INEQUALITY AND THE US IMPORT DEMAND FUNCTION

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Abstract

In this paper we build a model of trade in vertically differentiated products and find that income inequality can affect the demand for imports even in the presence of homothetic preferences. The empirical importance of changes in inequality on the demand for imports is then assessed by examining US data for the 1948-1996 period. Using the Johansen (1988) procedure we find that there is no evidence of a long run relationship of a standard imports equation (one including imports, income, and relative prices). However, once we include a measure of inequality in our VAR specification we find not only evidence for the existence of a cointegrating equation in imports, income, relative prices and inequality, but that the evolution of inequality has a large and positive influence on the demand for imports in the US. Moreover we find that our results are robust to alternative methods of estimating cointegration equations.

JEL Code: F13, H23, O24.

Keywords: inequality, US import demand, vertically differentiated products, cointegration.

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1 Introduction

Standard specifications of import demand functions are usually based on the imperfect substitutes model, in which imports and domestically produced goods are not perfect substitutes (see, for example, Armington (1969), Goldstein and Khan (1985), Rose (1991), Hooper and Marquez (1995)). In this model, the demand for imports is usually thought of as the result of a representative household's maximization of utility (which depends on the consumption of a"domestic" and an "imported" good) subject to a budget constraint.[1] The(aggregate) volume of imports is thus specified as an increasing function of aggregate income and of the ratio of domestic to imported goods prices.Implicit in this derivation of the import demand function is the idea that the distribution of income is not an important determinant of the demand for imports.

In the present paper we examine the -*ceteris paribus*- effects of changes in income inequality on the demand for imports.[2] We do this by using a model of trade in vertically-differentiated products in which household income determines the quality of goods demanded (Linder (1961), Flam and Helpman (1987)).[3] The domestic country is assumed to have

[1] In the case that the imported goods are intermediates used in domestic production, the demand for imports arises from profit maximization and it depends on relative prices and gross domestic product (e.g., Kohli (1982)).

[2] Although the influence of income inequality on macroeconomic outcomes has not been an active area of research in the field of open-economy macroeconomics, the same does not hold true for the field of international trade. Indeed, there is a large theoretical and empirical literature examining the effects of inequality on trade patterns in the presence of non-homothetic preferences (e.g. Markusen (1986), Hunter (1991), Francois and Kaplan (1996), Mitra and Trindade (2003). In addition to its focus, the present paper differs from this literature in that we examine the effects of inequality in a model with homothetic preferences in the presence of vertically differentiated products.

[3] Schott (2003) presents evidence that testifies to the importance of vertical intra-industry trade in the world economy. He finds that "... the relationship between unit values, exporter endowments and exporter production techniques supports the view that capital- and skill-abundant countries use their endowment advantage to produce vertically superior varieties, i.e. varieties that are relatively capital or skill intensive and possess added features or higher quality, thereby commanding a relatively high price" (Schott (2003),

comparative advantage and to export to the rest of the world (ROW)), high-quality (and high-price) varieties of the differentiated product, whereas it imports low-quality (and low-price) varieties that are consumed by low-income households. We show that mean-preserving changes in income inequality have an ambiguous effect on the demand for imports.

The flavour of the argument can be understood by the example of a hypothetical mean-preserving increase in income inequality. Let there be an income level such that all households with income up to this level (call it λ) maximize their utility (which depends on the quality of the vertically-differentiated product and the quantity of a homogeneous non-traded good) by purchasing low-quality, low-price imported varieties; similarly, households with incomes greater than λ consume the high-quality domestically produced varieties. Consider now a case in which the income of some households which intially had incomes greater than λ drops to a level below λ, whereas the incomes of some households (which initially were far greater than λ) rise further, so that the average income remains intact. The effect of these changes will be an increase in imports since the households for which income has dropped below λ will switch their demand to imported varieties, whereas the households whose incomes have increased will continue to consume domestically-produced varieties.[4] We trust that the reader will have by now thought of counterexamples in which a mean-preserving increase in inequality results in a reduction in the demand for imports - thus intuitively confirming the

p.658). Thus, along with Bowen et al. (1987) and Trefler (1995) he concludes that there is no evidence of endowment-driven specialization *across* products. Moreover, Grossman (1982) has attributed a significant role to vertical product differentiation regarding the size and interpretation of estimated price and income elasticities in international trade.

[4] The conclusion we draw from this example would remain intact had we assumed that, in addition to the income changes mentioned earlier, the income of low-income households declined as well. Some readers may regard the hypothetical changes in household incomes specified in this example as a rough approximation of the actual changes in the US income ditribution since the mid-seventies. (see, Acemoglu (2002) for a review of the evidence).

ambiguous effect of inequality on the demand for imports.

The theoretical ambiguity as to the effect of income inequality on the demand for imports is by no means an artifact of our assumption that the domestic country has comparative advantage in the production of high-quality varieties. Indeed, as section 2 of the paper makes clear, it would also be a feature of the model if the domestic country had comparative advantage in the production of low quality varieties. This further implies that the theoretical ambiguity would also be present if, as is the case for any country in the world in a multi-commodity setting, the domestic country's comparative advantage was in high-quality varieties for only a sub-set of the differentiated products, or if international trade was conducted in both homogeneous and differentiated goods.

In the empirical section of the paper we try to ascertain the influence of changes in income inequality on the demand for US imports. For this purpose, we investigate the existence of a long run relationship between real imports, real income, relative prices and inequality for the 1948-1996 period. Using the Johansen (1988) procedure, we fail to detect evidence of a standard imports equation (one including imports, income and relative prices) The picture changes when we include a measure of inequality in our VAR specification. In fact both the trace test and the maximim eigenvalue statistic support the existence of a cointegrating vector including imports, income, relative prices and inequality.[5] We also find our results to be robust to alternative methods of estimating cointegration equations, with all methods producing remarakably similar estimates of the cointegrating vector and

[5] Our finding about the importance of income heterogeneity in explaining the behavior of United States imports can be considered as complementary to the one advanced by Marquez (2000) in his effort to "solve" the Houthakker and Magee (1969) puzzle about the high income elasticity of US imports. Marquez argued (and provided the relevant evidence) that if immigrants retain their tastes for their native products, then an increase in immigration would increase the demand for imports.

providing estimates of the elasticity of imports with respect to inequality ranging from 0.8-1.2. Moreover, given that the efficiency of the various methods in small samples may differ considerably, we perform a small Monte Carlo experiment in order to assess their relative performance in small samples. We conclude that the Johansen procedure along with the Fully Modified Least Squares estimator of Phillips and Hansen (1990) seem to perform best both in terms of bias and variation. Interestingly, these two methods deliver the highest estimates of the inequality elasticity.

