The Puzzling Income Elasticity of US Imports

Jaime Marquez[1]

1 Introduction

"*It takes an estimate to beat an estimate.*"

I do not remember who said that but it aptly describes three decades of econometric analyses of US imports. The estimate to beat is that of Houthakker and Magee (1969) who, modeling US imports in terms of income and relative prices, report an income elasticity of 1.5. Their estimate implies that, in the absence of price increases, the United States will change from a largely self-sufficient economy to one that cannot pay for its imports. The literature's response to this puzzling estimate has been to modify the specification of Houthakker and Magee in various ways: allowing for simultaneity in the import market, recognizing dynamic adjustments and optimization, disaggregating imports across product and countries, removing measurement errors in official data, and differentiating between cyclical and secular forces.[2] Yet, three decades of econometric modeling of US imports along these lines yields estimated income elasticities much greater than one (figure 1).[3]

I resolve the elasticity puzzle by removing two undesirable features embodied in previous work: the substitution bias embodied in official import-price data and the representative-agent assumption. Substitution biases stem from the exclusion of prices of new products, especially those from developing countries, from official price measures. The representative-agent assumption facilitates modeling imports but the strength of immigration into the United States questions the usefulness of this assumption. That errors in price data and immigration matter for explaining trade is not new; what is new is using these two considerations to beat, so to speak, the estimate of Houthakker and Magee.

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[1] **Federal Reserve Board. Washington DC 20551; jaime.marquez@frb.gov** I would like to acknowledge first comments I received from Hendrik Houthakker and Steve Magee. I have also benefited from comments by Laura Adams, John Ammer, Bill Donnelly, Neil Ericsson, Jon Faust, Michael Ferrantino, Dale Henderson, Keith Head, William Helkie, Peter Hooper, Kishore Gawande, Linda Goldberg, David Gould, Jane Ihrig, Karen Johnson, Wolfang Keller, Peter Kennedy, Kala Krishna, Andrew Levin, Cathy Mann, Kathryn Morisse, J. David Richardson, Charles Pearson, Raymond Robertson, Wendy Takacs, Charles Thomas, Ralph Tryon, Kei-Mu Yi, and Joachim Zietz; also participants in the FRB International workshop, the Fall 1998 Midwest Economics Meetings at the University of Michigan, the Spring 1999 meetings of the System's Committee on International Economic Analysis, the US International Trade Commision, and Johns Hopkins' SAIS made useful remarks. The views in this paper are solely the responsibility of the author and should not be interpreted as reflecting the views of the Board of Governors of the Federal Reserve System or of any other person associated with the Federal Reserve System. I use PcGive Professional 9.1.

[2] Searching over the Social Sciences Citations Index indicates that Houthakker and Magee (1969) has near 300 citations from 1972 to the present; which is the largest number of citations from all of the articles in the spring issue of five journals: the *American Economic Review* (June), *Econometrica* (April), the *Journal of Political Economy* (May/June), the *Quarterly Journal of Economics* (May), and the *Review of Economics and Statistics* (May). I am grateful to Cathy Tunis for this search.

[3] For reviews of the literature see Magee (1975), Stern et al. (1976), Goldstein and Khan (1985), and Sawyer and Sprinkle (1996).

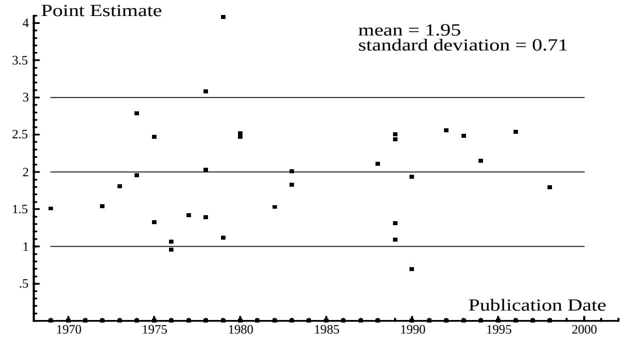


Figure 1: Selected Estimates of Income Elasticities – Marquez (1995, Appendix A).

Figures 2 and 3 show the roots of the elasticity puzzle and why it is hard to beat. First, the import-GDP ratios for imports consumer goods, producer goods, and services have increased; for aggregate imports, this ratio increases from less than five percent in 1960 to sixteen percent in 1997. If price effects were absent, then these increases would be reflected in income elasticities greater than one.

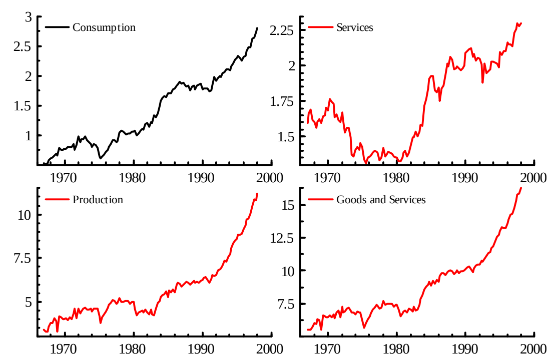


Figure 2: Import-GDP ratio, 1992 prices (percent). Source: Appendix A.

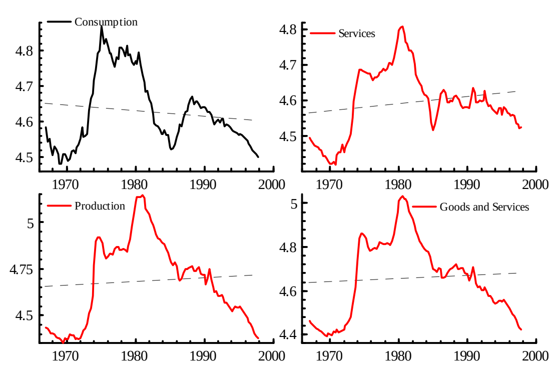


Figure 3: Import Prices Relative to GDP Deflator (logarithms). Source: Appendix A.

Price effects need not be zero but the available price data do not help in their identification. Specifically, after increasing in the 1970s, relative import prices decline to the level they had in the late 1960s and have, essentially, a flat trend. Econometrically, a flat trend means that relative prices contribute little to explaining the rise in the import-GDP ratio transferring to income the burden of explaining the rise in import-GDP, a transfer reflected in an income elasticity greater than one.

Section 2 elaborates on the puzzling nature of the income elasticity of US imports and describes previous methods addressing this puzzle. I show that, with one exception, those methods do not resolve the puzzle. The exception is the omission of prices of new products in official sampling techniques (Hooper and Richardson, 1991, Feenstra 1994). Mechanically, this omission induces an upward bias in the import-price data which understates the fall of import prices shown in figure 3. Thus addressing the resulting substitution bias would increase the role of relative prices in accounting for the upward trend in the GDP share of imports, diminish the burden on income to explain that trend, and thus lower the income elasticity. Section 3 follows Feenstra (1994) and modifies the conventional import demand model to address the substitution bias induced by the omission of new-products' prices.

Section 4 modifies the conventional model to remove the representative-agent assumption. I start with an individual's demand for imports depending *only* on income and relative prices. Then, consistent aggregation of the micro relations yields a macro equation with income, relative prices, *and* the variances of the distributions of imports, income, and relative prices. These variances embody individuals' heterogeneity which I model using the share of foreign-born individuals in the United States. What motivates this strategy is that immigration is an obvious source of heterogeneity accounting for about ten percent of the US population. Intuitively, if immigrants retain their tastes for their native products, then a ceteris paribus increase in immigration raises the demand for imports; ignoring this factor biases the estimated income elasticity.[4]

For parameter estimation, Section 5 uses the cointegration method of Johansen (1988) which recognizes dynamic adjustments, avoids simultaneity biases, and differentiates between secular and cyclical forces. Recognizing either the substitution biases in import prices or the heterogeneity induced by immigration point to unitary income elasticities (table 1). Excluding these factors and using the conventional model yields estimated income elasticities ranging from 1.6 for services to 2.6 for producer goods; these estimates exceed, by a significant margin, the income elasticity from the other two models. I evaluate the robustness of the estimates by testing parameter constancy and observational equivalence to trend and optimization models.

Table 1: Income Elasticities for US Imports–Alternative Models

Model	Services	Producer	Consumer	Aggregate
Substitution Bias (New Products' Prices)	1.56*	0.69*	1.41*	1.18*
Heterogeneity (Immigration)	1.21*	1.12*	1.31*	1.14*
Conventional (Income and Relative Prices)	1.60*	2.57*	1.94*	1.89*

* significant at the 5% significance level

[4] Gould (1994) and Head and Ries (1998) implement empirically gravity models with immigration but they are not concerned with, and do not solve, the elasticity puzzle–Houthakker and Magee (1969) is not even cited. Immigration may have indirect effects on imports that are summarized in Rybcyzinski's theorem (Caves 1967, pp.103, 116-18).

2 The Elasticity Puzzle

2.1 What is It?

Finding an estimated income elasticity for US imports in excess of one implies that, in the absence of price increases, the ratio of US imports to GDP will exceed one. Such a prediction is puzzling for several reasons. First, why should the GDP share of imports rise continuously whereas the GDP shares of consumption and investment are constant by comparison? Second, why should foreign creditors increase their holding of claims on US future production beyond the value of US output?[5]

From a practical standpoint, one may be tempted to dismiss the relevance of this elasticity puzzle because there are economies, such as Singapore, Malaysia, and Hong Kong, that have import-GDP ratios greater than one (Peebles and Wilson 1996, table 6.1, p. 160). Why should, then, an import-GDP ratio greater than one be puzzling at all? For two reasons. First, because the forces responsible for the import-GDP ratios in these entrepôt economies do not operate in the United States. Specifically, these economies have a poor resource base and have become centers for processing imports of raw materials for exports (Peebles and Wilson, 1996, p. 159). Thus exports, and not domestic demand, explain movements in imports with export proceeds paying for these imports. Indeed, the import-GDP ratio for imports destined to the domestic market is about 20 percent (Findlay and Welliz, 1993, p. 100). Second, because dismissing the relevance of the elasticity puzzle on the basis of this evidence confuses necessary with sufficient conditions: for given prices, an income elasticity greater than one implies a growing import-GDP ratio whereas an import-GDP ratio greater than one does not imply an income elasticity greater than one.

From a theoretical standpoint, one may be tempted to dismiss the relevance of this elasticity puzzle because the income elasticity need not be constant. Therefore the observed elasticity puzzle need not be a feature of the economy but rather a feature of the model used for parameter estimation. To emphasize this argument, denote m as imports and y as income, and compute the income elasticity as $\frac{dm}{dy}/\frac{m}{y}$. For a given marginal propensity to import, $\frac{dm}{dy}$, the income elasticity declines as the average propensity to import, $\frac{m}{y}$, rises. Models that do not assume constancy of elasticities have been implemented by Burgess (1974), Kohli (1978, 1991), Clarida (1994), Amano and Wirjanto (1997), and Marquez (1994, 1999) among others. These models are not widely used, however, because their predictive power is well below that of models with constant elasticities. Indeed the reason why the elasticity puzzle has received attention is because it is generated in models that have rather accurate forecast records (Helkie and Hooper, 1988; Cline, 1989; Hopper and Marquez, 1995). Thus to solve the elasticity puzzle at the expense of its essential feature–namely, the prediction of GDP being smaller than imports–is to solve it by assertion.

[5] Based on the mean of the point estimates of figure 1, appendix B shows that US imports will equal US GDP between 60 years and 100 years from now. The upper end of the range assumes an annual growth rate of per-capita income of 2% (Barro, 1997, p.47), an initial import-GDP ratio of 15%, and an income elasticity of 1.95; the lower end changes this elasticity to 2.66 which is the upper bound of the elasticity's 66% confidence interval. These calculations treat prices as fixed.

2.2 How has the Puzzle been Addressed?

The literature's approaches to solve the elasticity puzzle can be grouped in three categories: (1) separation of secular from cyclical forces; (2) relaxation of assumed price homogeneity; and (3) correction of biases in import-price data. These approaches share the view that the large income elasticity is the result of a misspecification bias.

