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Trade Elasticities for G-7 Countries

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Abstract: This paper reports the results of a project to estimate and test the stability properties of conventional equations relating real imports and exports of goods and services for the G-7 countries to their incomes and relative prices. We begin by estimating cointegration vectors and the error-correction formulations. We then test the stability of these equations using Chow and Kalman-Filter tests. The evidence suggests three findings. First, conventional trade equations and elasticities are stable enough, in most cases, to perform adequately in forecasting and policy simulations. Equations for German trade, as well as equations for French and Italian exports are an exception. Second, income elasticities of U.S. trade have not been shifting in a direction that will tend to ease the trend toward deterioration in the U.S. trade position. The income-elasticity gap for Japan found in earlier studies was not confirmed in this analysis. Finally, the price channel is weak, if not wholly ineffective, in the case of continental European countries.

Keywords: Exports, imports, elasticities, cointegration, error-correction modeling, G-7 countries

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1 Introduction and Summary

This paper reports the results of a project to estimate and test the stability properties of conventional equations relating real imports and exports of the G-7 countries to their incomes and relative prices. Our primary goal is to analyze the extent to which historical experience, as incorporated in these estimated equations, can be used as a reliable guide to future trends in real trade flows. Evidence of instability in estimated parameters may be indicative of changes in more fundamental factors affecting global trade and production, varying measurement error, omitted variables, and other forms of equation misspecification.

Our analysis follows partly in the tradition of Houthakker and Magee (1969), who pointed out that if a country's income elasticities of imports and exports differ significantly, and if that difference persists in the face of converging income growth rates across countries, then the country will experience a growing trade imbalance, ceteris paribus. Existing estimates of income elasticities of imports and exports suggest that at current trend rates of U.S. and foreign GDP growth, either the U.S. external imbalance will widen indefinitely or relative prices will have to adjust over time to keep it from doing so. However, are the existing estimates reliable enough to be treated as constants? If not, then such predictions are unwarranted. Our analysis also follows in the tradition of the literature on the price "elasticity pessimism" of the early 1950s, when analysts wondered whether trade flows were price elastic enough for exchange rate changes to engender desired adjustments of external balances.

We begin by estimating conventional trade equations for real imports and exports of goods and services of the G-7 countries using cointegration and error-correction estimation techniques. We then test the stability of these equations using a variety of Chow tests. Finally, we test the stability of the individual income and price elasticities in a simplified version of the conventional equations using Kalman filter estimation techniques. Our key findings with respect to equation and parameter stability are:

1. Based on the Chow tests for overall equation stability, we do not find evidence of pronounced or chronic instability of elasticities during the 1990s for the majority of formulations. Our results contrast with several earlier studies that found significant evidence of instability in U.S. and Japanese trade equations during the 1970s and 1980s.

2. We find a few exceptions to this pattern of stability, however. Equations for German trade, as well as equations for French and Italian exports, show substantial parameter instability around the time of German reunification. This instability is confirmed by estimates of static formulations based on the Kalman Filter technique.

Our findings about the magnitudes of elasticity estimates are also of interest:

1. The gap between the income elasticities of U.S. imports and exports proved quite robust. The income-elasticity gaps for Japan, France, Italy, and Germany appear to be quite small, with elasticity levels for both imports and exports generally centered around 1.5 for the European countries and a bit lower for Japan. This result for Japan is quite different from some studies, which found the income elasticity of exports to be well in excess of that for imports.

2. Estimated elasticities for the export equations are less reliable than those of imports. The construction of data for foreign prices and income, used in the export equations, rests on assumptions that might introduce aggregation biases.

3. Canadian, Japanese, U.K., and U.S. trade flows appear to be considerably more price responsive than those in continental Europe. Indeed, the estimated long-run price elasticities of imports and exports for Germany and France were found surprisingly low (in absolute value)–less than 0.5 in all cases–and jointly insufficient to meet the Marshall-Lerner condition required for successful external adjustment via the exchange rate channel.

In light of these findings, we draw the following implications:

1. Conventional trade equations and elasticities are stable enough, in most cases, to perform adequately in forecasting and policy simulations. Equations for German trade, as well as equations for French and Italian exports are an exception.

2. Income elasticities of U.S. trade have not been shifting in a direction that will tend to ease the trend toward deterioration in the U.S. trade position. The income-elasticity gap for Japan found in earlier studies was not confirmed in this analysis.

3. Our elasticity estimates suggest that the price channel is weak, if not wholly ineffective, in the case of continental European countries, although this result could also reflect the peculiarities of intra-European trade.

2 Previous Empirical Work

A substantial empirical literature exists on the estimation of income and price elasticities in international trade, much of it focussed on U.S. trade. Goldstein and Kahn (1985) survey empirical estimates of long-run income and price elasticities for imports and exports of major industrial countries; Hooper and Marquez (1995) also survey of price elasticities for trade in the United States, Japan, and Germany, while Marquez (1998) surveys estimates of income and price elasticities for U.S., Japanese, and Canadian imports.[1]

Chart 1 shows estimates of income and price elasticities for total merchandise exports and imports for G-7 countries from three representative studies: Cline (1989), Houthakker and Magee (1969), and Marquez (1990). These estimates are based on pre-1987 data and assume that the elasticities are constant over the sample studied. Estimates for the United States show income elasticities for imports centered around 2.0 and consistently higher than elasticities for exports. Estimates of the income elasticities for Japanese trade have an asymmetry in the opposite direction: the income elasticity for exports exceeds that of imports by a substantial margin. Estimated price elasticities, shown in the shaded areas of the chart, are smaller, in absolute value, than the income elasticities–especially for several European countries–and are less than one. Hooper and Marquez's survey indicates that price elasticities fall in a range of -1 to -1.5, on average, for U.S. and Japanese with European trade showing considerably less price responsiveness; Goldstein and Kahn report higher price elasticities for German imports based on data through the early 1970s.

There has been some limited research on the question of parameter stability in international trade equations, with most of the work focussing on U.S. import equations. Hooper (1978), Stern *et al.* (1979) using rolling-regression techniques, and Deyak, Sawyer, and Sprinkle (1989) and Zietz and Pemberton (1993), both using split-sample estimation techniques (Chow tests), found evidence of significant parameter instability in the U.S. import equation during the 1970s and/or 1980s; see also Maskus (1983). These studies found some evidence of an upward drift in the income elasticity of imports, but not consistently so across all of these studies. More often than not, an increase in the income elasticity tended to go hand in hand with a decrease in the price elasticity, and vice versa. Hooper (1978) also considered U.S. exports and found elasticity estimates to be much more stable than those for imports during the 1960s and 1970s. Ceglowski (1997) investigated the stability of Japanese import and export equations using both Chow tests and recursive estimation techniques and found significant evidence of instability in both during the mid-1980s.

3 Estimation: Model, Data, and Results

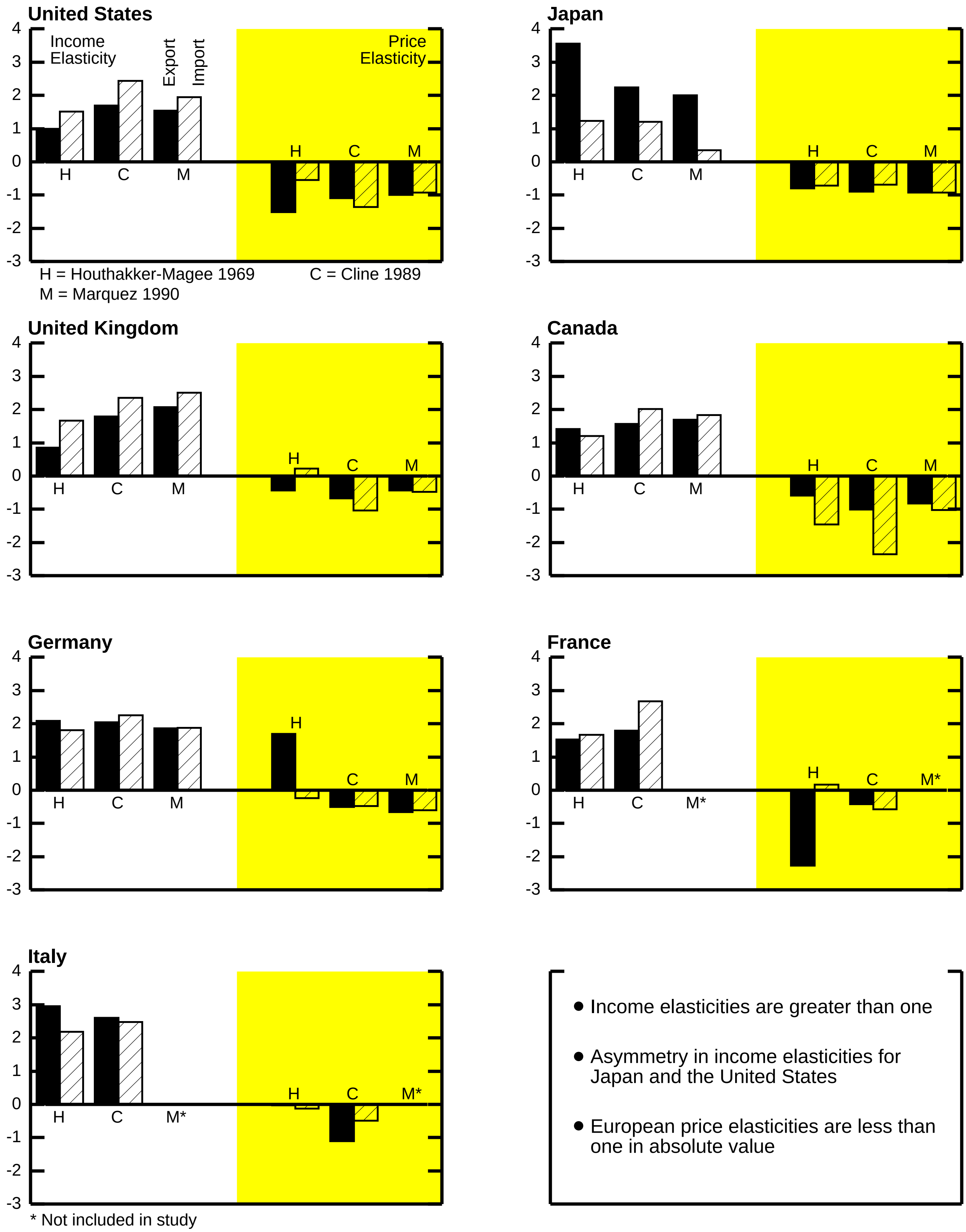
3.1 Model.

The model we use is the conventional treatment of trade flows as a function of real incomes and relative prices. The model assumes that price homogeneity holds and that the elasticities of imports and exports with respect to income and relative prices

[1] For an earlier review, see Stern et al. (1976); for a review focusing on U.S. elasticities see Sawyer and Sprinkle (1996).

Chart 1

Estimated Income and Price Elasticities for Trade: Selected Studies



are constant over time. Letting lower-case letters denote logarithmic values, we model country i's real imports of goods and services (m_i) in terms its real GDP (y_i) and the relative price of imports (rpm_i); real exports of goods and services (x_i) are determined by foreign income (fy_i) and the relative price of exports (rpx_i).

For estimation, our model consists of systems of equations for real imports and exports for each of the G-7 countries as well as their real GDPs and relative prices. These systems are then used to obtain cointegrating vectors (or long-run relationships) for real imports and exports.

The import system is

$$m_{i,t} = \sum_{j=1}^{n} \alpha_{ij} m_{i,t-j} + \sum_{j=1}^{n} \gamma_{ij} y_{i,t-j} + \sum_{j=1}^{n} \lambda_{ij} rpm_{i,t-j} + \varepsilon_{i,t} \tag{1}$$

$$y_{i,t} = \sum_{j=1}^{n} \tau_{ij} m_{i,t-j} + \sum_{j=1}^{n} \upsilon_{ij} y_{i,t-j} + \sum_{j=1}^{n} \phi_{ij} rpm_{i,t-j} + \zeta_{i,t} \tag{2}$$

$$rpm_{i,t} = \sum_{j=1}^{n} \theta_{ij} m_{i,t-j} + \sum_{j=1}^{n} \nu_{ij} y_{i,t-j} + \sum_{j=1}^{n} \chi_{ij} rpm_{i,t-j} + \varphi_{i,t} \tag{3}$$

where $(\varepsilon_{i,t}\ \zeta_{i,t}\ \varphi_{i,t})\tilde{}N(0, \Lambda_m)$.

The export system is

$$x_{i,t} = \sum_{j=1}^{n} \delta_{ij} x_{i,t-j} + \sum_{j=1}^{n} \kappa_{ij} fy_{i,t-j} + \sum_{j=1}^{n} \xi_{ij} rpx_{i,t-j} + \mu_{i,t} \tag{4}$$

$$fy_{i,t} = \sum_{j=1}^{n} \rho_{ij} x_{i,t-j} + \sum_{j=1}^{n} \eta_{ij} fy_{i,t-j} + \sum_{j=1}^{n} \iota_{ij} rpx_{i,t-j} + \sigma_{i,t} \tag{5}$$

$$rpx_{i,t} = \sum_{j=1}^{n} \omega_{ij} x_{i,t-j} + \sum_{j=1}^{n} \varsigma_{ij} fy_{i,t-j} + \sum_{j=1}^{n} \psi_{ij} rpx_{i,t-j} + \epsilon_{i,t} \tag{6}$$

where $(\mu_{i,t}\ \sigma_{i,t}\ \epsilon_{i,t})\tilde{}N(0, \Lambda_x)$.

Recognizing that the dynamic adjustment to various shocks can differ significantly between imports and exports and across countries, we also specify error-correction equations. These equations constrain the long-run elasticities to be the cointegrating vector and allow for a more complete specification of short-term income and price elasticities. For exports, the formulation is

$$\Delta x_{it} = \sum \alpha_{ij} \Delta x_{i,t-j} + \sum \tau_{ij} \Delta fy_{i,t-j} + \sum \rho_{ij} \Delta rpx_{i,t-j} + \Omega_i ECM_{ix,t-1}, \tag{7}$$

where ECM_{ix} is the difference between actual exports and their long-run value as predicted by the cointegrating relationship among x_i, fy_i, and rpx_i from (4)-(6) above. The "error-correction" coefficient Ω_i indicates the proportion of ECM_{ix} that is closed in the current period. For imports, the formulation is

$$\Delta m_{it} = \sum \varkappa_{ij} \Delta m_{i,t-j} + \sum \beta_{ij} \Delta y_{i,t-j} + \sum \pi_{ij} \Delta rpm_{i,t-j} + \Phi_i ECM_{im,t-1}, \qquad (8)$$

where ECM_{im} is the difference between actual imports and their long-run value as predicted by the cointegrating relationship among m_i, y_i, and rpm_i from (1)-(3). The "error-correction" coefficient, Φ_i, indicates the proportion of ECM_{im} that is closed in the current period.

3.2 Data

Quarterly data on real imports and exports of goods and services, GDPs, and the components of relative prices for each of the G-7 countries were obtained from national accounts statistics. Appendices E-K show the evolution for imports, exports, domestic GDP, foreign GDP, and relative prices for trade in each country.

Chart 2 shows that in all of the G-7 countries, real exports and real imports have risen as a share of real GDP over the past three to four decades. These shares are currently largest in Canada, a result of the sizeable increase in these ratios in that country since the 1960s. Differences in trade outcomes are reflected in the paths of current-account balances relative to GDP. As is clear from Chart 3, over the past twenty years, external imbalances for some of the G-7 countries have widened over fairly long intervals to reach significant size. For the United States, the tendency toward larger current account deficits was interrupted in the late 1980s. Canada and the United Kingdom have similarly shown episodes of widening deficits. For Japan, the current account surplus as a share of GDP reached a peak in the mid-1980s, but the surplus remains sizeable. The German surplus was at least temporarily eliminated by the effects of unification.

Ideally, price data should (1) have the same product coverage as the aggregate measure of trade being studied and (2) be consistent with the estimation model. For imports, our estimates rest on the conventional ratio of import prices, PM_i, to the domestic GDP deflator, PY_i:

$$rpm_i = \log(\frac{PM_i}{PY_i}).$$

Product coverage for this measure corresponds to the measure of trade being studied and, by allowing the share of imports in GDP to change over time, rpm_i incorporates changes in the country's openness to trade. At the same time, rpm_i might include products that are not consistent with the imperfect substitute model used here. For

Chart 2

Real Exports and Imports Relative to GDP

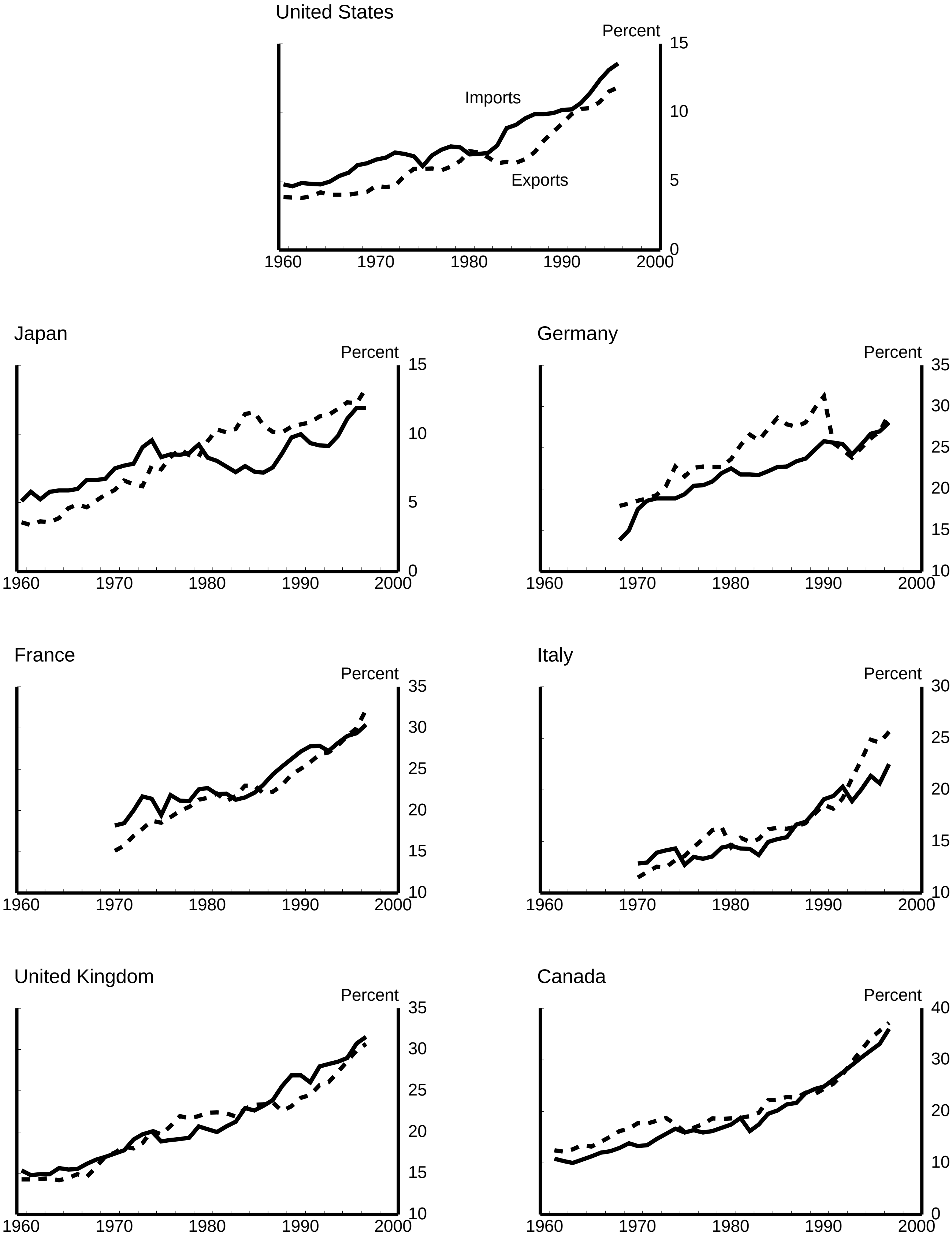
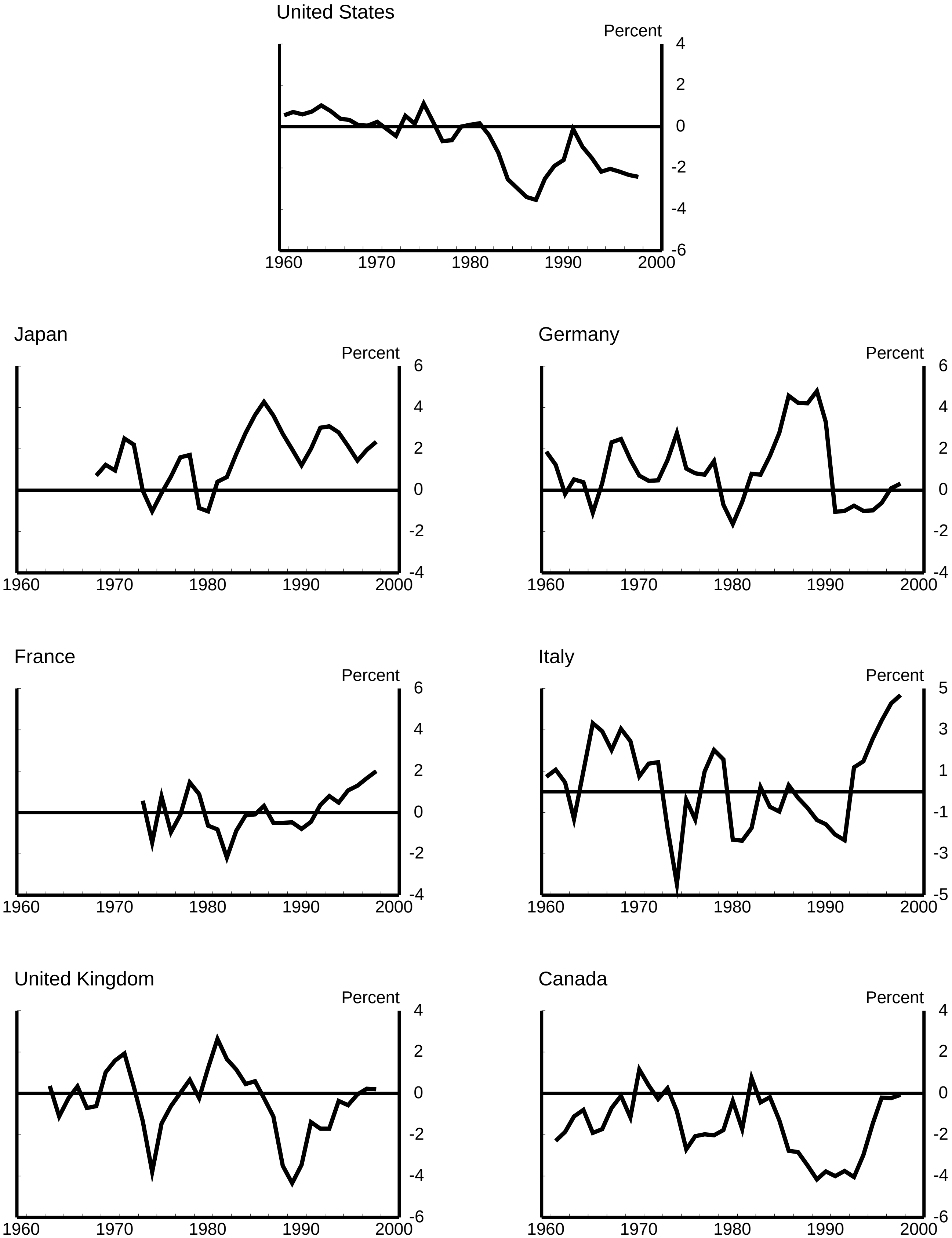