Our estimates suggest a significant impact of inequality on real imports. For example, according to our range of estimates (0.8-1.2), had inequality in the US remained at its 1975 level, imports in 1996 would have been lower between 12 and 19 percent of the fitted value (which is close to the actual value). The further rise in inequality since 1996 implies that had inequality in 2004 been at its 1975 level, the percentage decline in US imports in 2004 would have been even larger than in 1996, thus implying a very large improvement in the US current account deficit.[6]

The remainder of the paper is as follows: Section 2 develops the theoretical model showing the influence of income inequality on the demand for imports. The empirical analysis is presented and discussed in section 3. The last section concludes.

2 The model

We present a simple theoretical framework capable of illustrating the influence of income inequality on the demand for imports. The framework is akin to Katsimi and Moutos (2005),

[6] Although recent data on US income inequality exist, in our empirical analysis we use the longer data set available, which covers the 1944-1996 period.

which has in turn borrowed from Malley and Moutos(2002) and Flam and Helpman (1987). We will assume the existence of a small open economy, which produces (and consumes) two goods: a homogeneous non-traded good (X) and a vertically-differentiated product (Y) that is traded with the rest of the word (ROW). The model features two-way international trade in the vertically-differentiated good, with the domestic country producing (and exporting) a high-quality quality variety of good Y, and importing a low-quality variety of it.

2.1 Firms

Good X (the non-traded good) is a homogeneous good produced under perfectly competitive conditions in the domestic country with the use of labour services (L). We conceive of L as being the simple aggregate of effective labour services provided by perfectly substitutable workers with each of them possessing different units of effective labour.[7] We assume that firms pay the same wage rate per effective unit of labour - thus the distribution of talent across firms does not affect unit production costs. For simplicity, we assume that each unit of L produces one unit of the homogeneous good.under linear technology,

$$X = L \tag{1}$$

Using labour as the numeraire, we get that the price of the homogeneous, non-traded good is, $P_X = 1$. We assume that all prices in the domestic economy and in the ROW are expressed in a common currency (the exchange rate is fixed at unity).

[7] Alternatively, we could conceive of L as a function of the quantities of labour provided by imperfectly substitutable groups of workers, e.g., $L = f(L_S, L_U)$, where L_S and L_U stand for the effective units of skilled and unskilled labour. Under the interpretation adopted in the text, changes in (income) inequality can be the result of changes in the effective number of labour units each worker (cum household) is endowed with. Under the skilled-unskilled workers interpretation, changes in inequality can be the result of changes in the relative wage of skilled workers – the so-called skill premium. Although empirically the second interpretation may be more relevant (especially for the United States – see, for example, Acemoglu (2002)), it is analytically far simpler to consider the first case of perfectly substitutable workers with unequal endowments of effective labour units.

The vertically-differentiated good (Y) is produced by perfectly competitive firms in both the domestic country and the ROW. We assume that quality is measured by an index $Q > 0$, and that there is complete information regarding the quality level inherent in all varieties produced at home and abroad. Moreover, for simplicity,[8] we assume that there is only one variety offered by domestic firms, q, and only one variety offered by ROW firms, q^*,with $q > q^*$. We further assume that, in both the domestic country and the ROW, average costs depend on quality, and that each (physical) unit of a given quality is produced at constant cost. The dependence of average costs on quality is motivated by the fact that increases in quality – for a given state of technological capability – involve the "sacrifice" of an increasing number of personnel which must be allocated not only to the production of a higher number of features attached to each good (e.g., electric windows, air bags, ABS, etc. in the case of automobiles) but also to the development and refinement of these features as well.

We assume that the domestic country has comparative advantage in the production of the high quality variety of the differentiated good. This implies that the least cost producers of the variety with quality q are domestic producers (that is, $AC(q) < AC^*(q)$) , whereas the least cost producers for variety q^* are ROW producers (i.e., $AC(q^*) > AC^*(q^*)$). For simplicity, we set $P(q) = AC(q) = \gamma q$, and $P(q^*) = AC^*(q^*) = \gamma^* q^*$, with $\gamma, \gamma^* > 0$.Changes in γ, γ^* may, for example, occur either due to cost-changing process innovations, or due to changes in the macroeconomic environment (e.g.the exchange rate).

[8] Katsimi and Moutos (2005) present a model in which there is a continuum of domestic and foreign varieties offered to the domestic country consumers.

2.2 Households

All households are assumed to have identical preferences, and to be endowed with one unit of labour, which they offer inelastically. There are, however, differences in skill between households, which are reflected in differences in the endowment of each household's effective labour supply. This is in turn reflected in an unequal distribution of income across households. Following Rosen (1974) and Flam and Helpman (1987) we assume that the homogeneous good is divisible, whereas the quality-differentiated product is indivisible and households can consume only one unit of it. For simplicity, and in order to demonstrate that inequality can have an influence on the demand for imports even with homothetic preferences,[9] we write the utility function of household i as

$$U_i = Q_i X_i \tag{2}$$

where Q_i and X_i stand for the quality (either q or q^*) of the differentiated product and the quantity of the homogeneous good (respectively) consumed by household i.[10]

Let e_i stand for the endowment of effective labour units owned by household i. Since the wage rate per effective unit of labour is unity, e_i stands also for household income. Assume that there is a continuum of households, $i \in [0, 1]$, with Pareto distributed incomes. The Pareto distribution is defined over the interval $e \geq b$, and its CDF is

$$F(e) = 1 - (\frac{b}{e})^a \tag{3}$$

[9] An implication of Krugman's (1989) derivation of the import demand function, is that with homothetic preferences, changes in inequality will not have any effect on the demand for imports if trade is conducted in horizontally differentiated products, since changes in household income would not alter the proportion of spending that either poorer or richer households spend on imported varieties.

