International Finance Discussion Papers

Number 475

July 1994

The Constancy of Illusions or the Illusion of Constancies: Income and Price Elasticities for U.S. Imports, 1890-1992

Jaime Marquez

Abstract

Virtually all we know about the behavior of U.S. imports rests on studies estimating income and price elasticities with postwar data. But anyone examining the evolution of U.S. trade cannot avoid asking whether the postwar period provides enough information to characterize that behavior. From 1890 to 1940, the United States became an increasingly closed economy and experienced the most pronounced fluctuations in income and prices of this century. Is our current understanding of the behavior of U.S. imports consistent with those features of the U.S. economy? Being consistent with the distant past might not appear as relevant for forecasting, but the literature ignoring that past offers a range of elasticity estimates wide enough to suggest that the role of income and prices in determining imports is not known with any precision. This paper offers the first analysis of that role using data since 1890. Estimating the elasticities of the most popular model in the literature with 1890-1992 data, I find that income and prices do not affect imports whereas the opposite conclusion arises with postwar data. The difference in results stems from differences in the time-series properties of the data in the two samples. As an alternative, I consider several models consistent with both optimization and the time-series properties of the data. These models predict substantial secular changes in income and price elasticities and confirm the importance of optimization for characterizing the behavior of U.S. imports. What is new about the results is that only through optimization can one recognize the implications of the evolution of U.S. trade for estimating elasticities.

The Constancy of Illusions or the Illusion of Constancies:
Income and Price Elasticities for U.S. Imports, 1890-1992

Jaime Marquez[1]

Virtually all we know about the behavior of U.S. imports rests on studies estimating income and price elasticities with postwar data. But anyone examining the evolution of U.S. trade cannot avoid asking whether that period provides enough information for characterizing that behavior. From 1890 to 1940, the United States became an increasingly closed economy and experienced the most pronounced fluctuations in income and prices of this century. Are existing explanations of the behavior of U.S. imports consistent with those features of the U.S. economy? Being consistent with the distant past might seem irrelevant for practical applications, but the literature ignoring that past offers a large range of estimates suggesting that the role of income and prices in determining imports is not known with any precision. This paper examines that role using data since 1890 and offers three new findings.

First, the treatment of elasticities as constant parameters, even if valid for each study considered individually, is not valid when all studies are considered as a whole: As a collection, existing elasticity estimates are systematically influenced by the switch from fixed to floating exchange rates, even if the estimates from individual studies were constant. This paradox stems from the use of estimation samples covering short and non-overlapping periods that cannot detect secular changes in elasticities. Thus agreeing on a characterization of U.S. imports requires the longest span of data available. Second, when the sample includes the longest span of data, the estimates from the widely used log-linear formulation are effectively zero whereas the opposite conclusion arises with postwar data alone. Differences in the time-series properties of the data in the two samples account for this finding: the postwar sample exhibits cointegration among imports, expenditures, and relative prices whereas the full sample does not exhibit this property. Third, only models resting on optimizing

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[1] The author is a staff economist in the Division of International Finance. I have benefited from comments by William Barnett, David Bowman, Clive Granger, William Helkie, Dale Henderson, David Hendry, Jon Faust, Cathy Mann, William Melick, and seminar participants of our workshop series. The calculations use the following software: Limdep 5.0, PC-GIVE 7.0, and PC-FIML. I am grateful to David Hendry for allowing the use of a preliminary release of PC-FIML. The views expressed in this paper are the author's and should not be interpreted as reflecting those of the Board of Governors of the Federal Reserve System or other members of its staff.

behavior can incorporate both the evolution of U.S. trade and the properties of the associated time-series. The estimated income and price elasticities from these models fluctuate in response to changes in the composition of expenditures; models that ignore this response overstate the sensitivity of U.S. imports to changes in income and prices.

2. Second Thoughts

Most econometric analyses of U.S. imports treat income and price elasticities as constant parameters (Marquez, 1992). This assumption would be useful if the dispersion of estimates were small. But a survey of fifty years of econometric work (appendix A) reveals an unsettling dispersion of elasticities that range from -0.5 to -4.8 for price and from 0.8 to 4.0 for income (figure 1). In addition, the estimates contradict economic theory which predicts that if income and price elasticities are constant, then they must equal one and minus one (Deaton and Muellbauer, 1980b, p. 17).

The dispersion of estimates shown in figure 1 might arise from differences in modeling assumptions and sample periods. To quantify the relative importance of these possibilities, I model the elasticity estimates with a fixed-effect model in which the dependent variable is the ith study's elasticity estimate and the explanatory variables are the associated modeling assumptions (shown in appendix A):

(1) $\varepsilon_i = \alpha_0 + \alpha_1\text{Fixed} + \alpha_2\text{No-homogeneity} + \alpha_3\text{Annual} + \alpha_4\text{Static} + \alpha_5\text{Shiller} + \alpha_6\text{IV} + \alpha_7\text{Oil} + u_i$,

where ε_i = Long-run elasticity estimate of the ith study.

Fixed = Dummy variable equal to one if exchange rates are fixed in the sample.

No-homogeneity = Dummy variable equal to one if price homogeneity is absent.

Annual = Dummy variable equal to one if the sample's frequency of observation is annual.

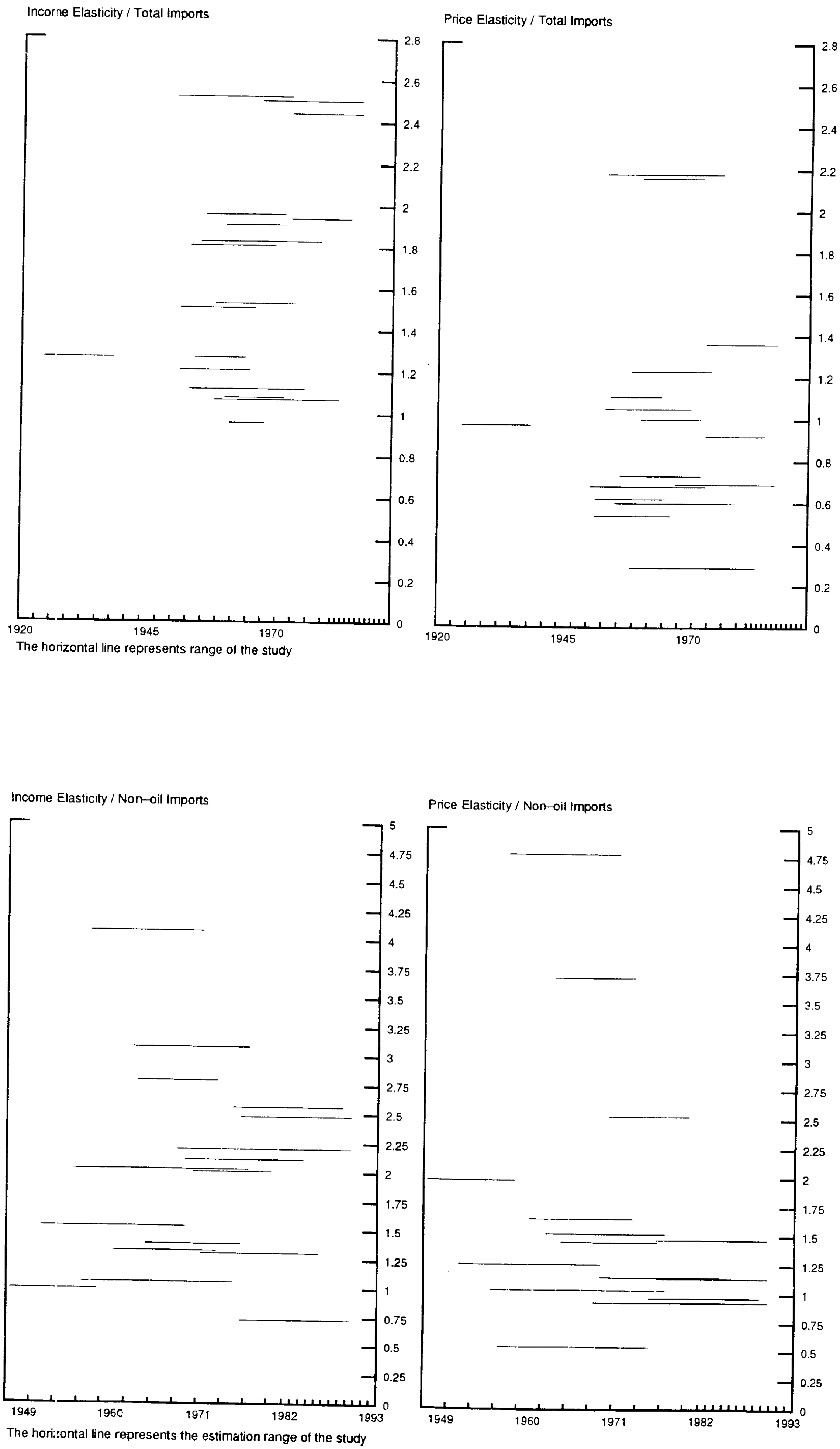
Static = Dummy variable equal to one if the estimates abstract from delayed adjustments.

Shiller = Dummy variable equal to one if dynamic adjustments follow a Shiller lag.

IV = Dummy variable equal to one if the estimation method recognizes simultaneity.

Oil = Dummy variable equal to one if imports include oil.

Figure 1: Income and Price Elasticities for U.S. Imports
selected studies



u_i = White noise disturbance.

Extending the list of explanatory variables is possible but would exhaust the degrees of freedom. The parameter α_0 is the prototype elasticity for studies that (1) use data for the period of floating exchange rates, (2) assume price homogeneity, (3) employ either semi-annual or quarterly data, (4) allow for lagged responses that do not involve Shiller lags, (5) apply ordinary least squares, and (6) exclude oil from the measure of imports.[2] The other parameters in (1) measure the extent to which alternative modeling assumptions change the prototype elasticity. For example, if α_1 is significantly different from zero then elasticity estimates based on data for the period of fixed exchange rates differ from the estimates based on data for the floating exchange-rate period.

I estimate the parameters of (1) using data from 33 studies listed in appendix A; the estimation method is weighted least squares where the weights equal the inverse of the estimated standard errors of the long-run elasticity estimates. The prototype estimates are 2.04 for income and -1.22 for prices (table 1). These estimates are not sensitive to either frequency of observation or estimation method but are sensitive to the treatment of price-homogeneity, the specification of dynamic adjustments, and the exchange-rate system.[3] Specifically, the estimate of α_1 in (1) is negative and significant which means that the estimated price elasticities based on data for the period of floating exchange rates are lower (in absolute terms) than the estimates based on data for the period of fixed exchange rates; estimated income elasticities exhibit the opposite pattern.