Chart 3

Current Account Balance Relative to GDP



Source: OECD database, dowloaded from OLIS example, rpm_{us} includes the price of oil imports; these imports are a perfect substitute for domestic oil which is produced abundantly in the United States.

For exports, we use the export price, PX_i, relative to the foreign GDP deflator, PYF_i:

$$rpx_i = \log(\frac{PX_i \cdot E_{\$/i}}{PYF_i}),$$

where $E_{\$/i}$ is price of the ith foreign currency in terms of US\$, expressed as an index. The numerator of rpx_i is the export price of the ith country expressed in US\$; the denominator of rpx_i is the foreign GDP deflator, from the standpoint of the ith country, also expressed in US\$; again, product coverage for rpx_i coincides with the measure of trade being studied. However, rpx_i does not fully reflect changes in openness because the weights used to construct the foreign GDP deflator are fixed by assumption.

Indeed, data for foreign aggregates for the ith country are constructed as geometric averages across all countries j (not including i) using j's share in i's exports as weights. For example, data for the foreign GDP deflator from the standpoint of country i are calculated as

$$PYF_i = \prod_{j \neq i} (\frac{PY_j}{E_{\$/j}})^{\omega_{ij}}, \sum_{j \neq i} \omega_{ij} = 1,$$

where ω_{ij} is the share of country j in i's nominal exports. These export shares come from the IMF's *Direction of Trade* Statistics for 1995 (see appendix A). Again, the shares used in weighting the components of these aggregates are fixed by assumption and thus do not fully reflect changes in the country composition of world trade. Data on real GDPs, GDP deflators, and nominal exchange rates for the non-G-7 countries were obtained from national sources. To study the sensitivity of the results, the appendices E-K report alternative specifications of the import and export equations in which we substitute for the relative price terms defined above the IMF's real effective exchange rate. Appendix B documents details on the IMF's real effective exchange rates.

The beginning point of the sample period over which the equations were estimated varied depending on the availability of data by country: the mid-1950s to early 1960s for Canada, Japan, the United Kingdom, and the United States, and around 1970 for Germany, France, and Italy; the end-point was either 1996.4 or 1997.1 in all cases.[2]

[2] We choose to use the longer sample period when available (rather than using a standard shorter sample period across all countries) in order to maximize the power of parameter-stability tests. We also considered shorter samples for the non-continental European countries to determine whether the choice of sample period affected the comparability of estimation results across countries. These estimates yielded results similar to those obtained when the full sample is used and are not discussed further below.

3.3 Results

We used the Johansen procedure to estimate the cointegrating vectors for imports and exports;[3] appendix D describes the procedure and appendices E-K show detailed results. In doing so, we examined the robustness of the results to lag lengths ranging from 2 to 9 quarters. Our criteria for choosing the lag length include the consistency of the signs of coefficients with theoretical priors and the degree of serial correlation in the residuals. The final selection of lag lengths for the cointegrating vectors varied across countries.[4]

The long-run elasticity estimates are shown in Table 1 (an '*' denotes statistical significance at the 5% level). Income elasticities for imports range from 0.9 (Japan) to 2.2 (the United Kingdom). The income elasticities for exports range from 0.8 (the United States) to 1.6 (Italy).

Table 1: Long-run Elasticities

	Income		Price	
	Exports	Imports	Exports	Imports
Canada	1.1*	1.4*	-0.9*	-0.9*
France	1.5*	1.6*	-0.2*	-0.4*
Germany	1.4*	1.5*	-0.3*	-0.06
Italy	1.6*	1.4*	-0.9	-0.4*
Japan	1.1*	0.9*	–1.0*	-0.3*
U.K.	1.1*	2.2*	-1.6*	-0.6
U.S.	0.8*	1.8*	-1.5*	-0.3*

The estimated price elasticities, particularly on the import side, are lower than those generally found in the literature. Specifically, price elasticities for continental European countries are in most cases noticeably lower than those of the other G-7 countries. Indeed, the results for France and Germany fail to meet the Marshall-Lerner condition.[5] One factor that may help to account for these lower estimates is that our measures of trade volume includes both oil and services (excluding factor income). The inclusion of oil lowers the estimated price elasticity of imports because of the delays in adjusting the capital stock. In addition, the inclusion of services

[3]As a first step, we used an Augmented Dickey-Fuller test to identify the time series properties of the data for estimation. In addition to a constant term and a trend, we included four lags for the change of the variable being examined. The results suggest (see appendix C) that the variables used in our regression analysis are all integrated of order one.

[4]Owing to the presence of lagged endogenous variables in these equations, the period of adjustment is generally longer than the lag length used in estimating these systems.

[5]Appendix L reports estimated elasticities for German and French trade with non-EU countries, along with several bilateral equations (involving German trade with France and with the United States). The price elasticity estimates obtained in these additional tests do not overturn the low-price elasticity results reported here.

may lower the price elasticity of both exports and imports to the extent that the substitutability between domestic and foreign services is lower than that for goods in general (conventional trade equations generally have focused on goods alone).

The short-run elasticities obtained by the error-correction formulations (7-8 above) are shown in table 2; details are presented in appendices E-K. Overall, there is a good deal less uniformity in the short-term income elasticities of the continental European countries than in the case of the cointegration results.

Table 2: Short-run Elasticities

	Income		Price	
	Exports	Imports	Exports	Imports
Canada	1.1*	1.3*	-0.5*	-0.1
France	1.8*	1.6*	-0.1	-0.1
Germany	0.5	1.0*	-0.1	-0.2*
Italy	2.3*	1.0*	-0.3*	-0.0
Japan	0.6	1.0*	-0.5*	-0.1
U.K.	1.1*	1.0*	-0.2*	-0.0
U.S.	1.8*	1.0*	-0.5*	-0.1

In addition, the short-term price elasticities are quite low–especially for the European countries.

4 Parameter Constancy

4.1 Chow Tests

We assessed the stability of the income and price elasticities with Chow tests. These tests analyze the stability of the overall equation by comparing the behavior of its residuals during alternative pairs of sub-samples. We implement this strategy in four steps. First, we split the sample in 1989 and use the first sub-sample to obtain initial elasticity estimates. Second, we use these initial estimates to forecast trade through 1994 using three forecast horizons that differ in length. Third, we test whether the forecast errors are, on average, statistically equal to zero for each horizon length. A rejection of this hypothesis means that income and price elasticities cannot be treated as constants for that particular choice of sample split. Finally, we 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 continues until all the observations are used. The result is a collection of tests of parameter constancy for each sample split starting in 1990.

For a given sample split, we consider three horizons all of them focusing on the 1990s. One involved a series of one-quarter-ahead comparisons. The second involved a series of h-period ahead comparisons, with the sample-period split at 1989.4 in all cases and the second subsample increasing in length from h=1 quarter to h=20 quarters. The third involved moving the sample split out one quarter at a time starting from 1989.4 and forecasting to the end of the estimation sample (1994.4); in this case, the horizon drops from h=20 quarters to h=1. The results are summarized in Table 3 below and the details are presented in appendices E-K.

The first three columns of the table give the results for the three sets of equations. The entries under these columns indicate whether the evidence of equation instability appeared to be strong (numerous failures of the Chow test) indicated by "+++", moderate (occasional failures) "++", weak (one or two failures at most) "+", or absent altogether, "0". The fourth column indicates the periods of greatest instability.

Inspection of the results reveals three findings. First, instability in the export equations is more frequent that in the import equation. One reason for this result is that the data for the relative price of imports allow for changes in a country's openness whereas the export data do not. Second, the only countries without any sign of instability in trade equations are Canada (before Nafta) and Japan. Finally, the system of cointegration equations tended to show slightly greater evidence of instability than either the single cointegration equation or the error-correction equation.

Table 3: Frequency of Violations of Equation-stability Hypothesis

Exports

	Cointegration		ECM	Dates of Instability
	Exports only	System		
Canada	0	0	0	-
France	+++	+++	+	1990-93
Germany	+++	+++	+	1990-94
Italy	0	++	+	1990-94
Japan	0	0	0	-
U.K.	0	+	+	1991
U.S.	+	0	+	1991, 1993

Imports

	Cointegration		ECM	Dates of Instability
	Imports only	System		
Canada	0	0	0	-
France	0	0	0	-
Germany	+++	+++	+	1990-94
Italy	0	+	0	1992
Japan	0	0	0	-
U.K.	0	0	0	-
U.S.	0	++	0	1991

4.2 Kalman Filter Tests

To assess the stability of the income and price elasticity estimates more specifically, we reestimated a simplified version of the import and export equations using the Kalman filter (KF) technique. This procedure allows the coefficients on income and relative prices to vary over the sample period by modeling them as random walks.[6] The KF technique was applied to the import equations beginning in 1980.1 and the export equations beginning in 1985.1.[7]

Because of complexities involved in applying the KF technique to either the system of cointegrating equations or the error-correction equations with their multiple lag coefficients, we applied it instead to a simple static version of the trade equations (i.e. with no lags). We start by comparing the full-sample estimates from the KF method using the static equations for each country to those obtained with the long-run coefficients estimated in the cointegration equations. On the whole, the static elasticities are quite similar in magnitude to the cointegration elasticities in most cases with the price elasticities of U.K. trade being the chief exception (table 4).

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[6] More specifically, our estimation using the Kalman filter entails estimating, for example, the following equation for exports:

$x_t = \alpha_t + \beta_t \cdot yf_t + \gamma_t \cdot rpx_t + u_t$

$\alpha_t = \alpha_{t-1}+$ updating rule depending on observations through t.

$\beta_t = \beta_{t-1}+$ updating rule depending on observations through t.

$\gamma_t = \gamma_{t-1}+$ updating rule depending on observations through t.

$u_t \tilde{} N(0, \sigma^2)$.

After obtaining initial estimates for the equation's coefficients using an initial sample of N observations, the Kalman filter augments the sample by one observation and computes the estimates for the N+1th period using an updating rule that incorporates all the observations through the N+1th period. The estimated elasticities the N+1th period become the initial conditions for the estimated elasticities for the N+2th period which the Kalman filter method obtains by increasing the sample by one observation to N+2 and applying the updating rule. This process continues until all the observations have been used for estimation.

[7] The latter starting point was used for the export equations because of the shorter sample period associated with foreign GDP.

Table 4. Sensitive of Long-run Elasticities to Estimation Method

Income Elasticities

	Exports		Imports	
	Cointegration	Static	Cointegration	Static
Canada	1.1	1.6	1.4	1.5
France	1.5	1.6	1.6	1.7
Germany	1.4	1.5	1.5	1.7
Italy	1.6	1.8	1.4	1.5
Japan	1.1	1.2	0.9	1.3
U.K.	1.1	1.3	2.2	1.7
U.S.	0.8	0.9	1.8	1.9

Price Elasticities

	Exports		Imports	
	Cointegration	Static	Cointegration	Static
Canada	-0.9	-0.5	-0.9	-0.9
France	-0.2	-0.1	-0.4	-0.3
Germany	-0.3	-0.2	-0.06	+0.1
Italy	-0.9	-0.3	-0.4	-0.3
Japan	-1.0	-0.8	-0.3	-0.1
U.K.	-1.6	-0.1	-0.6	-0.1
U.S.	-1.5	-1.4	-0.3	-0.4

We also examine the paths of the estimated income and price elasticities of imports and exports for each of the seven countries (Charts 4 and 5). The elasticities in the import equations, particularly the income elasticities, tend to be more stable over time than those in the export equations. Across countries, Germany's elasticities and, surprisingly, the U.S. export elasticities show the greatest instability, while Japan's show the least. With respect to possible trends in elasticities suggested by these results, in some cases, most notably Canada, the United States, and Japan, income elasticities of imports have been relatively stable in recent years while those of exports have declined. Japan's export elasticity has fallen from a level slightly above its import elasticity in the mid-1980s to be slightly below in the 1990s.[8] In Germany, Italy, and the United Kingdom, the opposite has been true, with the income elasticities of exports rising relative to those of imports.

[8]This finding is consistent with the earlier observation that our results showed a smaller elasticity differential in favor of exports than had been the case in earlier studies.

Chart 4

Kalman-Filter Income Elasticities

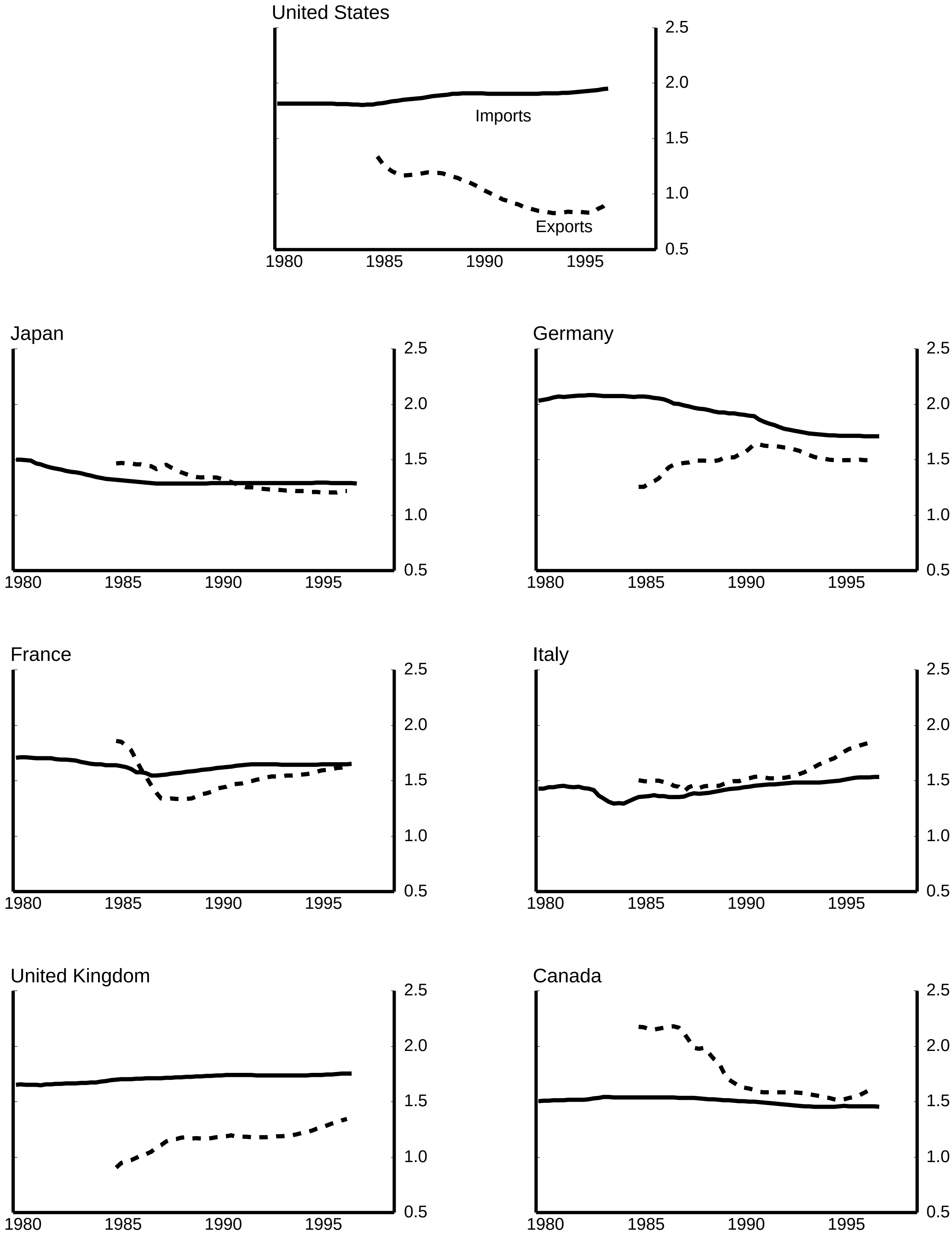
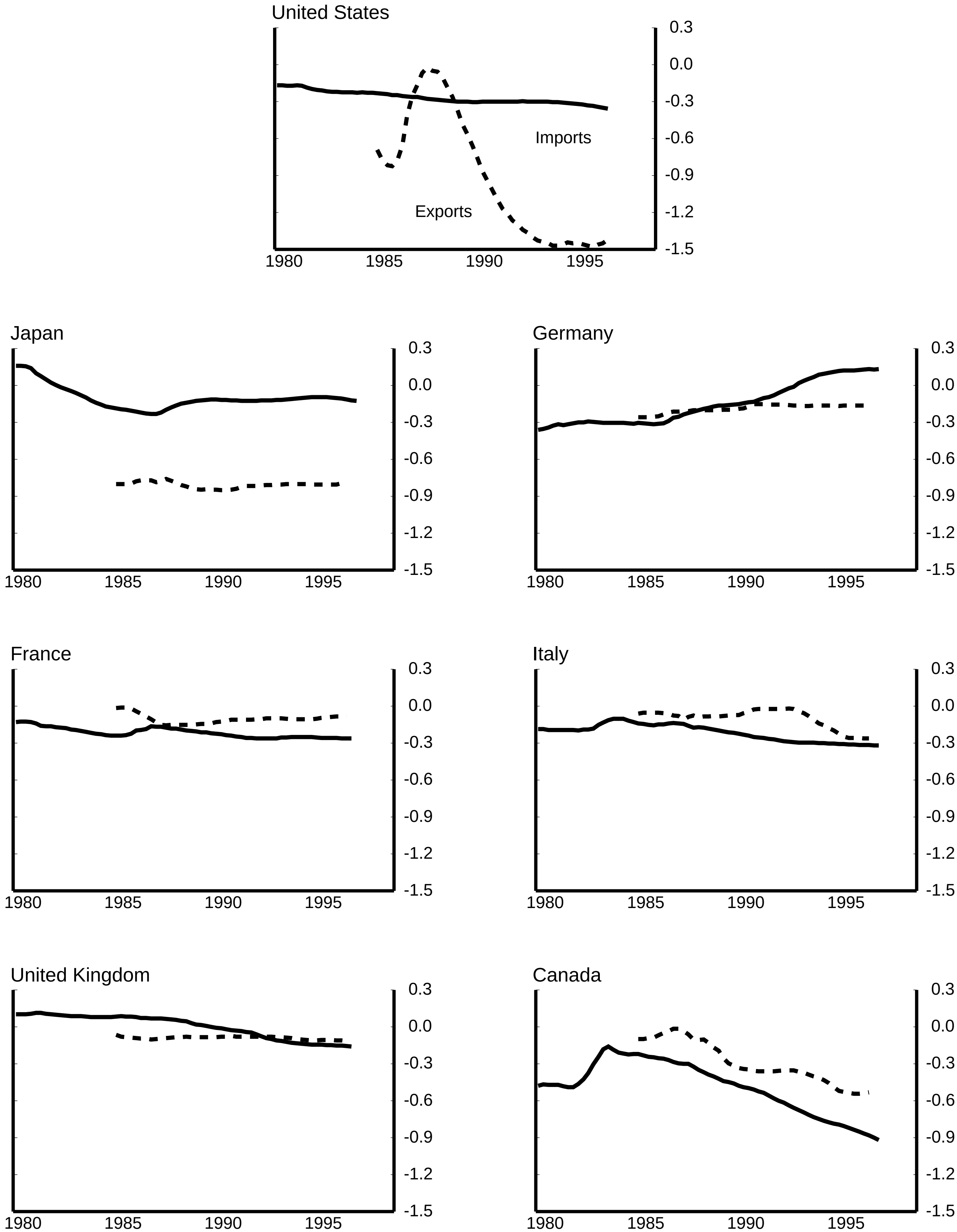


Chart 5

Kalman-Filter Price Elasticities



In the U.S. case, the movement in income elasticities (the import elasticity moving up slightly while the export elasticity has declined) has resulted in a widening of the gap between the estimated import and export elasticities. At the same time, however, the estimated price elasticity of U.S. exports has shown substantial instability (fluctuating between 0 and -1.4 over the past ten years), casting doubt on the reliability of the export equation altogether. More broadly, price elasticities tend more often than not to be increasing over time in (absolute) magnitude.

