlin model.<sup>12</sup>

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<sup>11</sup> This is equivalent to what we have termed the “pull effect” in the context of public procurement (Brühlhart and Trionfetti, 2004).

<sup>12</sup> Incidentally, we note here the difference between the effect of a tariff and that of the home bias. A tariff affects demand for the protected industry regardless of market structure, whereas the home bias affects demand only in the MC industry. A tariff on  $Y$ , for instance, would affect output in the  $Y$  industry of the importing country whereas home bias for  $Y$  does not.

In what follows, it will be useful to express the main variables in terms of the share of country  $A$ :

$$\eta_X \equiv n^A / (n^A + n^B), \quad \eta_Y \equiv Y^A / (Y^A + Y^B), \quad \lambda \equiv L^A / (L^A + L^B), \quad \text{and} \quad \varepsilon_S \equiv E_S^A / (E_S^A + E_S^B).$$

Differentiation of the system (1)-(6) yields the estimable equation:

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

where  $dh_S \equiv dh_S^A \equiv -dh_S^B > 0$  represents a change in relative home biases, and  $d\varepsilon_S > 0$  represents an idiosyncratic change in the size of expenditure.<sup>13</sup> The signs of the analytical expressions for the coefficients of (8) provide the formal basis of our discriminating criterion:

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

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

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

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

Hence:

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<sup>13</sup> 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.,  $dh_S^A = -dh_S^B$ . As discussed above, the HBE obtains also from independent shocks, such as  $dh_S^A > 0$  with  $dh_S^B = 0$ . 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 form of  $c_{2S}$  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.

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

In words: an increase in a country's degree of home bias for an MC good will result in an increase in the country's share of production of that good while an increase in the home bias for a PC good will have no impact on production. This discriminating criterion and its empirical implementation are the focus of our paper.<sup>14</sup>

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.<sup>15</sup>

#### 4. Robustness

The HBE-based discriminating criterion turns out to be robust to three important generalisations of the model: trade costs in both sectors or in none, inelastic sectoral factor supplies, and real-

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<sup>14</sup> 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 Sections 4.1 to 4.3, and/or if the home bias is modelled differently (see Section 4.4). Note also that, in the benchmark case, the coefficient  $c_{2Y}$  could be negative. The reason is that the larger is  $\alpha_X$ , the larger is the  $X$  sector compared to the  $Y$  sector. If  $X$  is large enough, its expansion due to  $d\lambda > 0$  may require so much labour moving into  $X$  that  $Y$  may shrink despite the increase in  $\lambda$ . Finally, in 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  $\eta_X = 1/2 + [(1+\phi)/(1-\phi)](\varepsilon_X - 1/2)$ .

<sup>15</sup> It is worth clarifying a possible source of confusion at this point. The HBE is different from the effect of an asymmetric tariff. An asymmetric tariff would affect specialisation in all sectors, by increasing the output of these sectors regardless of whether they are characterised by differentiated or homogeneous products. Asymmetric home bias, instead, affects the specialization only of sectors that feature product differentiation.

world geography with multiple non-equidistant countries. The HBE-based test is also robust to alternative ways of modelling the home bias.<sup>16</sup>

#### 4.1 Trade Costs

Davis (1998) has highlighted the importance of exploring the implications of trade costs in the PC sector. He showed that the necessary condition for the HME to disappear is that there be no trade in the  $Y$  sector: trade costs in  $Y$  must be high enough to eliminate all trade in  $Y$ .<sup>17</sup> In such a (rather extreme) case, the HBE-based criterion does not hold either. As long as trade costs in  $Y$  do not choke off all trade in  $Y$ , our criterion remains valid.

While it is straightforward that the home bias will continue to affect specialisation in the  $X$  sector as long as trade in  $Y$  is possible, we have to demonstrate that the home bias does not affect the  $Y$  sector even in the presence of trade costs in  $Y$ . This can be ascertained as follows. Assume that for each unit of  $Y$  shipped only a fraction  $\theta \in (0,1)$  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 four endogenous variables ( $p_Y^i, p_X^i, p_X^{*i}$ , and  $w^i$ ) and four independent equations must now be added to the system (1)-(6). The first three additional equations are equivalent to equations (1a)-(1c), which now become country specific. The fourth equation is akin to equation (7), which is no longer redundant. For clarity of

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<sup>16</sup> A further issue concerns the type of product differentiation. If a good is differentiated by country of origin (*à la* Armington) then the relationship between the share of output and of home-biased demand would be positive. This is consistent with the fact that the HBE test picks up product differentiation. The magnitude of the response of output to a given shock in home-biased demand is of lower in a setting featuring PC and Armington-type product differentiation than in a setting featuring MC without Armington-type product differentiation (apart from the distinction between home and foreign varieties). A formal derivation of this result can be provided on request.