Intuitively, suppose that the process generating the data is $\ln m_t = \ln y_t + \beta_3 \ln X_t$, where m is imports, y is real income, and X embodies the role of an omitted variable.[6] By assumption, the true income elasticity is unity and thus, for a given value of X_t, changes in income generate an import schedule with a slope of one (figure 4). Assuming that $\beta_3 > 0$, an increase in X_t raises imports for every level of income and thus shifts upwards the import schedule. I now assume that the process generating the data produces two observations, **a** and **b,** for two hypothetical values of y_t and X_t. These observations lie on two import schedules because the increase in X_t shifts up the import schedule.

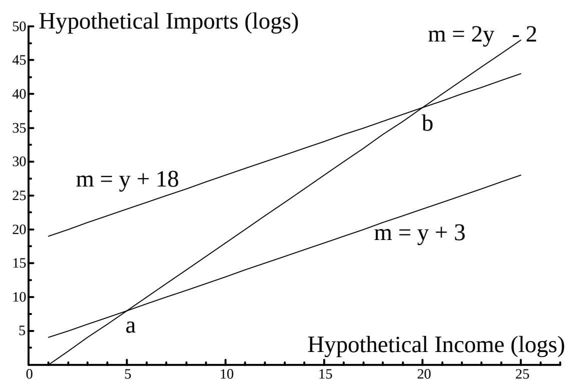


Figure 4: Model Misspecification and Estimated Income Elasticity

Suppose now that the econometric model used to estimate the income elasticity is $\ln m_t = \gamma \ln y_t + \varphi$, which excludes X. As figure 4 makes clear, the estimated income elasticity, $\widehat{\gamma}$, must exceed one. This upward bias in $\widehat{\gamma}$ stems from assuming that the increase in imports is due only to an increase in income–that is, from attributing to the income coefficient the effect of X on imports.

The rest of this section describes how previous work has addressed the puzzle–that is, the effort to characterize X. I start, however, by replicating the results of Houthakker and Magee (1969), a task not undertaken before. Such a replication is relevant to establish the robustness of their puzzling estimate.

2.2.1 Houthakker-Magee Redux

Houthakker and Magee assume that foreign and domestic products are imperfect substitutes for each other and that income and price elasticities are constant. Their formulation is

$$\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_2 \ln p_t + u_t, \beta_1 > 0, \beta_2 < 0,$$

[6] I set the relative price of imports to one and thus its logarithm is zero.

where m is merchandise imports; y is real GNP; p is the price of imports relative to the wholesale price index; and u is a white noise disturbance. Applying ordinary least squares with annual data over 1951-66, Houthakker and Magee obtain an estimate of 1.5 for β_1 (table 2). Following their steps, I virtually replicate their results and thus treat my estimates as theirs.[7] Given this treatment, their model's residuals are well behaved which rules out simple misspecifications of their model.[8]

Table 2: Income and Price Elasticities–1951-66

Elasticity Estimates	Original[a]	Replication
Income	1.51*	1.54*
Import Price/WPI	-0.54	-0.61
Time Trend	–	–
SER(%)	4.21	4.72
Residual Properties		
Serial Independence		0.26
Homoskedasticity		0.34
Normality		0.64

[a]Houthakker-Magee (1969), table 1; * statistically significant at the 5% level.

To study further the sensitivity of their results, I consider four modifications:

1. Including a time trend in their specification.
2. Modeling dynamic responses as

   $$\ln m_t = \beta_0 + \beta_{10} \ln y_t + \beta_{11} \ln y_{t-1} + \beta_{20} \ln p_t + \beta_{21} \ln p_{t-1} + \beta_3 \ln m_{t-1} + u_t,$$

   which is the formulation they applied to their quarterly data.
3. Expressing imports and income in per-capita terms.
4. Replacing the wholesale price index with the GNP deflator. Houthakker and Magee explicitly argue against using the GNP deflator because it includes services. Nevertheless this deflator ensures homogeneity of degree one in income and prices whereas using the wholesale price index violates that property.

These modifications do not eliminate the elasticity puzzle. If anything, they raise the income elasticity from 1.54 to at least 1.8 (table 3).[9]

[7]I do not replicate their results exactly because I am not using their data. Specifically, I use 1992 prices for real GNP whereas Houthakker and Magee use 1958 prices. I am, however, using their definitions (merchandise imports, real GNP, merchandise import price relative to wholesale production price), frequency (annual), period (1951-66), and source (*International Finance Statistics*).

[8]An entry below 0.05 means that the corresponding null hypothesis can be rejected at the 5% significance level. The test for Serial Independence is an F-test of the null hypothesis that the coefficients of an AR(1) for the residuals are jointly equal to zero; for quarterly data I use an AR(5). The homoskedasticity t-test is for the null hypothesis that the variance of the residuals is constant. The test of normality is a χ^2 test of the null hypothesis that the distribution of the residuals is normal; see Hendry and Doornik (1996) for details and references.

[9]As predicted by Houthakker and Magee, however, the estimated price elasticity with the GNP deflator is lower than the one using the WPI. Though not shown, the residuals from these formulations satisfy serial independence, homoskedasticity, and normality.

Table 3: Sensitivity of Elasticity Estimates–1951-66

Formulation Characteristics	Income	Price
WPI, Time trend, Aggregate Variables	2.13*	-1.15*
WPI, Per-capita Variables	1.82*	-0.93*
Dynamics, WPI, Per-capita Variables	1.76*	-1.07*
Dynamics, GDP Deflator, Per-capita Variables	1.80*	-0.55

* Significant at the 5% level.

A more radical change involves using an optimization model that relaxes the assumed constancy of income and price elasticities. For this I use the Rotterdam model developed by Barten (1968):

$$w_{t-1}\Delta \ln m_t = \mu \Delta \ln y_t + \pi \Delta \ln p_t, \mu > 0, \pi < 0,$$

where $w = \frac{m \cdot p_m}{y \cdot p_y}$ is the expenditure share of imports, μ is the percentage of a one-dollar increase in income devoted to imports, and π is the compensated price effect. The income elasticity is $\frac{\mu}{w_t}$ which varies inversely with the expenditure share of imports.[10] Based on annual data from 1951 to 1966, the (OLS) estimates are

$$\begin{array}{cccc} w_{t-1}\Delta \ln m_t = & -0.0699 & +7.396\Delta \ln y_t & -4.051\Delta \ln p_t \\ (se) & (0.05) & (1.54) & (0.91) \end{array}$$

	Null Hypothesis (p-value)	
$R^2 = 0.67; SER = 12.8\%$		
Sample: 1951-1966	Serial-Independence (0.56)	Normality (0.62)
	Homoskedasticity (0.57)	

The coefficient estimates are statistically significant and do not violate theoretical priors; the residuals satisfy serial independence, homoskedasticity, and normality. Interestingly, the estimated income elasticity is close to that of the log-linear model (figure 5) over 1951-1966. This finding suggests that the elasticity puzzle noted by Houthakker and Magee is not the result of ignoring optimization as modeled then.

The results also suggest that the log-linear formulation is a useful approximation to the Rotterdam model during 1951-66. If this usefulness held outside the estimation sample, then one would expect that subsequent studies using the log-linear approximation would have reported lower income elasticities given the increase in the import-GDP ratio in their estimation samples. Indeed, out-of-sample extrapolations of the income elasticity from the Rotterdam model exhibit an unmistakable trend to homotheticity (figure 5). Subsequent studies report, however, nothing like it as figure 1 above shows.

[10] Though the Rotterdam model was available in 1969, its use was not widespread. Thus I am not using it here to criticize the work of Houthakker and Magee (1969) but rather to see if it accounts for the elasticity puzzle.

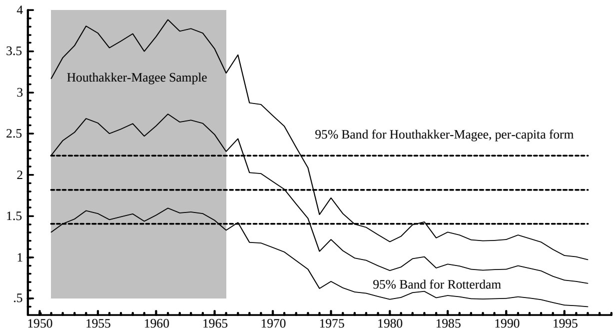


Figure 5: 95% Bands for Income Elasticities: Rotterdam and Log-linear Models

2.2.2 Separation of secular and cyclical forces

Interest in separating secular from cyclical forces stems from the differential response of imports to these forces. Secular forces, such as changes in comparative advantage, exert their influence gradually whereas cyclical forces, such as changes in inventories and production bottlenecks, exert their influences swiftly. Recognizing that published measures of income and prices embody secular and cyclical forces, Khan and Ross (1975) propose a method to let least squares differentiate their effects on imports. Specifically, they specifying the import-demand equation as

$$\ln m_t = \beta_0 + \beta_{1s} \ln y_t^s + \beta_{1c}(\ln y_t - \ln y_t^s) + \beta_2 \ln p_t,$$

where y_t^s represents secular (potential) income and $(\ln y_t - \ln y_t^s)$ measures the cyclical component as the deviation between actual and potential output. In the absence of published data for y_t^s, one may assume that $\ln y_t^s = \theta \cdot trend$ to obtain

$$\begin{aligned} \ln m_t &= \beta_0 + \beta_{1s} \cdot \theta \cdot trend + \beta_{1c} \cdot \ln y_t - \beta_{1c} \cdot \theta \cdot trend + \beta_2 \ln p_t \\ &= \beta_0 + \beta_{1c} \cdot \ln y + \beta_2 \ln p_t + \theta \cdot (\beta_{1s} - \beta_{1c}) \cdot trend. \end{aligned}$$

In this case, the trend is the X variable and if $\beta_{1s} - \beta_{1c} \neq 0$, then an equation that excludes the trend will induce a bias in the income elasticity. Moreover, the estimated income elasticity will embody cyclical effects and, therefore, it carries no implications for the long run: ignoring the trend overstates the relevance of the elasticity puzzle.

Though simple to implement, this approach restricts secular forces operating over income and prices to have the same effect on imports, as Haynes and Stone (1983) note. Their approach involves estimating (not with OLS)

$$\ln m_t^s = \beta_{0s} + \beta_{1s} \ln y_t^s + \beta_{2s} \ln p_t^s$$

$$\ln m_t^c = \beta_{0c} + \beta_{1c} \ln y_t^c + \beta_{2c} \ln p_t^c,$$

where the superscripts s and c denote the secular and cyclical observations which are generated by using a spectral decomposition of the original series. Their method lowers the estimated income elasticity from 1.9 to 1.5 (Haynes and Stone, 1983, table 3) which attenuates but does not solve the puzzle.

Why is this method not employed more widely? Because of the difficulties of integrating spectral methods into models to forecast ex-ante imports based on assumptions about future GDP and relative prices. Indeed, to integrate the spectral method into a forecasting model involves specifying, in advance, the mix of secular and cyclical factors associated with ex-ante paths for GDP and prices. Because this mix is not unique, having separate estimates for secular and cyclical factors is not particularly helpful. An alternative approach involves forecasting separately y_t^s and y_t^c, which is straightforward because these two variables are periodic functions of time. But the mechanistic character of this approach means that it cannot take into account recent developments that would influence a given GDP forecast.