4.3 Out-of-sample Prediction.

We also judged the reliability of the equations outside the estimation period. To that end, we used the error correction models to generate post-sample predictions over 1995-96; we summarize here the results which are presented in detail in appendices E-K. Given the 1995-96 predictions, we test the hypothesis of parameter constancy with the Chow test; see Hendry and Doornik (1996, page 222). Table 5 reports the probability for the validity null hypothesis of parameter constancy; entries below 0.05 indicate a rejection of this hypothesis at the five percent significance level. Based on the results, short-term predictions of exports and imports show no evidence of parameter instability during 1995-96.

Table 5: Out-of-Sample Chow Tests

	Exports	Imports
Canada	0.93	0.96
France	0.09	0.99
Germany	0.46	0.39
Italy	0.26	0.35
Japan	0.35	0.32
U.K.	0.32	0.95
U.S.	0.27	0.84

But to find support for the hypothesis of parameter constancy does not imply that the post-sample predictions are accurate. We thus implement a predictive-failure test; see Hendry and Doornik (1996, page 222). Table 6 reports the probability for the validity null hypothesis that the mean-squared forecast error is equal to the variance of the in-sample residuals; entries below 0.05 indicate a rejection of this hypothesis at the five percent significance level. We find no evidence of such failures.

Table 6: Predictive Failure Tests

	Exports	Imports
Canada	0.93	0.96
France	0.07	0.99
Germany	0.37	0.36
Italy	0.21	0.18
Japan	0.27	0.22
U.K.	0.22	0.96
U.S.	0.81	0.76

4.4 Summary of Results by Country

The above analysis consisted of three estimation techniques (performed using the same data set each time) and alternative stability tests for exports and imports for each of the G-7 countries. Table 7 provides a summary by country of our findings. For each country, the table reports (for imports on the left and exports on the right) the static income elasticity and comments on its stability based on the Kalman filter technique (first line), compared with the long-run (cointegration) income elasticity (second line). The third line describes the stability of the static price elasticity and the final line summarizes the Chow test results.

For the United States, our results for imports appear to be more robust than those for exports. We obtain an estimated income elasticity for imports that is nearly 2–a result commonly found in the literature. The Kalman filter results for the export price elasticity suggest that this technique is not succeeding in finding a useful estimate for that elasticity, perhaps because of the absence of lags in the estimating equation. The period of the very strong dollar in the mid-1980s is inducing swings in the estimated elasticity that may be contributing to the apparent movement in the estimated income elasticity, as well. For Japan, we obtained quite stable and generally consistent, although somewhat surprising, estimates of the income and price elasticities. Our results suggest that the income elasticities are about equal (close to 1) and thus that there is no elasticity gap tending toward surplus–contrary to earlier findings in the literature.

For Germany, our results point to parameter instability at the time of reunification; we obtain similar income elasticities (near 1.5) for exports and imports, but very low price elasticities and evidence of instability. For France and Italy, again the export equations appear to be somewhat less well estimated than the import equations. For both, the estimated elasticities for exports and imports are consistent across the techniques used and appear to be about equal (in the range of 1.5 -1.7).

For the United Kingdom, the estimated income elasticity for imports is quite high (close to that for the United States) and results for that equation suggest no instability. For exports, the Kalman filter technique suggests some upward movement in the income elasticity, but it is always somewhat lower than that for imports. Both price elasticities appear to be very low in the Kalman estimates, a result not consistent with the long-run cointegration results, especially for exports.

For Canada also, the results for exports appear less robust than those for imports. The income elasticity for imports is consistently estimated at about 1.5 and that equation is generally stable, although the price elasticity may be rising over time. The income elasticity for exports is not consistent across our techniques and appears to be shifting in the Kalman estimates. As a consequence, the long-run cointegration results suggest stable equations with the income elasticity for imports exceeding that for exports, while the Kalman technique reverses the elasticity gap and suggests instability in the export elasticity.

Table 7: Summary of Findings

Imports	Exports
United States	
Kalman η_y : Stable at 1.9	Kalman η_{yf}: downward trend to 0.9
LR η_y: same at 1.8	LR η_{yf} : = 0.8
Kalman η_{rpm} stable but low	Kalman η_{rpx}: unstable and unreliable
Chow: Generally stable	Chow: Generally stable
Japan	
Kalman η_y : Stable at 1.3	Kalman η_{yf}: trends from 1.5 to 1.2
LR η_y: same at 0.9	LR η_{yf} : = 1.1
Kalman η_{rpm} stable but low	Kalman η_{rpx}: stable at -0.8
Chow: stable	Chow: stable
Germany	
Kalman η_y : Trends from 2.1 to 1.7	Kalman η_{yf}: fluctuates around 1.5
LR η_y: same at 1.5	LR η_{yf} : very close at 1.4
Kalman η_{rpm} small increases, close to 0	Kalman η_{rpx}: stable but low
Chow: unstable	Chow: unstable
France	
Kalman η_y : stable at 1.7	Kalman η_{yf}: fluctuates around 1.6
LR η_y: very close at 1.6	LR η_{yf} : very close at 1.5
Kalman η_{rpm} : stable at -0.3	Kalman η_{rpx}: stable but low
Chow: stable	Chow: some instability
Italy	
Kalman η_y : stable at 1.5	Kalman η_{yf}: rises to 1.8
LR η_y: close at 1.4	LR η_{yf} : close at 1.6
Kalman η_{rpm} : stable	Kalman η_{rpx}: some movement but low
Chow: Stable	Chow: some instability
United Kingdom	
Kalman η_y : stable at 1.7	Kalman η_{yf}: rises to 1.3
LR η_y: stable at 2.2	LR η_{yf} : close at 1.1
Kalman η_{rpm} : stable around zero	Kalman η_{rpx}: stable but very low
Chow: stable	Chow: generally stable
Canada	
Kalman η_y : stable at 1.5	Kalman η_{yf}: declines to 1.6
LR η_y: stable at 1.4	LR η_{yf} : = 1.1
Kalman η_{rpm} : rise over time	Kalman η_{rpx}: rise over time
Chow: Generally stable	Chow: stable

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A Weights for Foreign Aggregates

The estimation of the elasticities for a country' exports needs data for foreign GDP and foreign GDP deflator. We constructed these foreign aggregates using a geometric mean with bilateral export shares for 1995 as weights; these shares are shown below.

Bilateral Export Shares

Importing\Exporting	Canada	France	Germany	Italy	Japan	U.K.	U.S.
Canada		0.009	0.008	0.012	0.023	0.020	0.201
France	0.007		0.116	0.118	0.014	0.095	0.023
Germany	0.012	0.167		0.160	0.045	0.123	0.036
Italy	0.007	0.096	0.075		0.009	0.049	0.014
Japan	0.045	0.024	0.027	0.028		0.031	0.103
U.K.	0.014	0.083	0.080	0.05	0.032		0.046
U.S.	0.778	0.063	0.074	0.075	0.285	0.120	
Other OECD	0.053	0.325	0.367	0.282	0.066	0.337	0.111
Mexico	0.004	0.003	0.005	0.003	0.008	0.002	0.072
NIEs	0.028	0.034	0.039	0.039	0.249	0.049	0.118
OPEC	0.012	0.034	0.023	0.033	0.041	0.039	0.031
ROW	0.040	0.163	0.187	0.200	0.229	0.134	0.245
Total	1.00	1.00	1.00	1.00	1.00	1.00	1.00

Source: *Direction of Trade*, International Monetary Fund.

B IMF's Real Exchange Rate

To study the sensitivity of the results, we consider alternative specifications of the import and export equations in which we substitute for the relative price terms defined above the IMF's real effective exchange rates, $reer$:[9]

$$reer_i = \log(\frac{NULC_i \cdot E_{\$/i}}{NULCF_i}),$$

$$NULCF_i = \prod_{j \neq i}(\frac{NULC_j}{E_{\$/j}})^{\varpi_{ij}}, \sum_{j \neq i} \varpi_{ij} = 1,$$

The numerator of $reer_i$ is the normalized unit labor cost of the ith country expressed in US\$; the denominator is the normalized unit labor cost for the rest of the

[9]Taken from *International Finance Statistics*.

ith's world, expressed in US$. The weights, ϖ_{ij}, are based on aggregate trade flows for *manufactured* goods. An increase in $reer_i$ represents a real appreciation of the domestic currency which stimulates, in theory, purchases of foreign products. We should, therefore, expect a negative price elasticity for exports and a positive price elasticity for imports.

This measure uses weights based on aggregate trade flows for manufactured goods only and thus the product coverage of this measure is narrower than the one corresponding to the measure of trade being studied–namely, goods and services. Moreover, given that observations on trade volumes are generated by deflating the nominal value of trade by the corresponding price index, reliance on the IMF's real effective exchange rate does not eliminate the need to use the conventional trade deflators. Finally, published data for the IMF's real effective exchange rates filter out short-run fluctuations with the Hodrick-Prescott filter. Thus the data are not ideally suited for estimating separately short and long run trade elasticities.

C Order of Integration

To determine the time-series properties of the variables, we used an Augmented Dickey-Fuller test. In addition to including a constant and a trend, we allowed for four lags for the change of the variable. The test results are below their corresponding significance levels which means that one cannot reject the hypothesis that these variables follow a random walk (integrated of order one).

Augmented Dickey-Fuller Test Results

	x	rpx	fy	m	rpm	y	*reer*
Canada	-1.92	-3.33	-2.48	-2.92	-2.12	-1.45	-2.92
France	-1.93	-2.13	-2.04	-3.15	-1.58	-2.30	-3.41
Germany	-1.93	-1.56	-1.69	-3.15	-2.59	-2.92	-2.23
Italy	-1.72	-2.10	-2.64	-2.74	-2.43	-1.42	-2.02
Japan	-2.51	-2.73	-2.43	-1.96	-2.52	-1.10	-2.51
United Kingdom	-1.43	-2.26	-2.22	-3.17	-2.39	-2.12	-2.23
United States	-1.74	-1.88	-2.47	-2.82	-1.13	-2.96	-1.88

D Experimental Design

Formulation To estimate the long-run elasticities that summarize the underlying trends in foreign trade, we apply the cointegration technique of Johansen (1988). This technique involves applying maximum likelihood to estimate the parameters of a system of equations:

$$\Delta z_t = \kappa + \sum_{i=1}^{n} \Gamma_i \Delta z_{t-i} + \alpha\beta' z_{t-1} + \epsilon_t, \epsilon_t \sim NI(0, \Omega), \tag{9}$$

where $z_t' = (\tau_t \; a_t \; p_t)$, κ is a 3x1 vector of intercepts, Γ_i is a 3x3 matrix of coefficients for short-run interrelations; τ_t is the measure of trade (exports or imports); a_t is the measure of income (foreign or domestic); p_t is the measure of relative prices (for exports or imports); and

$$\alpha\beta' = \begin{pmatrix} \alpha_{11} & ... & \alpha_{13} \\ ... & ... & ... \\ \alpha_{31} & ... & \alpha_{33} \end{pmatrix} \begin{pmatrix} \beta_{11} & ... & \beta_{31} \\ ... & ... & ... \\ \beta_{13} & ... & \beta_{33} \end{pmatrix}.$$

The elements of α measure the speed of adjustment and are known as loading coefficients; the vector $(\beta_{1i} \; \beta_{2i} \; \beta_{3i}) = \beta_i'$ characterizes the *ith* long-run relation among τ_t, a_t, and p_t.

Estimates derived from the Johansen procedure are known to be sensitive to the number of lags included in the system. In addition, estimates of income and price elasticities are also known to be sensitive to the choice of relative price. Thus we estimated the export and import systems using alternative lag distributions and measures of prices. To select the lag length, we implement the following selection criteria. First, consistency with economic theory, as reflected in the signs of the coefficients and the existence of one cointegrating vector. Though multiple cointegration vectors are not inconsistent with economic theory, their presence raises identification issues that will be addressed in future research. Second, consistency with statistical theory, as reflected in the absence of serial correlation of the residuals of the system. To determine the absence of serial correlation, we test whether the coefficients of a VAR(5) on each system's residuals are jointly zero; entries with an * denote a rejection of the hypothesis of serial independence. Entries with a † indicate our selection for the number of lags.

Based on the long-run elasticity estimates, we explain short run fluctuations in trade using an error-correction formulation:

$$\Delta\tau_t = \sum \alpha_i \Delta\tau_{t-i} + \sum \theta_i \Delta a_{t-i} + \sum \rho_i \Delta p_{t-i} + \Omega ECM_{\tau,t-1}.$$

For presentational purposes, we report $\frac{\sum \theta_i}{(1-\sum \alpha_i)}$, $\frac{\sum \rho_i}{(1-\sum \alpha_i)}$, and $\frac{\Omega}{(1-\sum \alpha_i)}$ and show charts with the 95% confidence bands for all of the model's coefficients using recursive estimation. We also examine the extent to which the residuals of the estimating equations are consistent with the estimation assumptions. The tests are

Serial Correlation: F-test of the null hypothesis that the coefficients of an AR(5) for the residuals are jointly equal to zero.

Homoskedasticity: t-test of the null hypothesis that the variance of the residuals is constant.

Normality: χ^2 Test of the null hypothesis that the distribution of the residuals is normal.

Functional Form: Reset test of the null hypothesis that the specification does not omit combinations of the predetermined variables.

Parameter Stability We used three variants of the Chow test all of which involve testing whether forecast errors are statistically significant. We report test-results for the individual trade equation and for the cointegrating system as a whole; we also report the same tests for the associated error-correction model from each trade flow. The forecast period is 1990-1994. The three tests are

One-step ahead F-tests (1up): Sequence of F-tests for the hypothesis that the vector of 1-step ahead forecast errors is zero.

Forecasts F-tests (Nup): Sequence of F-tests for the hypothesis that the vector of N-step ahead forecast errors is zero; the forecast horizon N increases from M, the last date of the initial sample, to the current observation t as the estimation sample increases.

Break-point F-tests (Ndn): Sequence of F-tests for the hypothesis that the vector of N-step ahead forecast errors is zero; the forecast horizon N decreases from T-M + 1 to 1 as the estimation sample increases where T is the total number of observations and M is the number of observations reserved as initial conditions.

For a fuller description of these tests, see Hendry and Doornik (1996).

Legend for Graphs: $\ell\mathtt{x}$ denotes the logarithm of real exports of the ℓth country; $\mathtt{d}\ell\mathtt{x}$ denotes the corresponding growth rate. $\ell\mathtt{m}$ denotes the logarithm of real imports of the ℓth country and $\mathtt{d}\ell\mathtt{m}$ represents the corresponding growth rate. The values for ℓ are c = Canada; e=U.K.; f=France; g=Germany; i=Italy; j=Japan, u=United States. As a way of examining the time-series properties of the series, we divide the full sample into six sub-samples and apply least squares to each sub-sample; the segments displayed along the line correspond to those regression lines.

E Canadian Trade

The evolution of Canada's foreign trade and its proximate determinants are displayed in Figure 1. Both export and import volumes grow over time along with foreign and domestic income. Relative prices for both trade flows show downward trends, but the

IMF's measure of the real effective exchange rate ($reer$) shows, in this case, volatility comparable to that of the conventional measure of relative prices based on deflators from the National Income Accounts.

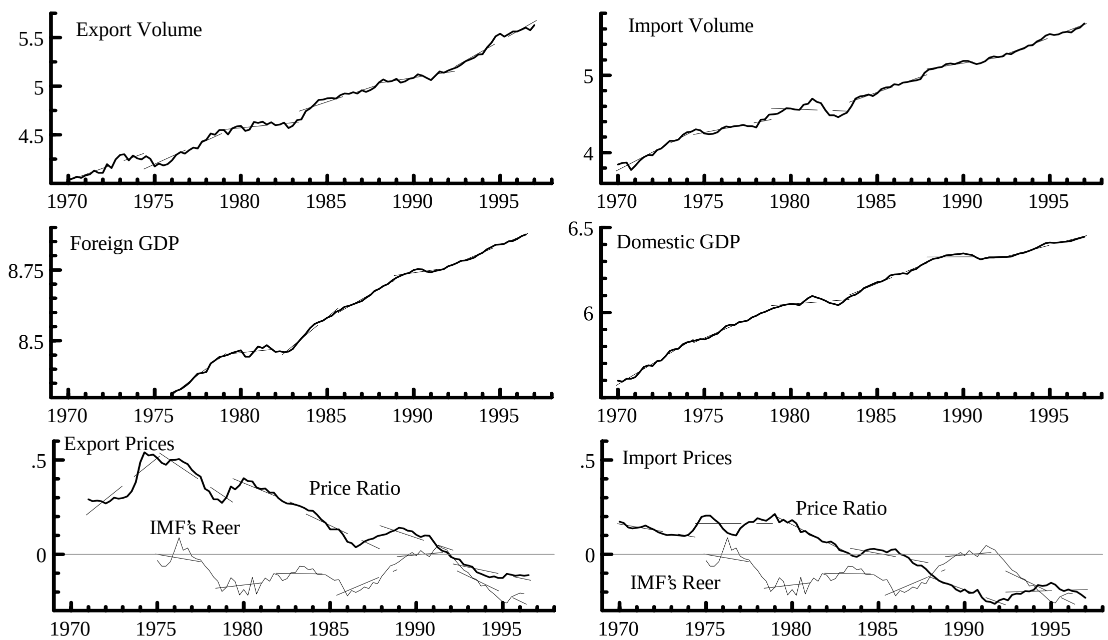


Figure 1: Trade, Income, and Prices: Canada

Canadian Exports

Long-run Forces The number of cointegration vectors is sensitive to both the measure of prices and to the number of lags included in the system. Measuring prices as the export deflator relative to the foreign GDP deflator (rpx) yields no cointegration for systems including fewer than nine lags. Including nine quarters yields, however, one cointegration vector which suggests a price elasticity of -1.1 and an income elasticity of 0.8. Measuring prices as the IMF's $reer$ reveals no cointegration for specifications including fewer than seven lags. Otherwise, the results reveal robust estimates of income and price elasticities in the order of 2.1 and -0.7, respectively. Unfortunately, the vector of residuals for these formulations exhibit serial correlation.