<sup>17</sup> The exact condition concerning relative trade costs that eliminate trade in  $Y$ , and the consequences that this condition has for the HME, are explored in Laussel and Paul (2007) and Crozet and Trionfetti (2007).

exposition and without loss of generality we consider the case where country  $A$  is the exporter of  $Y$ .<sup>18</sup> Then, equation (7) becomes

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

Once again,  $h_Y^i$  cancels out. This means that here too the home bias is irrelevant for international specialisation in the 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 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), and to solve them for  $n^A$  and  $n^B$ . Then, using our notation in terms of shares, we have the solution  $\eta_X = h_X^A \lambda / [h_X^A \lambda + h_X^B (1 - \lambda)]$ . It is easily verified that  $\partial \eta_X / \partial h_X^A > 0$  even in the absence of trade costs in  $X$ . By contrast, if neither country is home biased ( $h_X^A = h_X^B = 0$ ) and there are no trade

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<sup>18</sup> To assume that there exists a  $Y$ -exporting country implies assuming that demand for any of the varieties of  $X$  produced in the  $Y$ -exporting country is lower than for any of the varieties produced in the  $Y$ -importing country. This lower demand could be the consequence, for instance, of the  $Y$ -exporting country being smaller or less home biased than the other country. Given lower demand, if wages (and prices) were not lower too then the  $Y$ -exporting country would run an overall trade deficit. The fall in wages (and prices) and the change in the pattern of specialisation restore trade balance. Whatever the reason that generates the initial specialisation pattern whereby there exists a  $Y$ -exporting country, our point is to demonstrate that the discriminating criterion remains valid in such a case.

costs, the HME cannot provide a discriminating criterion. The reason is that, in this case, the solution of (5)-(6) 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 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  $Y$  (and  $Y$  only) 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 MC sector demands more labour, wages will rise because of the decreasing marginal productivity of labour in the PC sector.<sup>19</sup> Let  $\gamma$  denote the labour share of total costs in the production of  $Y$ . Given the production technology in the PC sector, labour demand from sector  $Y$  is  $(\gamma / w^i)^{1/(1-\gamma)} \bar{K}$ , and factor rewards are

$$w^i = \gamma (k_Y^i / l_Y^i)^{1-\gamma} \quad \text{and} \quad r^i = (1-\gamma) (\gamma / w^i)^{\gamma/(1-\gamma)}.$$

After some normalisations, the labour-market clearing conditions can be written as:

$$n^A = \bar{L} - (\gamma / w^A)^{1/(1-\gamma)} \bar{K}, \quad (3')$$

$$n^B = \bar{L} - (\gamma / w^B)^{1/(1-\gamma)} \bar{K}, \quad (4')$$

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<sup>19</sup> 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 simplicity, we assume that  $K^A = K^B = \bar{K}$  and  $L^A = L^B = \bar{L}$ .

The equation system resulting from this model is identical to that of the benchmark model in Section 3, except for the fact that equations (3')-(4') replace equations (3)-(4) and that wages are now given by the expressions above. Importantly, equations (5), (6), and (7) remain unaltered.

Totally differentiating the resulting system around its symmetric equilibrium, we again obtain our testing equation (8), where the coefficients now are as follows: <sup>20</sup>

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

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

It is clear by inspection that  $c_{1X} > 0$  since the denominator is negative for any combination of parameter values. The positive sign of  $c_{1X}$  means that the country that is relatively more home biased will, *ceteris paribus*, tend to specialise in the production of  $X$ . In sector  $Y$ , instead, the home bias is neutral since the parameters  $h_Y^i$  cancel out. These results confirm the validity of the HBE criterion even in the presence of imperfectly elastic labour supply.

The economic logic of these results is as follows. For sector  $Y$ , the rationale is as in the benchmark model: sector  $Y$  is insensitive to home-biased demand because any increase in demand for domestic  $Y$  caused by an increase in home-biased demand is immediately crowded out by a corresponding decrease in demand for domestic  $Y$  by unbiased consumers, as the latter are indifferent between domestic and foreign  $Y$ . The logic for the  $X$  sector is as follows. A shock

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<sup>20</sup> 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 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.