2.2.3 Relaxation of price homogeneity

Arguing that price homogeneity need not hold at the aggregate level, Murray and Ginman (1976) use

$$\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_{2y} \ln p_{yt} + \beta_{2m} \ln p_{mt}, \beta_{2y} > 0, \beta_{2m} < 0$$

where p_{yt} is the price for the domestic product and p_{mt} is the price for imports. To cast formulation in terms of omitted variables, I add and subtract $\beta_{2y} \ln p_{mt}$ and get

$$\begin{aligned} \ln m_t &= \beta_0 + \beta_1 \ln y_t + \beta_{2y} \ln p_{yt} + \beta_{2m} \ln p_{mt} + \beta_{2y} \ln p_{mt} - \beta_{2y} \ln p_{mt} \\ &= \beta_0 + \beta_1 \ln y_t + \beta_{2y} \cdot (\ln p_{yt} - \ln p_{mt}) + (\beta_{2m} + \beta_{2y}) \cdot \ln p_{mt} \\ &= \beta_0 + \beta_1 \ln y_t + \beta_{2y} \ln p_t + (\beta_{2m} + \beta_{2y}) \cdot \ln p_{mt}. \end{aligned}$$

In this case, $\ln p_{mt}$ is the X variable and if $\beta_{2m} + \beta_{2y} \neq 0$, then an equation that excludes $\ln p_{mt}$ will induce a bias in the income elasticity. Murray and Ginman reduce the income elasticity for US imports from 1.9 to 1.4 (Murray and Ginman, 1976, table 2). Stern, Baum, and Green (1979), however, also relax price homogeneity and find that the income elasticity is one. For models relaxing the price-homogeneity assumption, no technical considerations complicate ex-ante forecasting, but this approach is not popular because the predictions embody a violation of rational behavior.

2.2.4 Evidence with Recent Observations

To examine whether combining recent observations with previous methods lowers the estimated income elasticity of US imports, I apply OLS to

$$\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_{2y} \ln p_{yt} + \beta_{2m} \ln p_{mt} + \beta_3 \cdot trend$$

Based on with quarterly data over 1967-97, the conventional elasticity puzzle remains intact, a result robust to commodity disaggregation (table 4). Furthermore, the results imply that aggregate per-capita imports would grow at an annual rate of 3.6 percent (=0.89×4) even if income and relative prices were to remain, literally, fixed.

Table 4: Trend and Long-Run Elasticities for US Imports–1967-97*

	Services	Production	Consumption	Aggregate
Income β_1	1.53*	2.01*	3.47*	2.01*
Domestic Price β_{2y}	1.42*	-1.30*	1.35*	-0.47*
Import Price β_{2m}	-1.25*	0.06	-0.96*	-0.25*
Trend×100 β_3	-0.02	1.88*	-0.39*	0.89*

* Significant at the 5% level; both imports and income are expressed in per-capita terms.

Recognizing that a trend is not an ideal proxy for secular effects (the criticism of Haynes and Stones, 1983), I apply least squares to data filtered with the Hodrick-Prescott filter using a weight of 1600. The elasticity puzzle remains intact (table 5).

Table 5: HP Filter and Long-Run Elasticities for US Imports–1967-97*

	Services	Production	Consumption	Aggregate
Income	1.96*	2.69*	3.76*	2.67*
Relative Price	-1.30*	-0.33*	-0.71*	-0.44*

* Significant at the 5% level

Overall, then, efforts to address the elasticity puzzle have been insightful but not successful: estimated income elasticities based on recent observations exceed one by a significant margin. In the next two sections I solve the elasticity puzzle by addressing two factors: the substitution bias stemming from the omission of import prices of new products in official statistics and the heterogeneity of tastes for imports induced by immigration into the United States. Both factors were negligible for the sample of Houthakker and Magee, but have reached historical heights since then.

3 Imports and Prices of New Products

Helkie and Hooper (1988) note that data for import prices have deficiencies by failing to incorporate the prices of new products, especially those from developing countries which are, generally, lower than the prices of existing products.[11] This inadequacy in sampling induces a substitution bias in the import-price data which understates the fall of import prices shown in figure 3. Thus recording transaction prices properly would be reflected in relative-price declines greater than those shown in figure 3. Such additional declines would increase the role of prices in explaining the upward trend in the GDP share of imports, diminish the burden on income to explain that trend, and thus lower the estimated income elasticity. Helkie and Hooper argue that, until data-collection methods improve, one can model the upward bias in import prices as a function of increases in foreign supply. Thus they amend the Houthakker-Magee model as

$$\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_2 \ln p_t + \beta_3 \ln \left(\frac{K_{for,t}}{K_{us,t}} \right), \beta_3 > 0 \ ,$$

[11] Hooper and Richardson (1991) offer a collection of studies on the question of how to measure international prices and the consequences of measurement choices for practical questions.

where K_{us} is the US capital stock and K_{for} is the foreign capital stock. Their estimated income elasticity is, however, 2.1 which retains the puzzle (Helkie and Hooper 1988, table 2-4).

Feenstra (1994) takes up the task of improving the data and generating import prices that correct for the entrance of new products. He shows that the "true" import price equals the recorded import prices times a bias that varies over time (Feenstra 1994, page 159, Proposition 1):

$$p_{cm,t} = p_{m,t} \cdot \left(\frac{\lambda_t}{\lambda_{t-1}} \right)^{\frac{1}{\sigma-1}},$$

where

p_{cm} = correct price of imports–incorporates new products,
p_m = official price of imports–excludes new products,
λ = share of existing products relative to all products,
σ = elasticity of substitution among products, $\sigma > 1$.

Feenstra collects annual data over 1967-1987 on prices and quantities for selected imports and constructs the correct import-price indexes.[12] He finds significant biases relative to official data for import prices and, to emphasize the importance of these biases, Feenstra estimates the income elasticity for US imports with official and correct price data. His findings indicate a reduction of the income elasticity from 1.66 (official data), to 1.37 (correct data) for Athletic shoes, from 1.29 to 1.1 for Steel bars, and from 3.05 to 2.28 for TV receivers (Feenstra 1994, table 4).

These are important findings but they apply to a about 1.1% of US imports raising a question about their generality. To address this question, Feenstra and Shiells (1994) construct price data that correct for such biases for the aggregate of non-oil imports and apply least squares to

$$\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_2 \ln \left(\frac{p_{mt}}{p_{yt}} \right) + \frac{\beta_2}{(\sigma - 1)} \cdot \ln \left(\frac{\lambda_t}{\lambda_{t-1}} \right).$$

Based on quarterly data ending in 1988, they reduce the income elasticity from 2.6 to a statistically insignificant 1.7 (Feenstra and Shiells 1994, table 3). Again, the reduction of the income elasticity is large but the statistical insignificance is problematic.

Why is this method not employed more widely for ex-ante forecasting? Because the data needed to compute λ_t becomes available with delays measured in years. One can bypass this drawback by modeling historical observations on λ_t with variables that are updated frequently and thus relatively easy to extrapolate, which is what I propose here by combining the methods of Feenstra (1994) and Helkie and Hooper (1988). Specifically, I start with a specification using the correct relative price:

$$\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_2 \ln \left(\frac{p_{cm}}{p_y} \right)_t$$

[12] The products are men's leather athletic shoes, men's and boy's cotton knit shirts, stainless steel bars, carbon steel sheets, color TV receivers (over 17" in size), portable typewriters, gold bullion, and silver bullion (Feenstra 1994, table 1). The total value of these imports, in 1987, is \$4.7 billion.

Recognizing the lack of data for p_{cm}, I invoke Feenstra's result

$$p_{cm,t} = p_{m,t} \cdot \left(\frac{\lambda_t}{\lambda_{t-1}} \right)^{\frac{1}{\sigma-1}},$$

and reformulate the import equation as

$$\begin{aligned}
\ln m_t &= \beta_0 + \beta_1 \ln y_t + \beta_2 \ln \left(\frac{p_{mt} \cdot \left(\frac{\lambda_t}{\lambda_{t-1}} \right)^{\frac{1}{\sigma-1}}}{p_{yt}} \right) \\
&= \beta_0 + \beta_1 \ln y_t + \beta_2 \ln \left(\frac{p_{mt}}{p_{yt}} \right) + \beta_2 \ln \left(\frac{\lambda_t}{\lambda_{t-1}} \right)^{\frac{1}{\sigma-1}} \\
&= \beta_0 + \beta_1 \ln y_t + \beta_2 \ln \left(\frac{p_{mt}}{p_{yt}} \right) + \frac{\beta_2}{(\sigma-1)} \cdot \ln \left(\frac{\lambda_t}{\lambda_{t-1}} \right),
\end{aligned}$$

which is the formulation used by Feenstra and Shiells (1994). Following Helkie and Hooper, I now assume that

$$\ln \left(\frac{\lambda_t}{\lambda_{t-1}} \right) = \theta \cdot \ln \left(\frac{K_{for,t}}{K_{us,t}} \right), \theta < 0$$

where $\theta < 0$ means that an increase in foreign capital relative to US capital introduces new products and therefore *lowers* the expenditure share on existing products λ_t. With this assumption I get

$$\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_2 \ln \left(\frac{p_{mt}}{p_{yt}} \right) + \frac{\beta_2}{(\sigma-1)} \cdot \theta \cdot \ln \left(\frac{K_{for,t}}{K_{us,t}} \right),$$

where $\frac{\beta_2}{(\sigma-1)} \cdot \theta > 0$ because $\beta_2 < 0, \sigma > 1$, and $\theta < 0$.

I measure the foreign capital stock as a geometric weighted average of capital stock indexes of 14 developing countries, in real terms:

$$K_{for,t} = \prod K_{it}^{w_i}$$

where K_{it} is an index of the capital stock of the *ith* country and w_i is the share of US imports from the ith country; the w_i are normalized to sum to one. The countries (weights in percent) are Argentina (0.7%), Brazil (3.9%), China (12.9%), Chile (0.9%), Hong Kong (9.87%), Indonesia (2.75%), Korea (9.0%), Malaysia (5.6%), Mexico (23.5%), Philippines (2.5%), Singapore (7.5%), Taiwan (12.4%), Thailand (4.2%), and Venezuela (4.3%). For $K_{us,t}$ I use the non-residential capital stock constructed as the sum of producer durables and equipment.[13] Figure 6 below shows that the US capital stock has grown at a slower pace than the capital stock of these developing countries. The differential in growth rates embodies the rate at which the United States imports new products from these developing countries. One limitation of my approach, as currently implemented, is that the data for the capital stock of developing countries starts in 1984 which shortens the number of observations.

[13] The data for the capital stock come from the FRB/Global model; Brayton et al. (1997) and Levin et al. (1997).

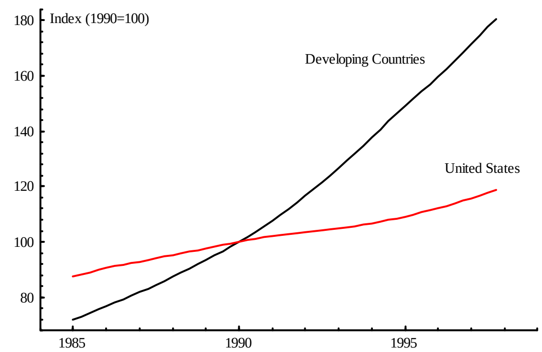


Figure 6: Capital Stocks–U.S. and Developing Countries.

I now focus on removing the assumption of the representative agent by assuming that agent heterogeneity stems from immigration.

4 Imports and Immigration

Priors That there might be an association between imports and immigration is evident in figure 7.[14] For the first half of this century both the import-GDP ratio and the share of foreign-born population decline; both trends are reversed in the postwar period.[15]

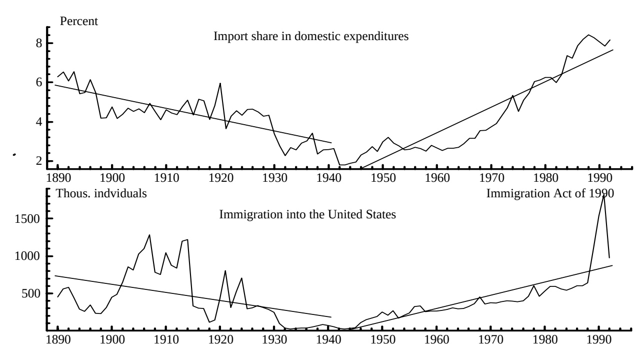


Figure 7: Immigration Flows and Non-oil Imports–United States, 1890-1992

This association has a regional counterpart (table 6). Indeed, the share of US imports and immigration from Asia increase from the smallest in 1970 to the largest in 1995. In contrast, the shares of US imports and immigration from Europe drop from the second largest in 1970 to the smallest in 1995. The US share of imports and immigration from North America change little in 25 years, despite the immigration surge in 1990.