Export Cointegration Results with rpx-Canada

	Number of lags Included							
	9†	8	7	6	5	4	3	2
Cointegration Vectors	1	0	0	0	0	0	0	0
Price-Elasticity	-1.08	ni	ni	ni	ni	ni	ni	ni
Income Elasticity	0.82	ni	ni	ni	ni	ni	ni	ni
Loading Coefficient	-0.27	ni	ni	ni	ni	ni	ni	ni
System's R^2	0.90	0.89	0.88	0.87	0.87	0.86	0.83	0.82
Exports' Serial Corr.	0.16	0.00	0.15	0.11	0.18	0.49	0.51	0.78
System's Serial Corr.	0.07	0.01*	0.01*	0.00*	0.08	0.11	0.02*	0.02*

Export Cointegration Results with $reer$-Canada

	Number of lags Included							
	9	8	7	6	5	4	3	2
Cointegration Vectors	1	1	1	0	0	0	0	0
Price-Elasticity	-0.71	-0.65	-0.68	ni	ni	ni	ni	ni
Income Elasticity	2.07	2.07	2.06	ni	ni	ni	ni	ni
Loading Coefficient	-0.02	0.20	0.23	ni	ni	ni	ni	ni
System's R^2	0.95	0.94	0.94	0.94	0.94	0.94	0.93	0.93
Export Serial Corr.	0.08	0.01*	0.45	0.06	0.85	0.32	0.45	0.66
System's Serial Corr.	0.03*	0.00*	0.02*	0.00*	0.01*	0.01*	0.15*	0.01*

The cointegration results using rpx are (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.271 & (0.21) \\ -0.209 & (0.05) \\ -0.173 & (0.13) \end{pmatrix} \begin{pmatrix} 1 & -0.823 & 1.085 \\ (na) & (0.108) & (0.104) \end{pmatrix} \begin{pmatrix} x \\ fy \\ rpx \end{pmatrix}, 1978.2\text{-}1994.4$$

which shows significant long-run elasticities for income and prices but an insignificant error-correction term. The 95% confidence bands for the out-sample predictions are shown in figure 2. Recursive Chow tests are shown in figure 3 and, though the parameter-constancy tests suggest stable elasticities through 1993, the results point to a significant change in those elasticities starting in 1994, as evidenced by the crossing of the horizontal dashed line denoting a rejection of the hypothesis of parameter constancy at the five-percent level..

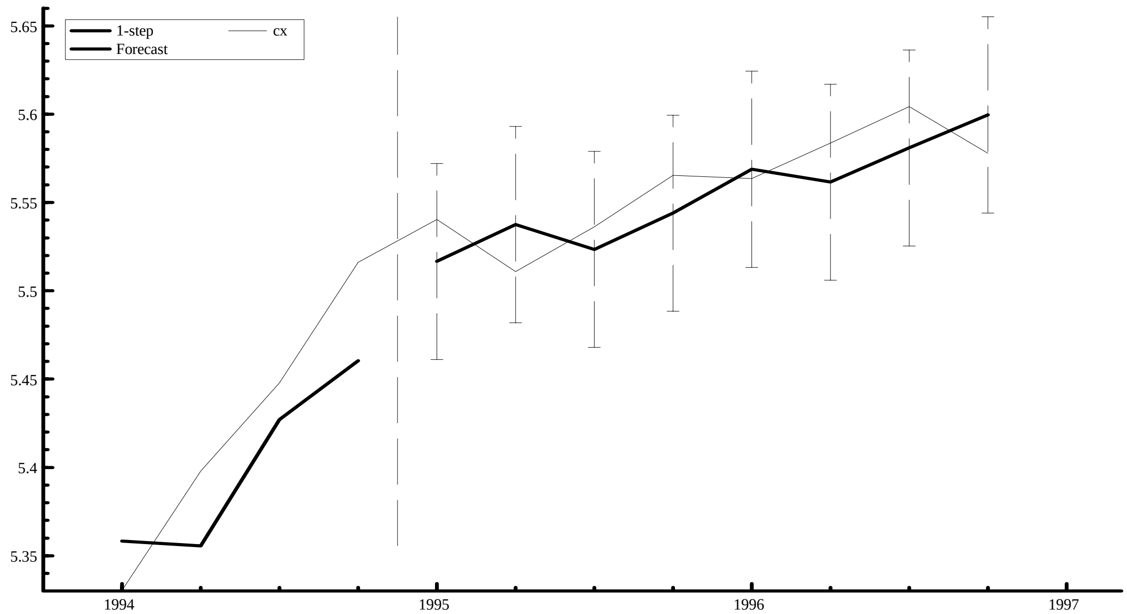


Figure 2: Predictive-Accuracy-Cointegration-Exports-Canada

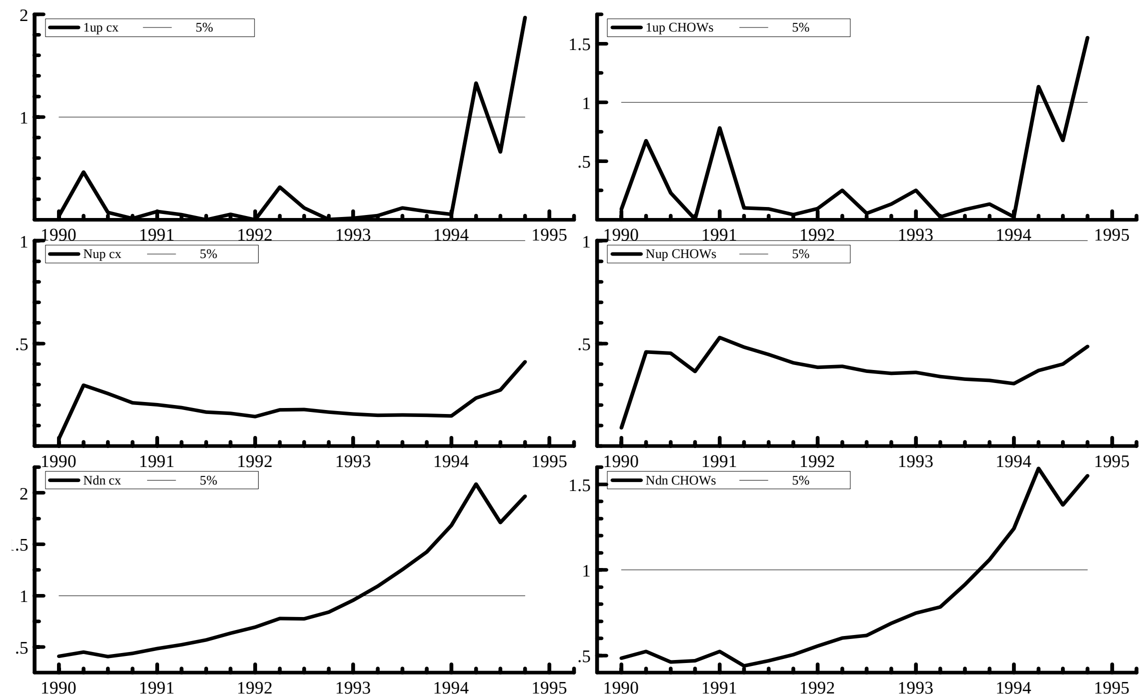


Figure 3: Chow Tests-Cointegration-Exports-Canada

One factor that might account for this breakdown is the introduction of NAFTA early in 1994. To examine this possibility, we re-estimate the system with data through 1993 and test for the stability of the long-run elasticities. The re-estimation results, shown below, reveal that the exclusion of the post-NAFTA period from estimation changes significantly the estimated speed of adjustment, both in statistical significance and economic interpretation. The re-estimated income elasticity is a bit higher and the opposite holds for the re-estimated price elasticity:

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.799 & (0.24) \\ -0.294 & (0.07) \\ -0.114 & (0.18) \end{pmatrix} \begin{pmatrix} 1 & -1.054 & 0.851 \\ (na) & (0.067) & (0.066) \end{pmatrix} \begin{pmatrix} x \\ fy \\ rpx \end{pmatrix}, \text{1978.2-1993.4}$$

The post-1993 predictions show that actual exports are significantly higher than what exports would have been in the absence of NAFTA (figure 4).

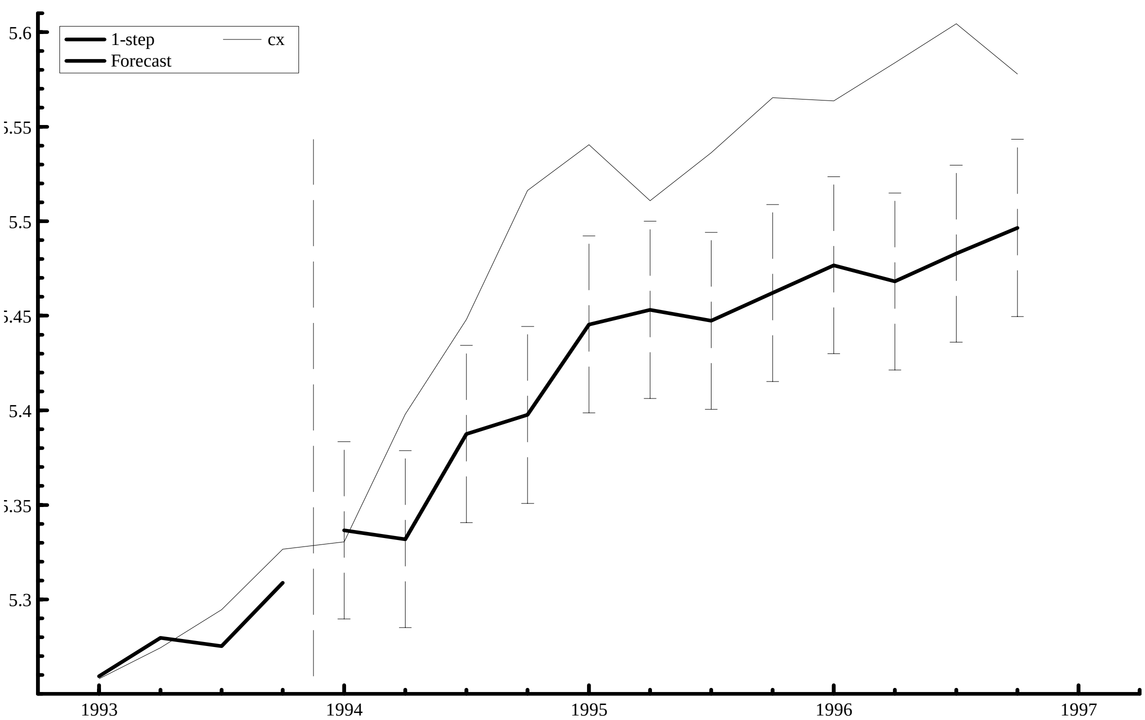


Figure 4: Predictive Accuracy - Cointegration pre-1994 - Exports-Canada

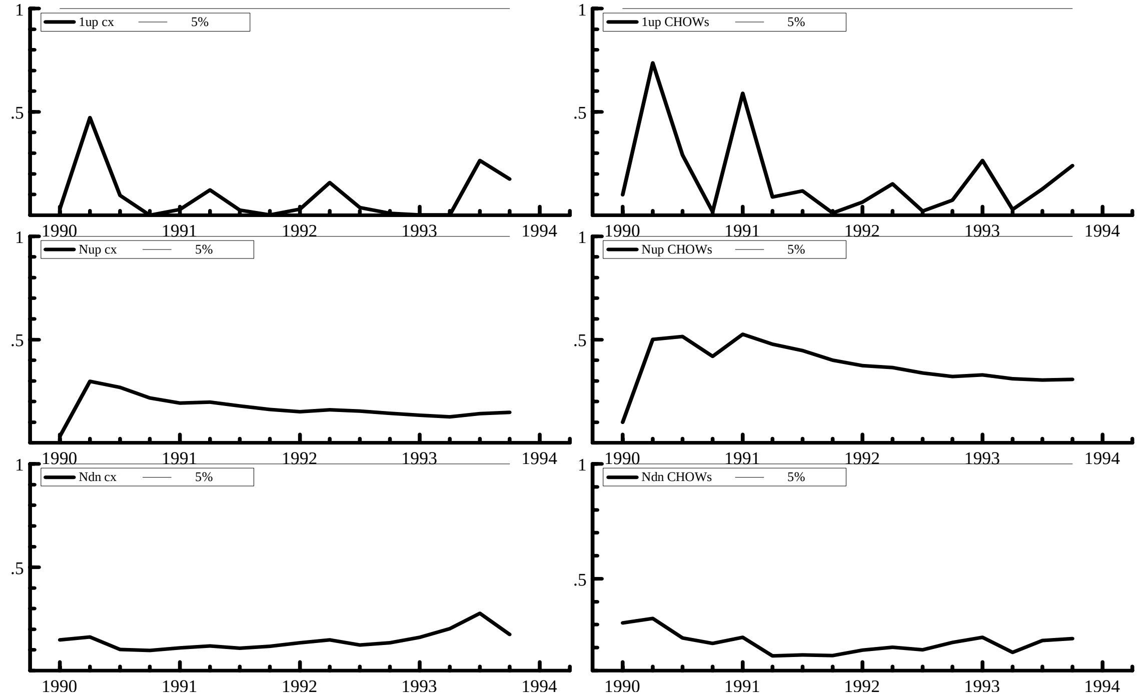


Figure 5: Chow Tests - Cointegration pre Nafta - Exports - Canada

Chow tests based on the pre-NAFTA estimates show no evidence of parameter instability (figure 5). This result confirms the suspicion that the earlier failures of this test stem from the introduction of NAFTA.

Short-run Forces To explain short-run fluctuations, we estimate the parameters of an error-correction model including a NAFTA dummy variable that takes a value of one starting in 1994.1. The estimation results are

$$\begin{array}{ccccc} \Delta x_t = & +1.0665\Delta fy_t & -0.47532\Delta rpx_t & +0.0405NAFTA_t & -0.2018ECM_{x,t-1} \\ (se) & (0.34) & (0.15) & (0.013) & (-0.075) \end{array}$$

where $ECM_x = x - 1.054 \cdot fy + 0.851 rpx + intercept$ and se stands for standard error.

$R^2 = 0.51; SER = 2.51\%$	Null Hypothesis (p-value)	
Sample: 1976.2-1994.4	Serial-Independence (0.25)	Normality (0.53)
	Homoskedasticity (0.07)	Func. Form (0.90)

The model explains about half of the variability of the growth rate of exports and the empirical distribution of the residuals satisfies the assumptions maintained for estimation. The predictions of the error-correction model do not show a tendency to be one-sided though the variable being predicted is the growth rate (figure 6).

Finally, the Chow tests (figure 7) and coefficient estimates (figure 8) from the error correction model support parameter constancy.

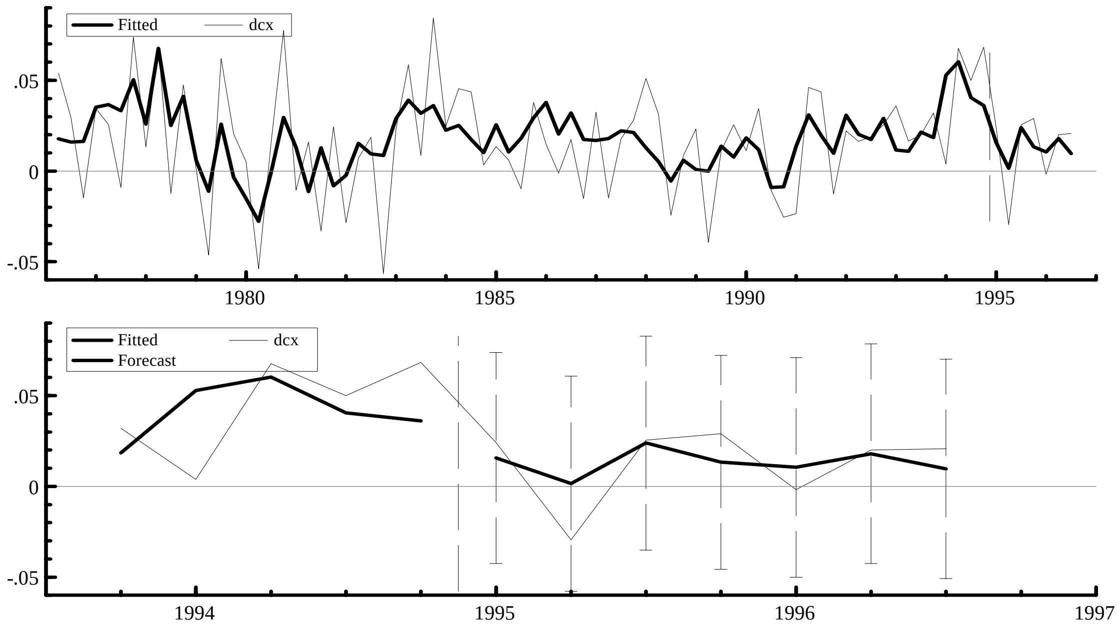


Figure 6: Predictive Accuracy-ECM-Exports-Canada

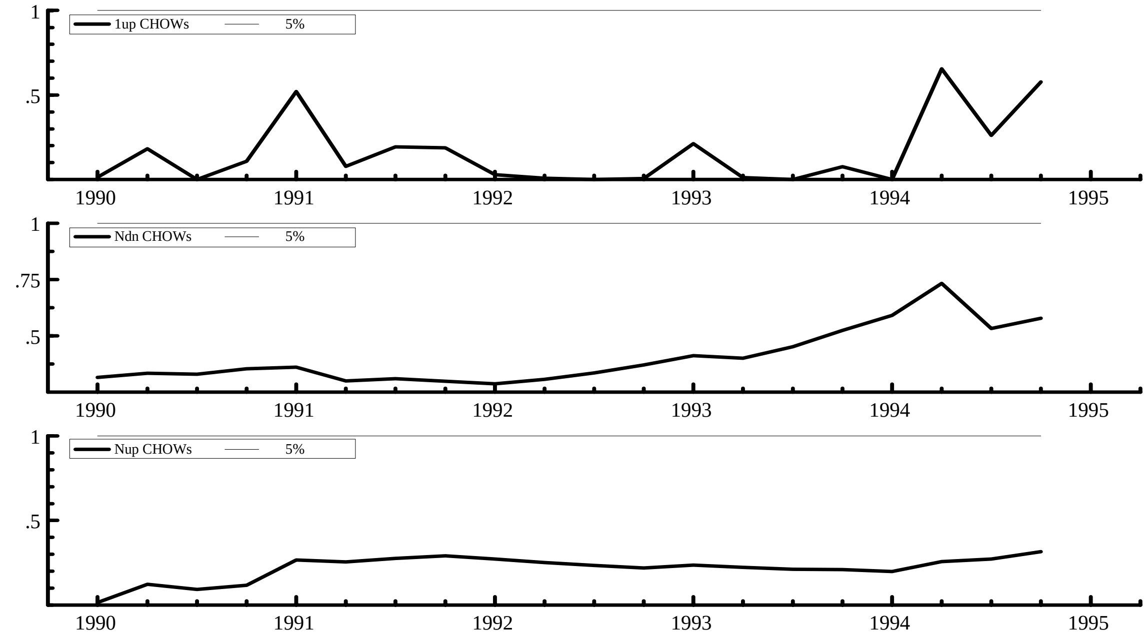


Figure 7: Chow Tests - ECM- Exports-Canada

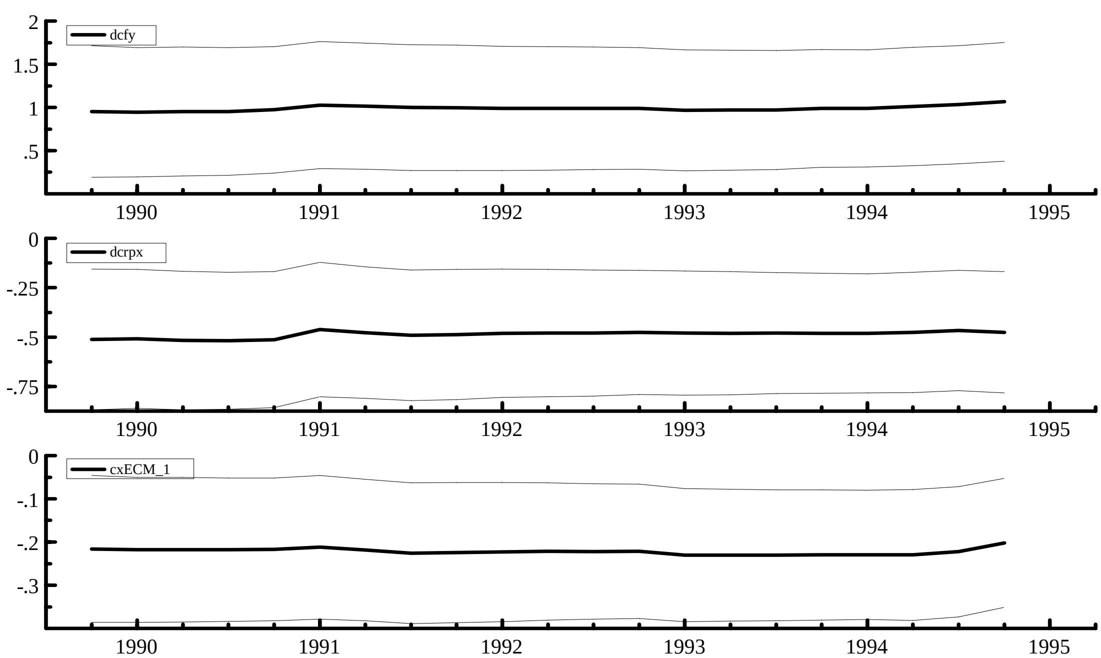


Figure 8: 95% Bands for ECM Coefficients - Exports-Canada

Canadian Imports

Long-run Forces Measuring prices as the import price relative to the GDP deflator (*rpm*) yields estimated elasticities for Canadian imports that are not robust to the choice of lag. The absence of cointegration is the exception rather than the rule and the cointegration coefficients vary considerably to changes in the number of lags except for systems with 8 and 6 lags. Measuring prices as the IMF's *reer* reveals no cointegration as the rule; for the two cases with cointegration the results point to either unreasonably large or invalidly signed estimates.