$dh_X^A > 0$  causes an increase in demand for the varieties produced in country  $A$  and a decline in demand for varieties produced in  $B$  like in the benchmark model and for the same reason. In response to the excess demand in  $A$  and to the excess supply in  $B$  the value of output must increase in  $A$  and decline in  $B$ . The value of output can change in three possible ways (from the viewpoint of country  $A$ ): (1) if labour supply is perfectly elastic, the excess demand is entirely absorbed by an increase in output; (2) if labour supply is imperfectly inelastic, the excess demand is absorbed partially by wages and partially by an increase in output; and (3) if labour supply is perfectly inelastic, the excess demand is entirely absorbed by an increase in wages. The first case is considered in Section 3.4. The second case is the one analysed in this sub-section. 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 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*

The HME derived in the two-country model extends 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 problem empirically, and an extended theoretical and empirical analysis is provided by Behrens *et al.* (2004).

We can show that our discriminating criterion remains valid in an asymmetric multi-country world. In the multi-country model, the HME may not exist, because the adding-up constraint on the distribution of shocks to countries' expenditure shares can result in a less than one-for-one relation between expenditure shares and output shares in the MC sector.<sup>21</sup>

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 consider a shock  $dh^i > 0$ , *ceteris paribus*. Such a shock will cause an increase in the share of output of country  $i$  for the reasons discussed in section 3.1 above. As noted there, the economic logic of the discriminating criterion is independent of the number of countries and of the structure of bilateral trade costs.

To see this more formally we write the market-clearing conditions for an arbitrary number of countries,  $M$ . These are:

$$p_X x = \sum_{j=1}^M \frac{\phi_{ij} p_X^{1-\sigma}}{(p_X^j)^{1-\sigma}} (1-h_X^j) E_X^j + \frac{\phi_{ii} p_X^{1-\sigma}}{(p_X^i)^{1-\sigma}} h_X^i E_X^i, \quad i = 1, \dots, M. \quad (5')$$

Non-equidistance implies that  $\phi_{ij} \neq \phi_{ik}$  for any  $j \neq k$ . We retain the assumption made in the literature that  $\phi_{ii} = 1$ , although this is not essential. Price indices associated with unbiased

demand contain varieties produced in all countries:  $P_X^i = p_X \left[ \sum_{j=1}^M \phi_{ij} n^j \right]^{\frac{1}{1-\sigma}}$ . Price indices associated

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<sup>21</sup> 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 MC sectors in such a general model. We consider that relationship as a complement to the HBE test in our empirical analysis.

with home-biased demand, however, contain only domestic varieties and are identical to those of

the two-country model:  $P_X^{hi} = p_X [n^i]^{\frac{1}{1-\sigma}}$ . Note that  $P_X^{hi} > P_X^i$ , since the former contains a smaller number of varieties than the latter. Now consider the shock  $dh_X^i > 0$ . Inspection of equation (5') shows that this shock will cause an increase in the demand for any variety

produced in country  $i$  by  $\left( -\frac{p_X^{1-\sigma}}{(P_X^i)^{1-\sigma}} E_X^i + \frac{p_X^{1-\sigma}}{(P_X^{hi})^{1-\sigma}} E_X^i \right) dh_X^i > 0$ . The positive sign is ensured by

the fact that  $P_X^{hi} > P_X^i$ . Inspection of (5') also shows that the shock  $dh_X^i > 0$  will also cause a

decline in the demand equal to  $\left( \frac{\phi_{ij} p_X^{1-\sigma}}{(P_X^i)^{1-\sigma}} E_X^i \right) dh_X^i$  for any of the varieties produced in any other

country  $j \neq i$ . In order to restore the equilibrium in all markets it is necessary that supply increases in country  $i$  and declines in all other countries:  $n^i$  must increase and  $n^j$  must decline for all  $j \neq i$ , and country  $i$ 's share of output must increase. In conclusion, country  $i$ 's share of output is related positively to its own degree of home bias and negatively to the degree of home bias of all other countries. This holds regardless of the number of countries and of the geographic allocation of bilateral trade costs. The HBE-based criterion, therefore, is robust to a generalisation of the model that allows for multiple non-equidistant countries.