[14] I use the Bureau of Census' definition of immigrants: Those nonresident aliens admitted to the United States for permanent residence. This definition excludes nonresident aliens coming to the United States for a temporary period or those foreign nationals living in the United States with the intention of becoming permanent residents but waiting to meet the eligibility criteria.

[15] Data for non-oil imports, per-capita expenditures, and relative prices come from Marquez (1999). Data for immigration come from Mitchell (1998), pages 93-96.

Table 6: U.S. Bilateral Imports and Immigration Shares[a] (%)

	1970	1975	1980	1985	1990	1995
Immigration from:						
Europe	31.1	19.2	13.6	11.1	7.3	17.8
Asia	25.4	34.3	44.5	46.4	22.0	37.2
North America[b]	34.6	38.0	31.1	31.9	62.3	32.1
Imports from:						
Europe	30.6	23.4	21.1	24.7	22.7	19.9
Asia	23.2	22.8	25.8	36.9	38.2	42.4
North America	30.6	24.5	21.3	24.6	24.1	27.4

[a] Source: Appendix A; [b]Canada and Mexico

Modeling My goal is to translate these observations into a formal model of imports showing that an increase in the share of foreign-born raises imports for given income and relative prices. I start by assuming a log-linear import demand for the ith individual depending *only* on income and relative prices:

$$\ln m_{it} = \beta_{0i} + \beta_{1i} \ln y_{it} + \beta_{2i} \ln p_{it} + u_{it}, \tag{1}$$

where β_{0i} is the foreign-product bias which is negative if the *ith* individual prefers to "buy American." Other formulations are available but I focus here on the log-linear model to avoid confusing the roles of model choice and immigration in accounting for the elasticity puzzle.[16]

Adding across individuals and scaling by population N_t yields

$$\frac{\sum_i^{N_t} \ln m_{it}}{N_t} = \frac{\sum_i^{N_t} \beta_{0i}}{N_t} + \frac{\sum_i^{N_t} \beta_{1i} \ln y_{it}}{N_t} + \frac{\sum_i^{N_t} \beta_{2i} \ln p_{it}}{N_t} + \frac{\sum_i^{N_t} u_{it}}{N_t}. \tag{2}$$

If the distributions of income and price effects are symmetric–that is, if

$$\lim_{N_t \to \infty} \left(\frac{\sum_i^{N_t} (\beta_{1i} - \beta_1) \ln y_{it}}{N_t} \right) = 0,$$

$$\lim_{N_t \to \infty} \left(\frac{\sum_i^{N_t} (\beta_{2i} - \beta_2) \ln p_{it}}{N_t} \right) = 0,$$

[16] The formal derivation of (1) is due to Senhadji (1998) who uses Houthakker's (Houthakker 1960) Addilog utility function in an intertemporal, stochastic environment. The ith individual seeks to maximize $E_0 \sum_{t=0}^{\infty} \frac{u(d_t, m_t)}{(1+\delta)}$ subject to

$$b_{t+1} = (1+r)b_t + (e_t - d_t) - p_t m_t$$
$$e_t = (1-\rho)\overline{e} + \rho e_{t-1} + \theta_t, \ \theta_t \tilde{} N(0, \sigma_\theta)$$
$$u(d_t, m_t) = \frac{A_t d_t^{1-\alpha}}{1-\alpha} + \frac{B_t m_t^{1-\beta}}{1-\beta},$$

where d_t is the demand for domestic goods, m_t is the demand for imports, b_{t+1} is net foreign wealth, δ is the discount rate, r is the interest rate, e_t is the endowment, p_t is the price of imports relative to the price of domestic products, A_t and B_t are preferences parameters. Solving the first-order conditions for optimization yields (1) above.

where $\beta_j = \left(\frac{\sum_i^{N_t} \beta_{ji}}{N_t} \right)$, then

$$\frac{\sum_i^{N_t} \ln m_{it}}{N_t} = \frac{\sum_i^{N_t} \beta_{0i}}{N_t} + \left(\frac{\sum_i^{N_t} \ln y_{it}}{N_t} \right) \beta_1 + \left(\frac{\sum_i^{N_t} \ln p_{it}}{N_t} \right) \beta_2 + u_t, \tag{3}$$

where $u_t = \frac{\sum_i^{N_t} u_{it}}{N_t}$.

Equation (3) is hard to implement empirically because data for the means of the logarithms, as opposed to data for the logarithm of the mean, are not available. Thus I assume that the *logarithms* of relative prices, of per-capita imports, and per-capita income are normally distributed which implies that (Klein 1962, p. 155)

$$\frac{\sum_i^{N_t} \ln m_{it}}{N_t} = \ln \left(\frac{\sum_i^{N_t} m_{it}}{N_t} \right) - \frac{\sigma_{mt}^2}{2} = \ln(m_t) - \frac{\sigma_{mt}^2}{2}, \tag{4}$$

$$\frac{\sum_i^{N_t} \ln y_{it}}{N_t} = \ln \left(\frac{\sum_i^{N_t} y_{it}}{N_t} \right) - \frac{\sigma_{yt}^2}{2} = \ln (y_t) - \frac{\sigma_{yt}^2}{2}, \tag{5}$$

$$\frac{\sum_i^{N_t} \ln p_{it}}{N_t} = \ln \left(\frac{\sum_i^{N_t} p_{it}}{N_t} \right) - \frac{\sigma_{pt}^2}{2} = \ln (p_t) - \frac{\sigma_{pt}^2}{2}, \tag{6}$$

where σ_{mt}^2, σ_{yt}^2, and σ_{pt}^2 are the variances of the distributions of the *logarithms* of per-capita imports, income, and relative prices, respectively. Substituting (4), (5), and (6) into (3) yields

$$\ln(m_t) = \frac{\sum_i^{N_t} \beta_{0i}}{N_t} + \beta_1 \ln y_t + \beta_2 \ln p_t + \frac{\sigma_{mt}^2}{2} - \beta_1 \frac{\sigma_{yt}^2}{2} - \beta_2 \frac{\sigma_{pt}^2}{2} + u_t. \tag{7}$$

Previous work with log-linear models sets $\beta_{0i} = \beta_0$ and $\sigma_{mt}^2 = \sigma_{yt}^2 = \sigma_{pt}^2 = 0$. These settings amount to assuming the representative-agent model in a situation where the usefulness of representativeness is being questioned. But bypassing this assumption is hard because data for $\frac{\sum_i^{N_t} \beta_{0i}}{N_t}$, σ_{mt}^2, σ_{pt}^2, and σ_{yt}^2 are hard to get. As an alternative, I model these moments in terms of observable magnitudes.

First, following the evidence from table 6 above, I assume that increases in the share of foreign-born in US population raises the mean of the foreign-product bias:

$$\frac{\sum_i^{N_t} \beta_{0i}}{N_t} = \beta_{00} + \beta_{01} \ln(I_t) + \epsilon_{\beta_{0,t}}, \tag{8}$$

where I_t is the share of foreign-born in US population and $\epsilon_{\beta_{0,t}}$ is a random term. Intuitively, if immigrants retain their tastes for their native products, then a ceteris paribus increase in immigration raises the demand for imports. The term $\frac{\sum_i^{N_t} \beta_{0i}}{N_t}$ represents the average propensity to buy foreign products which, in the presence of immigration, increases in response to a re-orientation of preferences for given prices and income, however measured.

Second, I assume that the bulk of immigrants come from countries with per-capita income sufficiently different from that of the United States so that an increase in immigration raises the dispersion of the income distribution σ_{yt}^2 :[17]

$$\frac{\sigma_{yt}^2}{2} = \sigma_{y0} + \sigma_{y1} \ln(I_t) + \epsilon_{\sigma_y,t}, \tag{9}$$

where ϵ_{σ_y} is a random term. Empirical support for equation (10) rests on two sources. First, Borjas, Freeman, and Katz (1991) find that immigration contributes to the dispersion of wages. Second, the Gini coefficient (a measure of σ_{yt}^2) and the share of foreign-born population are closely related (figure 8).[18]

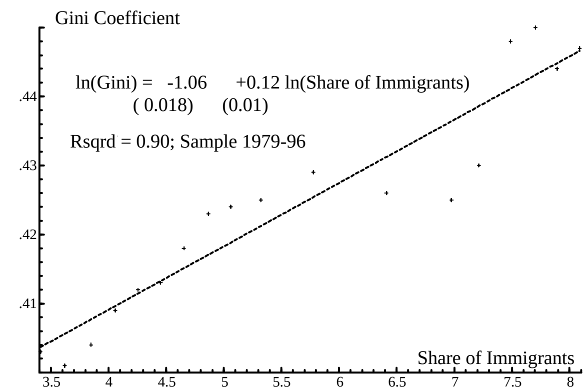


Figure 8: Gini Coefficient and Immigration

Third, I postulate that an increase in immigration changes the composition of the typical consumption basket and thus raises the variance of per-capita imports:

$$\frac{\sigma_{mt}^2}{2} = \sigma_{m0} + \sigma_{m1} \ln(I_t) + \epsilon_{\sigma_m,t}, \tag{10}$$

where ϵ_{σ_m} is a random term. The consumption surveys reported in Gould (1994) support this assumption.

Finally, following Gould (1994), I assume that the foreign-born have, compared to the native-born, an advantage in terms of information about prices, language, customs, and regulations of foreign products. This information differential translates into a widening of the distribution of relative prices for imports paid by US importers:

$$\frac{\sigma_{pt}^2}{2} = \sigma_{p0} + \sigma_{p1} \ln(I_t) + \epsilon_{\sigma_p,t}, \tag{11}$$

where ϵ_{σ_p} is a random term.

[17] For evidence of a dispersion of median income across immigrants by country, see the Statistical Abstract of the United States (1986, table No. 40).

[18] Because the Gini coefficient and the share of immigrants are restricted to take positive values, I report in the graph the regression using logarithms of these variables. The data source for the Gini coefficient is US Census Bureau, Historical Income Tables: Experimental Measures, Table RDI-5 available at http://www.census.gov/hhes/income/histinc/rdi05.html

Substituting (8)-(11) into (7) yields

$$\begin{aligned} \ln m_t &= \beta_0 + \beta_1 \ln y_t + \beta_2 \ln p_t + \beta_3 \ln(I_t) + \epsilon_t \\ \epsilon_t &= u_t + \epsilon_{\sigma_{m,t}} - \beta_1 \epsilon_{\sigma_y,t} - \beta_2 \epsilon_{\sigma_p,t} + \epsilon_{\beta_{0,t}}, \end{aligned} \quad (12)$$

which is the Houthakker-Magee model augmented to allow for the role of immigration with $\beta_0 = \beta_{00} + \sigma_{m0} + \sigma_{y0} + \sigma_{p0}$ and $\beta_3 = \sigma_{m1} - \beta_1 \sigma_{y1} - \beta_2 \sigma_{p1} + \beta_{01}$. Thus the modeling strategy retains identifiability of the income and price elasticities but not of the immigration effect. Nevertheless, if $\sigma_{p1} = \sigma_{y1}$, $\beta_2 = -1$, and $\beta_1 = 1$, then $\beta_3 = \sigma_{m1} + \beta_{01} > 0$. Alternatively, if $\frac{\sigma_{m1} + \beta_{01} + \beta_2 \sigma_{p1}}{\sigma_{y1}} > \beta_1$ then $\beta_3 > 0$.