Import Cointegration Results with *rpm*-Canada

	Number of lags Included							
	9	8†	7	6	5	4	3	2
Cointegration Vectors	0	1	0	1	1	1	1	1
Price-Elasticity	ni	-1.01	ni	-0.83	3.35	-2.83	-17.33	0.33
Income Elasticity	ni	1.36	ni	1.24	0.74	1.60	4.02	1.17
Loading Coefficient	ni	-0.20	ni	-0.06	-0.00	-0.01	0.00	-0.01
System's R^2	0.95	0.94	0.94	0.94	0.94	0.94	0.93	0.93
Import's Serial Corr.	0.25	0.04*	0.16	0.15	0.36	0.11	0.08	0.55
System's Serial Corr.	0.60	0.14	0.30	0.53	0.58	0.51	0.25	0.13

Import Cointegration Results with *reer*-Canada

	Number of lags Included							
	9	8	7	6	5	4	3	2
Cointegration Vectors	1	2	1	0	0	0	0	0
Price-Elasticity	55.70	ni	-0.92	ni	ni	ni	ni	ni
Income Elasticity	3.61	ni	2.45	ni	ni	ni	ni	ni
Loading Coefficient	-0.00	ni	-0.08	ni	ni	ni	ni	ni
System's R^2	0.96	0.96	0.96	0.96	0.95	0.95	0.95	0.95
Import's Serial Corr.	0.79	0.94	0.28	0.01*	0.31	0.21	0.41	0.40
System's Serial Corr.	0.49	0.58	0.78	0.17	0.75	0.10	0.07	0.04

The cointegration results using *rpm* are (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.206 & (0.05) \\ -0.043 & (0.02) \\ 0.018 & (0.03) \end{pmatrix} \begin{pmatrix} 1 & -1.357 & 1.005 \\ (na) & (0.071) & (0.174) \end{pmatrix} \begin{pmatrix} m \\ y \\ rpm \end{pmatrix}, \text{ 1963.1-1994.4}$$

which shows significant long-run elasticities for income and prices and a significant error-correction term. The out-sample predictions based on these estimates are shown in figure 9 and the recursive Chow tests are shown in figure 10. Though the parameter-constancy tests suggest stable elasticities through 1993, the results point to a break-down in those elasticities starting in 1994.

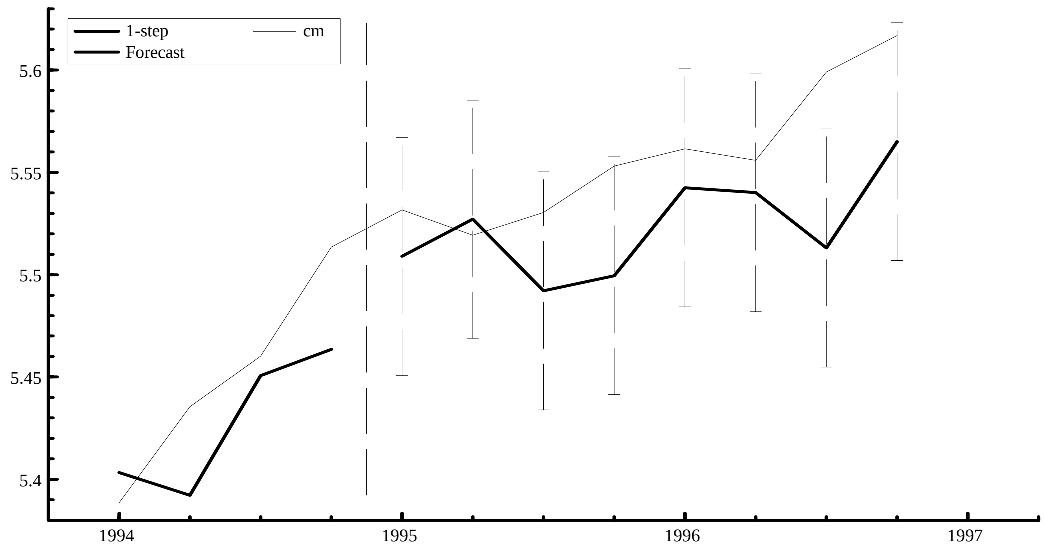


Figure 9: Predictive Accuracy - Cointegration - Imports - Canada

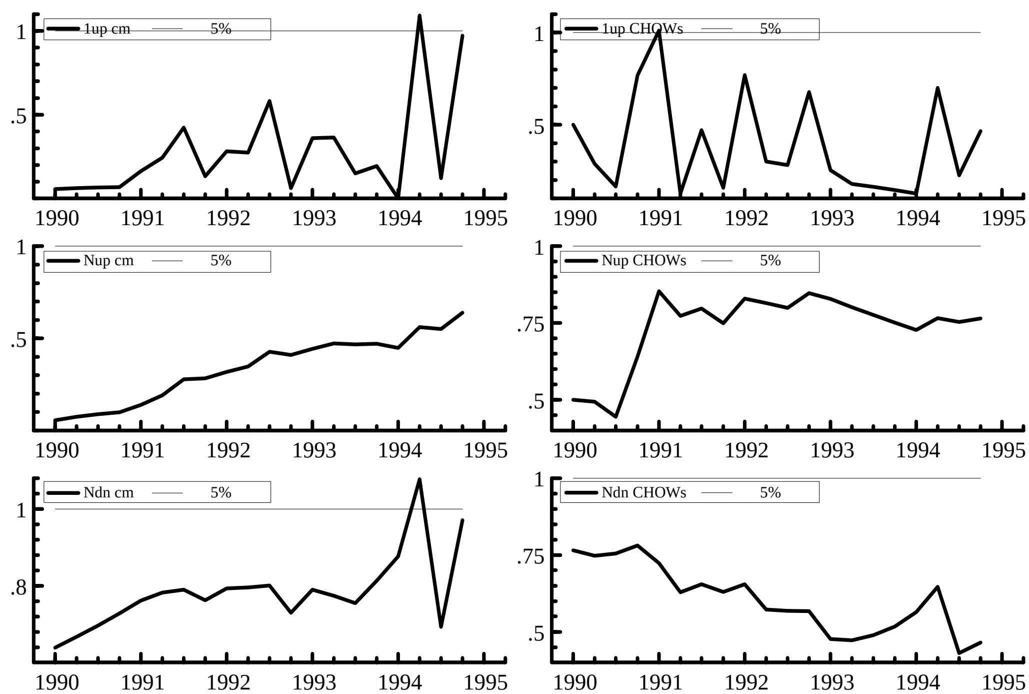


Figure 10: Chow Tests-Cointegration-Imports-Canada

As in the case of exports, we examine whether this breakdown stems from the introduction of NAFTA early in 1994. To examine this possibility, we reestimated the system with data through 1993 and tested for the stability of the long-run elasticities. The cointegration results are

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.327 & (0.07) \\ -0.068 & (0.02) \\ 0.019 & (0.04) \end{pmatrix} \begin{pmatrix} 1 & -1.392 & 0.845 \\ (na) & (0.045) & (0.114) \end{pmatrix} \begin{pmatrix} m \\ y \\ rpm \end{pmatrix}, \text{ 1963.1-1993.4.}$$

The results reveal that the exclusion of the post-NAFTA period from estimation raises significantly the estimated speed of adjustment. The re-estimated income elasticity remains unchanged but the price elasticity increases (in absolute terms) a bit. Figure 11 shows that the exclusion of the NAFTA period produces a systematic underprediction of imports during the NAFTA period and this observation confirms the belief that the trade agreement might have changed income and price elasticities for Canadian trade. Figure 12 shows indeed that prior to this trade agreement, these elasticities could be treated as constants, as reflected in insignificant Chow tests.

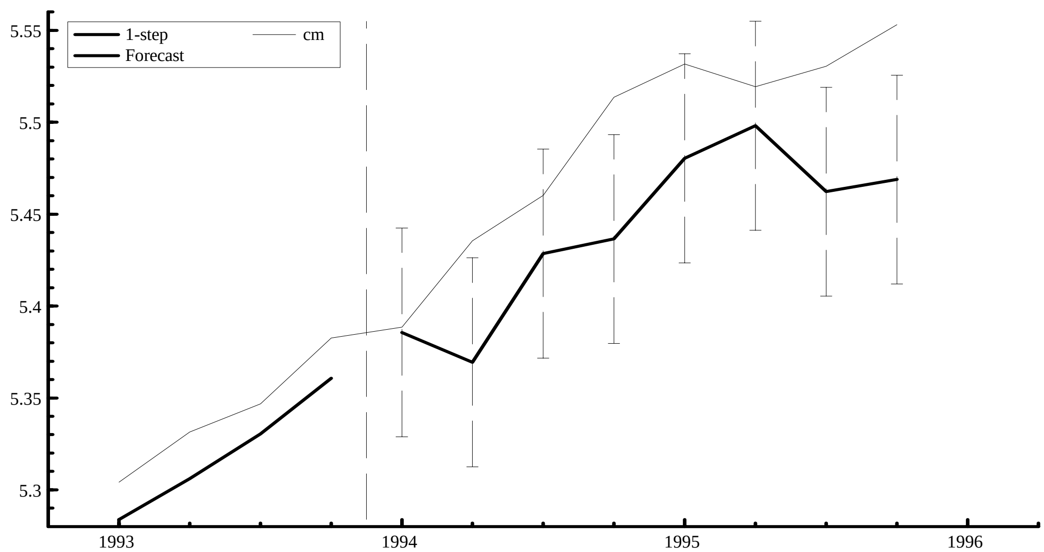


Figure 11: Predictive Accuracy-Cointegration-Imports pre–Nafta-Canada

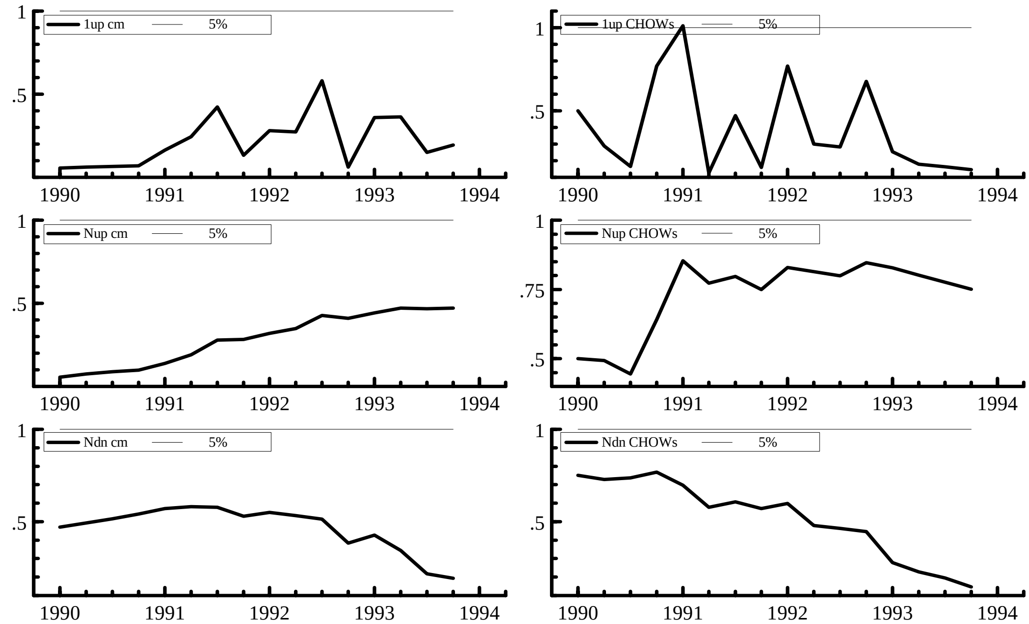


Figure 12: Chow Tests-Cointegration-Imports pre-Nafta-Canada

Short-run Forces To explain short-run fluctuations, we estimate the parameters of an error-correction model including a NAFTA dummy variable that takes a value of one starting in 1994.1 until the end of the sample. The results are

$$\begin{array}{lllll} \Delta m_t = & +1.2576\Delta y_t & -0.1352\Delta rpm_t & +0.03165 NAFTA_t & -0.1059 ECM_{m,t-1} \\ (se) & (0.21) & (0.16) & (0.015) & (0.037) \end{array}$$

where $ECM_m = m - 1.392 \cdot dy + 0.845 rpm + intercept.$

$R^2 = 0.44; SER = 2.77\%$	Null Hypothesis (p-value)	
Sample: 1961.2-1994.4	Serial-Independence (0.97)	Normality (0.01*)
	Homoskedasticity (0.79)	Func. Form (0.38)

The model explains just under half of the variability of the growth rate of imports and the empirical distribution of the residuals satisfies the assumptions maintained for estimation except normality. Unlike the predictions of the cointegration system, the predictions of the error-correction model do not show a tendency to be one-sided (figure 13) though the variable being predicted is the growth rate. Finally, the Chow tests (figure 14) and coefficient estimates (figure 15) from the error correction model support parameter constancy.

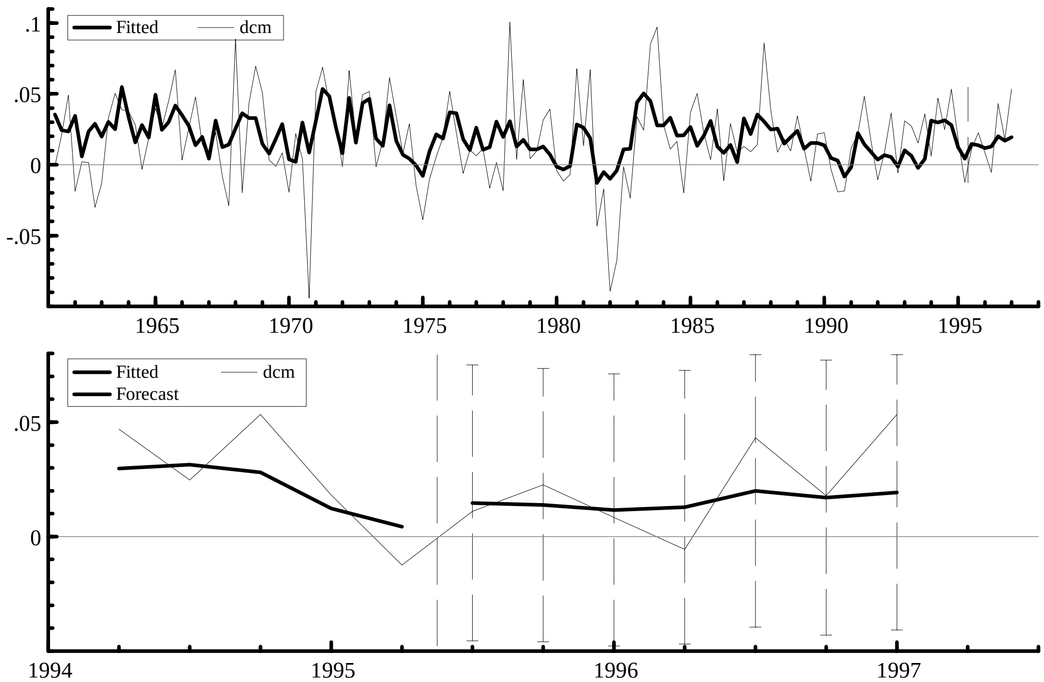


Figure 13: Predictive Accuracy-ECM-Imports-Canada

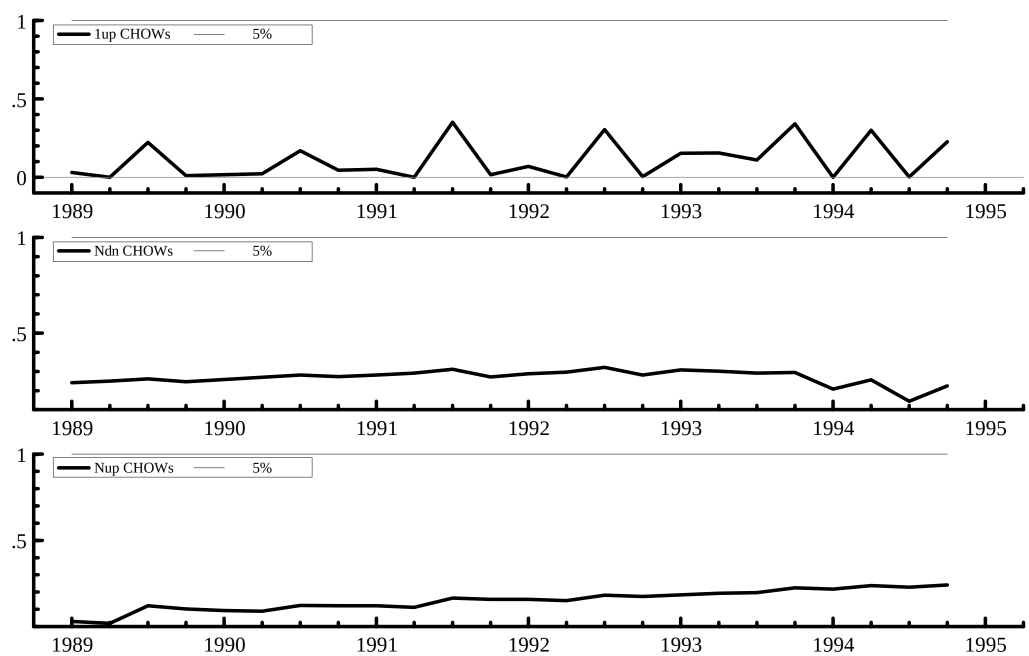


Figure 14: Chow Tests-ECM Coefficients - Imports Canada

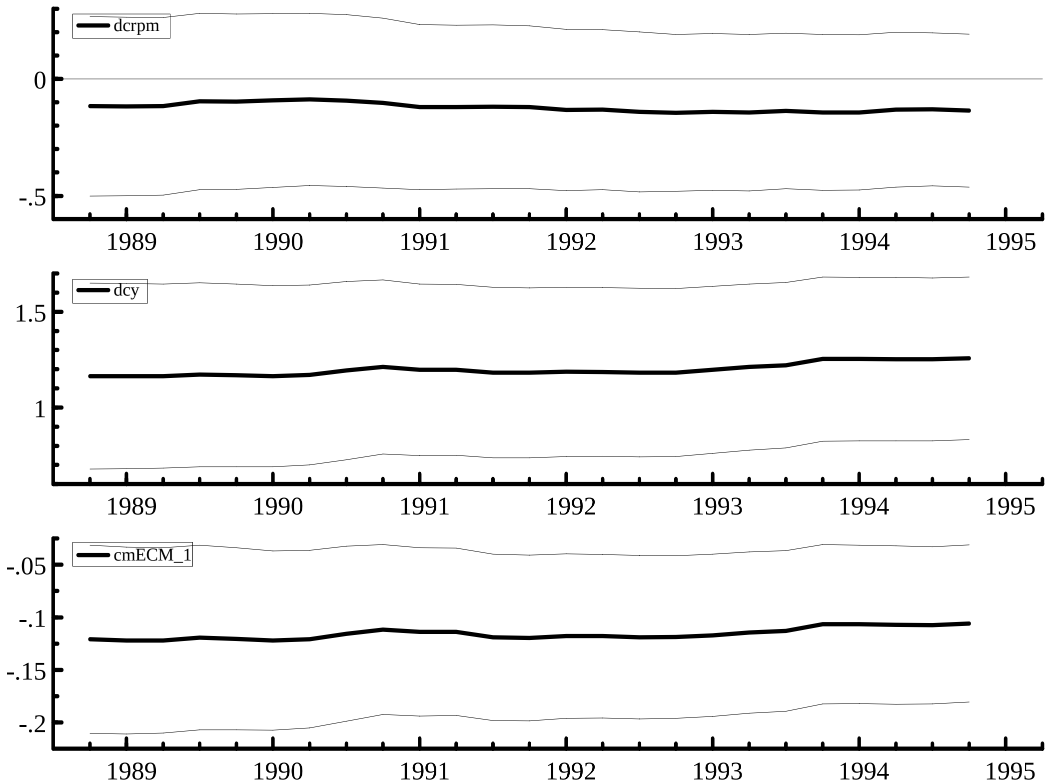


Figure 15: 95% Bands for ECM Coefficients-Imports-Canada

F French Trade

The evolution of France's foreign trade and its proximate determinants are displayed in figure 16. Both export and import volumes grow over time along with foreign and domestic income. Relative prices for imports show a downward trend to a greater extent than exports' relative prices. In addition, the IMF's measure of the real effective exchange rate shows smaller volatility than the conventional measure of relative prices based on deflators from the National Income Accounts.