This finding indicates that even if the assumed constancy of elasticities were correct for each study considered individually, it is not correct when all studies are considered as a whole. This paradox stems, as figure 1 suggests, from the use of small samples with a minimal of overlap; not one

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[2] The Shiller lag is singled out even though only one study (Wilson and Takacs, 1979) used this technique because the elasticity estimates appear to be outliers.

[3] Several studies that report ordinary least squares (OLS) results note that in preliminary testing, simultaneous-equation estimation yields nearly the same results as OLS estimation--see, for example, Geraci and Prewo (1982) and Helkie and Hooper (1988).

Table 1
Estimation Results from Fixed-Effect Model of Trade Elasticities

$$\varepsilon_i = \alpha_0 + \alpha_1 \text{ Fixed} + \alpha_2 \text{ No-homogeneity} + \alpha_3 \text{ Annual} + \alpha_4 \text{ Static} + \alpha_5 \text{ Shiller} + \alpha_6 \text{ IV} + \alpha_7 \text{ Oil} + u_i$$

Dep. Variable →	Income Elasticity		Price Elasticity	
Ind. Variable ↓	α	t-stat	α	t-stat
Intercept	2.04	12.06	-1.22	-4.59
Fixed	-0.51	-2.10	-0.96	-3.11
No-homogeneity	0.09	0.36	-0.74	-2.45
Annual	-0.29	-0.98	0.03	0.10
Static	0.14	0.46	0.70	2.20
Shiller	2.47	4.20	-1.86	-4.33
IV	-0.20	-0.17	0.58	0.70
Oil	0.13	0.71	0.70	2.99
R^2	0.59		0.91	
S.E.R.	0.47		0.54	

ε_i = Estimate of the long-run elasticity of the ith study.
Fixed = Dummy variable equal to one if exchange rates are fixed in the estimation sample.
No-homogeneity = Dummy variable equal to one if price homogeneity is absent.
Annual = Dummy variable equal to one if the sample's frequency of observation is annual.
Static = Dummy variable equal to one if the estimation abstracts from delayed adjustments.
Shiller = Dummy variable equal to one if dynamic adjustments are modeled with Shiller lags.
IV = Dummy variable equal to one if the estimation method recognizes simultaneity.
Oil = Dummy variable equal to one if imports include oil.

study covers the postwar period as whole, much less the whole evolution of U.S. trade. Short and discontinuous samples do not have enough information for detecting parameter instability arising from secular forces. Equation (1), however, effectively combines all of these sub-samples and uncovers the instability concealed by sample selection. Overall, the results suggest that a necessary condition for achieving consensus on the role of income and prices in determining U.S. imports is the use of the longest span of data available.

3. Hidden History

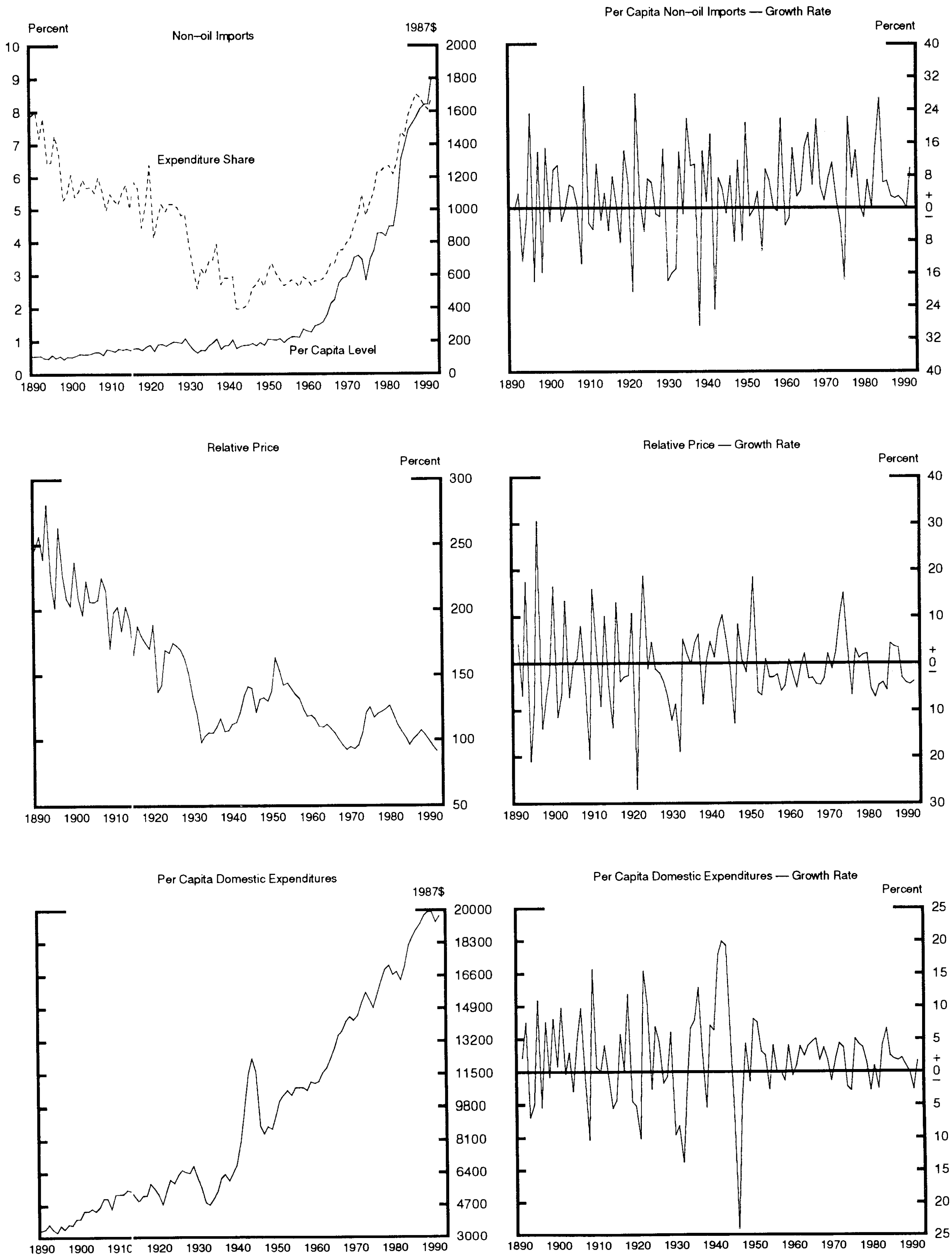
Inspecting the evolution of U.S. openness since 1890 reveals several features pertinent for characterizing the behavior of U.S. imports (figure 2). First, the United States became an increasingly closed economy during the prewar period with the share of imports in expenditures declining from eight percent to two percent; this pattern stands in sharp contrast to the experience of the postwar period. Second, the volatility of expenditures and relative prices during the prewar period dwarfs that of the postwar period; neither the oil-price shocks nor the exchange-rate fluctuations of 1970-80 induced changes in relative prices comparable to those occurring in the beginning of this century. Finally, the U.S. population has been aging with the share of the population of at least sixty-five years of age increasing from 3.8 percent in 1890 to 12.2 percent in 1992 (figure B1, appendix B). Aging might affect preferences with a corresponding effect on purchases.[4]

Given these data, I use OLS to estimate the parameters of a widely used log-linear formulation assuming imperfect substitutability between domestic and foreign products:

$$\beta_1(L)\ln(m^*_t) = \beta_0 + \beta_2(L)\ln(y^*_t) + \beta_3(L)\ln(p^*_{mt}/p^*_{dt}), \tag{2}$$

[4] Branson (1980), Dornbusch and Fischer (1986), and Lipsey (1994) document the evolution of U.S. openness but do not quantify the role of income and prices in explaining U.S. imports. Appendix B discusses the construction of the series, displays the evolution of their components, and reports the associated data sources. Real expenditures equal real GNP plus real imports of goods and services minus real exports of goods and services. The relative price of imports equals the ratio between the non-oil tariff-adjusted price of non-oil imports and the GNP deflator.

Figure 2: Imports, Relative Prices, and Expenditures
United States, 1890–1992



where $\beta_i(L)$ is a polynomial in the lag operator L ($i > 0$) and the symbol "*" denotes variables demographically scaled.[5] Because these variables are not observable, I follow Pollak and Wales (1992, pp. 76-78) and express them as $m^*_t = m_t a_t^{-\xi}$, $y^*_t = y_t a_t^{-\xi}$, and $p^*_{it} = p_{it} a_t^{\xi}$ for i=m,d, where m_t is per-capita imports; a_t is the share of U.S. population of at least sixty-five years of age; p_{mt} is the price of non-oil imports adjusted by non-oil tariffs; p_{dt} is the price of domestic products; y_t is per-capita real domestic expenditure, and ξ is the demographic-scaling parameter. Given these assumptions, (2) becomes

$$(3) \qquad \beta_1(L)\ln(m_t) = \beta_0 + \beta_2(L)\ln(y_t) + \beta_3(L)\ln(p_{mt}/p_{dt}) + \xi(\beta_1(L) - \beta_2(L))\ln(a_t).$$

The long-run elasticities are $\beta_2(1)/\beta_1(1)$ for income and $\beta_3(1)/\beta_1(1)$ for prices; if $\beta_1(1) = 0$ then imports lack a steady state; I include three dummy variables in (3) to control for world-war disruptions.[6]

Based on 1890-1992 data, the long-run elasticity estimates of (3) lack statistical significance and are, relative to existing studies, extraordinarily large (table 2).[7] Moreover, the estimated coefficient for the lagged dependent variable is nearly one and the magnitude of the other coefficient estimates suggests that the equation can be re-written in logarithmic differences. Overall, this formulation indicates that imports do not have a steady state. Starting the estimation period in 1960, however, reverses these findings: long-run elasticities are significantly different from zero, the coefficient for the lagged dependent variable is far from one, and the estimates replicate those found in the literature (which violate the theoretical benchmarks); population aging, as modeled here, plays no role in determining U.S. imports.

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[5] The logarithmic formulation is the most common specification for modeling U.S. imports (see Goldstein and Khan 1985 and Marquez 1992); for demographic scaling see Pollak and Wales (1992).

[6] Using several test statistics, I cannot reject the hypothesis that the variables in (3) are integrated of order one which means that the regression is balanced; appendix C documents these results. Note that (3) is an unrestricted Autoregressive Distributed Lag and its long-run parameters are identical to those given by the Error-Correction formulation (see Banerjee et al., 1993, section 2.5).