Comments on Equation (12)

1. Equation (12) generalizes previous work by removing the assumption of a representative agent. Reinstating that assumption involves setting $\sigma_m^2 = \sigma_y^2 = \sigma_p^2 = \beta_{01} = 0$ changing eq. (12) to

$$\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_2 \ln p_t + u_t,$$

which is the log-linear model without immigration.[19]

2. Immigration matters for modeling imports even if data for σ_{mt}^2, σ_{yt}^2, and σ_{pt}^2 were available. In that case one would re-express the variables as

$$m_t^* = m_t e^{-\frac{\sigma_{mt}^2}{2}}; \; y_t^* = y_t e^{-\frac{\sigma_{yt}^2}{2}}; \; p_t^* = p_t e^{-\frac{\sigma_{pt}^2}{2}}$$

and then apply least squares to

$$\ln(m_t^*) = \frac{\sum_i^{N_t} \beta_{0i}}{N_t} + \beta_1 \ln y_t^* + \beta_2 \ln p_t^* + u_t.$$

To the extent that the foreign-product bias, $\frac{\sum_i^{N_t} \beta_{0i}}{N_t}$, is not fixed in the presence of immigration, there is a re-orientation of preferences for given prices and income, however measured. Modeling this foreign bias with equation (12) would yield

$$\ln m_t = \beta_{00} + \beta_1 \ln y_t + \beta_2 \ln p_t + \beta_{01} \ln(I_t) + u_t + \epsilon_{\beta_{0,t}}.$$

3. Given that little is known a priori about the distributions of $\epsilon_{\sigma_{m,t}}$, $\epsilon_{\sigma_y,t}$, $\epsilon_{\sigma_p,t}$, and $\epsilon_{\beta_{0,t}}$, there is no presumption that $\epsilon_t = u_t + \epsilon_{\sigma_{m,t}} - \beta_1 \epsilon_{\sigma_y,t} - \beta_2 \epsilon_{\sigma_p,t} + \epsilon_{\beta_{0,t}}$,is white noise. Thus I emphasize testing the properties of ϵ_t to ensure that inferences do not violate the maintained assumptions.

[19] Notice that this case differs from the one where the entire population consists of immigrants for in that case $I_t = 1$ and $\beta_{01} = 0$ giving $\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_2 \ln p_t + \epsilon_t$,which is *not* observationally equivalent to the case of the representative agent because $var(\epsilon_t) \neq var(u_t)$.

I construct data for the share of foreign-born population, I, as F/N where F is the stock of foreign-born residents and N is resident population. Data for the stock of foreign-born residents are constructed as $F_t = f_t + (1 - \omega_t)F_{t-1}$, where f_t is the flow of immigrants and ω_t is the mortality rate of immigrants. I assume an ω of 8.9 per 1000 (annual rate), the same as that of the native population. As a benchmark for foreign-born resident population, I use $F_{1970} = 3.322$ million permanent residents (Statistical Abstract 1986, table 127). I use the 1970 benchmark because, prior to 1953, the official definition of immigrant changed many times undermining the comparability of the values across periods. Based on these assumptions, figure 9 shows the upward trend in the flow of immigrants which has raised the share of foreign-born residents in total US population to more than eight percentage points by 1997. Appendix C deals with measurement errors stemming from illegal immigrants and the choice of benchmark.

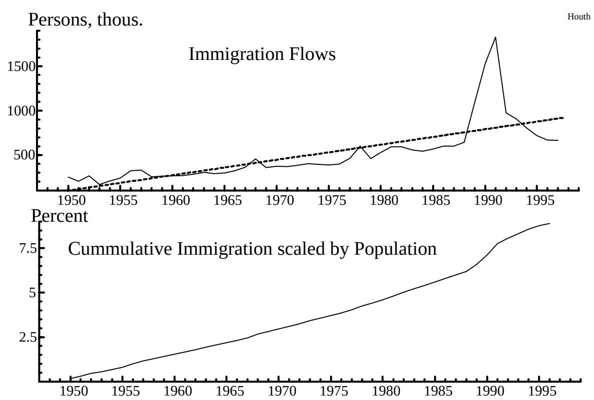


Figure 9: Immigration Flow (top) and Population Share of Foreign-born (bottom)

5 Econometric Analysis

5.1 Method

For parameter estimation, I use the cointegration method of Johansen (1988) which avoids simultaneity biases and differentiates secular from cyclical effects. I apply this method to two models:

Price-Bias Model

$$\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_2 \ln p_t + \beta_3 \ln \left(\frac{K_{for,t}}{K_{us,t}} \right)$$

Immigration Model

$$\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_2 \ln p_t + \beta_3 \ln I_t$$

Implementing Johansen's method involves applying maximum likelihood to

$$\Delta z_t = \sum_{i=1}^{n} \Gamma_i \Delta z_{t-i} + \alpha\theta' z_{t-1} + \epsilon_t, \epsilon_t \sim NI(0, \Omega) \tag{13}$$

where $z_t' = (\ln m_t \ \ln y_t \ \ln p_t \ \ln X_t\ )$; n is the number of lags; m represents per-capita imports; y is per-capita real GDP; p is the price of imports relative to the GDP deflator; X_t is either $\left(\frac{K_{for,t}}{K_{us,t}}\right)$ or I_t; Γ_i is a 4x4 matrix of coefficients for cyclical effects; and $\alpha\theta'$ embodies the secular effects. Specifically,

$$\alpha\theta' = \begin{pmatrix} \alpha_{11} & ... & \alpha_{14} \\ ... & ... & ... \\ \alpha_{41} & ... & \alpha_{44} \end{pmatrix} \begin{pmatrix} \theta_{11} & ... & \theta_{41} \\ ... & ... & ... \\ \theta_{14} & ... & \theta_{44} \end{pmatrix},$$

where the elements of α measure the speed of adjustment and are known as loading coefficients; the vector $(\theta_{1i}...\theta_{4i}) = \theta_i'$ characterizes the *ith* secular (long-run) relation among $\ln m_t$, $\ln y_t$, $\ln p_t$, and $\ln X$. I determine the number of long-run relations with the tests of Johansen and Juselius (1990): $\lambda_{i,\max}$ for testing that there are i long-run relations and $\lambda_{i,trace}$ for testing that there are at most i long-run relations. Recognizing that estimates derived from the Johansen procedure are sensitive to the number of lags (n), I consider values of n from 2 to 16 and reject lag lengths that yield multiple cointegration vectors or violate priors from economic theory.

5.2 Results

Based on quarterly data through 1997, the results indicate that the estimated income elasticity for aggregate imports is 1.2 for the price-bias model and 1.1 for the immigration model (table 7); these estimates are statistically significant. The estimated price elasticity is -1.2 for the price-bias model and -0.5 for the immigration model; these estimates are statistically significant. The effects of price biases and immigration on aggregate imports are positive and significant.[20] Finally, the residuals of both models satisfy normality, serial independence, and homoskedasticity. Overall, the results indicate that the elasticity puzzle of the last three decades is the result of ignoring either the substitution bias induced by excluding prices for new products or the growing heterogeneity of individuals associated with immigration.

But which of these two formulation should be used? One can address this question by forming an augmented model including both factors and testing which of the two is the significant one. As an alternative, I exploit the theoretical implications from optimization for models with constant elasticities noted by Lau (1986, page 1527):

> "We conclude that (local) summability alone implies that the system of consumer demand functions must take the form:
>
> $$\ln X_i = \alpha_i - \ln p_i + \ln M_i$$
>
> which is no longer flexible. For this system, the own price elasticity is minus one, the cross-price elasticities are zeroes, and the income elasticity is unity for the demand function of each and every commodity."

[20] The coefficient for immigration is 0.32 which, when combined with the observations on per-capita imports and the share of foreign-born population, explains about two-fifths of the recorded increase of per-capita imports of goods and services over 1971-97. The change in imports is given by $m_{97}/m_{71} = (y_{97}/y_{71})^{1.147}(p_{97}/p_{71})^{-0.481}(I_{97}/I_{71})^{0.328}(other factors)$. Thus, substituting I get $3.955 = (1.608)^{1.147}(1)^{-0.481}(4.316)^{0.328}(other factors)$. The entry in the text is $\frac{(4.316)^{0.328}}{3.955}$.

Table 7: Long-Run Elasticities for US Imports–New-Product Prices and Immigration

	Aggregate		Production		Consumption		Services	
Income	1.18	1.14	0.69	1.12	1.41	1.31	1.56	1.21
	(0.24)	(0.06)	(0.23)	(0.04)	(0.13)	(0.06)	(0.22)	(0.06)
Price	-1.18	-0.48	-0.80	-0.60	-1.75	-1.10	-2.07	-1.14
	(0.21)	(0.14)	(0.14)	(0.10)	(0.29)	(0.14)	(0.48)	(0.13)
Price Bias	0.35	-	0.62	-	0.13	-	0.02	-
	(0.12)		(0.12)		(0.10)		(0.11)	
Immigration	-	0.32	-	0.26	-	0.73	-	0.35
		(0.08)		(0.05)		(0.04)		(0.03)
Tests								
Independence	0.39	0.52	0.80	0.62	0.04*	0.62	0.59	0.76
Normality	0.36	0.93	0.19	0.00*	0.83	0.00*	0.08	0.21
Homoskedasticty	0.87	0.05	0.71	0.61	0.98	0.61	0.36	0.52

Entries in parentheses are standard errors.

Reliance on this theoretical implication implies that the price-bias model is the relevant one for aggregate imports given that it has unitary elasticities. But what holds true for the aggregate need not hold true for the components and, thus, there is a natural interest in evaluating the potential informational losses induced by aggregation. To this end, table also reports estimation results for the three components of aggregate imports.

Imports of producer goods The estimated income elasticity is 0.7 for the price-bias model and 1.1 for the immigration model; these estimates are statistically significant. The estimated price elasticity is -0.8 for the price-bias model and -0.6 for the immigration model; these estimates are statistically significant. Both the price-bias and immigration variables have positive and significant effects on imports. The residuals of the price-bias model satisfy normality, serial independence, and homoskedasticity; the residuals for the immigration model violate normality. Overall, the price-bias model is the relevant one because it is the only one that satisfies the hypothesis of unitary elasticities.

Imports of consumer goods The estimated income elasticity is 1.4 for the price-bias model and 1.3 for the immigration model; these estimates are statistically significant. The estimated price elasticity is -1.8 for the price-bias model and -1.1 for the immigration model; these estimates are statistically significant. The price-bias variable is not significant whereas the immigration variable is significant; the relative high value (0.7) reflects the persistence of tastes for foreign products by the foreign-born. The residuals of the price-bias model lack serial independence whereas the residuals of the immigration model lack normality. Overall, the immigration model is the relevant one because it is the only one with unitary elasticities and the price-bias variable has no effect.

Imports of services The estimated income elasticity is 1.6 for the price-bias model and 1.2 for the immigration model; these estimates are statistically significant. The estimated price elasticity is -2.1 for the price-bias model and -1.1 for the immigration model; these estimates are statistically significant. The price-bias variable is not significant whereas the immigration variable is significant. The residuals of both models satisfy normality, serial independence, and homoskedasticity. Overall, the immigration model is the relevant one because it yields unitary elasticities and the price-bias variable has no significant effect.

Thus the analysis for disaggregated imports reveals there is no silver bullet that resolves the elasticity puzzle: the price-bias model is suitable only for imports of producer goods. For imports of consumption and services, the immigration model is the relevant one. How much is lost by ignoring this information and using instead the estimates for aggregate imports? Imports of production goods account for two-thirds of aggregate imports (figure 10). Thus using the price-bias model for explaining aggregate imports involves misrepresenting the behavior of 1/3 of US imports, which is not a trivial magnitude. In other words, both factors are needed to explain US imports.