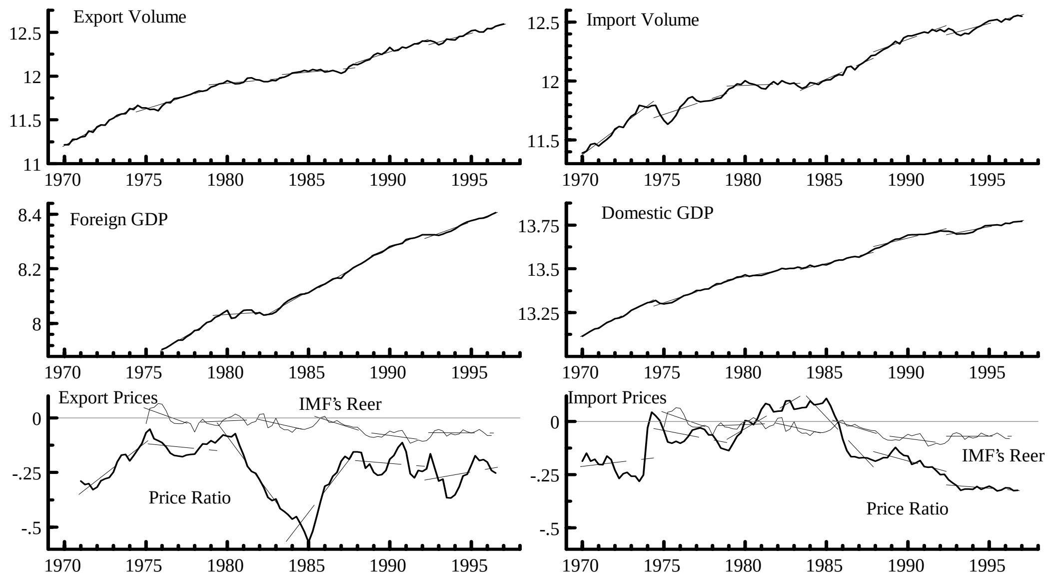


Figure 16: Trade, Income, and Prices: France

French Exports

Long-run Forces Measuring prices as *rpx* yields cointegration for formulations including lags in excess of four quarters but the price elasticity is positive. For lags fewer than four quarters, the cointegration results point to estimates of income and price elasticities of 1.5 and -0.3, respectively. Measuring prices with the *reer* yields no cointegration for systems including more than three lags. For the two cases with cointegration, the results point to relatively high price elasticities (in excess of two in absolute value) and income elasticities of one. Thus the choice of measure of relative price yields a rather different characterization of French exports.

Export Cointegration Results with *rpx*–France

	Number of lags Included							
	9	8	7	6	5	4	3	2†
Cointegration Vectors	1	1	1	1	1	2	1	1
Price-Elasticity	0.62	0.65	0.43	0.58	4.12	ni	-0.30	-0.21
Income Elasticity	1.80	1.80	1.72	1.78	3.16	ni	1.47	1.49
Loading Coefficient	-0.19	-0.15	-0.24	-0.14	-0.01	ni	-0.01	-0.06
System's R^2	0.95	0.95	0.94	0.93	0.92	0.92	0.90	0.89
Export's Serial Corr.	0.55	0.13	0.48	0.40	0.05	0.11	0.07	0.01*
System's Serial Corr.	0.17	0.11	0.17	0.24	0.31	0.74	0.48	0.17

Export Cointegration Results with *reer*–France

	Number of lags Included							
	9	8	7	6	5	4	3†	2
Cointegration Vectors	0	0	0	0	0	0	1	2
Price-Elasticity	ni	ni	ni	ni	ni	ni	-3.42	-2.94
Income Elasticity	ni	ni	ni	ni	ni	ni	0.952	0.96
Loading Coefficient	ni	ni	ni	ni	ni	ni	-0.06	-0.08
System's R^2	0.94	0.93	0.93	0.92	0.92	0.91	0.91	0.90
Export's Serial Corr.	0.17	0.42	0.92	0.89	0.75	0.40	0.41	0.15
System's Serial Corr.	0.35	0.07	0.42	0.31	0.69	0.66	0.73	0.56

ni: indicates that the elasticities are not identified.

The detailed cointegration results using *rpx* are (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.069 & (0.04) \\ -0.071 & (0.01) \\ -0.047 & (0.07) \end{pmatrix} \begin{pmatrix} 1 & -1.491 & 0.214 \\ (na) & (0.079) & (0.090) \end{pmatrix} \begin{pmatrix} x \\ fy \\ rpx \end{pmatrix}, \text{ 1976.3-1994.4}$$

whereas the cointegration results using *reer* are

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.072 & (0.03) \\ -0.037 & (0.01) \\ -0.042 & (0.02) \end{pmatrix} \begin{pmatrix} 1 & -0.952 & 3.423 \\ (na) & (0.205) & (0.886) \end{pmatrix} \begin{pmatrix} x \\ fy \\ reer \end{pmatrix}, \text{ 1976.4-1994.4}$$

The results indicate that reliance on the conventional *rpx* understates, by a substantial margin, the price responsiveness of French exports relative to the IMF's *reer*. Moreover, the error-correction term is not statistically significant for *rpx* which gives a statistical edge to the formulation using *reer*. Figures 17 and 18 below compare the ex-post predictions using the two measures of relative prices; figures 19 and 20 report the various Chow tests for each measure of relative prices. Inspection of this evidence gives a slight edge to the system using *reer* as the measure of relative prices. Though reliance on this measure results in slightly higher forecast errors, rejections of the Chow tests are considerably fewer than using the *rpx* measure.

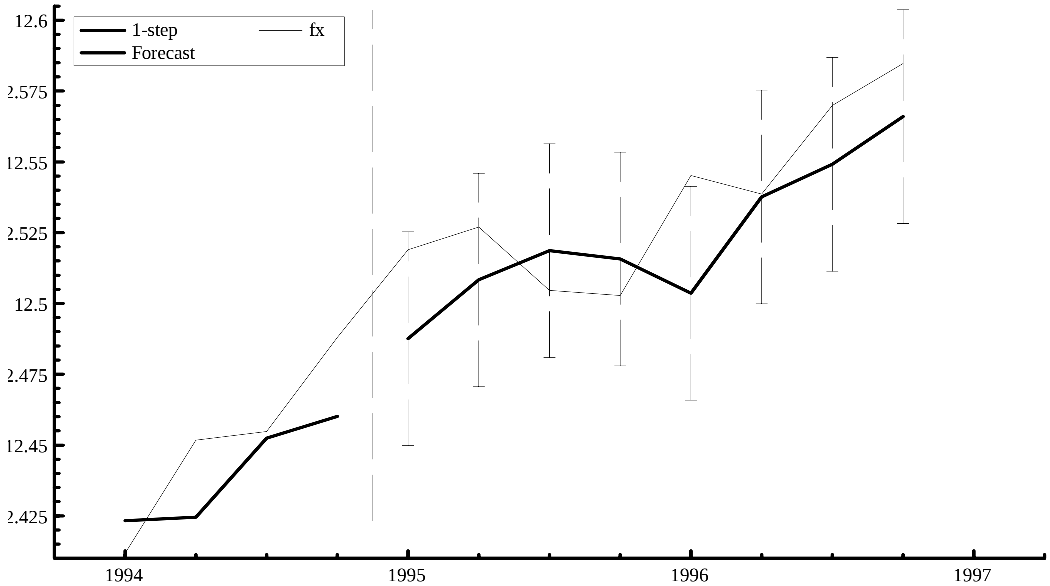


Figure 17: Predictive Accuracy-Cointegration-Exports with rpx-France

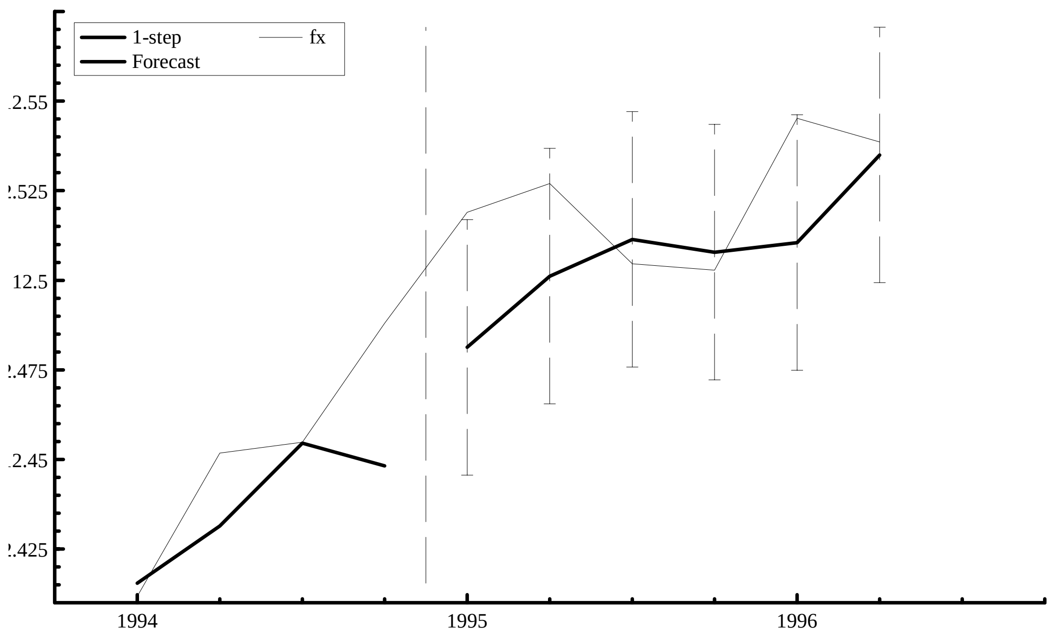


Figure 18: Predictive Accuracy-Cointegration-Exports with $reer$-France

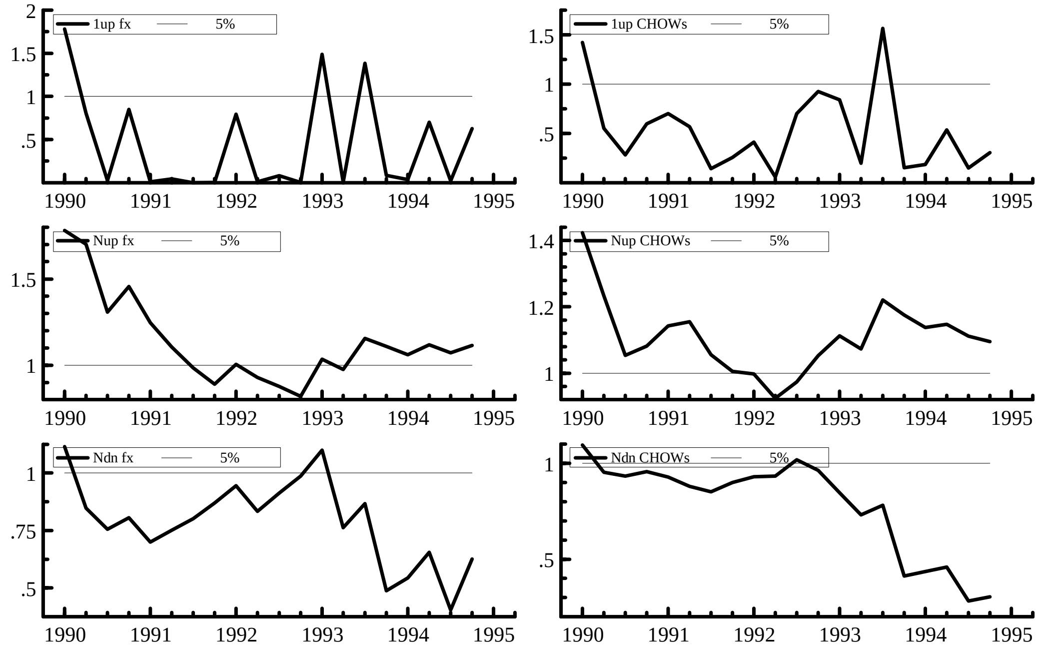


Figure 19: Chow Tests-Cointegration-Exports with rpx-France

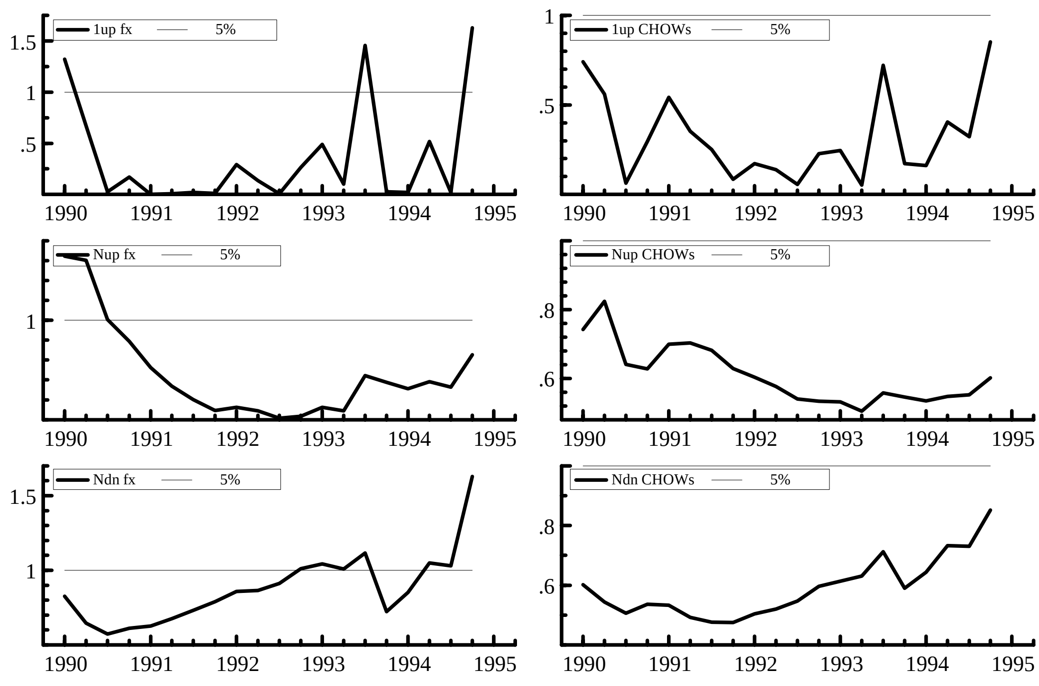


Figure 20: Chow Tests-Cointegration-Exports with $reer$-France

Short-run Forces To explain short-run fluctuations, we estimate the parameters of an error-correction model including two dummy variables for 1990.1 and 1990.2 to control for the German re-unification effects on French exports. The estimation results using rpx are

$$\underset{(se)}{\Delta x_t =} \quad \underset{(0.43)}{+1.82\Delta fy_t} \quad \underset{(0.06)}{-0.092\Delta rpx_t} \quad \underset{(0.04)}{-0.004ECM_{x,t-1}},$$

where $ECM_x = x - 1.491 \cdot fy + 0.214 \cdot rpx + intercept.$

$R^2 = 0.52; SER = 1.69\%$	Null Hypothesis (p-value)	
Sample: 1976.3-1994.4	Serial-Independence (0.21)	Normality (0.78)
	Homoskedasticity (0.84)	Func. Form (0.82)

The estimation results using $reer$ are

$$\underset{(se)}{\Delta x_t =} \quad \underset{(0.29)}{+0.994\Delta fy_t} \quad \underset{(0.08)}{-0.224\Delta reer_t} \quad \underset{(0.02)}{-0.078ECM_{x,t-1}},$$

where $ECM_x = x - 0.952 \cdot fy + 3.423 \cdot reer + intercept.$

$R^2 = 0.54; SER = 1.51\%$	Null Hypothesis (p-value)	
Sample: 1976.3-1994.4	Serial-Independence (0.94)	Normality (0.24)
	Homoskedasticity (0.68)	Func. Form (0.15)

Both formulations explain about half of the variability of the growth rate of exports and the empirical distribution of the residuals satisfies the assumptions maintained for estimation. Notice, however, that the error-correction coefficient for the formulation using rpx is not significantly different from zero. This result weakens the evidence on cointegration for the system relying on rpx.

Unlike the predictions of the cointegration system, the predictions of the error-correction model do not show a tendency to be one-sided (figures 21 and 22) though the variable being predicted is the growth rate. Notice, however, that the out-of-sample predictions of the formulation using rpx are worse than those of $reer$.