[10] We implicitly assume that there is a fixed (and common across households) disutility of work effort which enters additively in the utility function. We also assume that the lowest ability household gets a higher level of utility (due to consumption) from working rather than from sitting idle.

where $a > 1$.Parameter b stands for the lowest income (ability) in the population, and parameter a determines the shape of the distribution (higher values of a imply greater equality). The mean of the Pareto distribution is equal to

$$\mu = \frac{ab}{a-1}. \tag{4}$$

The budget constraint of a household depends on whether it consumes the domestic or the foreign variety of the differentiated product. The budget constraint of a household which buys the domestically-produced variety is,

$$e_i(1-t) = X_i + \gamma q \tag{5}$$

whereas the budget constraint of a household buying the imported variety is,

$$e_i = X_i + \gamma^* q^* \tag{6}$$

where t stands for the (linear) income tax rate, and τ for the ad-valorem tariff rate.[11] As a result, the utility maximizing demand for the homogeneous good if the household chooses to consume the domestically-produced variety is,

$$X_i^D = e_i - \gamma q \tag{7}$$

whereas if the household chooses to consume the ROW-produced variety the demand for X is,

$$X_i^F = e_i - \gamma^* q^*. \tag{8}$$

[11] We assume that for all relevant values of the tariff rate τ, it will never be possible for domestic producers to supply to the domestic market the variety q^* at a lower price than the (inclusive of the tariff) price at which the ROW producers can sell the good to domestic consumers.

In deriving the above we have assumed that for all households income is high enough to generate positive demands for both goods. The resulting indirect utility functions in the two cases are then,

$$V_i^D = (e_i - \gamma q)q \tag{9}$$

$$V_i^F = (e_i - \gamma^* q^*)q^* \tag{10}$$

Household i will buy a foreign produced variety if $V_i^F > V_i^D$. We note that $\vartheta(V_i^D - V_i^F)/\vartheta e_i > 0$, i.e. the difference between V_i^D and V_i^F is increasing in household income. This implies that only households with large incomes will be willing to buy the high-quality variety which is domestically produced, whereas low-income households will find it optimal to consume the low-quality variety which is imported from the ROW. In Figure 1, high income households face the budget constraint BC1 and achieve higher utility at point 1 (by consuming the domestically produced variety) rather than at point 2 (which is associated with the foreign-produced variety). On the other hand, low income households bace the budget constraint BC3 and prefer to consume the imported variety (point 3) rather than the domestically produced one (point 4). Finally, there exist households with income λ, depicted by BC2, which are indifferent between the domestically produced and the imported varieties (points 5 and 6).

Let λ denote the income of a household that is indifferent between consuming the domestically produced variety and the foreign variety, i.e., for this household it holds that

$$V^D = (\lambda - \gamma q)q = (\lambda - \gamma^* q^*)q^* = V^F. \tag{11}$$

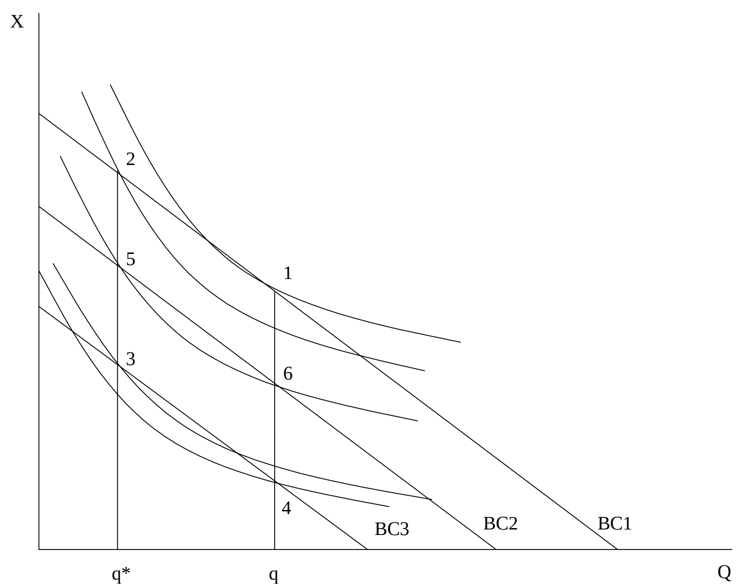

Figure 1: The relationship between income and inequality

We term λ the dividing level of income (ability). Solving for λ we find that

$$\lambda = \frac{\gamma q^2 - \gamma^*(q^*)^2}{q - q^*}. \tag{12}$$

Equation (12) indicates that the value of λ is independent of both parameters (a and b) describing the distribution of income. It depends only on domestic and ROW costs and the associated quality levels.

The Pareto distribution implies that the proportion of households with incomes smaller or equal to λ (that is, the proportion of households which choose to consume the foreign-produced variety), is equal to $1 - (b/\lambda)^a$. Thus, the real value (volume) of total imports is

$$M = [1 - (b/\lambda)^a]\, \gamma^* q^*. \tag{13}$$

Given our interest in the effect of mean preserving changes in income inequality, and the independence of λ from changes in a and b, we can use equation (13) to find the effect of changes in a while adjusting b (the lowest income in the population) so as to keep average income ($= ba/(a-1)$) constant.[12] Letting $\overline{\mu}$ denote the given level of average income, we find that

$$\frac{\vartheta M}{\vartheta a} = (M - \gamma^* q^*) \left[\ln(\frac{(a-1)\overline{\mu}}{a\lambda}) + \frac{1}{a-1} \right]. \tag{14}$$