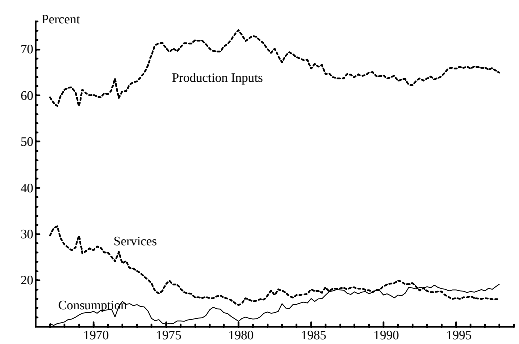


Figure 10: Commodity Composition of Aggregate Imports – Nominal Shares

What remains to be established is whether controlling for simultaneity, dynamics, and business cycles but omitting variables reflecting price-biases and immigration yields unitary income elasticities. Figure 11 compares the estimated income elasticities across categories for three models: price-bias, immigration, and conventional ($z_t' = (\ln m_t \ \ln y_t \ \ln p_t)$). The estimated income elasticity from the conventional model ranges from 1.6 for services to 2.6 for producer goods and it exceeds, by a significant margin, the income elasticity from the other two models. This increase in estimates is what figure 4 predicts: as long as one relies on official data for imports, prices, and income, the most popular formulation of imports will be misspecified and yield a biased income elasticity in excess of one.

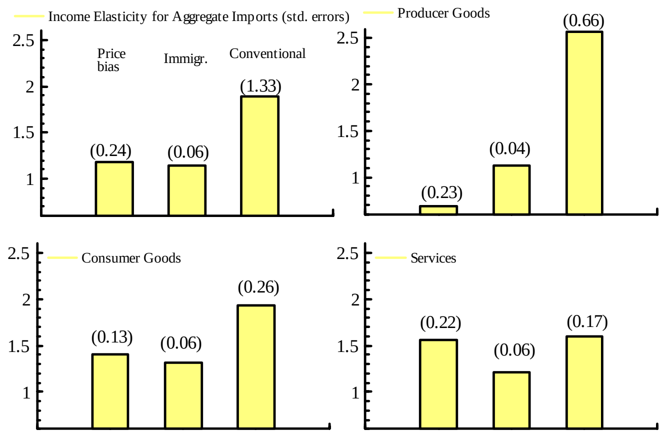


Figure 11: Estimated Income Elasticities–Sensitivity to Disaggregation and Omitted-variable Bias (std. errors).

5.3 Sensitivity of Results

Parameter Constancy I test parameter constancy with the Chow test. Specifically, I start by splitting the sample in 1983.4, and using the first sub-sample to obtain initial elasticity estimates. Second, I use these estimates to forecast imports over the post-estimation sample. Third, I test whether the forecast errors are jointly equal to zero with an F-test. Finally, I extend the first sub-sample by one quarter, update the elasticity estimates, and recompute the forecast tests. This process of moving forward the sample split one quarter at a time continues until all the observations are used. The result is a time-series of F-tests from 1984 to 1997. For the model with the price-bias variable, the first sample-split is in 1989.4 because of the short span available for this variable.

I test the whether forecast errors are zero over three horizons. The first horizon is the one-quarter-ahead prediction and is denoted as *1up*. The second horizon starts with the initial sample split (1983.4) and ends with the last observation of the *current* sample split. Because the current split moves forward one quarter at a time, this horizon increases from 1 quarter to 57 quarters; the test is denoted *Nup*. The third forecast horizon starts with the current sample split and ends in 1997.4, the last date. Because the current split increases one quarter at a time, the forecast horizon declines from 57 quarters to 1 quarter as the estimation sample expands; the test is denoted as *Ndn*. Figures 12-19 report the Chow tests results, which I scale by their 5% significance level; a crossing of the horizontal line means a rejection of the hypothesis of parameter constancy for that sample split. The left panels report results for the import equation and the right panels for the system as a whole (denoted as CHOW); see Hendry and Doornik (1996) for further details.

I cannot reject parameter constancy for imports of consumption, production goods, and the aggregate of goods and services. Finding parameter constancy at the aggregate level is in contrast to the findings of Hooper (1978) and Stern *et al.* (1979), who use rolling-regression techniques, and to those of Deyak, Sawyer, and

Sprinkle (1989) and Zietz and Pemberton (1993), who use Chow tests. These studies exclude immigration and do not recognize the role of new products' prices. Thus their results could reflect a misspecification bias. For imports of services, I detect parameter instability and further work on this category is needed.

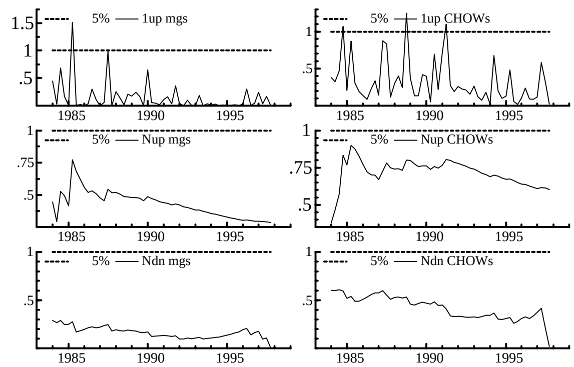


Figure 12: Chow Tests for Aggregate Imports–Immigration

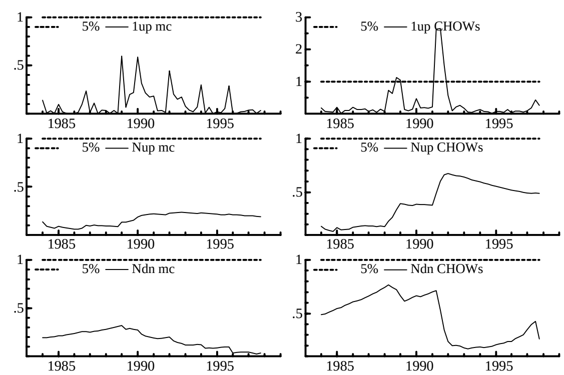


Figure 13: Chow Tests for Imports of Consumption Goods – Immigration

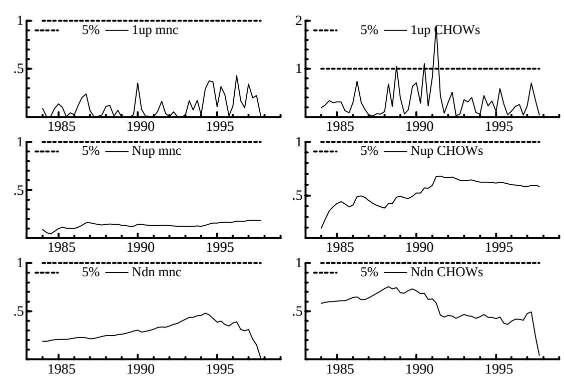


Figure 14: Chow Tests for Imports of Production Goods–Immigration

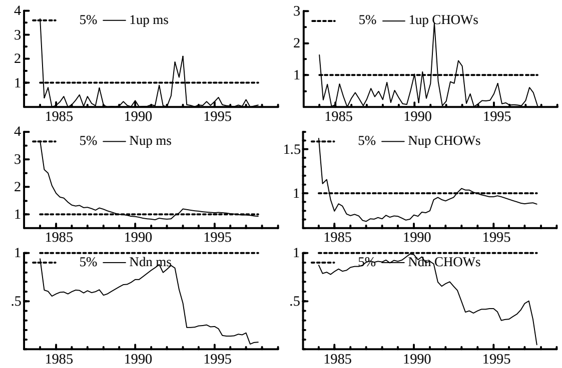


Figure 15: Chow Tests for Imports of Services–Immigration

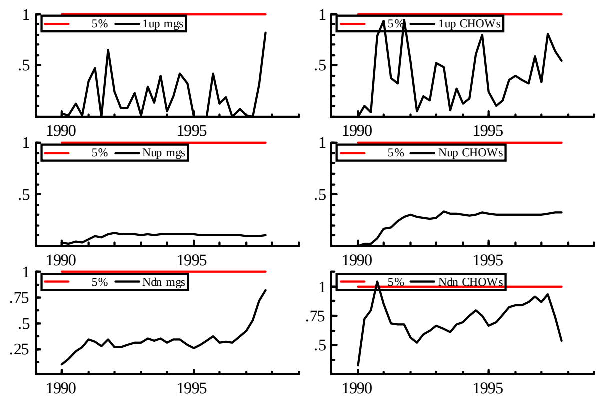


Figure 16: Chow Tests for Aggregate Imports–New-Product Prices

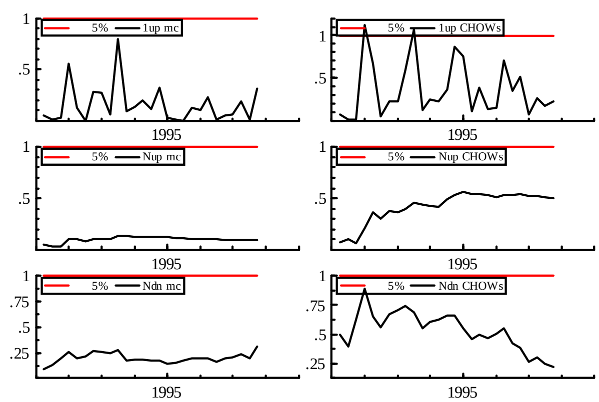


Figure 17: Chow Tests for Consumer Imports–New-Product Prices

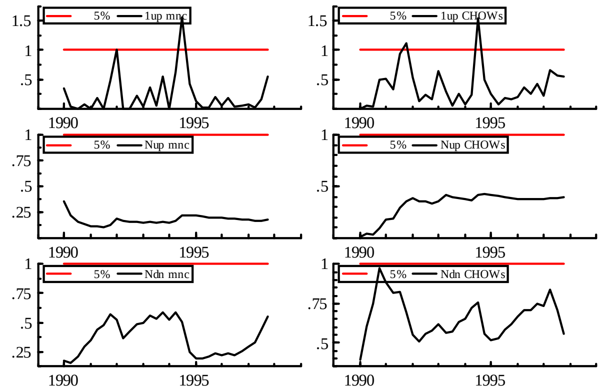


Figure 18: Chow Tests for Producer Imports–New-Product Prices

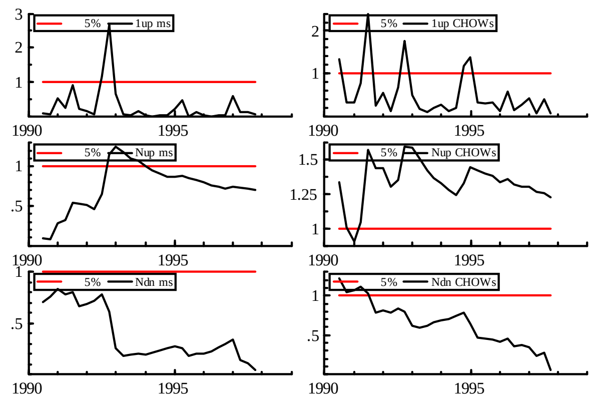


Figure 19: Chow Tests for Service Imports–New-Product Prices

5.3.1 Observational Equivalence to a Trend

The analysis developed thus far indicates that if one relaxes the assumption of a representative agent, recognizes the role of new products' prices, and uses Johansen's cointegration method, then one gets unitary income elasticities for imports of services, consumption, production goods, and the aggregate of these three categories. But the steady increases in $\left(\frac{K_{for,t}}{K_{us,t}}\right)$ (figure 6 above) and in the population share of foreign-born (figure 9 above), suggest the possibility that the results are observationally equivalent to those of a trend.

To this end, I re-estimate the parameters replacing $\ln I_t$ and $\ln\left(\frac{K_{for,t}}{K_{us,t}}\right)$ with a trend term which I then treat as an exogenous variable in the system; table 8 reports the cointegration results for each lag. With one exception, the results based on a time trend can be grouped into four categories: (1) no cointegration, (2) unique cointegration with implausible results, (3) unique cointegration with the elasticity puzzle intact, and (4) multiple cointegration.