[7] To select the lag structure, I begin with a formulation having four lags for imports, income, relative prices and test for zero restrictions at each lag length.

The difference in results in the two samples is not due to heteroskedasticity, serial correlation in the residuals, or failure to meet the functional-form test.[8] Rather, the difference in results stems from differences in the time-series properties of the data in the two samples. For 1890-1992 data, the Engle-Granger cointegration test (Engle and Granger, 1987) indicates that these variables do not have a long-run relation among themselves--that they are not cointegrated--whereas estimates using 1960-1992 data reverse this finding; Clarida (1994) also finds cointegration for postwar data. Thus the results suggest that (2) has a long-run relation if the estimates do not use data covering a long run.

The cointegration results of table 2 treat, however, income and prices as given and I examine the importance of this treatment with Johansen's technique (Johansen, 1988). This technique uses the reduced form of a dynamic, simultaneous system for imports, expenditures, and relative prices:

$$\Delta x_t = \varphi + \Gamma_1 \Delta x_{t-1} + \ldots + \Gamma_3 \Delta x_{t-3} + \Gamma x_{t-4} + \delta D_t + u_t \,, \tag{4}$$

where x_t is a 3x1 vector with the three variables of interest; D is a vector of exogenous variables; Γ is a 3x3 matrix of coefficients; and u_t is a 3x1 vector of serially independent and jointly normal disturbances. The exogenous variables included in (4) are the war dummies and the share of population of 65 years of age and older; statistical tests for these data reveal that four years accounts for delayed adjustments.

Johansen shows that the rank of Γ gives the number of cointegrating relations--that is, the number of equations for which a long-run formulation exists. To determine the rank of Γ, Johansen estimates Γ with a maximum likelihood method, computes the associated eigenvalues, and shows that the rank of Γ equals the number of non-zero eigenvalues. Johansen and Juselius (1989) offer a statistic to test whether the *i*th eigenvalue is zero ($\lambda(i)_{max}$) and another statistic to test whether the sum of *r* eigenvalues is zero ($\lambda(r)_{trace}$). Based on 1890-1992 data, both statistics indicate that the rank of Γ is

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[8] I test homoskedasticity using Engle's ARCH test (Engle,1982). I test serial independence by applying an F-test to the hypothesis that all the coefficients of an AR(4) for the residual are zero. The test for functional form uses Ramsey's statistic. These tests are implemented and documented in Doornik and Hendry (1992).

Table 2

Sensitivity of Income and Price Elasticities to Sample Period:Log-linear Formulation

$$\ln(m_t) = \beta_0 + \beta_1 \ln(m_{t-1}) + \beta_{20}\ln(y_t) + \beta_{21}\ln(y_{t-1}) + \beta_{30}\ln(p_t) + \beta_{31}\ln(p_{t-1}) + \beta_4\ln(a_t) + \beta_5\ln(a_{t-1})$$

	1890-1992		1960-1992	
Variables	β	t-stat	β	t-stat
Intercept	-0.718	-0.890	-6.580	-2.544
$\ln(m_{t-1})$	0.995	43.167	0.503	3.024
$\ln(y_t)$	1.461	10.464	2.385	7.361
$\ln(y_{t-1})$	-1.381	-9.516	-1.077	-2.058
$\ln(p_t)$	-0.234	-2.597	-0.470	-2.742
$\ln(p_{t-1})$	0.195	2.092	0.140	0.613
$\ln(a_t)$	-0.455	-0.464	0.783	0.573
$\ln(a_{t-1})$	0.371	0.378	-0.250	-0.195
L.R. Income Elasticity	16.69	0.23	2.632	7.475
L.R. Price Elasticity	-8.28	-0.20	-0.664	-3.965
L.R. Age Elasticity	-17.57	-0.21	1.073	1.941
R^2	0.99		0.99	
SER	0.075		0.037	
Serial Independence[a]	0.18		0.21	
Homoskedasticity[b]	0.43		0.21	
Skewness[c]	-0.32		-0.01	
Excess Kurtosis[d]	2.05		-0.62	
Functional Form[e]	0.10		0.09	
Cointegration[f]	-0.70		-4.97**	

Notes: m is per-capita, real non-oil imports; y is per-capita, real domestic expenditures; p is the ratio of the tariff-adjusted non-oil import price to the GNP deflator. The regression includes three dummy variables for three war years: 1918, 1942, and 1946.

[a]Significance level for rejecting the hypothesis that residuals are serially independent.

[b]Significance level for rejecting the hypothesis that residuals are homoskedastic.

[c]Skewness of the empirical distribution of the residuals; this entry is zero for a normal distribution.

[d]Excess Kurtosis of the empirical distribution of the residuals; this entry is zero in normal distributions.

[e]Significance level for rejecting the choice of functional form using the RESET test statistic.

[f]Engle-Granger test of cointegration: a ** denotes significant at the 1 percent level.

zero (table 3): the procedure cannot identify a single long-run relation among the logarithms of imports, price, and expenditures just as found in table 2.

Table 3
Cointegration Tests for Imports, Relative Prices, and Expenditures: United States, 1890-1992

Measurement →	logarithms		log-differences	
Null Hypothesis ↓	$\lambda(i)_{max}$	$\lambda(r)_{trace}$	$\lambda(i)_{max}$	$\lambda(r)_{trace}$
Rank ≥ 0	16.3	23.46	35.7 **	74.58**
Rank ≥ 1	7.15	7.17	21.74**	38.88**
Rank ≥ 2	0.02	0.02	17.15**	17.15**

Note: A "**" denotes statistical significance at the one percent level. Both $\lambda(i)_{max}$ and $\lambda(r)_{trace}$ include an adjustment for degrees of freedom equal to the product of the number of variables and the number of lags. The critical values for $\lambda(i)_{max}$ are 21, 14.1, and 3.8; the critical values for $\lambda(r)_{trace}$ are 29.7, 15.4, and 3.8.

This failure to find cointegration does not involve, however, a rejection of the imperfect substitute model embodied in (2). Indeed, the solution to the first-order conditions for optimization relates changes in purchases to changes in income and relative prices suggesting that changes in these variables should be cointegrated. Cointegration results for differenced data support this theoretical prediction (table 3) and suggest that modeling U.S. imports in terms of expenditures and relative prices should use data in first differences. The remaining question is how.

4. Model Formulation and Results

I assume that individuals determine their spending on domestic and foreign products, d and m respectively, by maximizing a utility function u(d,m) subject to $p^*_m m^* + p^*_d d^* = p^*_{yt} y^*_t$. Differentiating the first order conditions for maximizing *any* utility function and solving the associated system for quantities in terms of prices yields

(5) $$w_t d\ln(m^*_t) = \mu_t d\ln(y^*_t) + \pi_t d\ln(p^*_t), \quad \mu_t > 0 \quad \pi_t < 0 ,$$

where w_t is the share of imports in expenditures; μ_t is the marginal response of spending on imports to changes in expenditures; π_t is the compensated (Slutsky), own-price effect, and $p^*_t = (p^*_{mt} / p^*_{dt})$.[9] Equation (5) explains how much of the change in the nominal import share stems from changes in import volume. In the absence of further restrictions, however, (5) cannot be rejected by the data and thus is not suitable for empirical work. One way of bypassing this limitation involves assuming that μ_t and π_t are constant, which gives the Rotterdam model:

$$w_t d\ln(m^*_t) = \mu d\ln(y^*_t) + \pi d\ln(p^*_t) + u_{rt}, \tag{6}$$

where u_{rt} is the disturbance induced by the assumed constancy of μ and π. The Rotterdam model is appealing because it is linear in the parameters and, with the exception of integrability, it satisfies the properties of consumer demand theory globally.[10] Treating μ and π as autonomous, however, amounts to approximating (5) and the quality of this approximation depends on the assumed constancy of μ and π (Barnett, 1984; Byron 1984); I evaluate this assumption through parameter-constancy tests.

An alternative to the assumed constancy of μ and π involves approximating the utility function with a specific formulation and finding the exact solution to the optimization problem. The Almost Ideal model of Deaton and Muellbauer (1980a) follows this approach and the corresponding solution is

$$dw_t = \delta d\ln(y^*_t) + \gamma d\ln(p^*_t) + u_{at}, \tag{7}$$

which explains the change in the nominal import share.[11] Equation (7) is appealing because it rests on a specific utility function, is linear in the parameters, and satisfies integrability; this model, however, meets the concavity requirement only locally.

The income elasticities are μ/w_t for the Rotterdam model and $(1+\delta/w_t)$ for the Almost Ideal

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[9] Expenditures shares are invariant to demographic scaling. For the derivation of equation (5) see Barten (1964), Theil (1965), and Barnett (1979); see Kohli (1991) for an approach that uses cost minimization by firms.

[10] A necessary condition for the Rotterdam model to meet the integrability condition is the presence of cointegration among the levels of imports, expenditures, and prices. (As noted before, the logarithms of these variables are not cointegrated.) I apply the Johansen technique and find that there is one cointegration vector which suggests the existence of a relation in the levels of these variables.

[11] The solution of the Almost Ideal System is $w_t = \delta\ln(y^*_t) + \gamma\ln(p^*_t)$ but this formulation fails cointegration tests. Thus I use equation (17) of Deaton and Muellbauer (1980a, p.317).

model; the compensated own-price elasticities are π/w_t for the Rotterdam and $(-1+w_t+\gamma/w_t)$ for the Almost Ideal model. These expressions underscore that elasticities from optimizing models respond to changes in the composition of expenditures; these elasticities will be constant if either w_t is constant, which contradicts the data, or if parameter changes offset exactly changes in w_t, which I test.

Expressing (6)-(7) in terms of observables yields

(8) Rotterdam: $w_{t-1}\Delta\ln(m_t) = \mu\Delta\ln(y_t) + \pi\Delta\ln(p_t) - \xi\mu\Delta\ln(a_t) + \xi w_{t-1}\Delta\ln(a_t) + u_{rt}$,

(9) Almost Ideal: $\Delta w_t = \delta\Delta\ln(y_t) + \gamma\Delta\ln(p_t) - \delta\xi\Delta\ln(a_t) + u_{at}$,

where the use of w_{t-1} instead of w_t reflects the switch from infinitesimal changes to discrete differences (Theil 1971, p.371) and the variables are expressed in logarithmic differences as suggested by the cointegration evidence of table 3 above. Demographic scaling induces both overidentifying restrictions and lagged effects in the Rotterdam model but not in the Almost Ideal formulation.