#### 4.4. An alternative way of modelling the home bias.

An alternative and reasonable way of modelling the home bias is to insert different weights for the set of domestic and foreign varieties in the CES sub-utility function. The HBE remains valid when the home bias is modelled 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  $\psi^i$  and  $\omega^i$  are weights, and  $h_X^i \equiv (\psi^i / \omega^i)^\sigma \in (0, \infty)$  represents the home bias. This is the approach chosen by Head and Ries (2001), Hummels (2001), and Combes *et al.* (2005). Expressing the equilibrium equations (5) and (6) in terms of  $\eta_X$  and  $\lambda$ , 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:

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

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

The coefficient  $c_{1X}$  is positive, which confirms the validity of the discriminating criterion based on the HBE.<sup>22</sup>

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<sup>22</sup> The mathematical derivation of these results is based on standard utility maximisation and available from the authors - we omit them here to save space. 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  $\sqrt{h_X^A h_X^B} > \phi > (h_X^A + h_X^B)/2$ .

We can mention a further robustness check. The discriminating criterion does not depend on the use of a utility function where the upper tier is Cobb-Douglas and the lower tier is CES. It is easy to show that our criterion remains valid if both the upper tier and the lower tier of the utility function are CES. The logic of this result is the same as that discussed above. A shock to  $h_X^i$  affects demand for goods produced in country  $i$  and country  $j$  in opposite directions, whereas a shock to  $h_Y^i$  leaves demand for output of  $Y$  in either country unchanged.<sup>23</sup>

#### 4.5. Conclusion to this section

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, (c) to the assumption of  $M > 2$  asymmetrically spaced countries, and (d) to alternative ways of modelling the home bias.

## 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).

### 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

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<sup>23</sup> A formal proof is available from the authors.

exercise should not prejudice our inference in stage two.<sup>24</sup> Specifically, we estimate the following regression equation:

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

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

|                |   |  |
|----------------|---|--|
| $LOGIM_S^{ij}$ | = | 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 $j$ 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,                                 |
| $SAMECTRYDUM$  | = | 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 $j$ share a common official language, and zero otherwise,  |

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<sup>24</sup> 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.

$SPKLANGDUM =$  dummy equal to one if  $i$  and  $j$  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

$\varepsilon =$  a potentially heteroskedastic stochastic error term.

The object of our interest is  $\beta_1$ , 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, legal 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 effects of trade barriers 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.

Three practical difficulties remain. First, one has to find a measure of “trade within countries”, second, the distance variable has to be defined for intra-country trade, and, third, an estimator needs to be chosen based on appropriate assumptions about the error term  $\varepsilon$ .

Following Wei (1996), we define trade within countries as output minus exports.<sup>25</sup> 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 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 (*DIST*).<sup>26</sup> 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 ( $DIST\_DISC^{ii} = \ln(0.67\sqrt{Area^i / \pi})$ ).

We estimate the gravity equation using the Poisson pseudo-maximum-likelihood approach as suggested by Santos Silva and Tenreyro (2004). This estimation method addresses the problem of conditional heteroskedasticity inherent in standard least-squares regression of log-linearised

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<sup>25</sup> Hence,  $LOGIM^i = \log(Output^i - \sum_j Exports^{ij})$ .

<sup>26</sup> See [www.cepii.fr/anglaisgraph/bdd/distances.htm](http://www.cepii.fr/anglaisgraph/bdd/distances.htm).



versions of multiplicative models: since the mean of the log of the original (multiplicative, mean-independent) error term depends on the variance of the original error, the log error will be correlated with all regressors that are correlated with the variance of the original error. In gravity models this is clearly relevant, as the variance of trade flows generally increases with their expected size and thus with all regressors that predict the size of trade flows. Since the Poisson estimator implies that the conditional variance is proportional to the conditional mean of the error, it offers a suitable approach in this context. Poisson estimation holds the further attraction that it allows consideration of zero-trade flows as well as non-zero observations.<sup>27</sup>

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

We begin by running equation (11) on the entire data set. The baseline results, using the CEPII distance measure, are shown in Table 1, column 1. As usual, the gravity model provides a good fit to the data, and all statistically significant coefficients have the expected signs. Our estimates suggest significant home bias: on average, purchases from national sources are predicted to be 3.2 ( $=e^{1.15}$ ) times larger than purchases from sources the same distance away and unaffected by tariffs, NTBs or language barriers but located in a different country. This estimate may seem large, but it 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 2 of Table 1), the mean estimated home bias exactly doubles, to a factor of 6.4 ( $=e^{1.85}$ ). Given the unavoidably imprecise nature of distance measures, one must

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<sup>27</sup> Note that to estimate the parameters of equation (11) via Poisson, the dependent variable considered is bilateral imports (and not their log), while the non-binary right-hand side variables enter in logs.

therefore 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.<sup>28</sup> Thus, we obtain individual home-bias estimates per country-industry pair, which we call *IDIOBIAS* and which provide the key ingredient to our testing equation.<sup>29</sup>

## 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:

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

where  $i$  again denotes countries,  $S$  denotes industries, and:

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<sup>28</sup> To be precise, we estimate equation (11) separately for each sector (thus dropping the industry fixed effects), retaining the *DIST* measure of bilateral distance (as in column 1 of Table 1), and we interact *HOMEDUM* with the importer fixed effects.