Table 8: Import Elasticities and Trend

Lags	Services				Production Goods			
	Rank[a]	Income	Price	Trend[b]	Rank	Income	Price	Trend
16	1	3.79	-6.56	-1.61	2	0.72	-0.17	0.94
15	0	2.37	-3.73	-0.67	0	0.71	-0.13	0.94
14	0	-1.26	+3.68	2.35	0	0.70	-0.13	1.00
13	1	7.69	-14.68	-4.94	2	0.58	+0.21	0.76
12	1	6.86	-12.90	-4.46	2	0.67	-0.11	1.16
11	1	3.17	-5.39	-1.34	2	0.62	-0.01	1.21
10	1	-17.49	+36.68	16.96	2	0.57	-0.12	1.37
9	2	1.69	-2.34	-0.01	2	0.56	-0.04	1.30
8	1	1.78	-2.51	-0.19	2	-3.84	+3.88	14.15
7	1	1.84	-2.64	-0.26	1	7.43	-11.92	-9.07
6	1	2.31	-3.57	-0.65	1	2.85	-3.99	-1.74
5	1	2.30	-3.54	-0.64	1	3.39	-4.92	-3.02
4	1	1.80	-2.56	-0.21	2	0.23	+0.65	1.94
3	2	1.57	-2.09	0.00	2	0.62	-0.03	1.34
2	2	1.45	-1.85	1.01	2	1.52	-1.58	-0.4

Lags	Consumption Goods				Aggregate Goods and Services			
	Rank	Income	Price	Trend	Rank	Income	Price	Trend
16	0	-3.37	+7.82	4.32	0	0.84	-0.32	0.86
15	0	6.87	-13.03	-3.95	0	0.81	-0.26	0.91
14	0	2.12	-3.32	0.06	0	0.78	-0.26	0.10
13	0	-3.45	8.18	4.06	2	0.69	-0.02	1.00
12	1	-0.83	2.73	2.24	2	0.75	-0.13	1.03
11	1	-1.79	4.72	2.94	0	0.69	-0.04	1.14
10	2	0.43	0.12	1.34	2	0.74	-0.00	1.05
9	2	0.74	-0.52	1.15	2	0.31	-0.00	1.58
8	2	0.79	-0.62	1.13	2	0.81	-0.24	0.93
7	2	0.96	-0.99	1.05	2	14.20	-24.0	21.0
6	1	0.97	-1.01	1.06	2	0.94	-0.50	0.74
5	2	1.05	-1.18	1.00	2	0.77	-0.19	1.00
4	2	1.09	-1.25	0.95	2	0.75	-0.16	1.00
3	2	1.17	-1.41	0.89	2	0.79	-0.23	1.00
2	2	1.17	-1.40	0.89	2	0.03	+1.17	2.20

[a] For cases where the rank differs from one, I report the estimates associated with the largest eigenvalue as a reference only. [b] quarterly growth rate in percentage points.

The exception is imports of consumption goods with six lags. For that case, I get unitary income elasticities using a time trend instead of foreign-population:

$$\begin{array}{cccc} \ln m_t = & +0.97 \ln y_t & -1.01 \ln p_t & +1.01 Trend \\ (se) & (0.15) & (0.30) & (0.001) \end{array}$$

Residual Properties: Null Hypotheses (p-value)	
Serial-Independence (0.00*)	Normality (0.00*)
Homoskedasticity (0.13)	*Sample:1967-97*

This evidence weakens the role of immigration in solving the elasticity puzzle. But before accepting this alternative model, note its limitations. First, it has serially correlated residuals which, by itself, breaks a formal observational equivalence. Second, assuming that prices and income are, literally, fixed, the results imply that imports of consumption will automatically increase their GDP share from 2.5 percent to 100 percent in 94 years (appendix B). In other words, this alternative model translates the income-elasticity puzzle into the trend puzzle.

5.3.2 An Almost Ideal Model for Imports

Given that optimization models can avoid the elasticity puzzle, why not use them instead of the log-linear model? A fully satisfactory answer to this question is beyond the scope of this paper but this section documents the results from using the Almost Ideal model of Deaton and Muellbauer (1980) to explain US aggregate imports:

$$w_t = \alpha + \beta \ln y_t + \delta \ln p_t, \beta \gtreqless 0, \delta \gtreqless 0$$

where $w = \frac{m \cdot p_m}{y \cdot p_y}$, β gives the response of the expenditure share to an increase in income and δ measures the corresponding effect of an increase in relative prices. The income elasticity is $1+\frac{\beta}{w}$ and the price elasticity is $(-1+w-\frac{\delta}{w})$; note that an increase in the expenditure share devoted to imports lowers the income elasticity.

Based on quarterly data for imports of goods and services, the results I get for the price-bias model are

$$\begin{array}{lllll} w_t = & -0.62 & +7.53 \ln y_t & -0.75 \ln p_t & +1.63 \ln (K_{for,t}/K_{us,t}) \\ (se) & (0.32) & (3.17) & (2.18) & (1.32) \end{array}$$

$R^2 = 0.78; SER = 0.45\%$
Sample: 1967-97

Null Hypothesis (p-value)	
Serial-Independence (0.00*)	Normality (0.03*)
Homoskedasticity (0.00*)	

and the results for the immigration model are

$$\begin{array}{lllll} w_t = & -0.61 & +6.37 \ln y_t & +3.04 \ln p_t & +2.10 \ln I_t \\ (se) & (0.21) & (1.88) & (0.32) & (0.50) \end{array}$$

$R^2 = 0.94; SER = 0.48\%$
Sample: 1967-97

Null Hypothesis (p-value)	
Serial-Independence (0.00*)	Normality (0.01*)
Homoskedasticity (0.00*)	

According to the results, neither model has residuals consistent with the assumptions needed for inference and the price-bias variable is not significant. This finding does not mean that optimization models are immune to biases from price data but rather that the my modeling of the substitution bias might not be the best one. For what the immigration model is worth, all the variables are significant and its implied income elasticity declines from about 2.5 to a bit above one with an unmistakable trend to homotheticity (figure 20).

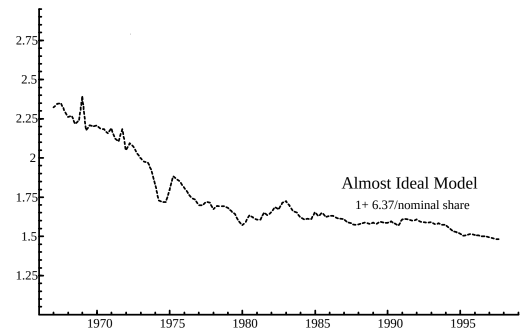


Figure 20: Income Elasticity for Almost Ideal Model–Immigration Model

Thus this formulation avoids the elasticity puzzle, recognizes the role of immigration, and is consistent with optimization but its statistical properties are not satisfactory. Indeed, the Chow tests reject parameter constancy (figure 21).

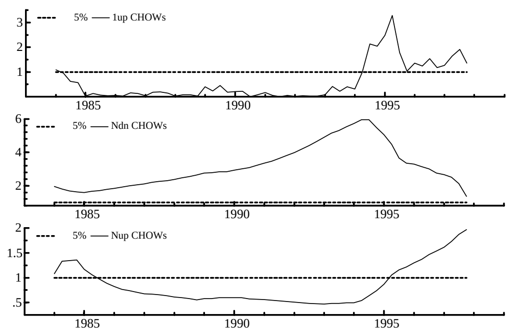


Figure 21*:* Chow Tests for Almost Ideal System–Immigration Model

Finally, this formulation has large prediction errors (figure 22). For example, a prediction error of one percentage point for the import share translates, for 1997, into a ten percent error in the associated level of nominal merchandise imports.

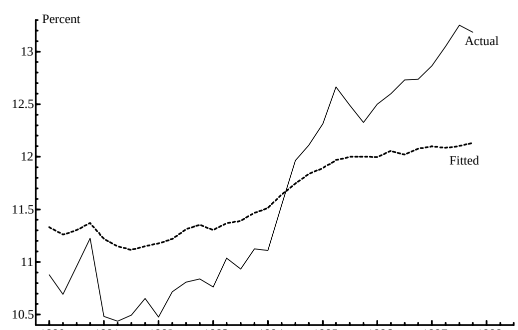


Figure 22: Actual and Predicted Values for Almost Ideal System

5.4 Historiography: Immigration and Houthakker-Magee

Given that by 1969 trade theory had already established that immigration could affect imports, one cannot but help wondering whether the estimated income elasticity of Houthakker and Magee are sensitive to the inclusion of immigrants. Based on annual data for merchandise imports and prices over 1951-1966, I get

$$\begin{array}{cccc} \ln m_t = & +1.91 \ln y_t & -1.39 \ln p_t & +1.97 \ln I_t \\ (se) & (0.24) & (0.51) & (1.06) \end{array}$$

	Null Hypothesis (p-value)	
$R^2 = 0.98; SER = 4.33\%$		
Sample: 1951-1966	Serial-Independence (0.35)	Normality (0.48)
	Homoskedasticity (0.56)	

The results yield the conventional elasticity puzzle along with a rather large, and barely significant, coefficient estimate for the share of foreign born. Thus using the sample dates and data definitions, I find that the estimates of Houthakker and Magee are not sensitive to the inclusion of immigration in their formulation.

The difference between these results and those of table 7 stems from differences in the volatility of the explanatory variables. Specifically, identifying a separate role for prices and immigration with pre-1966 data is hard because these variables fluctuated little compared to income (table 9). As a result, fluctuations in income have a disproportionately large role in accounting for the fluctuations in imports giving rise to an income elasticity greater than one.

Table 9: Unconditional Standard Deviations

Dates	$\ln p$	$\ln y$	$\ln I$	$\ln m$
1951-66	0.0719	0.1539	0.0557	0.2788
1967-97	0.1778	0.1565	0.5423	0.4191

But post-1966 fluctuations in income decline relative to those of prices and immigration factors. As a result, the burden on income fluctuations to explain fluctuations in imports declines which lowers the estimated income elasticity.

6 Conclusions

Existing estimates of the income elasticity suggest that, in the absence of price increases, US imports will eventually exceed US income. That the United States will change from being a largely self-sufficient economy to one that cannot cover its imports has received a great deal of attention but the ensuing three decades of methodological improvements in modeling and estimation have returned even greater estimated income elasticities. This paper resolves this puzzling prediction by removing the representative-agent assumption and addressing the substitution bias embodied in official import prices stemming from their exclusion of new products. Based on Johansen's cointegration method, I get unitary income elasticities for imports of services, consumption, production goods, and the aggregate of these three categories.

The analysis also reveals that there is no silver bullet resolving the elasticity puzzle: the price-bias model is suitable only for imports of producer goods. For imports of consumption and services, the immigration model is the relevant one. In other words, both factors are needed to explain US imports.

I also get my share of disappointing results. Not all the test statistics are pristine, immigration effects could be confounded with a trend for imports of consumption, and there is parameter instability for imports of services. Though I offer explanations for these anomalies, my explanations have limitations of their own. A fuller analysis would involve finding a formulation that recognizes optimizing considerations, addresses the substitution bias embody in official import prices, allows for the heterogeneity induced by immigration, and whose predictive power exceeds that of the immigration-augmented log-linear formulation. I tried to do that with the Almost Ideal Demand model but failed. Until then, the log-linear approximation correcting for biases in prices and heterogeneity will, despite its shortcomings, help in conducting predictions without the elasticity puzzle.

A Data: Sources, Methods, and Properties

A.1 Imports

The data sources for income, imports, and prices are listed below:

Variables	Current Prices	1992 Prices
GDP	SCB[a]; Table 1.1, Line 1	SCB; Table 1.2, Line 1
Imports of		
Goods and Services	SCB; Table 4.3, Line 26	SCB; Table 4.4, Line 27
Consumer Goods ex. Autos	SCB; Table 4.3, Line 38	SCB; Table 4.4, Line 39
Production Goods	Aggregate-Cons.-Services	Aggregate-Cons.-Services
Services	SCB; Table 4.3, Line 44	SCB; Table 4.4, Line 45

[a] Survey of Current Business, US Dept. of Commerce, Bureau of Economic Analysis.