The sign of $\vartheta M/\vartheta a$ is ambiguous, since $\ln((a-1)\overline{\mu}/a\lambda) = \ln(b/\lambda) < 0$.[13]

In order to understand the reason for this ambiguous effect consider first the result of a rise in a while holding b constant. In this case the rise in a (which implies a reduction in inequality) is associated with a reduction in average income (ability) and in the proportion of households with income greater than λ (i.e. the households buying the domestically produced variety). As a result, the proportion of households choosing to buy domestically produced goods decreases and imports increase (see also equation (11)). Given our wish to examine the effects of mean preserving changes in income inequality, an increase in a must be paired with an increase in b in order to keep μ constant. A -ceteris paribus- increase in the scale parameter b (which impies a rise in the lowest income in the population, as well as a rise in average income) implies that there will be fewer households below any given level of λ, thus decreasing the proportion of households buying the imported variety. This implies that the CDFs representing the two income distributions will be intersecting, with the one associated

[12] As can be easily seen from equation (13) a rise in equality (with b given) results in a rise in imports , i.e.

$$\frac{\vartheta M}{\vartheta a} = - \left[\frac{b}{\lambda}\right]^a \gamma^* q^* \ln \left[\frac{b}{\lambda}\right] > 0,$$

This results because the rise in a causes a fall in average income and a corresponding rise in the proportion of households wishing to consume imported varieties.

[13] Note also that equation (13) implies that $M - \gamma^* q^* < 0$.

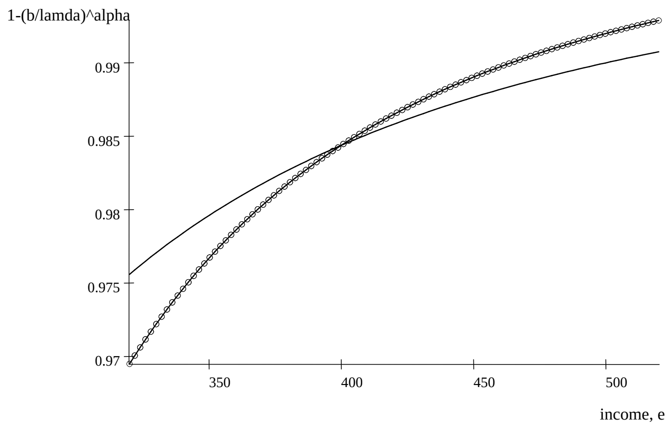

Figure 2: Inequality and the CDF

with higher values of a and b, crossing from below the one represening lower values of these parameters. Figure 2 focuses on the intersection of two alternative CDFs . The solid curve depicts the CDF for $a = 2$ and $b = 50$, whereas the dotted bold curve represents the CDF for the same average income ($\mu = 100$) for $a = 3$ and $b = 100$. Thus, if the value of λ is lower (higher) than the level of income at which the two CDFs intersect, a rise in a from 2 to 3 (accompanied by a rise in b from 50 to 100) will reduce (increase) the proportion of households wishing to buy imported varieties, and imports will decrease (increase).

The theoretical ambiguity as to the effect of a mean preserving increase in inequality on the volume of imports which exists in the present model is also a feature of more complicated models (e.g. in models allowing for a continuum of varieties to be offered by domestic and ROW producers, or for the presence of imported intermediate inputs). It would also be present if the domestic country had comparative advantage in the production of the low-quality variety. This can be easily verified by noting that in this case equation (12) would be modified to $M = (b/\lambda)^a \gamma^* q^*$, since in this case the imported variety would be bought by households with income greater than (or equal to) λ.

However, changes in actual income distributions may not be as "smooth" as described by varying the parameters of theoretical distributions. Consider, for example, the case of a rise in inequality which involves the reduction of the income of some households -which initially had incomes slightly larger than λ- to less than λ, and the concurrent rise of the incomes of households with incomes significantly greater than λ so that average income says constant. Our analysis would then predict an unambiguous effect on the demand for imports; since the households whose incomes have been reduced to less than λ will switch their demand from the domestically-produced variety to the foreign-produced one, the demand for imports will increase.[14] Since it is also easy to construct other hypothetical examples in which a rise in inequality results in a fall in import demand, we proceed with the empirical examination of this issue.

3 Econometric Analysis

3.1 Empirical Literature Review and Data

We aim at analyzing empirically the impact of US inequality on the US demand for imports. Most empirical studies on the macroeconomic determinants of the demand for imports estimate a standard real import demand function according to which imports depend on real income and relative prices. A large body of empirical literature has estimated price and in-

[14] The households whose income rises and remains higher than λ will continue to consume the domestically produced variety.

come elasticities of imports and much of it focused on US trade.[15] More recent papers have attempted to find evidence of a long run relationship (cointegration) between the levels of imports, income and relative prices (or the real exchange rate). The results are mixed. Rose and Yellen (1989) and Meade (1992) fail to find evidence of cointegration for the 1960-87 period. Johnston and Chinn (1996) find a cointegrating relationship by excluding agricultural products and fuels for the 1973-95 period whrereas Chinn (2005b) obtains evidence of a cointegrating relationship only when excluding computers. Boyd et al. (2001) obtain a long run import demand function for the 1975-95 period but they impose the restriction that the income elasticity of imports should equal the income elasticity of exports with the opposite sign. Finally, Hooper et al. (1998) find evidence for a cointegrating relationship among real imports, real income and relative prices for the 1960-1994.

Another strand of this literature challenges the conventional wisdom by arguing that the standard imports demand function may be misspecified due to the ommission of other determinants of a long run imports equation. Along these lines, Marquez (2000, 2002) provides evidence for a cointegrating imports equation for the 1967-1997 period by including either the share of immigrants in the population or the ratio of the foreign capital stock to the US capital stock. The inclusion of immigration is based on his argument that if immigrants retain their tastes for their native products, then an increase in immigration would increase the demand for imports. On the other hand, as argued by Helkie and Hooper (1988), the inclusion of the relative capital stock is a measure of an existing upward bias in import prices. This bias is the result of the failure of import prices to incorporate the prices of new products

[15] For surveys of literature on this topic see Goldstein and Kahn (1985) and Sawyer and Springle (1996).

which are most of the times lower that the prices of existing products.