<sup>29</sup> Estimated coefficients on *HOMEDUM* are relative to an omitted base country. Since our panel is not perfectly balanced as country coverage varies across industries, the benchmark is not identical across industries and intersectoral comparisons of estimated home biases are not feasible. While, given the dependence on the omitted reference country, absolute values of mean country coefficients do not have a direct interpretation either, their ranking carries relevant information (which provides the basis for our empirical identification of HBEs). We find the strongest average home biases for France, followed by Venezuela and Egypt. The smallest average home biases are estimated for Sri Lanka, followed by Honduras and the Netherlands.

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

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

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

According to our discriminating criterion, industries with estimated  $c_{1S}$  of zero conform with the PC model, whereas industries with positive estimated  $c_{1S}$  conform with the 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). Our model features more goods than factors. Hence, factor endowments need not appear in the reduced-form testing equation (8).<sup>30</sup>

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

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<sup>30</sup> 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), any bias due to omitted endowment variables should be limited.

definition of *NetExpenditure* includes expenditure from sources that use the industry's output for final consumption, and exclude expenditure from those sources that use the output as intermediate inputs.<sup>31</sup>

Third,  $\nu_s^i$  is likely to be heteroskedastic, as the variance of errors may well be positively correlated with the size of countries.<sup>32</sup> Our significance tests are therefore based on heteroskedasticity-consistent standard errors in order to minimise the risk of wrongly attributing sectors to the 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 replacement, 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 this 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 in

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<sup>31</sup> See the Appendix 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 international trading opportunities. 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 evolution over time 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 (2001) 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  $c_1$ .

<sup>32</sup> A Breusch-Pagan test on the pooled model strongly rejects the null of constant error variance.

all of the specification we estimate. Hence, we report OLS coefficients but base hypothesis tests on bootstrap error distributions.<sup>33</sup>

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 such a generalisation would be erroneous for the relationship between expenditure shares and output shares would be erroneous. Instead, they point out that, in an asymmetric  $M$ -country 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)

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

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

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

(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 MC provides the appropriate model. We therefore estimate both versions of the testing equation.

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<sup>33</sup> 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 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 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.<sup>34</sup> In both cases, the estimated coefficients on *IDIOBIAS* are positive, but statistical significance is only found when we include country fixed effects. Taken at face value, this is contradictory evidence: while the pooled runs are consistent with the PC model and reject the MC model, the estimations featuring country fixed effects are consistent with the MC model and reject the PC model.<sup>35</sup> The same patterns are found in regressions (3) and (4), where we use spatially filtered expenditure shares. The coefficients on spatially filtered expenditure shares being statistically significantly smaller than unity in those runs, however, consistently rejects the 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, suggesting consistently insignificant HBEs and thus favouring the PC model. Purging expenditure measures of intermediate expenditure in order to avoid simultaneity bias is therefore confirmed as an important component of our estimation strategy.

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<sup>34</sup> Industry fixed effects are redundant, because *IDIOBIAS* represents deviations from industry means, and the other variables represent industry shares.

<sup>35</sup> The estimations with country fixed effects should be interpreted with caution, given the limited intersectoral comparability of *IDIOBIAS* (see footnote 26).

The results of Table 2 point towards a rejection of the MC model, but they paint an inconsistent picture. By imposing equal coefficients across sectors, however, 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 0.55 and 0.99. Coefficient estimates on *IDIOBIAS* are in the expected positive or insignificant range for all industries.

At the 90-percent confidence level, we find that seven of the 17 sectors conform with the MC paradigm.<sup>36</sup> The allocation of sectors looks plausible, as it comprises all the machinery and engineering industries (ISIC 382-384) plus textiles, clothing and footwear, 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 MC prediction account for some 45 percent of sample output.

Finally, Table 4 reports the corresponding results for the specification that includes spatially filtered expenditure shares, *EXPENDISHARE\_BLOT*. This equation fits the data less well, with

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<sup>36</sup> Note that statistical significance can be compatible with unusually large standard errors since bootstrap confidence intervals need not be symmetric.