Data for import prices for each category are constructed as deflators by dividing the current-price value by the corresponding 1992-price value. Data for relative prices for imports are calculated as the ratio of a given import price (1992=100) to the U.S. GDP chain-weighted price index (1992=100).[21]

A.2 Bilateral Immigration and Trade

The sources are

1970 and 1975: Statistical Abstract of the United States, 1980, table 135.

1980: Statistical Abstract of the United States, 1984, table 126.

1985: Statistical Abstract of the United States, 1987, table 8.

1990: Statistical Abstract of the United States, 1992, table 8.

1995: Statistical Abstract of the United States, 1997, table 8.

Bilateral Imports of the United States: Direction of Trade Tape, International Monetary Fund.

A.3 Population and Immigration

N = Resident population, mid-period (SCB Table 2.1)

I = share of immigrants in population. Original data are annual and expressed as the flow of immigrants. The source is the Statistical Abstract of the United States, published by the Bureau of the Census, US Department of Commerce. The specific issues that I used are 1986 (table 127), 1990 (table 5), and 1997 (table 4). The 1997 publication reports data on the flow of immigrants up to 1995 and the growth rates for net immigration for 1996 and 1997. I use these two growth rates to estimate immigration flows for 1996-97. Given the time series for these flows, I then construct the stock of immigrants. For this, I

[21] WARNING: The Survey of Current Business' definitions are not a perfect match for an economist's definition. For example, the Survey's definition of imports of consumption merchandise excludes imports of Autos, Parts, and Trucks. Simply adding imports of Autos, Parts, and Trucks to imports of consumption is not a solution as it mixes production and consumption products.

use the Census' benchmark for 1970 of 3.322 million immigrants (Statistical Abstract 1986, table 127). I also assume that the mortality rate for immigrants is the same as that of the US born population 8.9 per 1000 per year. With this assumed mortality rate and the benchmark value, I construct an annual series for the foreign-born population as

$$F_t = f_t + (1 - 0.0089)F_{t-1},$$

where f_t is the flow of immigrants (the observations in the top panel of figure 9), F_t is the stock of foreign born, and $F_{1970} = 3.322$. I then splice the series for F_t to obtain the corresponding quarterly series. The average of quarterly figures equals the annual value. Having obtained a quarterly series for the stock of immigrants, I scale the values by N above to obtain the series I.

A.4 Time-series Properties

To determine the time-series properties of the variables, I use an Augmented Dickey-Fuller test with five lags with and without drift. The evidence suggests that one can reject the hypothesis that the logarithms of the levels of these variables are stationary.

Variable	Augmented Dickey-Fuller With Drift[a]	Augmented Dickey-Fuller With No Drift[b]
Agg. Imports	-2.05	0.62
Cons. Imports	-3.23	-0.55
Prod. Imports	-1.51	-0.80
Service Imports	-1.77	0.14
Rel. Price. Agg. Imports	-1.25	-1.41
Rel. Price Cons. Imports	-2.42	-2.23
Rel. Price Prod. Imports	-1.10	-1.31
Rel. Price Service Imports	-1.68	-1.90
Per-Capita GDP	-3.42	-0.41
Share of Foreign Born	-0.57	-1.69
Stock of Foreign Equip	0.75	-2.62

[a]5% value = -2.886; [b]5% value = -3.448

B Illiquidity Date and Income Elasticity

I calculate the "illiquidity" date by assuming that per-capita income grows following

$$y_t = y_0(1 + \widehat{y})^t,$$

where y_0 is the initial condition for income and $\widehat{y}$ is its constant, annual growth rate. I also assume that per-capita imports behave according to

$$\ln m_t = \beta_0 + \beta_1 \ln y_t + \beta_2 \ln p_t.$$

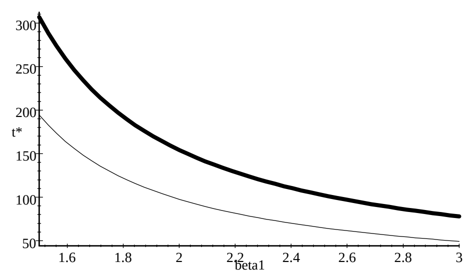
For given prices, this assumption implies that

$$m_t = m_0(1 + \widehat{y}\beta_1)^t,$$

where m_0 is the initial condition for imports. Then I solve for the value of t, t^*, which equates imports and income: $m_0(1 + \widehat{y}\beta_1)^{t*} = y_0(1 + \widehat{y})^{t*}$:

$$t^* = \frac{\ln \frac{m_0}{y_0}}{\ln \frac{(1+\widehat{y})}{(1+\widehat{y}\beta_1)}}.$$

The figure below displays two schedules relating the illiquidity date to alternative values of both the income elasticity β_1 and initial trade shares $\frac{m_0}{y_0}$. Specifically, I use two values of $\frac{m_0}{y_0}$: 5% (thick line) and 15% (thin line); all the calculations assume that $\widehat{y} = 0.02$. The calculations illustrate the inverse association between t^* and β_1; similarly, an increase in the initial import share lowers t^* for every value of β_1. Note that, regardless of the value of $\frac{m_0}{y_0}$, if $\beta_1 = 1$ then the import-GDP ratio remains constant and thus t^* approaches infinity.



Income Elasticity and Illiquidity Date

The illiquidity date for a model with trend is

$$t^* = \frac{\ln \frac{m_0}{y_0}}{\ln \frac{(1+\widehat{y})}{(1+\widehat{y}\beta_1+\beta_4)}},$$

where β_4 is the annualized autonomous growth rate (coefficient on trend x 4). Using 1997 as the initial condition and assuming a trend rate of 1% per quarter along with fixed prices and income implies that the illiquidity date is

$$t^* = \frac{\ln(0.025)}{\ln\left(\frac{(1+0.0)}{(1+0.0+0.04)}\right)} = 94 \; years.$$

C Measurement Errors in Immigration Data

Observations on the share of foreign-born are subject to many sources of errors. I address here the implications of only two of them: illegal immigrants and initial conditions.

C.1 Illegal Immigrants

By construction, the measure of F_t excludes illegal immigrants implying that the true value of I_t is unobserved. One could avoid this limitation by estimating the flow of illegal immigrants, adding it to f_t, and recomputing F_t. But existing estimates of the number of illegal immigrants are imprecise suggesting that the gains from an increased coverage might be lost to the resulting increase in imprecision. For example, Fix and Passel (1994), as reported in Littman (1998, p. 16), have done this estimation for 1980-92; the mean of their lower bound is 2.5 million illegal residents and the mean for their upper bound is 3.5 million.

As an alternative, I recognize that the actual (but unobserved) share of foreign-born is identical to the sum of the legal and illegal foreign-born population: $I_t = I_t^r + I_t^u$, where I_t^r is the share of foreign-born population residing legally, which is observed, and I_t^u is the share of foreign-born population residing illegally, which is unobserved. I now take advantage of a key feature of US immigration: Regional concentration by country of origin (Lapham, 1993). This feature suggests that existing legal immigrants develop a network of relations (job, housing, family) that attracts illegal aliens. Because developing such a network takes time, I model the share of foreign-born population residing illegally as a distributed lag of the share of foreign-born population residing legally:

$$I_t^u = \left[\sum_{j=1}^{\ell} \iota_j I_{t-j}^r\right] = \iota(L) I_t^r.$$

Combining this assumption with the identity relating the share of foreign-born to the sum of legal and illegal foreign-born population gives

$$I_t = I_t^r + I_t^u = I_t^r + \iota(L) I_t^r = I_t^r [1 + \iota(L)],$$

which is what I use in the empirical analysis.

C.2 Initial Conditions

Even if the number of illegal immigrants were negligible, data for the foreign-born population are subject to other measurement error in the initial condition. To that end, I construct a measure of foreign-born population assuming that immigrants prior to 1820 were a negligible portion of the native-born population. These data show that immigration into the United States has not been uniform over time (figure C1). The immigration waves of the late 1800s and early 1900s were followed by periods of low immigration which reached by 1945 the lowest value in the 20th century. Since 1950, however, immigration increases steadily reaching a historical record during 1990-91 with the adoption of the Immigration Act of 1990 (Littman 1998, p.16). I choose 1820 because it is the first year for which immigration data are available. Given this consideration, I cumulate the flow of immigrants since 1820 using the mortality rates for the native population.

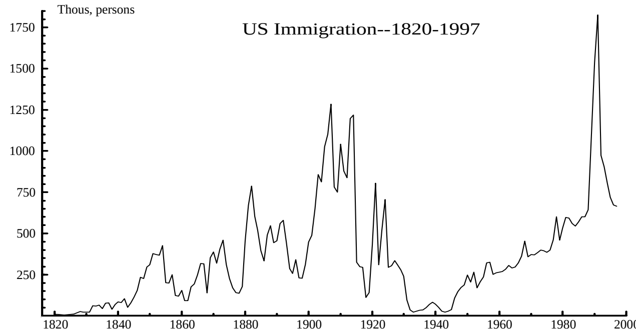


Figure C1: Immigration into the United States–1820-1997

Visual inspection of the two alternative series suggests that changing the benchmark to 1820 raises the stock of foreign-born population but it does so without changing the time profiles of the two series (figure C2). In addition, the gap between the two series diminishes over time because the effects from immigration flows dominate the importance of differences in initial conditions.

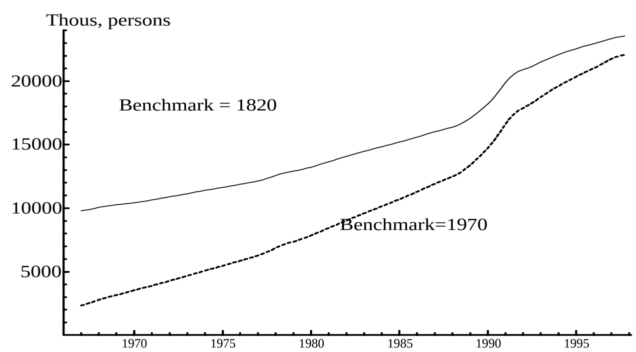


Figure C2: Foreign-Born Population–Alternative Benchmarks

Using the measure of foreign-born population with an 1820 benchmark, I apply the Johansen method using quarterly data over 1967-97 and get:

$$\begin{array}{cccc} \ln m_t = & +0.96 \ln y_t & -0.76 \ln p_t & +0.98 \ln I_t \\ (se) & (0.19) & (0.28) & (0.39) \end{array}$$

Residual Properties: Null Hypothesis (p-value)	
Serial-Independence (0.10)	Normality (0.24)
Homoskedasticity (0.08)	*Sample:1967-97*

The estimated income elasticity is positive, significant, and virtually equal to one; the estimated price elasticity is -0.75 and significant. The effect of immigration on imports is positive, significant, and larger than the effect shown in table 4 above. This result is not surprising given the smaller *increase* recorded by the alternative measure of foreign population. From a statistical standpoint, the residuals satisfy the estimation assumptions and the recursive Chow tests support parameter constancy (figure C1). Thus using immigration data benchmarked in 1820 does not alter the finding that the income elasticity for aggregate imports is one.

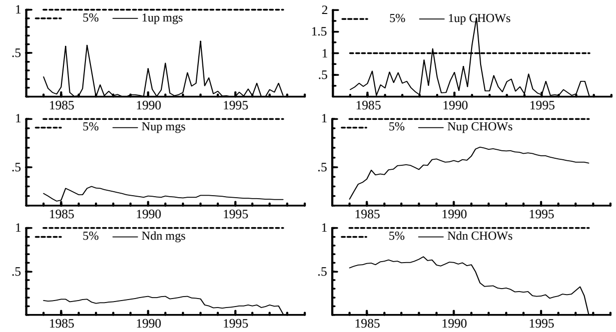


Figure C3: Chow Tests for Aggregate Imports–Cointegration

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