For estimation, I include in (8)-(9) the war dummies and an intercept.[12] Finding a statistically significant intercept means that imports change even if income, prices, and population's age profile were unchanged, a sign of misspecification. Based on 1890-1992 data, the estimated income and price effects are significant and the intercept is insignificant (table 4).[13] The residuals do not show autocorrelation or heteroskedasticity but their distribution is more concentrated than the normal distribution; the Engle-Granger test cannot reject the hypothesis that the variables in (8)-(9) are cointegrated. Only the Almost Ideal model fails the functional-form test suggesting that explaining nominal shares involves additional terms.

I examine parameter constancy with three formulations of the Chow test: a sequence of F-tests for the hypothesis that the one-period ahead forecast error is zero (1↑); a sequence of F-tests for the

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[12] I use an Augmented Dickey-Fuller to test whether $w_{t-1}\Delta\ln(m_t)$ is stationary and the corresponding statistic is -4.09 suggesting that it is; the ADF regression has a constant, a trend, and four lags.

[13] To examine the validity of the overidentifying restriction in the Rotterdam model, I re-estimate (8) with nonlinear least squares. The estimates (t-stats) are 0.0652 (9.37) for μ, -0.0147 (-3.38) for π, and 0.261 (0.9) for ξ; the standard error of the regression is 0.0039. Thus the main implication of imposing the overidentifying restriction is that population appears to have no role.

Table 4
Income and Price Effects for U.S. Imports, 1890-1992
Alternative Models

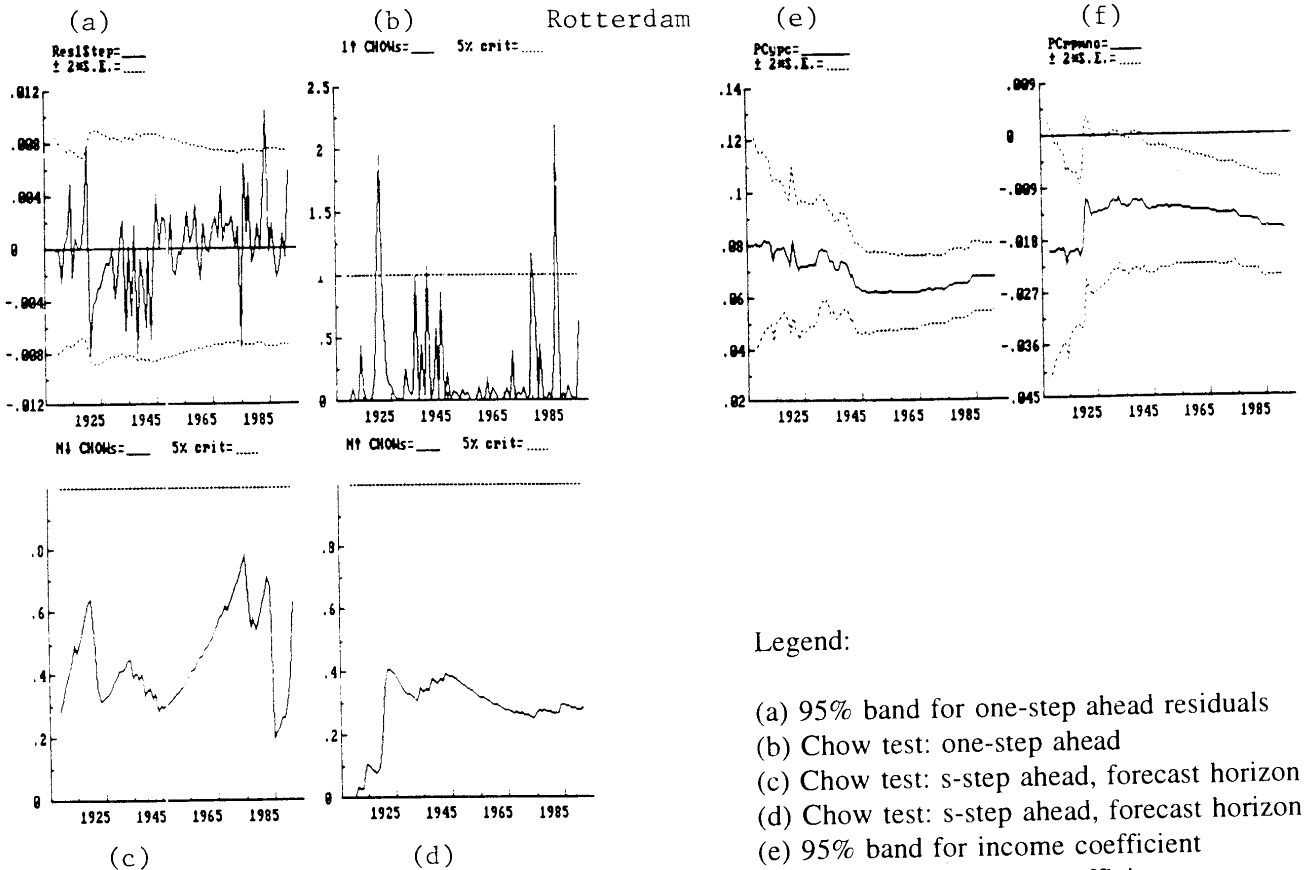
Variables	Rotterdam: $w_{t-1}\,\Delta\ln(m_t)$		Almost Ideal: $\Delta\ln(w_t)$	
Intercept x 100	-0.0124 (0.53)		0.0196 (-0.32)	
$\Delta\ln(y_t)$	0.0679 (10.29)	0.0678 (10.38)	0.0134 (2.10)	0.0132 (2.09)
$\Delta\ln(p_t)$	-0.0161 (-3.73)	-0.0160 (-3.67)	0.0391 (9.97)	0.0393 (9.52)
$\Delta\ln(a_t)$	-0.202 (-2.39)	-0.207 (-2.54)	0.0313 (0.72)	0.0206 (0.75)
$w_{t-1}\,\Delta\ln(a_t)$	5.212 (2.87)	5.154 (2.89)		
R^2	0.57	0.60	0.57	0.57
SER x 100	0.37	0.37	0.36	0.36
AR(4)[a]	0.23	0.22	0.17	0.18
ARCH[b]	0.66	0.66	0.42	0.43
Skewness[c]	-0.16	-0.16	-0.33	-0.33
Excess Kurtosis[d]	1.38	1.41	1.28	1.33
Functional Form[e]	0.79	0.75	0.00	0.00
Cointegration[f]	-4.97**	-5.01**	-3.93*	-4.01*

Notes: m is per-capita, real non-oil imports; y is per-capita, real domestic expenditures; p is the ratio of the tariff-adjusted non-oil import price to the GNP deflator. The regression includes three dummy variables for three war years: 1918, 1942, and 1946.

[a]Significance level for rejecting the hypothesis that residuals are serially independent.
[b]Significance level for rejecting the hypothesis that residuals are homoskedastic.
[c]Skewness of the empirical distribution of the residuals; this entry is zero for a normal distribution.
[d]Excess Kurtosis of the empirical distribution of the residuals; this entry is zero in normal distributions.
[e]Significance level for rejecting the choice of functional form with the RESET test statistic.
[f]Engle-Granger test of cointegration: a * denotes significant at the 5% level; ** means 1% level.

Figure 3: Parameter-Constancy Tests, 1909-1992

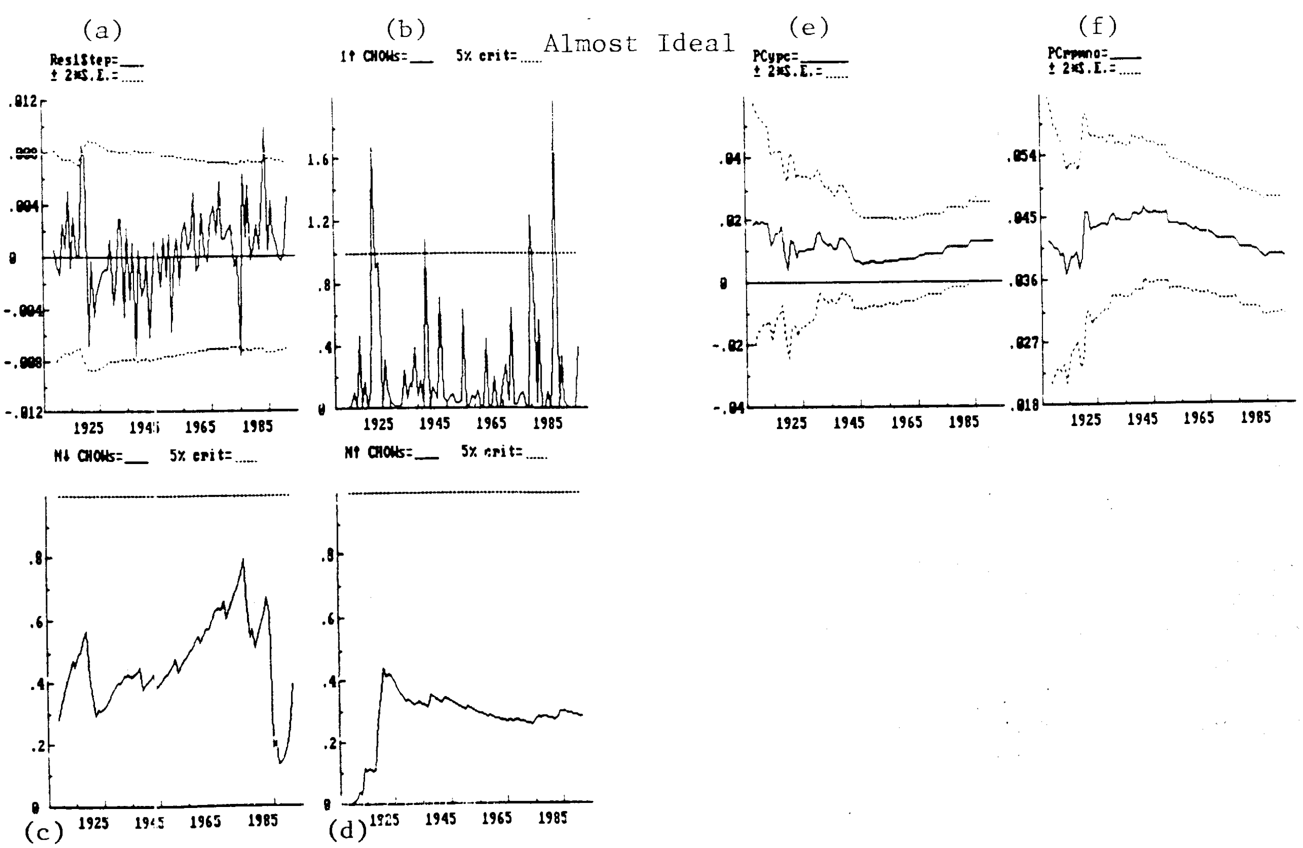
Rotterdam



Legend:

(a) 95% band for one-step ahead residuals
(b) Chow test: one-step ahead
(c) Chow test: s-step ahead, forecast horizon declines
(d) Chow test: s-step ahead, forecast horizon increases
(e) 95% band for income coefficient
(f) 95% band for price coefficient

Almost Ideal



hypothesis that the out-of-sample forecast errors are jointly equal to zero where the forecast horizon decreases from N to 1 (N↓); and a sequence of F-tests for the hypothesis that the out-of-sample forecast errors are jointly equal to zero where the forecast horizon increases from 1 to N (N↑). I perform these tests over 1909-92 by applying recursive least squares to (8)-(9); this method also gives a sequence of parameter estimates and associated standard errors.[14] Based on these tests, the two models exhibit a remarkable degree of parameter constancy (figure 3).[15]

The 95 percent confidence intervals for the elasticity estimates implied by these models reveals several features of interest (figure 4). First, income elasticities are positive, significant, and fluctuate over time; own-price elasticities are significant, fluctuate over time, and with the exception of the Almost Ideal model, are negative. This model violates the negativity condition from 1930 to 1970 and I treat this evidence as an argument to not use it here. Second, fluctuations in the income and price elasticities of the Rotterdam model replicate the timing of instability in elasticity estimates reported previously:

> The results reveal significant upward shifts in estimates of both income and price elasticities of U.S. non-oil imports during the early 1960s, followed by significant downward shifts in those estimates during the early 1970's. [Hooper, 1978, p.3]

> Previous investigations of imports had suggested evidence of structural change in the mid-1960s on the basis of split samples but without an unambiguous dating as to when the change may have occurred. We used a series of tests that permitted the data to indicate the presence of structural change. Our results give some support to the earlier findings that change occurred in the mid- to late 1960s. [Stern, Baum, and Green, 1979, p.191]

> This paper provides strong evidence that parameter change occurred in US import demand during the shift to flexible exchange rates and the first oil-price shock. [Ziets and Pemberton, 1993, p. 664]

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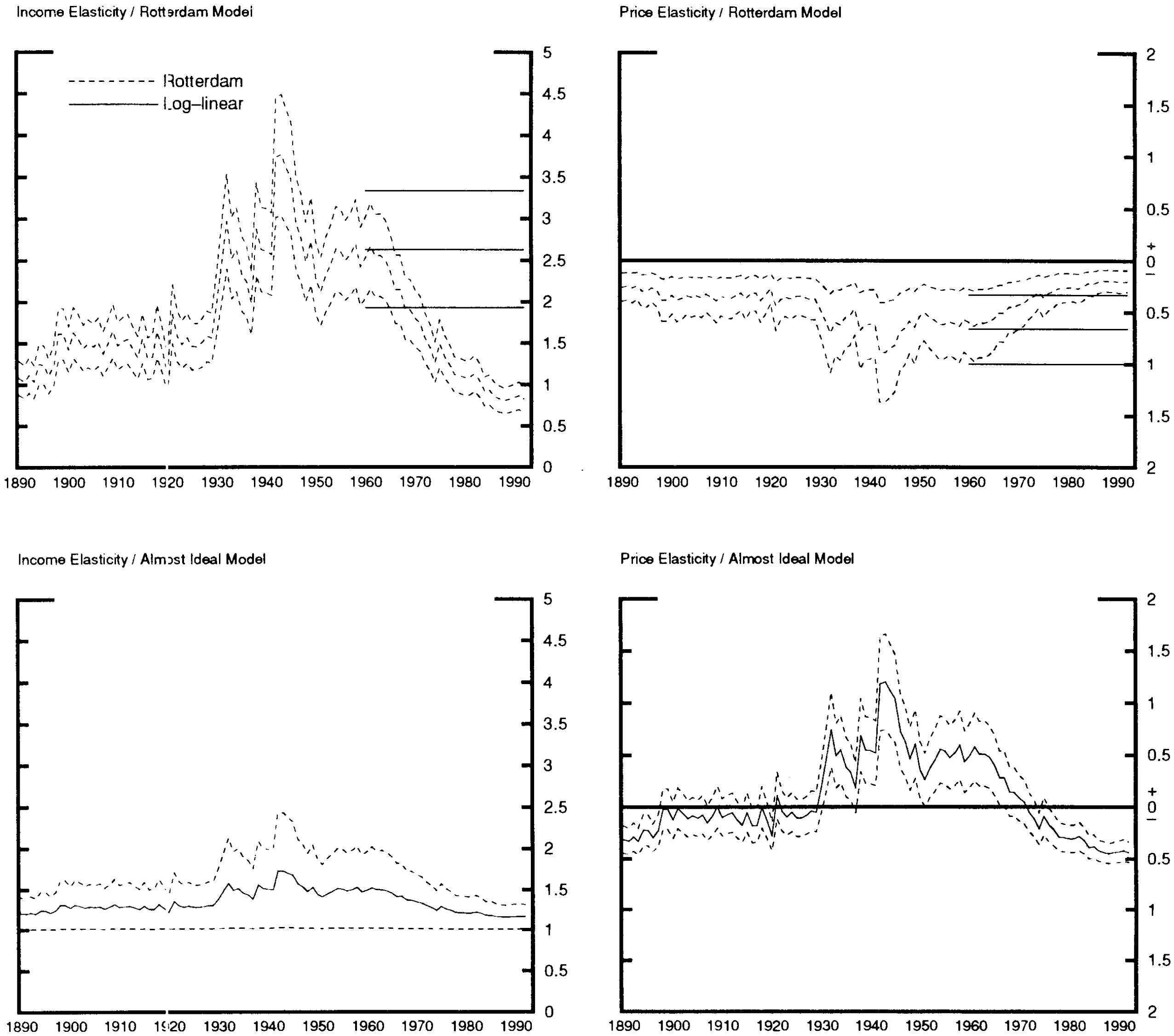
[14] I reserve eighteen observations to serve as initial conditions. Thus for the first sub-period (1891-1909), the value of N is 83; for the second sub-sample (1891-1910), the value of N is 82, and so on.

[15] To examine whether income and prices can be taken as given, I implement the superexogeneity test of Engle, Hendry, and Richard (1983). This procedure involves testing parameter constancy for the equations explaining income and relative prices in terms of exogenous variables. I use government purchases and currency in circulation, both in real and per-capita terms, as instruments and reject the hypothesis of parameter constancy in the equations for income and relative prices. This result, combined with the finding of parameter constancy in (8) and (9), suggests that treating income and prices as exogenous generates no loss of information for estimating the parameters of these two equations.

Figure 4: 95% Confidence Bands for Income and Price Elasticities
United States, 1890–1992



But fluctuations in the elasticities from the Rotterdam model arise from optimization and not from unspecified structural changes.

Third, the elasticity estimates of the log-linear model overstate the sensitivity of U.S. imports to changes in income and prices relative to the estimates from the Rotterdam model. For 1992, the income elasticity of the log-linear model is 2.6 versus 0.8 for the Rotterdam model; price elasticities also differ substantially but the large standard error for the estimate of the log-linear model makes the difference statistically insignificant. Overall, these results confirm the importance of optimization for characterizing the behavior of U.S. imports. What is new about the results is that only through optimization can one recognize the evolution of U.S. trade for estimating elasticities.[16]

5. Conclusion

Treating elasticities as constant parameters and estimating them with sub-samples of the postwar period are two features unifying fifty years of econometric work on U.S. imports. These two features, however, do not provide an adequate understanding of the role of income and prices in determining these imports: the dispersion of elasticities is substantial and the estimates are unstable. A necessary condition for reaching an agreement on the determinants of U.S. imports is the use of all the data available. Otherwise, fluctuations in elasticities arising from the use of different sub-samples precludes arriving at a definitive view of the role of income and prices in determining these imports. But when all the data available are used to estimate the parameters of the popular log-linear model, the results misrepresent the role of income and prices because that model does not treat adequately the time-series properties of the data.

---

[16] To examine the role of tradeables and non-tradeables, I respecify the Rotterdam model as

$$w_{t-1}\Delta\ln(m_t) = \mu\Delta\ln(y_t) + \pi_1\Delta\ln\,(p_{mt}/p_{gt}) + \pi_2\Delta\ln\,(p_{st}/p_{gt}) - \xi\mu\Delta\ln(a_t) + \xi w_{t-1}\Delta\ln(a_t) + u_{\pi t}\,, \quad (8')$$

where p_{gt} is the price of domestically produced goods and p_{st} is the price of domestically rendered services; data for these prices are not available prior to 1930. The estimated compensated, own-price effect (π_1) is -0.0143, which is close to the estimate reported in table 4 and confirms the low substitutability between foreign and domestic tradeables. This conclusion ignores, however, the difference between single-equation and system estimation with cross-equation restrictions; further work is thus required to verify this result.

As an alternative, I estimate the parameters of formulations that use all the data available, recognize the associated time-series properties, and rest on optimization. Recognizing the evolution of the U.S. economy need not be central to estimating income and price effects if the associated estimates are used for short-term forecasting. This practice, however, confuses understanding with predicting, contributes to the large dispersion of estimates found in the literature, and leaves unanswered a key question: If elasticities are constant why are they reestimated so often?

Appendix A: Chronology of Elasticity Estimates for Selected Studies of U.S. Imports

This appendix lists papers reporting econometric estimates of trade elasticities for single-equation models of U.S. aggregate import demand. By design, the appendix excludes studies examining the structure of U.S. trade on the basis of factor content or using non-parametric methods; papers reporting econometric estimates for relatively small components of U.S. imports are not included in this survey. Stern et al. (1976), Thursby and Thursby (1984), Goldstein and Khan (1985), Kohli (1991), and Marquez (1992) provide reviews of the literature that are not limited by these considerations. Table A1 shows the studies used in this paper and their main characteristics which are listed below.