$R^2$ s in the range 0.05 to 0.45, which is hardly surprising given the inevitably increased scope for measurement error from the constructions of *EXPENDISHARE\_BLOT*. Our estimated coefficients, however, again look largely plausible. 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 90-percent criterion for our test based on *IDIOBIAS*, six sectors conform with the MC prediction, “petroleum products” being the only (and rather unlikely) sector exhibiting the HBE here but not in the estimations reported in Table 3. It is interesting to note that the one-to-one relationship predicted to hold between *OUTPUTSHARE* and *EXPENDISHARE\_BLOT* by the MC model of Behrens *et al.* (2004) seems to fit best for the engineering sectors (ISIC 382-384) and for “other manufactures”, which also have statistically positive coefficients on *IDIOBIAS* throughout. While the coefficients on *EXPENDISHARE\_BLOT* range from 0.40 to 0.59 for all other industries, they lie between 0.76 and 0.87, and are not statistically significantly different from unity, for ISIC sectors 382-390. Together with our findings of Table 3, we therefore conclude that these four sectors, accounting for some 40 percent of sample output, fit the predictions of the MC model best.

Despite imperfect overlap of sector classifications, we can make a rough comparison of our set of MC sectors to those identified in previous studies. Davis and Weinstein (1999) found significant HMEs for three of these four sectors (the exception being “other manufactures”) . All four sectors also featured among the set of MC industries identified by Trionfetti (2001).<sup>37</sup> The

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<sup>37</sup> Davis and Weinstein (1999) and Trionfetti (2001) have identified larger sets of industries as conforming with the MC model than ours, but these sets included some rather implausible candidates such as pulp and paper (by both Davis and Weinstein, 1999, and Trionfetti, 2001), and rubber and plastic (by Trionfetti, 2001). Based on international rather than regional data, Davis and Weinstein (2003) found only one three-digit sector (textiles) as exhibiting a significant HME, and they considered it “somewhat disappointing that sectors like electrical machinery and transportation equipment do not have point estimates that exceed unity”.



four sectors for which we find correspondence with the MC model can therefore be viewed as representing the core set of MC-consistent industries emerging from this literature.

## **6. Conclusions**

We develop and apply an empirical test to distinguish two paradigms of international trade theory: a model featuring homogeneous products and perfect competition (PC), and a model featuring differentiated products and monopolistic competition (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 MC paradigm, but not if it is characterised by 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 and use these estimates to apply our test separately for 17 manufacturing industries. The results suggest that the engineering industries (machinery, precision engineering and transport equipment), plus “other manufacturing”, which together account for some 40 percent of manufacturing output value in our sample, best conform with the predictions of the 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 Sample**<sup>1</sup>  
(dependent variable = value of bilateral imports)

|                        | (1)                 | (2)                 |
|------------------------|---------------------|---------------------|
| <i>HOMEDUM</i>         | 1.152**<br>(0.239)  | 1.851**<br>(0.212)  |
| <i>DIST</i>            | -0.331**<br>(0.046) |                     |
| <i>DIST_DISC</i>       |                     | -0.371**<br>(0.052) |
| <i>ADJACENCYDUM</i>    | 1.004**<br>(0.144)  | 0.828**<br>(0.157)  |
| <i>TARIFF</i>          | -0.215**<br>(0.026) | -0.207**<br>(0.026) |
| <i>NTB</i>             | 0.281<br>(0.319)    | 0.260<br>(0.317)    |
| <i>COLONYDUM</i>       | 0.232*<br>(0.116)   | 0.294*<br>(0.116)   |
| <i>SAMECOUNTRYDUM</i>  | 0.125<br>(0.223)    | 0.329<br>(0.216)    |
| <i>OFFLANGDUM</i>      | -0.132<br>(0.170)   | -0.364*<br>(0.163)  |
| <i>SPOKLANGDUM</i>     | 0.429**<br>(0.144)  | 0.461**<br>(0.144)  |
| Importer fixed effects | yes                 | yes                 |
| Exporter fixed effects | yes                 | yes                 |
| Industry fixed effects | yes                 | yes                 |
| Year fixed effects     | yes                 | yes                 |
| Observations           | 112,010             | 112,010             |
| Pseudo R <sup>2</sup>  | 0.914               | 0.914               |

<sup>1</sup> Poisson estimation. 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 99% (95%) confidence level, based on heteroskedasticity-consistent standard errors.