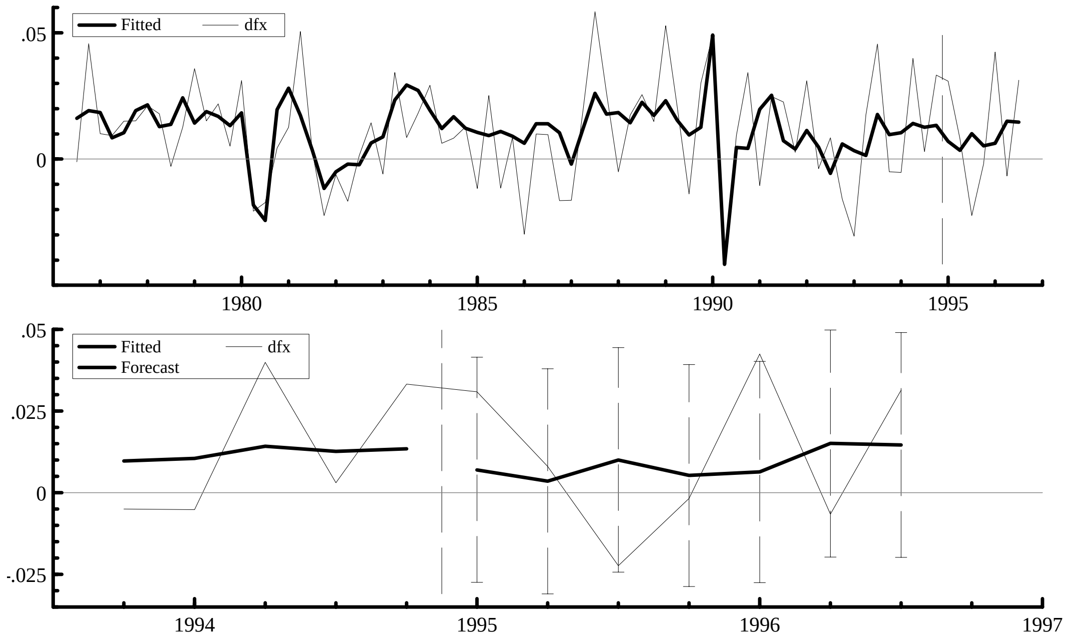


Figure 21: Predictive Accuracy-ECM Coefficients-Exports with rpx-France

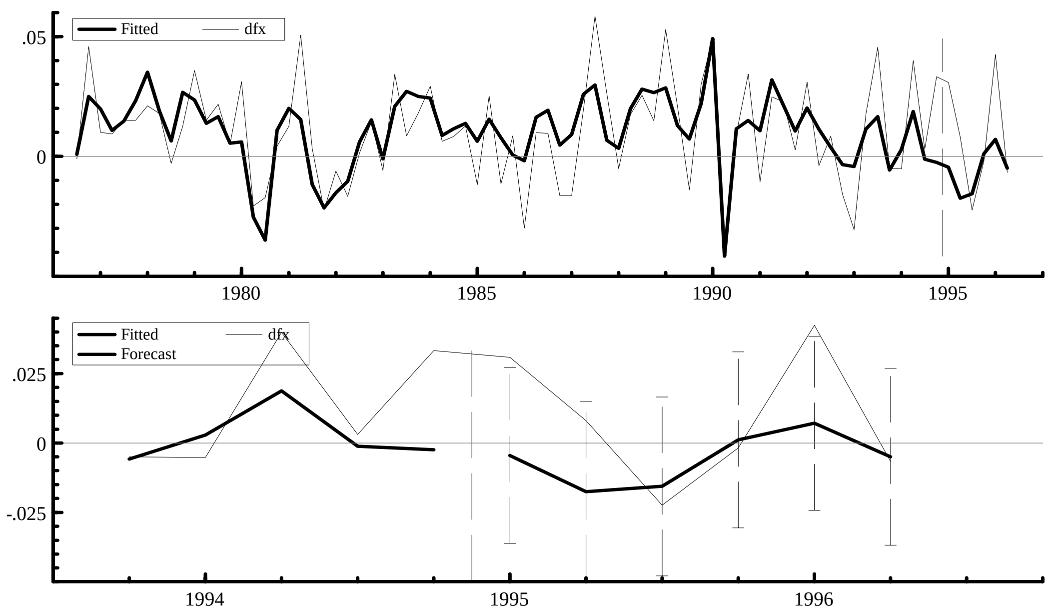


Figure 22: Predictive Accuracy-ECM Coefficients-Exports with $reer$-France

Finally, the Chow tests and the 95% confidence bands for the coefficient estimates support the hypothesis of parameter constancy (figures 23-26).

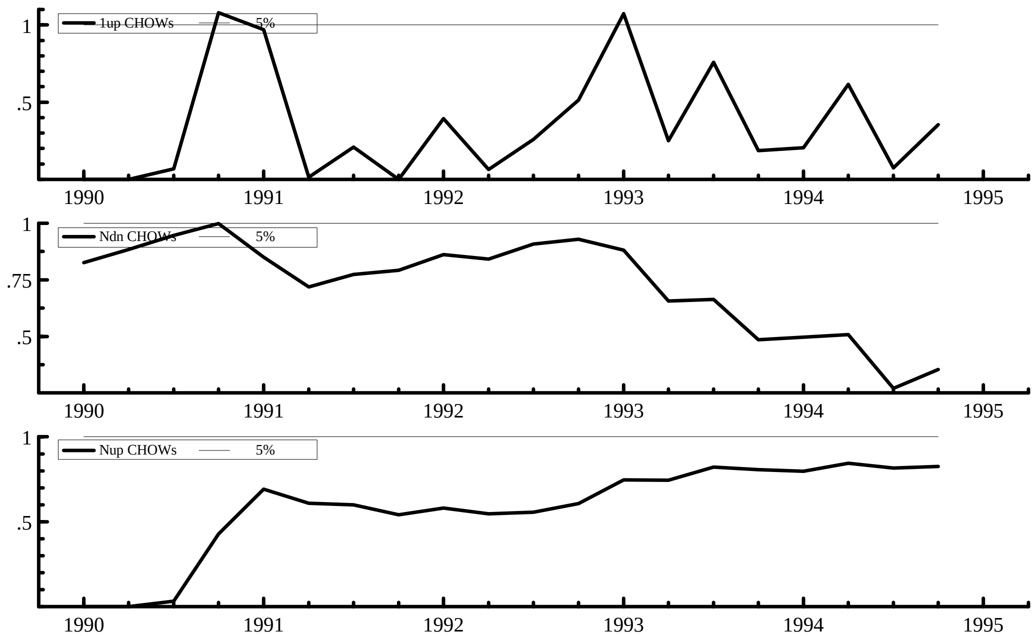


Figure 23: Chow Tests-ECM Coefficients-Exports with rpx-France

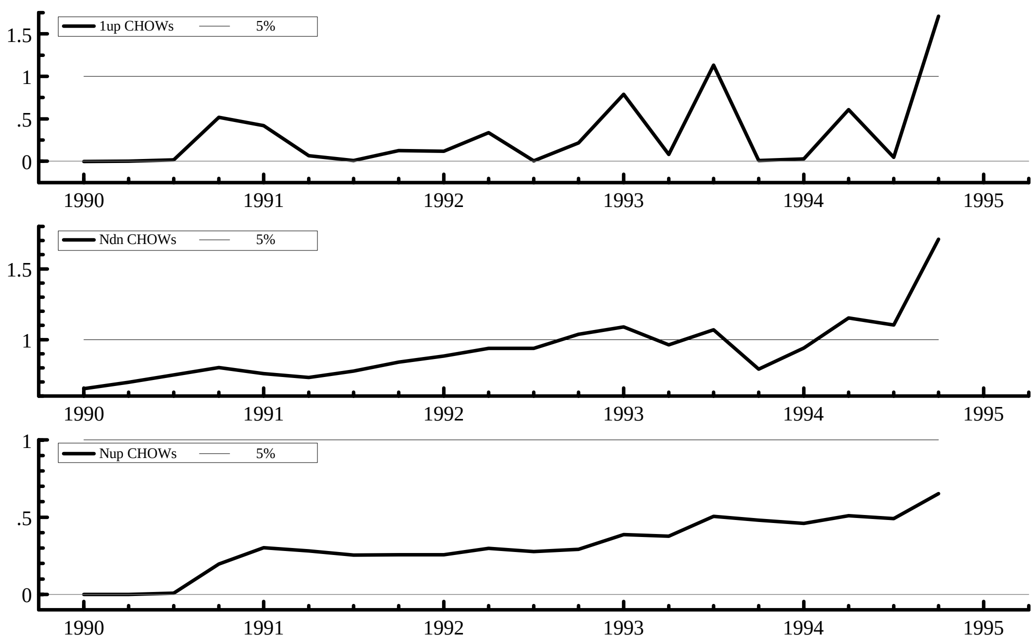


Figure 24: Chow Tests-ECM Coefficients-Exports with $reer$-France

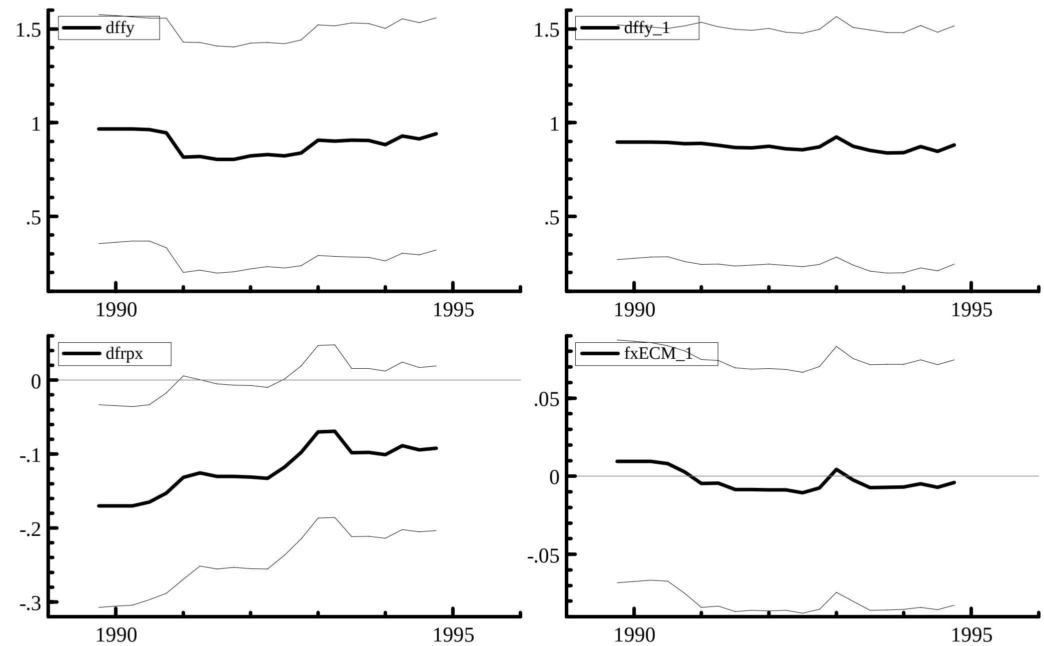


Figure 25: 95% Bands for ECM Coefficients-Exports with rpx-France

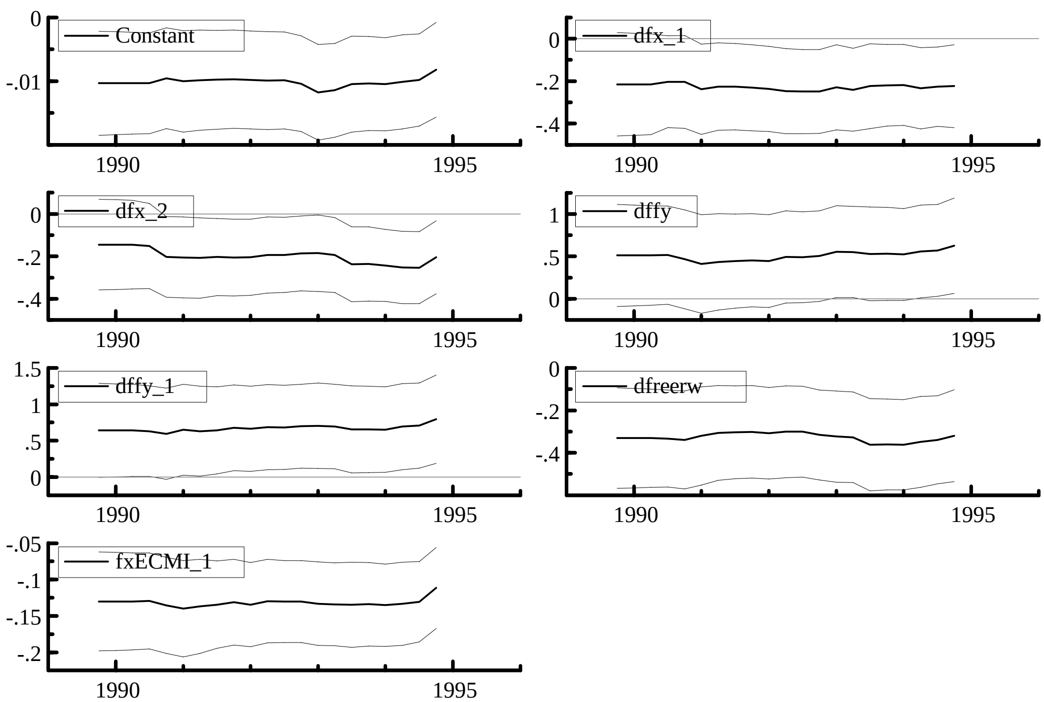


Figure 26: 95% Bands for ECM Coefficients-Exports with $reer$-France

French Imports

Long-run Forces Measuring relative prices as the import price relative to the GDP deflator yields no cointegration for systems using more than four lags. For systems exhibiting cointegration, the price elasticities are in the order of -0.4 and the income elasticities are in the order of 1.6. Measuring relative prices with the IMF's real effective exchange rate reveals that, for the two cases where there is cointegration, income and price elasticities are large enough to question their reliability.

Import Cointegration Results with *rpm*-France

	Number of lags Included							
	9	8	7	6	5	4	3†	2
Cointegration Vectors	0	0	0	0	0	1	1	1
Price-Elasticity	ni	ni	ni	ni	ni	-0.35	-0.37	-0.38
Income Elasticity	ni	ni	ni	ni	ni	1.61	1.59	1.59
Loading Coefficient	ni	ni	ni	ni	ni	-0.61	-0.60	-0.47
System's R^2	0.92	0.92	0.91	0.91	0.90	0.89	0.89	0.88
Import's Serial Corr.	0.00*	0.02*	0.08*	0.32	0.16	0.18	0.08	0.03*
System's Serial Corr.	0.09	0.17	0.07	0.00*	0.10	0.02*	0.12	0.01*

Import Cointegration Results with *reer*-France

	Number of lags Included							
	9	8	7	6	5	4	3	2
Cointegration Vectors	0	0	0	0	0	1	1	2
Price-Elasticity	ni	ni	ni	ni	ni	5.08	30.9	ni
Income Elasticity	ni	ni	ni	ni	ni	0.97	13.5	ni
Loading Coefficient	ni	ni	ni	ni	ni	-0.00	0.00	ni
System's R^2	0.91	0.91	0.90	0.90	0.90	0.90	0.88	0.88
Import's Serial Corr.	0.08	0.04*	0.02*	0.33	0.57	0.04*	0.15	0.40
System's Serial Corr.	0.01*	0.18	0.13	0.22	0.11	0.04*	0.02*	0.13

ni: indicates that the elasticities are not identified.

The cointegration results using *rpm* are (standard errors in parentheses):

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.601 & (0.09) \\ 0.091 & (0.13) \\ -0.086 & (0.03) \end{pmatrix} \begin{pmatrix} 1 & -1.595 & 0.369 \\ (na) & (0.026) & (0.032) \end{pmatrix} \begin{pmatrix} m \\ y \\ rpm \end{pmatrix}, \text{ 1972.2-1994.4.}$$

The out-of-sample predictions from the cointegration system reveal a reasonable degree of predictive accuracy (figure 27). Moreover, Chow tests for the cointegration system cannot reject the hypothesis of parameter constancy for the long-run elasticities (figure 28).

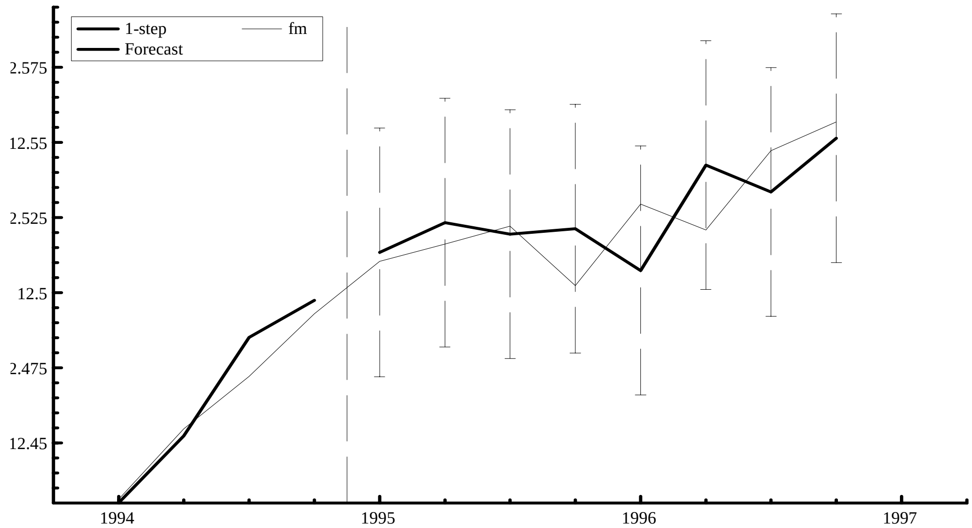


Figure 27: Predictive Accuracy-Cointegration-Imports-France

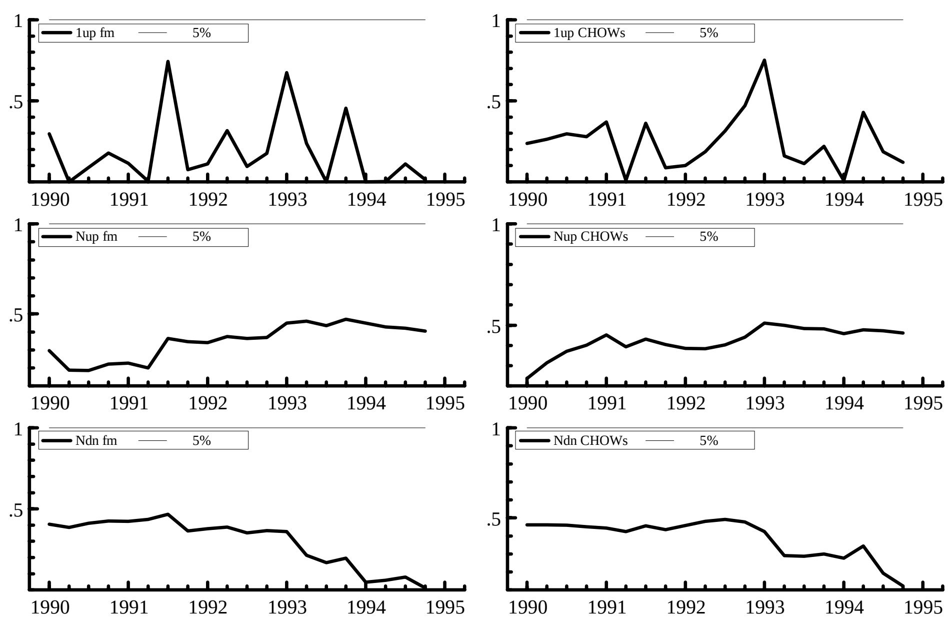


Figure 28: Chow Tests - Cointegration - Imports - France

Short-run Forces To explain short-run fluctuations, we estimate the parameters of an error-correction model:

$$\begin{array}{lccc} \Delta m_t = & +1.645\Delta y_t & -0.0559\Delta rpm_t & -0.3352ECM_{m,t-1} \\ (se) & (0.40) & (0.14) & (0.07) \end{array}$$

where $ECM_m = m - 1.5946 \cdot y + 0.36898 \cdot rpm$.

$R^2 = 0.67; SER = 1.63\%$	Null Hypothesis (p-value)	
Sample: 1971.3-1994.4	Serial-Independence (0.21)	Normality (0.74)
	Homoskedasticity (0.24)	Func. Form (0.51)

The model explains two-thirds of the variability of the growth rate of imports and the empirical distribution of the residuals satisfies the assumptions maintained for estimation. Like the predictions of the cointegration system, the predictions of the error-correction model are not one-sided (figure 29). Finally, the Chow tests (figure 30) and the 95% confidence bands for the ECM's coefficient estimates (figure 31) suggest a remarkable degree of constancy over this period.