Estimator:	ILS: Indirect Least Squares NLS: Nonlinear Least Squares OLS: Ordinary Least Squares
Price Behavior:	Exogenous: Prices are taken as given for estimation. Endogenous: Prices are not taken as given for estimation.
Dynamic Structure:	DL: Distributed lags. ECM: Error-correction model. Koyck: Lagged dependent variable is the only lag. PDL: Polynomial distributed lag. RL: Rational lag. Shiller: Shiller lags. Static: No allowance for lags.
Homogeneity:	Yes: Estimating equation maintains homogeneity of degree zero in prices. No: In the absence of homogeneity, the price elasticity corresponds to the estimated coefficient on the foreign-price, whether or not it is combined with an exchange-rate term.
Price Data:	Multilateral: Price data do not differentiate across trading partners. Bilateral: Price data differentiate across trading partners.
Data Frequency:	A: Annual; Q: Quarterly; S: Semi-annual.
Trade Data:	Total: Measure of imports includes oil imports. Non-oil: Measure of imports excludes oil imports.
*	Author's aggregation of individual elasticities using trade shares.
i	Author's imputation of standard error. If the study does not report standard errors but indicates that the elasticities are significant, then I impute a t-statistic of 2. If the study does not give a sense of how significant are the elasticities, then I impute a t-statistic of one.

Table A1
Chronology of Elasticity Estimates: Selected Studies for the United States

Study	Estimator/ Price Behavior	Dynamic Structure/ Homogeneity	Price Data/ Frequency; Sample	Trade Data	Elasticity Estimates	
					Income (t-stat)	Price (t-stat)
Chang (1946,table 4)	OLS/Exog.	Statis/Yes	Multilateral/ A;1924-38	Total	1.27 (1.00i)	-0.97 (-1.00i)
Krause (1962,table 3)	OLS/Exog.	Static/Yes	Cross-sec./ A;1947-58	Nonoil	1.00 (1.00i)	-1.98 (-4.13)
Kreinin (1967,table 3)	OLS/Exog.	DL/Yes	Multilateral/ A; 1954-64	Total	1.27 (16.3)	-1.11 (-6.94)
Hein (1968,pp.709)	OLS/Exog.	DL/Yes	Multilateral/ A;1951-65	Total	1.21 (7.56)	-0.62 (-3.10)
Houthakker and Magee (1969, table 1)	OLS/Exog.	Static/Yes	Multilateral/ A;1951-66	Total	1.51 (12.1)	-0.54 (-1.60)
Magee (1972,pp. 8-9)	OLS/Exog.	Static/Yes	Bilateral/ A;1951-69	Nonoil	1.54 (1.00i)	-1.26 (-1.00i)
Taplin (1973, tables 1-2)	OLS/Exog.	Static/Yes	Multilateral/ A;1953-70	Total	1.81 (2.00i)	-1.05 (-2.00i)
Clark (1974,pp.220-8)	OLS/Exog.	PDL/Yes	Multilateral/ Q;1963-73	Nonoil	2.79 (51.2)*	-3.72 (-7.40)*
Miller and Fratiani (1974, table 1)	OLS/Exog.	Koyck/Yes	Multilateral/ Q;1956-72	Total	1.96 (2.00i)	-0.73 (-1.00i)
Ahluwalia and Hernandez-Cata (1975, table 1)	ILS/Endog.	DL/No	Multilateral/ Q;1960-73	Nonoil	1.33 (14.77)	-1.65 (-5.64)
Khan and Ross (1975, table 1)	OLS/Exog.	Static/Yes	Multilateral/ S;1960-72	Total	1.91 (4.87)	-1.00 (-1.90)
Hooper (1976, table 2)	OLS/Exog.	DL/Yes	Multilateral/ Q;1956-75	Nonoil	1.06 (2.00)	-0.54 (-5.32)
Murray and Ginman (1976, table 2)	OLS/Exog.	Static/No	Multilateral/ Q;1961-68	Total	0.96 (3.80)	-1.05 (-1.60)
Khan and Ross (1977, table 2)	OLS/Exog	Koyck/Yes	Multilateral/ Q;1960-72	Total	1.42 (5.68)	-2.16 (-2.00)
Deppler and Ripley (1978, tables 11, 13, 14, and 16)	OLS/Exog.	DL/Yes	Multilateral/ S;1964-76	Nonoil	1.39 (2.24)*	-1.45 (-6.00)*
Hooper (1978, table 3)	OLS/Exog.	Static/Yes	Multilateral/ Q;1955-77	Nonoil	2.03 (10.65)	-1.04 (-3.90)
Wilson and Takacs (1979, table 6)	OLS/Exog.	Shiller/No	Multilateral/ Q;1957-71	Nonoil	4.08 (8.66)	-4.78 (-1.00)

Table A1 (continued)
Chronology of Elasticity Estimates: Selected Studies for the United States

Study	Estimator/ Price Behavior	Dynamic Structure/ Homogeneity	Price Data/ Frequency; Sample	Trade Data	Elasticity Estimates	
					Income (t-stat)	Price (t-stat)
Lawrence (1978, table 6)	OLS/Exog.	DL/Yes	Multilateral/ S;1962-77	Nonoil	3.08 (27.00)	-1.52 (-4.70)
Stern, Baum, and Green (1979, table 2)	OLS/Exog.	DL/No	Multilateral/ Q;1953-76	Total	1.12 (3.24)	-2.18 (-3.41)
Goldstein, Khan, and Officer (1980, table 3)	OLS/Exog.	Static/Yes	Multilateral/ A;1950-73	Total	2.52 (6.37)	-0.68 (-3.30)
Geraci and Prewo (1982, table 1)	OLS/Exog.	Koyck/Yes	Bilateral/ Q;1958-74	Total	1.53 (10.2)	-1.23 (-2.20)
Haynes and Stone (1983, table 1)	OLS/Exog.	Static/No	Multilateral/ Q;1955-79	Total	1.83 (9.42)	-0.61 (-3.16)
Warner and Kreinin (1983, table 2)	OLS/Exog.	PDL/No	Multilateral/ Q;1970-80	Nonoil	2.01 (2.00i)	-2.53(-2.00i)
Helkie and Hooper (1988, table 4)	OLS/Exog.	PDL/Yes	Multilateral/ Q;1969-84	Nonoil	2.11 (5.30)	-1.15 (-10.0)
Cline (1989, table 4A.3)	OLS/Exog.	DL/Yes	Multilateral/ Q;1973-87	Total	2.44 (2.00i)	-1.36 (-1.00i)
Deyak, Sawyer, and Sprinkle (1989, table 1)	OLS/Exog.	Koyck/Yes	Multilateral/ Q;1958-83	Total	1.07 (4.60)	-0.29 (-1.00)
Krugman (1989, table 3)	OLS/Exog.	DL/Yes	Multilateral/ A;1971-86	Nonoil	1.31 (2.98)	-0.93 (-3.10)
Moffet (1989, table 5)	OLS/Exog.	PDL/No	Multilateral/ Q;1967-87	Total	2.50 (9.07)	-0.69 (-7.40)
Lawrence (1990, table 8)	OLS/Exog.	PDL/Yes	Multilateral/ S;1976-90	Nonoil	0.73 (3.00)	-1.47 (-14.3)
Marquez (1990, table 2)	OLS/Exog.	RL/Yes	Multilateral/ Q;1973-85	Total	1.94 (4.97)	-0.92 (-4.80)
Blecker (1992, table A-1)	OLS/Exog.	PDL/Yes	Multilateral/ Q;1975-89	Nonoil	2.56 (3.00)	-0.97 (-2.00)
Zietz and Pemberton (1993, table 5)	OLS/Exog.	ECM-DL/ Yes	Multilateral/ Q;1976-90	Nonoil	2.48 (40.9)	-1.14 (-15.5)
Clarida (1994, p.306)	NLS/Exog.	ECM-DL/ Yes	Multilateral/ Q;1968-90	Nonoil	2.15 (1.00i)	-0.95 (-1.00i)

Appendix B: Data Sources

The main data sources are the *Survey of Current Business* (Survey) and the *Historical Statistics of the United States: Colonial Time to 1970* (Historical Statistics) assembled by the Bureau of Economic Analysis of the U.S. Department of Commerce; alternative sources are explicitly indicated whenever they are used. The Bureau of Economic Analysis offers data over 1929-92 with "real" variables expressed in 1987 prices. For the period 1890-1928, the Historical Statistics offers a comparable database but uses a different base year to deflate nominal magnitudes. Thus to obtain data for the 1890-1928 period in 1987 prices, I extrapolate backwards the series from the Survey using the growth rates of the series of the Historical Statistics.[17]

Data Series

Figure B1 shows annual data for the level of U.S. real GNP, population, per-capita GNP, and the growth rates of these variables since 1890.[18] According to the data, the expansion of GNP has been largely sustained and the volatility in the growth rate of GNP declined considerably in the postwar period; the largest one-year decline in GNP occurs in 1946 because of the cutback in defense spending. Total U.S. population shows also a sustained expansion.[19] Easterlin (1980) argues that fluctuations in the growth rate of population up to 1940 arose from changes in the immigration rate; since 1940, changes in the growth rate of population stem from changes in the fertility rate induced by changes in the ability of a given cohort to provide comparable living standards to their offsprings. The figure also shows that the share of population with at least 65 years of age remained largely unchanged at 4 percent from 1890 to 1920 and increased steadily to reach 12 percent by 1992.

Figure B2 shows the evolution of the deflators for GNP and non-oil imports, the tariff rate, and the relative price of imports since 1890. The data reveal substantial price instability from 1890 to 1940 including the deflationary pressures of the 1930s. Since 1973, increases in the GNP deflator have been interrupted by the recessions of 1980 and 1990. Up to 1945, non-oil import prices show fluctuations as large as those of the GNP deflator especially during the WWI period. The decline in import prices in 1920 is the largest decline over the last century. Finally, figure B2 shows the evolution of the total and non-oil tariff rate since 1890. I measure these rates as the ratio between the level of duties and nominal merchandise imports excluding tariffs. The tariff rate shows wide swings prior to 1944 and table B1 has a brief chronology of the main pieces of tariff legislation.

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[17] Romer (1989) reports estimates of U.S. real GNP (and its deflator) for the period 1869-1929 in 1982 prices; thus differences in the choice of base year are still present. Nevertheless, the correlation between the series that I use and those reported by Romer, for 1890-1929, are 0.994 for real GNP and 0.998 for the GNP deflator.

[18] I use GNP as a measure of income instead of GDP because data for GDP are not readily available for the 19th century and the early part of the 20th century.

[19] The data on population include Armed Forces overseas. Data for population including Armed Forces prior to 1930 are not available except for 1917-19 which appear in footnote 1 of page 8 of Historical Statistics. I adjust the growth rate to include Armed Forces overseas for 1917-19.