The main empirical implication of our theoretical model is that inequality may be an important determinant of the demand for imports. As a result, ommitting the level of inequality may be one reason why most previous studies failed to provide strong evidence of a stable long run import demand function. Our purpose is to enrich the commonly used empirical specification by including a measure of inequality. Specifically, in line with the most recent research in this topic, we use the Johansen (1988, 1991) procedure in order to investigate the existence of a long-run relationship between imports, income, relative prices and inequality. We expand on this traditional specification since -unlike our stylized model- international trade is conducted not only in vertically differentiated goods but in horizontally differentiated and homogeneous goods as well.

Our analysis is based on annual data since there are no higher frequency data for inequality. We model US real imports of goods and services (IM) as a function of US real GDP (Y), the relative price of imports (RP) and inequality (IN), where all variables are in logs.[16] Our measure of ineq uality is taken from the revised version of World Income Inequality Dataset (WIID) constructed by Deininger and Squire (1996). This data set is to our knowledge the most complete and reliable source of inequality data and it provides alternative estimates for the US GINI coefficient. We measure inequality, IN with the GINI coefficient that covers

[16] Note that in equation (13), the effect of changes in income and relative prices are captured through changes in the parameters a, b, q, q^{*}, γ,and γ^{*}.In this respect it is important to note that in our theoretical analysis labour is assumed to be the only domestically owned factor of production. Nevertheless, since household consumption choices are made on the basis of total household income, rather than income derived from the sale of the household's labour services alone, care must be taken to control for the other sources of income. Also, the presence of not only final consumption goods but of intermediate inputs as well as homogeneous and horizontally differentiated consumption goods in the actual import data necessitates the inclusion of a variable measuring aggregate domestic activity. We use domestic GDP to control for the influence of the above concerns.

the longest period (1944-1996) constructed by Brandolini (1998). Real imports, IM and real GDP, Y are both in 2000 chain weighted dollars from the Federal Reserve Bank of St. Louis. Following Hooper et al (1998) we use the price of imports over the GDP deflator, RP from the International Financial Statistics as our main measure of competitiveness. All variables exist after 1948. In many empirical studies competitiveness is measured with some exchange rate index such as the real effective exchange rate or some trade weighted exchange rate. Therefore, in an alternative specification we also use the real effective exchange rate, RER from the OECD Economic Outlook as a measure of competitiveness. However, given that real exchange rate data exist only after 1970 this alternative specification reduces our sample from the 1948-1996 period to the 1970-1996 period.

3.2 Estimation and Testing Procedure

First, we test the unit root hypothesis for each of the individual component of the vector stochastic process $\{Z_t\}$, where $Z_t' = (IM_t \; Y_t \; RP_t \; IN_t)$. Standard unit root tests of Dickey and Fuller (1981) and Phillips and Perron (1988?) fail to reject the unit root null for all the four series under consideration. Note that this evidence is robust to the choice of the lag-length in the Dickey-Fuller regressions and the choice of the bandwidth parameter in the context of the Phillips-Perron non-parametric procedure. Therefore, we proceed by assuming that the process $\{Z_t\}$ consists of $I(1)$ components. Then we move on to multivariate analysis within the Johansen (1998, 1991) cointegration framework. We take the following steps: (i) Since the Johansen procedure is based on the estimation of a VAR(p) model, we first, we choose the optimal lag length of the VAR. (ii) In the context of the Vector Error Correction (VEC) representation of VAR(p), we test for cointegration by using the trace and the maximum eigenvalue statistic. (iii) Having determined the cointegration rank, we re-estimate the VEC model with the cointegration rank restriction imposed on the long-run matrix of the model. In this framework, we estimate both the long-run and the short-run dynamics of the system. More specifically, let us assume that the stochastic process $\{Z_t\}$, where $Z_t' = (IM_t \ Y_t \ RP_t \ IN_t)$, is generated by the following VAR(p) model

$$Z_t = A_0 + \sum_{i=1}^{p} A_i Z_{t-i} + U_t \tag{15}$$

whose VEC representation takes the form:

$$\Delta Z_t = A_0 + \Pi Z_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta Z_{t-i} + U_t \tag{16}$$

with $U_t \sim NI(0, \Omega)$. The process $\{Z_t\}$ is cointegrated if the matrix Π is of reduced rank, that is when $r(\Pi) = r < 4$ in our case. The rank of Π describes the number of the cointegrating vectors in the system. If the matrix Π is of full rank, that is $r(\Pi) = r < 4$ then the VAR(p) is stable VAR in levels and there are no unit roots in the system. Note that this case contradicts the assumption that each of the four series is I(1). Finally, if $r(\Pi) = 0$ then the number of unit roots in the system is equal to four, and the series are not cointegrated. Let us assume that $r(\Pi) = 1$. In such a case, the long-run matrix Π can be decomposed into

$$\Pi = \mathbf{cb}'$$

where $\mathbf{c}$ and $\mathbf{b}$ are (4×1) vectors. In such a case, the system (16) becomes

$$\Delta Z_t = A_0 + \begin{bmatrix} c_{11} \\ c_{21} \\ c_{31} \\ c_{41} \end{bmatrix} \begin{bmatrix} b_{11} & b_{21} & b_{31} & b_{41} \end{bmatrix} Z_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta Z_{t-i} + U_t$$

It can be seen that the vector **b** contains the long-run parameters of the system, whereas the vector **c** contains the adjustment coefficients of each of the four variables IM_t Y_t RP_t IN_t to the disequilibrium error of the of the previous period.