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

|                           | (1)              | (2)                            | (3)                            | (4)                            | (5)                            | (6)              |
|---------------------------|------------------|--------------------------------|--------------------------------|--------------------------------|--------------------------------|------------------|
| <i>IDIOBIAS</i>           | 0.059<br>(0.081) | 0.241**<br>(0.085)             | 0.265<br>(0.210)               | 0.279**<br>(0.087)             | 0.022<br>(0.042)               | 0.045<br>(0.043) |
| <i>EXPENDISHARE</i>       | 0.948<br>(0.044) | 0.423 <sup>##</sup><br>(0.110) |                                |                                |                                |                  |
| <i>EXPENDISHARE_BLOT</i>  |                  |                                | 0.202 <sup>##</sup><br>(0.060) | 0.129 <sup>##</sup><br>(0.076) |                                |                  |
| <i>EXPENDISHARE_GROSS</i> |                  |                                |                                |                                | 0.938 <sup>##</sup><br>(0.022) | 1.104<br>(0.116) |
| Country fixed effects     | no               | yes                            | no                             | yes                            | no                             | yes              |
| <i>N</i>                  | 623              | 623                            | 623                            | 623                            | 625                            | 625              |
| <i>R</i> <sup>2</sup>     | 0.867            | 0.930                          | 0.044                          | 0.918                          | 0.969                          | 0.975            |

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

**TABLE 3: Industry-by-Industry Estimation of the Discriminating Criterion, Two-Country Model <sup>1</sup>**  
(dependent variable = *OUTPUTSHARE*)

| ISIC<br>(Rev. 2) | Description                                    | <i>IDIIBIAS</i>                 |                  | <i>EXPENDISHARE</i>             |                   | R <sup>2</sup> | N  | Size share <sup>2</sup> |
|------------------|--|---------------------------------|------------------|---------------------------------|-------------------|----------------|----|-------------------------|
|                  |  | <i>coefficient</i> <sup>3</sup> | <i>std error</i> | <i>coefficient</i> <sup>4</sup> | <i>std. error</i> |                |    |                         |
| 311              | Food products                                  | -0.611                          | 0.592            | 0.955                           | 0.083             | 0.985          | 34 | 0.109                   |
| 313/4            | Beverages, tobacco                             | 0.685                           | 0.632            | 0.921                           | 0.037             | 0.997          | 37 | 0.032                   |
| 321              | Textiles                                       | 0.292**                         | 0.189            | 1.090                           | 0.009             | 0.957          | 37 | 0.033                   |
| 322              | Clothing                                       | 0.298**                         | 0.209            | 0.753                           | 0.210             | 0.951          | 30 | 0.014                   |
| 323/4            | Leather, footwear                              | 0.548**                         | 0.511            | 0.556                           | 0.455             | 0.552          | 35 | 0.007                   |
| 331/2            | Wood products                                  | -0.016                          | 0.968            | 0.942                           | 0.140             | 0.975          | 31 | 0.030                   |
| 341/2            | Paper products, publishing                     | -0.132                          | 1.885            | 0.971                           | 0.407             | 0.921          | 37 | 0.073                   |
| 351/2/5/6        | Chemicals                                      | 0.096                           | 1.928            | 1.199                           | 0.259             | 0.942          | 37 | 0.129                   |
| 353/4            | Petroleum products                             | -0.178                          | 0.289            | 1.139                           | 0.163             | 0.977          | 36 | 0.034                   |
| 361/2/9          | Non-metall. mineral prods                      | -0.186                          | 1.271            | 0.904                           | 0.292             | 0.720          | 37 | 0.033                   |
| 371              | Iron, steel                                    | -0.500                          | 0.640            | 1.245                           | 0.399             | 0.692          | 37 | 0.037                   |
| 372              | Non-ferrous metals                             | 0.120                           | 0.207            | 1.268                           | 0.246             | 0.909          | 36 | 0.019                   |
| 381              | Fabricated metal products                      | 0.540                           | 1.010            | 1.171                           | 0.418             | 0.769          | 38 | 0.054                   |
| 382              | Non-electrical machinery                       | 0.135*                          | 0.155            | 1.089                           | 0.256             | 0.946          | 37 | 0.115                   |
| 383/5            | Electrical machinery,<br>precision engineering | 0.148*                          | 0.178            | 1.064                           | 0.303             | 0.922          | 37 | 0.137                   |
| 384              | Transport equipment                            | 0.267**                         | 0.335            | 0.906                           | 0.245             | 0.941          | 37 | 0.130                   |
| 390              | Other manufactures                             | 0.455**                         | 0.265            | 0.771                           | 0.574             | 0.865          | 33 | 0.012                   |

<sup>1</sup> Estimation with OLS; bootstrap standard errors (5,000 iterations with replacement). Constant term included in all regressions but not reported.