Figure B1: GNP and Population
United States, 1890–1992

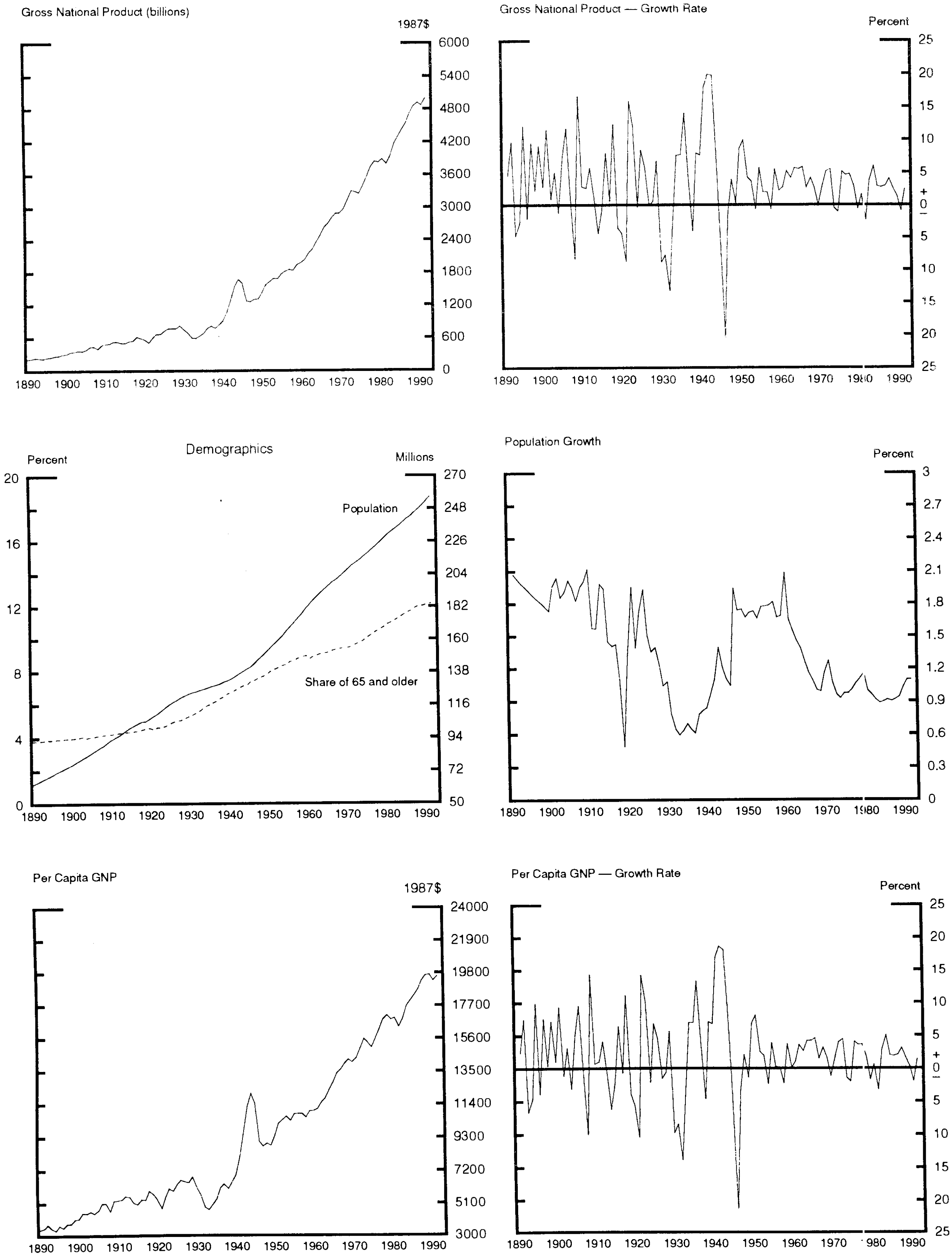
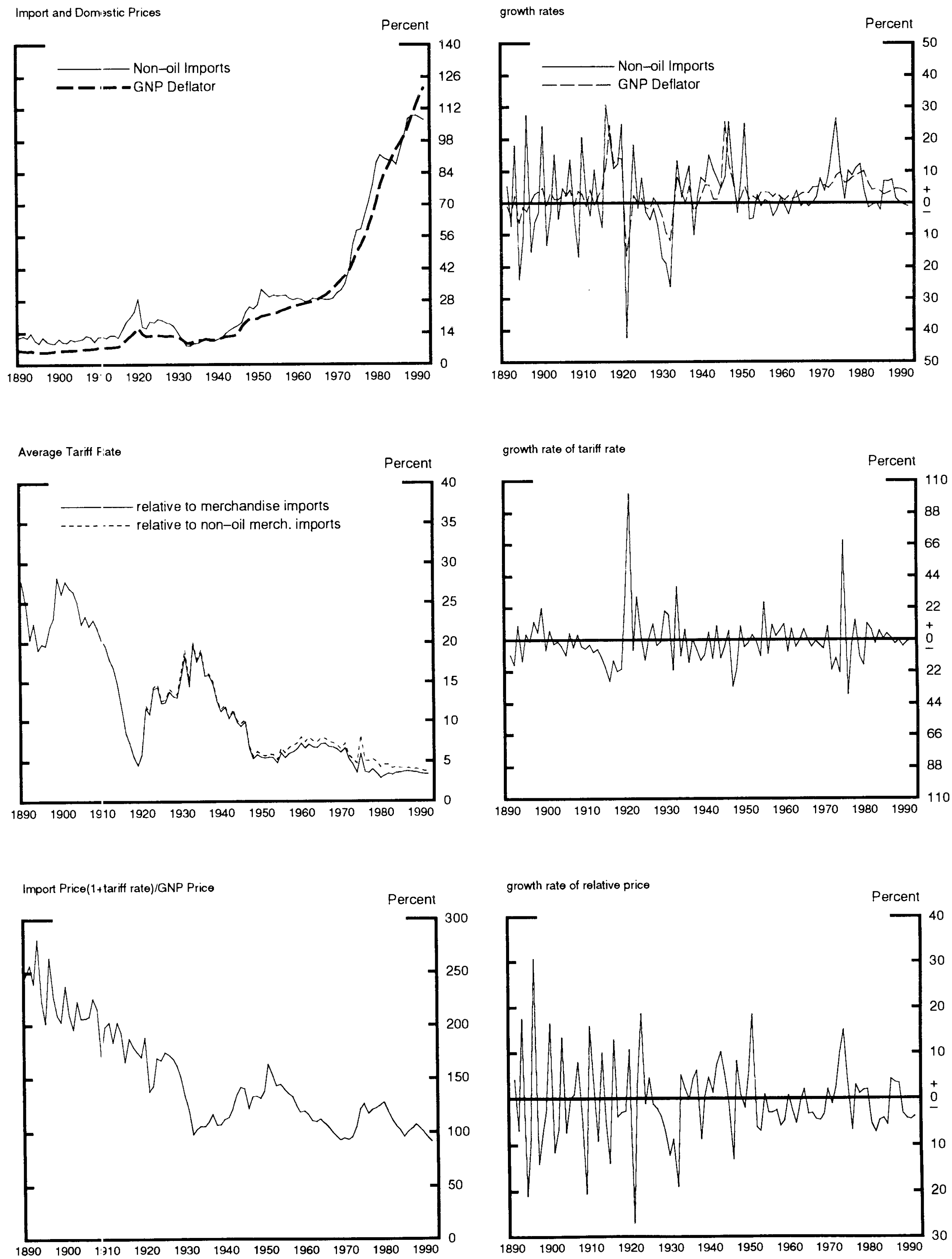


Figure B2: Import and Domestic Prices

United States, 1890–1992



Sources

1. *Real GNP in 1987 Prices*:
1929-1992: Survey December 1992, table 2.
1890-1928: Backward extrapolation using growth rates from Historical Statistics, p. 224, series F-3.

2. *Total Resident Population*:
1929-1992: Survey, table 2.1.
1890-1928: Backward extrapolation using growth rates from Historical Statistics, p. 8, series A-7.

3. *Merchandise Imports, Current Prices*:
1929-1992: Survey, table 4.1.
1890-1928: Backward extrapolation using growth rates from Historical Statistics, p. 889, series U-219.

4. *Merchandise Imports, 1987 Prices*:
1929-1992: Survey, table 4.2.
1890-1928: Backward extrapolation using growth rates from Historical Statistics, p. 893, series U-237.

5. *Imports of Petroleum and Petroleum Products, Current Prices*:
1965-1992: Survey, table 3.B-U.S. Merchandise Trade.
1908-1964: Backward extrapolation using growth rates from Historical Statistics, p. 900, series U-316.
1890-1907: Volume of oil imports are negligible (Historical Statistics, series M-140) and set to zero.

6. *Imports of Petroleum and Petroleum Products, 1987 Prices*:
Imports of Petroleum and Products in current prices (#5) deflated by the associated price index (#10 below).

7. *Non-oil Imports, Current Prices:*
Difference between total imports in current prices (#3) and oil imports in current prices (#5).

8. *Non-oil Imports, 1987 Prices*:
Difference between total imports in 1987 prices (#4) and oil imports in 1987 prices (#6).

9. *GNP Deflator*:
1929-1992: Ratio between nominal and real GNP; for data on nominal GNP: Survey December 1992, table 1.
1890-1928: Backward extrapolation using growth rates from Historical Statistics, p. 224, series F-5.

10. *Oil-import Price Deflator*:
1967-1992: Ratio between current-price and 1987-price data for fuel imports (BOP basis) from the U.S. Commerce Department, Merchandise Trade Statistical Release.
1947-1966: Grows at the rate of the U.S. domestic price of oil (Producer Price Index Press Release, Bureau of Labor Statistics).
1890-1946: Grows at the rate of the price of domestic petroleum production ($/barrel): Historical Statistics, p. 593, series M-139.

11. *Non-oil Import Price Deflator*:
Ratio of non-oil imports in current prices (#7) to non-oil imports in 1987 prices (#8).

12. *Custom Duties*:
1929-1992: Survey, table 3.2.
1890-1928: Backward extrapolation using growth rates from Historical Statistics, p. 1105, series Y-344.

13. *Export of Goods and Services in Current Prices*:
1929-1992: Survey December 1992, table 1.

1890-1928: Backward extrapolation using growth rates from Historical Statistics, p. 891, series U-225 (index of export volume) times U-226 (index of export price).

14. *Export of Goods and Services in 1987 Prices*:
1929-1992: Survey December 1992, table 2.
1890-1928: Backward extrapolation using growth rates from Historical Statistics, p. 891, series U-225 (index of export volume).