3.3 Results

We use the Johansen (1988), and Johansen and Juselius (1990) procedure in order to test for cointegration and to determine the number of long-run relations. We choose the 2 lag specification for our VAR since the 1 lag specification suffers from serial correlation. Our results are reported in Table 1. We first examine whether in the absence of the inequality variable there is cointegration among real imports, real GDP and relative prices. As can be seen from column (1), there is no evidence of cointegration among IM, Y and RP. The inclusion of inequality as an additional determinant of the volume of imports provides us with evidence of cointegration according to both the trace test and the maximal eigenvalue statistics - see column (2).[17] All reported coefficients (the elements of the cointegration vector **b**) are significant. The income elasticity of imports is 1.6, lower than the estimates reported by Chinn (2005a). The relative price sensitivity is 0.15 and has the expected negative sign. Inequality has a positive and significant effect on imports.

[17] The cointegrating equation reported in Table 1 does not include a time trend. Nevertheless, even if we include a time trend in the regressors of column (2), we still get a cointegrating vector according to the trace test.

TABLE 1: Import Cointegration Results I

Long Run	US real imports, IM		
Coefficients, ζ_i	(1)	(2)	(3)
Cointegr. vectors:			
trace test	0	1	1
max. eigenvalue	0	1	1
Y	1.980	1.614	1.11
	(0.080)	(0.181)	(0.076)
RP	−0.345	−0.155	
	(0.227)	(0.046)	
IN		1.220	3.03
		(0.189)	(0.289)
$REER$			−0.11
			(0.04)
constant	-11.00	-12.38	-14.05
lag	2	2	1
N	46	46	25
Error correction coefficients			
IM		−0.638	−0.873
		(0.181)	(0.362)
Y		−0.169	−0.070
		(0.09)	(0.125)
RP		0.592	
		(0.22)	
IN		0.167	−0.075
		(0.06)	(0.089)
$REER$			−0.602
			(0.230)

Notes: Standard errors in parentheses.

Column (3) of Table 1 reports estimates using the real exchange rate, $REER$ as an alternative measure of competitiveness. Again, a long run relationship is detected only after the inclusion of inequality, IN among the determinants of the volume of imports in a VAR(1) specification. In this case, the income elasticity is close to unity whereas inequality has an even stronger impact on imports. However, as in Hooper et al (1998), we obtain an incorrect sign for the real exchange rate elasticity. The error correction coefficients of real imports reported in the last two columns of Table 1 are negative and significant under both specifications. This indicates that in the presence of disequilibrium the volume of imports gradually adjusts towards its long-run value. Finally, the residuals of both models satisfy homoskedasticity, and normality. However, the residuals of model (3) suffer from serial correlation.

In order to get an idea for the importance of inequality in shaping the evolution of US imports, we depict in Figure 3 the fitted values of imports derived from the long run imports equation shown in column (2) of Table 1 (series 1), whereas series 2 represents the fitted values of imports which would obtain had inequality remained constant at its 1975 level. Series 3 depicts the actual evolution of US imports. According to our estimates, had inequality remained at its 1975 level, the fitted value of imports in 1996 would have been 19% lower than the fitted value of imports derived by using the actual level of inequality for 1996. The further rise in inequality since 1996 depicted by more recent but shorter data sets (see Current Population Survey, U.S. Census Bureau), implies that had inequality in 2004 been at its 1975 level, the percentage decline in US imports in 2004 would have been even larger than in 1996, thus implying a very large improvement in the current account deficit.

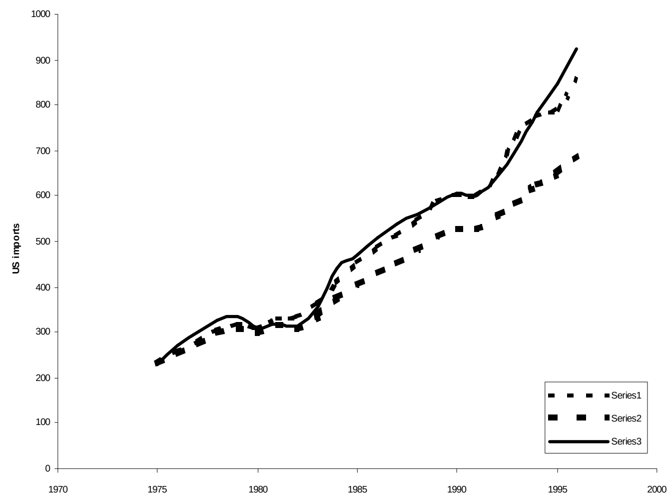

Figure 3: US imports

3.4 Robustness

In this section we address the following questions:

(i) How robust are our empirical results to the choice of the cointegration estimation method? In other words, how different would our results be if we adopted other asymptotically efficient cointegration estimators?

(ii) The Johansen cointegration method is asymptotically optimal. However, in samples as small as ours (46 observations) it has been reported that the Johansen method as well as other asymptotically equivalent methods suffer from small sample bias [see Hargreaves, (1994), Inder (1993) and Gonzalo (1994)]. This bias depends on the dynamics of the system. For example in the context of the triangular model of cointegration of Phillips (1991) this bias depends on the Granger causality structure between the cointegration error and the error that drives the regressor and the serial correlation properties of the former.

To address these questions we take the following two steps: First, since our previous results indicate a single cointegrating vector, we estimate our model with two other asymptotically efficient single equation cointegration methods. Second, we perform a small Monte Carlo experiment to assess the relative performance of the alternative estimators for a sample equal to that used in the estimation and a Data Generation Process which resembles as closer as possible the one that is likely to have given rise to the observed data.

3.4.1 Alternative cointegration methods

As far as cointegration estimators are concerned we consider, apart from the Johansen procedure (JOH) described in the previous section, the following estimators: (i) The simple OLS which is *not* asymptotically efficient but is included as a benchmark. (ii) the Autoregrssive Distributed Lag estimator (ARDL) suggested by Pesaran and Y. Shin (1999) (see also Phillips and M Loretan (1991) for a version of ARDL) (iii) the semi-parametric Fully Modified Least Squares (FMLS) estimator of Phillips and Hansen (1990) The difference between FMLS and ARDL(p,q,k) lies in the way the 'long-run correlation' and 'endogeneity' cointegration effects are accounted for. In particular, FMLS generates estimates of the nuisance parameters present in the asympotic distribution of OLS non-parametrically, whereas ARDL eliminates the nuisance parameters from the limiting distributions by estimating a full dynamic model including lags and leads of the variables in the system (see Pesaran and Shin (1999), and Panopoulou and Pittis (2004)).