Coefficients and standard errors reported with respect to (*IDIIBIAS* \* 1,000).

<sup>2</sup> Share in total output of sample countries.

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

<sup>4</sup> ## (#) denotes rejection of H0: coeff. = 1 at 95% (90%) confidence level (one-tail test), based on bias-corrected bootstrap confidence intervals.

**TABLE 4: Industry-by-Industry Estimation of the Discriminating Criterion , *M*-Country Model <sup>1</sup>**  
(bootstrapped OLS; dependent variable = *OUTPUTSHARE*)

| ISIC<br>(Rev. 2) | Description                                    | <i>IDIOBIAS</i>                 |                  | <i>EXPENDISHARE_BLOT</i> <sup>2</sup> |                   | <b>R<sup>2</sup></b> | <b>N</b> |
|------------------|--|---------------------------------|------------------|---------------------------------------|-------------------|----------------------|----------|
|                  |  | <i>coefficient</i> <sup>3</sup> | <i>std error</i> | <i>coefficient</i> <sup>4</sup>       | <i>std. error</i> |                      |          |
| 311              | Food products                                  | -13.760**                       | 11.465           | 0.414 <sup>##</sup>                   | 0.539             | 0.232                | 34       |
| 313/4            | Beverages, tobacco                             | -1.730                          | 5.737            | 0.577 <sup>##</sup>                   | 0.410             | 0.341                | 37       |
| 321              | Textiles                                       | 0.108                           | 2.175            | 0.532 <sup>#</sup>                    | 0.276             | 0.169                | 37       |
| 322              | Clothing                                       | 0.699**                         | 0.693            | 0.450 <sup>##</sup>                   | 0.300             | 0.115                | 31       |
| 323/4            | Leather, footwear                              | 0.325                           | 0.635            | 0.754                                 | 0.335             | 0.451                | 36       |
| 331/2            | Wood products                                  | -8.411**                        | 8.099            | 0.435 <sup>#</sup>                    | 0.493             | 0.148                | 33       |
| 341/2            | Paper products, publishing                     | -0.315                          | 11.748           | 0.452                                 | 0.851             | 0.048                | 37       |
| 351/2/5/6        | Chemicals                                      | -0.407                          | 11.998           | 0.511                                 | 0.814             | 0.075                | 37       |
| 353/4            | Petroleum products                             | 1.013*                          | 0.900            | 0.441 <sup>##</sup>                   | 0.601             | 0.118                | 35       |
| 361/2/9          | Non-metall. mineral prod                       | -0.655                          | 3.807            | 0.512                                 | 0.447             | 0.134                | 37       |
| 371              | Iron, steel                                    | 0.644                           | 8.074            | 0.549                                 | 0.524             | 0.162                | 37       |
| 372              | Non-ferrous metals                             | 0.227                           | 0.593            | 0.403 <sup>##</sup>                   | 0.529             | 0.078                | 36       |
| 381              | Fabricated metal products                      | 1.018                           | 6.295            | 0.587                                 | 0.648             | 0.081                | 37       |
| 382              | Non-electrical machinery                       | 0.843**                         | 0.800            | 0.869                                 | 0.735             | 0.216                | 37       |
| 383/5            | Electrical machinery,<br>precision engineering | 0.853**                         | 1.121            | 0.780                                 | 0.595             | 0.158                | 37       |
| 384              | Transport equipment                            | 0.615*                          | 1.798            | 0.756                                 | 0.707             | 0.200                | 37       |
| 390              | Other manufactures                             | 1.445**                         | 0.774            | 0.843                                 | 0.527             | 0.309                | 34       |

<sup>1</sup> Estimation with OLS; bootstrap standard errors (5,000 iterations with replacement). Constant term included in all regressions but not reported.

Coefficients and standard errors reported with respect to (*IDIOBIAS* \* 1,000).

<sup>2</sup> Expenditure share weighted by spatial trade-freeness matrix *à la* Behrens *et al.* (2004); see text for details.

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

<sup>4</sup> ## (#) denotes rejection of H0: coeff. = 1 at 95% (90%) confidence level (two-tail test), based on bias-corrected bootstrap confidence intervals.



## Appendix: 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):

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

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](http://www.cepii.fr/anglaisgraph/bdd/distances.htm).

Tariff data are from CEPII's MAcMap database (see [www.cepii.fr/anglaisgraph/bdd/macmap.htm](http://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](http://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.