15. *Import of Goods and Services in Current Prices*:
1929-1992: Survey December 1992, table 1.
1890-1928: Backward extrapolation using growth rates from Historical Statistics, p. 890, series U-219.

16. *Import of Goods and Services in 1987 Prices*:
1929-1992: Survey December 1992, table 2.
1890-1928: Backward extrapolation using growth rates from Historical Statistics, p. 893, series U-237.

17. *Domestic Expenditures in Current Prices*:
GNP in 1987 prices (#1) times GNP deflator (#9) + Imports of Goods and Services in current prices (#15) - Exports of Goods and Services in current prices (#13).

18. *Domestic Expenditures in 1987 Prices*:
GNP in 1987 prices (#1) + Imports of Goods and Services in 1987 prices (#16) - Exports of Goods and Services in 1987 prices (#14).

19. *Price Deflator for Domestic Expenditures*:
Ratio between Domestic Expenditures in current prices (#17) and 1987 prices (#18).

20. Population Age:
1988-1992: Statistical Abstract of the United States, 1993, U.S. Department of Commerce, Bureau of the Census, p. 14 (table 14) and p. 33 (table 35).
1900-1987: Liesner (1989), table US.9 p. 92-93.
1890-1899: Linear interpolation using the 1890 and 1900 values reported in Liesner (1989).

21. *Price of Domestic Tradeables*:
1947-1992: Ratio of current and constant price expenditure on domestic goods: Survey, table 1.3 (current prices) and table 1.4 (constant prices).
1930-1946: Backward extrapolation using growth rates from Historical Statistics, p. 198, series E-16.

22. *Price of Domestic Non-Tradeables*:
1947-1992: Ratio of current and constant price expenditure on domestic services: Survey, table 1.3 (current prices) and table 1.4 (constant prices).
1930-1946: Backward extrapolation using growth rates from Historical Statistics, p. 198, series E-17.

23. *Government Purchases in Current Prices*:
1929-1992: Survey, table 1.1.
1890-1928: Backward extrapolation using growth rates from Liesner (1989), table US.1, p. 74.
Real government purchases equal nominal government purchases (#23) deflated by the GNP deflator (#9).

24. *Currency in Circulation*:
1959-1992: Board of Governors of the Federal Reserve System, Weekly Statistical Release, table 4.
1890-1958: Historical Statistics, p. 992, series X-410.

Table B1: Chronology of U.S. Legislation on External Tariffs

Date	Description
Pre-1861	Average tariff rate is 20%; the maximum rate is 24% (p. 370)
1861	Morril Act: Higher taxes to finance Civil War; by 1864 the average tariff rate reaches 47% (p. 370)
1872	Tariff rates are lowered by 10%, but raised again in 1875 (p. 370)
1890	McKinley Act raises the level of tariff rates to 50% (p. 371)
1897	Tariff rates are raised to almost 60%, except for raw materials and semifinished products (p. 371)
1913	Underwood-Simmons bill lowers tariff rates and increases the list of items exempt from tariffs. The average tariff rate is 25%.
1921	Emergency Tariff Act (p. 662).
1922	Fordney-McCumber Act raises average tariff rate to 33% (p. 663).
1930	Hawley-Smoot Act raises average tariff rate to 40% (p. 663).
1934	Congress gives the President authority to negotiate tariffs p. 663).
1945	President's power to negotiate tariffs is expanded (p. 663).
1947	Creation of the GATT (p. 665).
1962	Trade Expansion Act gives power to the President to lower tariffs (p. 665).
1971	Nixon Administration announces a 10% surcharge on dutiable imports (p. 662).

Source: Robertson (1973); page numbers refer to this publication.

Appendix C: Order of Integration

I apply the procedure of Dickey and Pantula (1987) to determine the order of integration of the logarithms of per-capita, real domestic expenditures, y_t ; per-capita, real non-oil imports, m_t ; and the tariff-adjusted relative price, p_t , using annual observations for 1890-1992. This procedure tests sequentially the hypotheses of multiple unit-roots, single unit-root, and stationarity while removing the assumption of at most one unit-root maintained by the Dickey-Fuller test.

Sequential Testing for Multiple Unit Roots

I assume that a given variable z_t evolves over time according to

(C1) $$z_t = \mu + \gamma T + (\rho_1 + \rho_2)z_{t-1} - (\rho_1\rho_2)z_{t-2} + e_t ,$$

where μ is a drift parameter; γ is a deterministic-trend parameter; both ρ_1 and ρ_2 are the roots of the dynamic process; T denotes time; and e_t is a white-noise disturbance.[20] Re-arranging (C1) gives

(C2) $$\Delta^2 z_t = \mu + \gamma T + \alpha_1 \Delta z_{t-1} + \alpha_2 z_{t-1} + e_t ,$$

where $\alpha_1 = (\rho_1\rho_2 - 1)$ and $\alpha_2 = (\rho_1 - 1)(\rho_2 - 1)$. If z_t has two unit roots ($\rho_1 = \rho_2 = 1$), then $\alpha_1 = \alpha_2 = 0$. If z_t has one unit root ($\rho_1 = 1$, $\rho_2 < 1$) then $\alpha_1 < 0$; $\alpha_2 = 0$. Noting that $\alpha_2 = 0$ in both the null and the alternative hypotheses, Dickey and Pantula propose a two-step testing strategy. The first step sets $\alpha_2 = 0$, applies least squares to $\Delta^2 z_t = \mu + \gamma T + \alpha_1 \Delta z_{t-1} + e_t$, and tests the null hypothesis of two unit roots ($\alpha_1 = 0$) versus the alternative of only one unit root ($\alpha_1 < 0$). If the null hypothesis is rejected, then the order of the series is at most one. The second step applies least squares to $\Delta^2 z_t = \mu + \gamma T + \alpha_1 \Delta z_{t-1} + \alpha_2 z_{t-1} + e_t$ and tests the null hypothesis of one unit root ($\alpha_2 = 0$) versus the alternative of stationarity ($\alpha_2 < 0$).

Test Results

The estimates of α_1 are negative and significantly less than zero for all three series (table C1). Thus the data reject the hypothesis of two unit roots for the three series and suggest that their order of integration is at most one. The estimates of α_2 are not significantly different from zero for imports and relative prices -- that is, these two variables are integrated of order one. For domestic expenditures, however, the estimate of α_2 is significantly less than zero which suggests that this series is stationary. Finding that the logarithm of per-capita U.S. domestic expenditures in real terms is stationary contradicts the generally accepted view that the main U.S. economic aggregates are non-stationary (Stock and Watson, 1988). Previous findings, however, normally exclude from their samples the WWII period and do not use the Dickey-Pantula sequential procedure adopted here.

Excluding 1941-1946 from the sample gives estimates of α_1 from the first step that are negative and significantly less than zero for all three series (table C1) which suggests the absence of multiple unit roots; the estimates of α_2 are not significantly different from zero which suggests that the associated series are non-stationary. Thus, abstracting from the period 1941-1946, the data for the logarithms of per-capita domestic expenditure, per-capita non-oil imports, and relative prices are integrated of order 1.

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[20] Equation (C1) extends section 4.5.1 of Banerjee et al. (1993) by adding both a drift and a deterministic trend to the process explaining z.

Table C1
Test Results for Order of Integration for Imports, Expenditures, and Relative Prices: United States,1890-1992

Sample: 1892-1992	Step 1		Step 2	
	α_1	$t(\alpha_1)$	α_2	$t(\alpha_2)$
ln(m)	-1.32	-13.8*	-0.02	-0.7
ln(y)	-0.77	-7.8*	-0.21	-4.0*
ln(p)	-1.17	-11.8*	-0.17	-2.8

Sample: 1892-1940	Step 1		Step 2	
	α_1	$t(\alpha_1)$	α_2	$t(\alpha_2)$
ln(m)	-1.38	-10.1*	-0.41	-2.7
ln(y)	-1.01	-6.8*	-0.26	-2.6
ln(p)	-1.30	-9.2*	-0.44	-3.0

Sample: 1949-1992	Step 1		Step 2	
	α_1	$t(\alpha_1)$	α_2	$t(\alpha_2)$
ln(m)	-1.13	-7.4*	-0.30	-2.8
ln(y)	-0.88	-5.7*	-0.32	-2.9
ln(p)	-0.67	-4.5*	-0.17	-2.4

Notes:

The critical values for the full sample are -3.73 (T=100) and -3.66 (asymptotic); for the sub-samples, the critical values are -3.8 (T=50) and -3.66 (Banerjee et al., 1993, p.103, c-block of table 4.2). An entry with * means that the data reject the associated null hypothesis. The estimation samples allow for lags.

Step 1: Estimate $\Delta^2 z_t = \alpha_0 + \alpha_1 \Delta z_{t-1} + \alpha_3$ Time $+ e_t$
Test H_0: two unit roots ($\alpha_1 = 0$) versus H_1: one unit root ($\alpha_1 < 0$)

Step 2: Estimate $\Delta^2 z_t = \alpha_0 + \alpha_1 \Delta z_{t-1} + \alpha_2 z_{t-1} + \alpha_3$ Time $+ e_t$
Test H_0: one unit root ($\alpha_2 = 0$) versus H_1: stationarity ($\alpha_2 < 0$)

To assess the robustness of these conclusions, I report the results from the Augmented Dickey-Fuller (ADF) formulation:

(C3) $$\Delta z_t = \mu + \gamma T + \alpha_1 \Delta z_{t-1} + \ldots + \alpha_4 \Delta z_{t-4} + \beta z_{t-1} + e_t.$$

The ADF procedure applies least squares to (C3) and tests the null hypothesis that β=0. If β is not significantly different from zero, then z_t has a unit root. All of the estimated t-statistics for β are above the ADF critical value (-3.73) which suggests that the data cannot reject the hypothesis that β = 0 (table C2).

Table C2
Augmented Dickey-Fuller Test, 1890-1992

Variables	ln(m)	ln(y)	ln(p)
Test Statistic	-1.83	-3.17	-2.76

Several conclusions emerge from these tests. First, the data for the logarithms of the series used in this analysis appear to be non-stationary. Nevertheless, the inclusion of the WWII period in the estimation sample induces stationarity in the data for the logarithm of domestic expenditure because of the magnitude of the 1946 contraction in real GNP. Second, the unambiguous support to non-stationarity provided by the ADF test results stands in contrast to the evidence from the Dickey-Pantula procedure. As noted by Banerjee et al. (1993), the choice of test statistic matters for establishing the order of integration. From the standpoint of this paper, I will treat the order of integration of these series as being equal to one.

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