Table 2 presents the results for these alternative cointegration estimators. These results should be compared with those reported in the second column of Table 1. This comparison reveals that our results are robust across the different methods: not only all coefficients are significant and of the same sign independently of the estimation method used, but also the income and relative price elasticities vary very little across all estimation methods. We also observe that the inequality elasticity of imports increases from 0.8 to 1.2 when the Johansen and FMLS are used. However, even under the lower elasticity, the effect of the rise in US inequality on US imports would still be very large - it would imply that imports in 2004 would have been lower by about 12% of their 1996 value (25% of their 2004 value) had inequality remained at its 1975 level.

TABLE 2 : Import Cointegration Results II

	US real imports, IM		
Method	OLS	ARDL*	FMLS**
Y	1.707 (0.02)	1.747 (0.046)	1.683 (0.019)
RP	-0.170 (0.045)	-0.252 (0.100)	-0.244 (0.041)
IN	0.784 (0.166)	0.861 (0.346)	1.24 (0.155)
constant	-11.58 (0.530)	-12.18 (1.123)	-13.04 (0.488)
N	49	48	48

Notes: Standard errors in parentheses,

* The Schwartz Order Selection Criterion suggested one lag of IM and one lag of Y. No time trend is included.

** Bartlett weights have been used. A truncation lag of 8 has been selected. 8 is the lag that, according to the Monte Carlo experiment of the following section, eliminates the bias for this parameters configuration.

Still since the rise of the GINI elasticity is realtively large when either the JOH or FMLS procedures are employed, a natural question to ask is which estimate we trust. This question cannot be answered by appealing to asymptotic arguments, since all the three estimators (JOH, ARDL, FMLS) are asymptotically equivalent. Therefore, in order to assess the relative performance of the alternative estimators, we proceed to Monte Carlo simulations. In the following section we run a small Monte Carlo experiment for a sample equal to that used in the estimation (49 observations) and a Data Generation Process which resembles as close as possible the one that is likely to have given rise to the observed data.

3.4.2 A Monte Carlo experiment

We shall assess the performance of these estimators in the context of the triangular model of cointegration suggested by Phillips (1991). In our case and assuming that the cointegration error and the errors that drive the regressors follow a VAR(1) model, we have

$$y_t = c + \mathbf{b}^\top \mathbf{x}_t + u_{1t} \tag{17}$$

$$\mathbf{x}_t = I_3 \mathbf{x}_{t-1} + \mathbf{e}_t$$

$$\begin{bmatrix} u_{1t} \\ \mathbf{e}_t \end{bmatrix} = \begin{bmatrix} a_{11} & \mathbf{a}_{12}^\top \\ \mathbf{a}_{21} & A_{22} \end{bmatrix} \begin{bmatrix} u_{1t} \\ \mathbf{e}_{t-1} \end{bmatrix} + \begin{bmatrix} \nu_{1t} \\ \boldsymbol{\nu}_t \end{bmatrix}$$

and

$$\begin{bmatrix} \nu_{1t} \\ \boldsymbol{\nu}_t \end{bmatrix} \sim NIID \begin{bmatrix} 0 \\ \mathbf{0} \end{bmatrix}, \begin{bmatrix} \sigma_{11} & \boldsymbol{\sigma}_{12}^\top \\ \boldsymbol{\sigma}_{12} & \Sigma_{22} \end{bmatrix} \tag{18}$$

where $\mathbf{x}_t = [x_{1t}, x_{2t}, x_{3t}]^\top$, $\mathbf{b} = [b_1, b_2, b_3]^\top$, $\mathbf{e}_t = [e_{1t}, e_{2t}, e_{3t}]^\top$, $\boldsymbol{\nu}_t = [\nu_{1t}, v_{2t}, v_{3t}]^\top$. In the context of our empirical model, y_t denotes real imports, IM, x_1 denotes real GDP, Y, x_2 denotes relative prices, RP and x_3 denotes inequality, IN.

One can make the following remarks regarding the estimators used in our analysis as opposed to the OLS estimator:

(i) The presence of nuisance parameters (cointegration effects) in the asymptotic distribution of the OLS estimator can be due to either (a) Granger causality from $\mathbf{e}_t$ to u_{1t} ($\mathbf{a}_{12} \neq \mathbf{0}$), and/or (b) Granger causality from u_{1t} to $\mathbf{e}_t$ ($\mathbf{a}_{21} \neq \mathbf{0}$), and/or (c) contemporaneous correlation between $\mathbf{e}_t$ and u_{1t} ($\boldsymbol{\sigma}_{12} \neq \mathbf{0}$). In other words, if $\mathbf{a}_{12}$, $\mathbf{a}_{21}$ and $\boldsymbol{\sigma}_{12}$ were all zero vectors then the OLS estimator would be the optimal estimator for estimating $\mathbf{b}$.[18]

(ii) The asymptotically efficient estimators, namely JOH, ARDL and FMLS basically deal with the nuisance parameters of the OLS estimator asymptotically. However, in the presence of a small sample some remaining effects may be manifested in biases produced even by JOH, ARDL and FMLS.

(iii) The previous remarks suggest that different estimates among JOH, ARDL and FMLS may arise depending on the relative ability of each estimator to remove the cointegration effects 'relatively fast'. Moreover, if these effects were present only in specific location of the above system, then these estimators would differ only with respect to the corresponding parameter. For example, if only e_{3t} were either temporally or contemporaneously correlated with u_{1t} then the estimators are likely to produce different estimates of only say b_3.

Next, we callibrate the above model using our data. This gives us estimates of a_{11}, $\mathbf{a}_{12}, \mathbf{a}_{21}, A_{22}, \sigma_{11}, \boldsymbol{\sigma}_{12}, \Sigma_{22}$. These estimates allow us to simplify our Monte Carlo design, by moving to a lower dimensional model where we have only one regressor. This is due to the fact that our estimates suggest Granger causality and (negative) contemporaneous correlation

[18] Some further corrections would be necessary for estimating its standard error if $a_{11} \neq 0$.

mainly between u_{1t} and e_{3t}. As a result, we adopt the following DGP:

$$y_t = \theta x_t + u_{1t} \tag{19}$$

with $\theta = 1$,

$$x_t = x_{t-1} + e_t$$

$$\begin{pmatrix} u_{1t} \\ e_t \end{pmatrix} = \begin{pmatrix} 0.7 & -0.21 \\ 0.30 & 0.30 \end{pmatrix} \begin{pmatrix} u_{1t-1} \\ e_{t-1} \end{pmatrix} + \begin{pmatrix} \nu_{1t} \\ \nu_{2t} \end{pmatrix} \tag{20}$$

and

$$\begin{pmatrix} \nu_{1t} \\ \nu_{2t} \end{pmatrix} \tilde{} NIID \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix} \begin{pmatrix} 0.0014 & -0.00019 \\ -0.00019 & 0.00031 \end{pmatrix} \right] \tag{21}$$

Regarding the assesment of our estimators, all estimators of θ are compared on the basis of the following three statistics:

1) Bias, computed as:

$$\widehat{\theta} - \theta_0$$

where:

$$\widehat{\theta} = \sum_{i=1}^{r} \widehat{\theta}_i / r$$

$i = 1, ..., r$ and r is the number of replications and $\theta_0 = 1$.

2) Average standard error, astde($\widehat{\theta}$)

$$astde(\widehat{\theta}) = \sqrt{\sum_{i=1}^{r} \left(\widehat{\theta}_i - \overline{\theta} \right)^2 / r}$$

3) Average root mean squared error, rmse($\widehat{\theta}$), computed according to the previous formula in which $\overline{\theta}$ has been replaced by θ_0.

TABLE 3 : Monte Carlo Results

	Mean bias	Standard error	Root mean sq. error
	Panel A: Sample size= 49		
Estimator			
OLS	−0.0995	0.2035	0.0513
ARDL	−0.0318	0.2266	0.0523
JOHANSEN	−0.0115	0.2157	0.0466
FMLS	−0.0232	0.1038	0.0113
	Panel B: Sample size:=490		
Estimator			
OLS	−0.0090	0.0200	0.0005
ARDL	−0.0004	0.0187	0.0003
JOHANSEN	−0.0015	0.0131	0.0002
FMLS	−0.0022	0.0124	0.0002
	Panel C: Sample size=4900		
Estimator			
OLS	−0.0010	0.0020	0.000
ARDL	−0.0001	0.0018	0.000
JOHANSEN	−0.0002	0.0013	0.000
FMLS	−0.0002	0.0013	0.000
Number of replications: 1000			

The results of 1000 replications of the above model are presented in Table 3. Panel

A reports simulations results for a sample size of 49 observations. As expected, the OLS appears to be the worst estimator of all, since it exhibits the largest bias and variation. On the other hand, the Johansen specifications consistently outperforms the ARDL procedure in terms of the bias, the standard deviation and root mean square error. However, the fully modified estimator outperforms the ARDL dynamic specification in terms of variation. JOH and FMLS exhibit the lowest bias and variation of all estimators, which implies that it is more likely to be closer to the true value with these two estimators than with any other estimator. In terms of mean bias, the ARDL procedure fairs well in comparison to the simple OLS, but appears to be about three times worse than the Johansen procedure. Thus our results strongly support the superiority of the fully modified estimator and the Johansen estimator for estimation and inference on θ. These procedures appear to be the best since they minimize the corresponding biases and the variation. Note that these procedures also imply the highest inequality elasticity of imports.

Finally, we investigate the effect of the sample size on the estimators' performance. Panel B and C report the Monte Carlo results when our sample increases by a factor of 10 in panel B and by a factor of 100 in panel C. As expected, the bias becomes negligible for all the estimators as our sample increases, with only the OLS bias remaining relatively high (OLS does not account for the cointegration effects even asymptotically, although it is super consistent). Moreover, the standard deviation (and the root mean squared error) is almost the same for all estimators. These results are consistent with the relevant asymptotic theory. Indeed, our results show that the bias for all estimators decrease at a rate close to T (instead of $\sqrt{T}$). For example, the bias of the OLS and FMLS at T=1 is -0.0995 and -0.0232 respectively, whereas at T=10 the bias decreases to -0.0090 and -0.0022 respectively and at T=100 the bias decreases further to -0.001 and -0.0002 respectively.

4 Concluding Remarks

The present paper explains our finding that US income inequality has a significant influence on the US demand for imports on the basis of a model in which trade is conducted in vertically-differentiated products. However, one could advance alternative explanations for this finding. For example, if one assumes that preferences are non-homothetic, and imports have a higher income elasticity than domestically produced goods, then changes in inequality can affect the demand for imports even if trade is conducted in homogeneous goods. Given our objective to improve on the standard specifications of the *aggregate* import demand function we regard the existence of alternative channels for the influence of inequality on the demand for imports as a plus; after all, despite the increasing importance of vertically-differentiated products in world trade, the share of international trade that is conducted in either homogeneous goods or in horizontally-differentiated products remains significant.

In this paper, in line with Rose and Yellen (1989), Meade (1992), Johnston and Chinn (1996) and Chinn (2005b), we find no evidence for the existence of a long run relationship between agrregate imports, income and competitiveness in the US. However, the addition of US income inequality as a determinant of the aggregate demand for imports improves the picture significantly. Using US data for the 1948-1996 period we find not only that there is a stable long run relationship between aggregate imports, income relative prices and inequality, but that the influence of inequality is quantitatively very important as well. This result appears robust accross alternative methods of estimating cointegration equations. Moreover, Monte Carlo simulations suggest that the methods delivering the highest inequality impact on imports are those with the best performance in small samples.

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