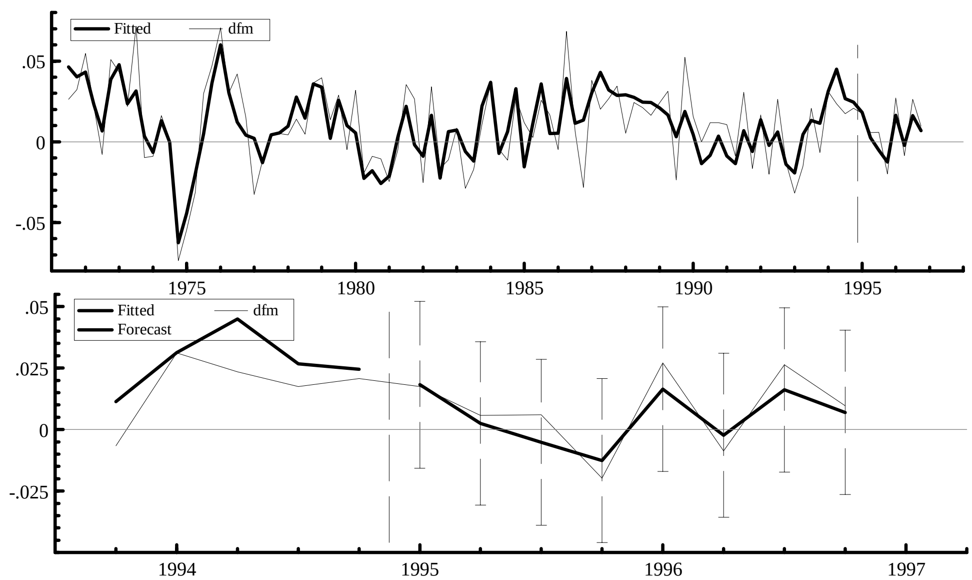


Figure 29: Predictive Accuracy-ECM Imports - France

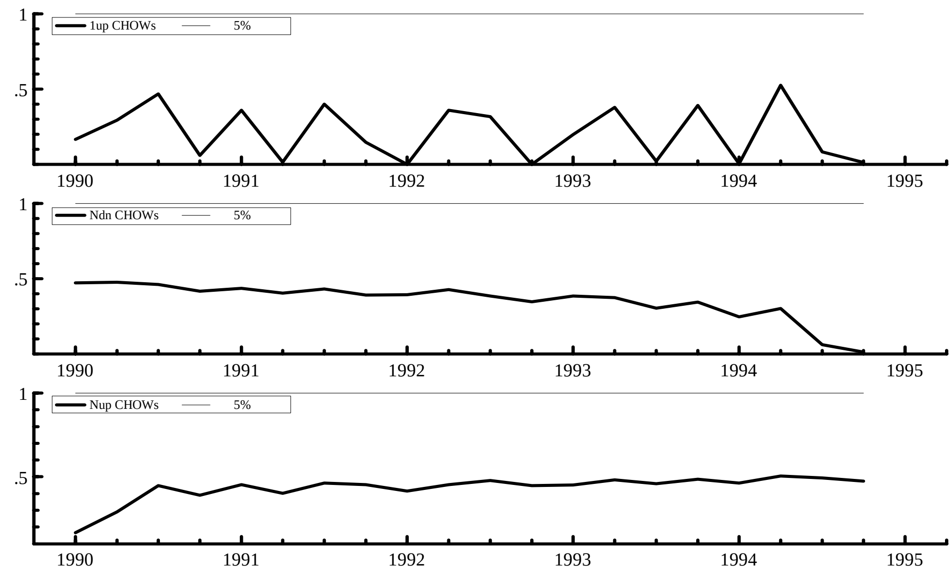


Figure 30: Chow Tests-ECM Imports - France

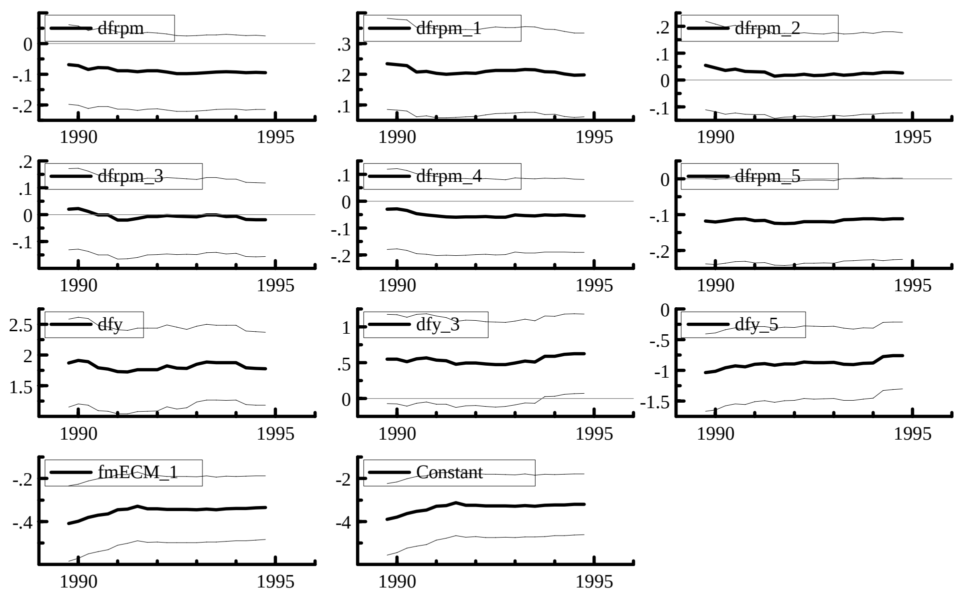


Figure 31: 95% Bands for ECM Coefficients-Imports-France

G German Trade

The evolution of Germany's foreign trade and its proximate determinants are displayed in figure 32. Both export and import volumes grow over time along with foreign and domestic income. Relative prices for imports show a downward trend whereas exports' relative prices show less of a trend. In addition, the IMF's real effective exchange rate shows smaller volatility than the conventional measure of relative prices based on deflators from the National Income Accounts.

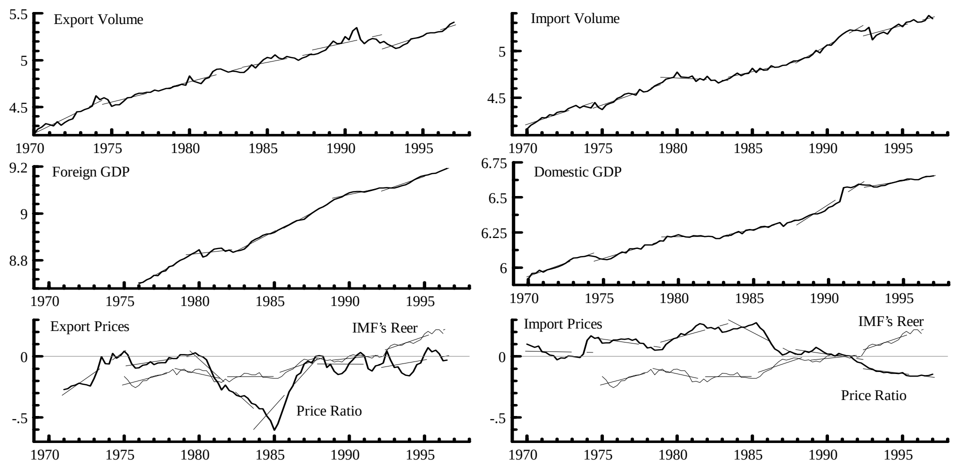


Figure 32: Trade, Income, and Prices: Germany

German Exports

Long-run Forces Measuring prices with *rpx* yields cointegration vectors that are not robust to the choice of lag. Specifically, systems using fewer than eight lags and more than two lags do not exhibit cointegration. For the cases where cointegration exists, the estimated price elasticity ranges from -0.22 to -0.27 and the income elasticity is centered on 1.5. Measuring prices with the IMF's *reer* reveals cointegration in three instances and the associated price elasticities are about minus one and the income elasticities are about two. These formulations are statistically satisfactory and provide a rather different characterization of the behavior of German exports.

Export Cointegration Results with *rpx*-Germany

	Number of lags Included							
	9	8	7	6	5	4	3	2†
Cointegration Vectors	1	1	0	0	0	0	0	1
Price-Elasticity	-0.27	-0.22	ni	ni	ni	ni	ni	-0.26
Income Elasticity	1.62	1.64	ni	ni	ni	ni	ni	1.43
Loading Coefficient	-0.38	-0.36	ni	ni	ni	ni	ni	-0.10
System's R^2	0.97	0.97	0.96	0.96	0.96	0.96	096	0.96
Export's Serial Corr.	0.58	0.70	0.64	0.11	0.46	0.45	0.73	0.28
System's Serial Corr.	0.99	0.93	0.84	0.37	0.57	0.86	0.99	0.99

Export Cointegration Results with *reer*-Germany

	Number of lags Included							
	9	8	7	6	5	4	3†	2
Cointegration Vectors	0	1	0	0	0	1	1	2
Price-Elasticity	ni	-0.90	ni	ni	ni	-1.08	-1.13	ni
Income Elasticity	ni	2.08	ni	ni	ni	2.17	2.23	ni
Loading Coefficient	ni	-0.98	ni	ni	ni	-0.34	-0.30	ni
System's R^2	0.95	0.95	0.94	0.94	0.94	0.94	0.93	0.93
Export's Serial Corr.	0.64	0.48	0.57	0.10	0.62	0.92	0.95	0.94
System's Serial Corr.	0.67	0.33	0.13	0.47	0.64	0.80	0.75	0.97

ni: indicates that the elasticities are not identified.

The detailed cointegration results using *rpx* are (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.100 & (0.05) \\ 0.037 & (0.07) \\ -0.048 & (0.01) \end{pmatrix} \begin{pmatrix} 1 & -1.433 & 0.255 \\ (na) & (0.135) & (0.097) \end{pmatrix} \begin{pmatrix} x \\ fy \\ rpx \end{pmatrix}, \text{ 1978.2-1994.4}$$

whereas the cointegration results using *reer* are

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.318 & (0.11) \\ 0.072 & (0.03) \\ -0.083 & (0.07) \end{pmatrix} \begin{pmatrix} 1 & -2.234 & 1.129 \\ (na) & (0.116) & (0.155) \end{pmatrix} \begin{pmatrix} x \\ fy \\ reer \end{pmatrix}, \text{ 1978.1-1994.4}$$

The results indicate that reliance on *rpx* understates, by a substantial margin, the responsiveness of German exports to changes in both income and prices compared to reliance on *reer*.

Inspection of the ex-post prediction errors for exports suggests that reliance on *rpx* gives smaller forecast errors than using *reer* (figures 33 and 34):

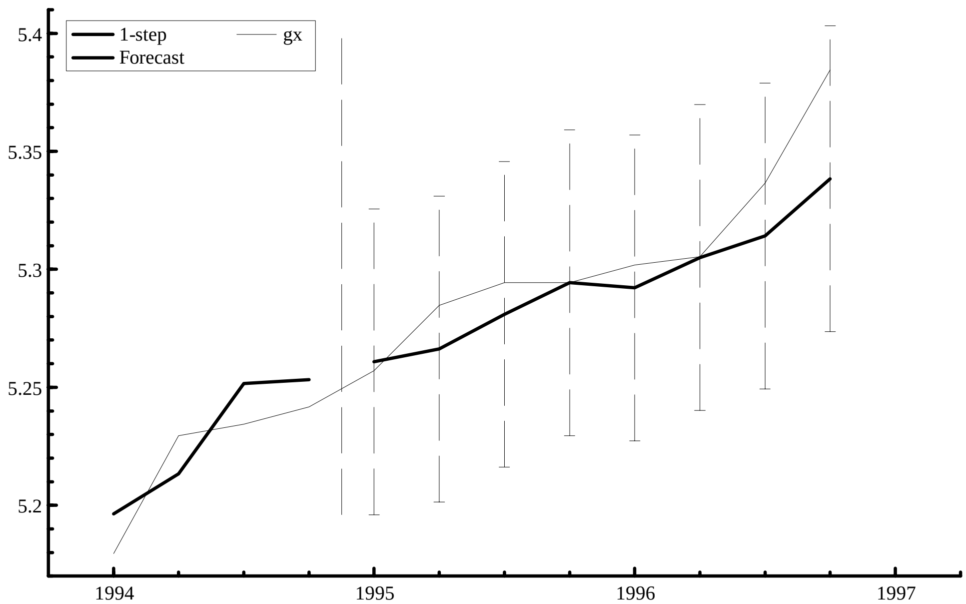


Figure 33: Predictive Accuracy-Cointegration-Exports with rpx-Germany

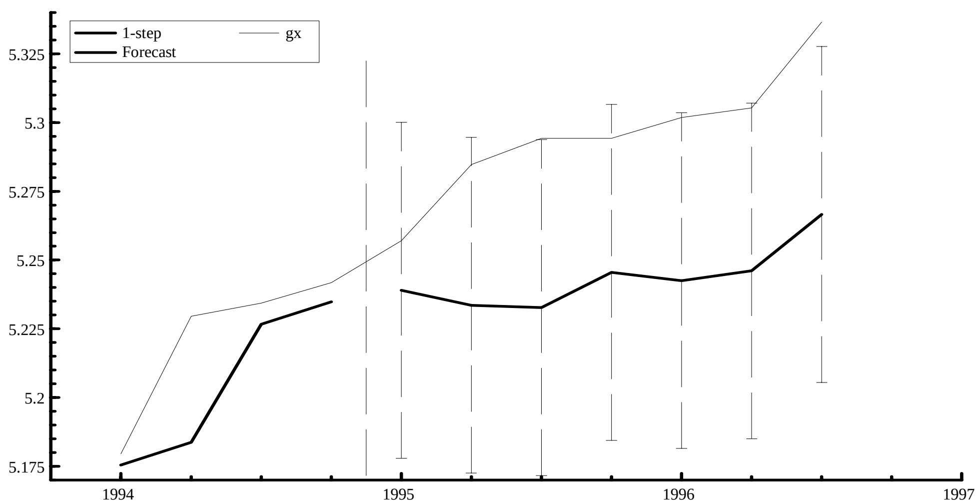


Figure 34: Predictive Accuracy-Cointegration-Exports with $reer$-Germany

Unfortunately, reliance on either measure for relative prices suggests parameter instability (figure 35 and 36).

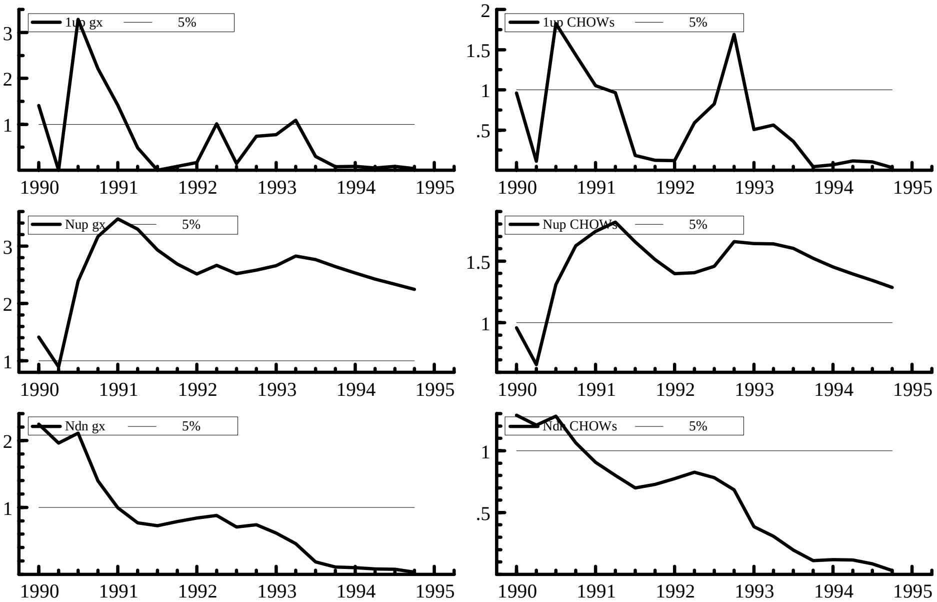


Figure 35: Chow Tests-Cointegration-Exports with rpx-Germany

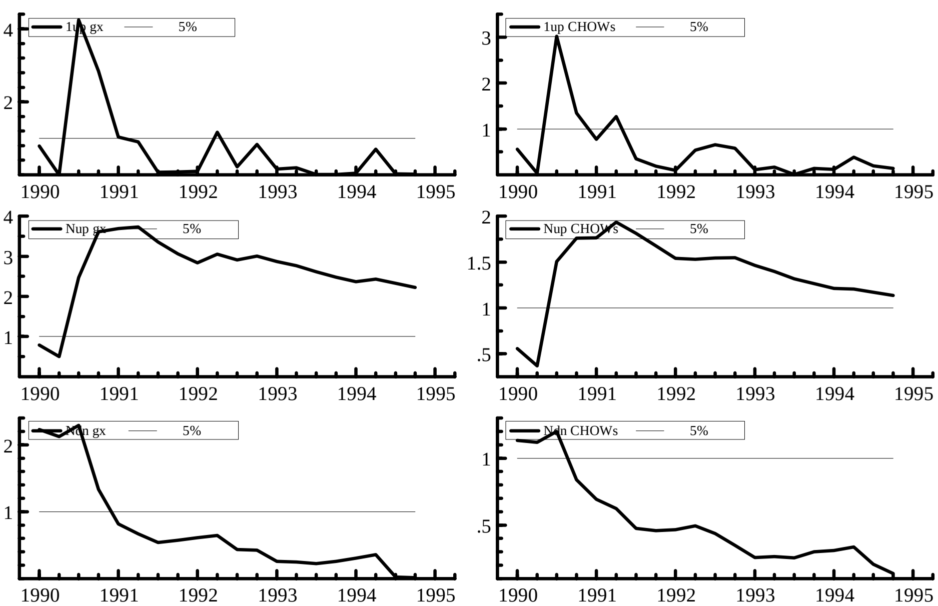


Figure 36: Chow Tests-Cointegration-Exports with $reer$-Germany

Short-run Forces To explain short-run fluctuations, we estimate the parameters of an error-correction model including dummy variables for 1990.1, 1990.3, and 1990.4 to control for the German re-unification effects. The estimation results using

rpx are

$$\underset{(se)}{\Delta x_t =} \underset{(0.41)}{+0.547\Delta fy_t} \underset{(0.06)}{-0.053\Delta rpx_t} \underset{(0.07)}{-0.476ECM_{x,t-1}},$$

where $ECM_x = x - 1.433 \cdot fy + 0.255 \cdot rpx + intercept.$

$R^2 = 0.57; SER = 2.30\%$	Null Hypothesis (p-value)	
Sample: 1976.2-1994.4	Serial-Independence (0.19)	Normality (0.09)
	Homoskedasticity (0.50)	Func. Form (0.06)

The estimation results using $reer$ are

$$\underset{(se)}{\Delta x_t =} \underset{(0.42)}{+1.429\Delta fy_t} \underset{(0.16)}{-0.167\Delta reer_t} \underset{(0.07)}{-0.47ECM_{x,t-1}},$$

where $ECM_x = x - 2.234 \cdot fy + 1.129 \cdot reer + intercept.$

$R^2 = 0.57; SER = 2.37\%$	Null Hypothesis (p-value)	
Sample: 1977.2-1994.4	Serial-Independence (0.47)	Normality (0.17)
	Homoskedasticity (0.69)	Func. Form (0.29)

Both formulations explain about half of the variability of the growth rate of exports and the empirical distributions of their residuals satisfy the assumptions maintained for estimation.

The predictions from both models are one-sided but the prediction errors of the model using rpx are smaller than those of the model using $reer$ (figures 37 and 38).

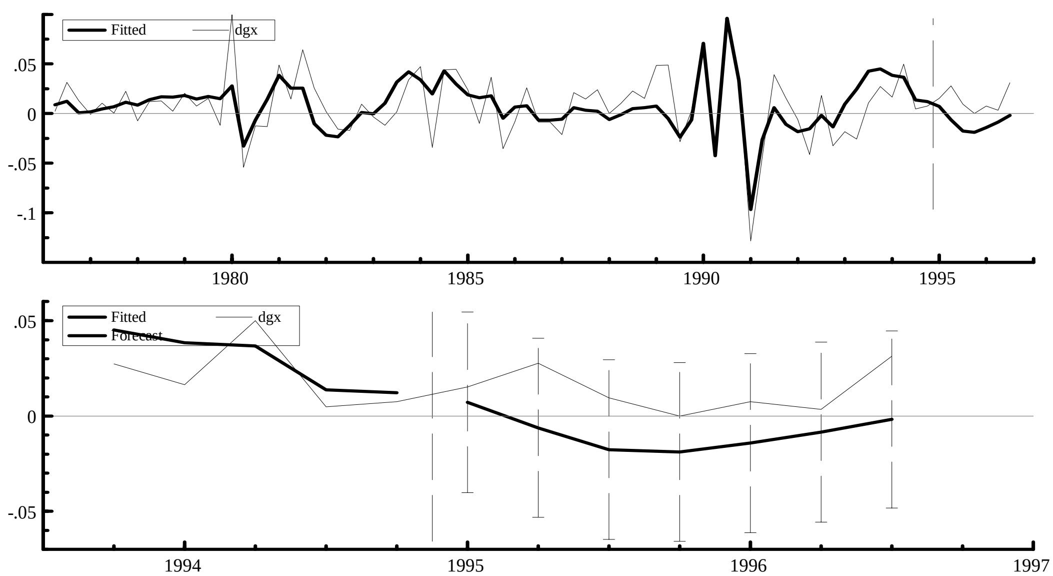


Figure 37: Predictive Accuracy-ECM-Exports with rpx-Germany

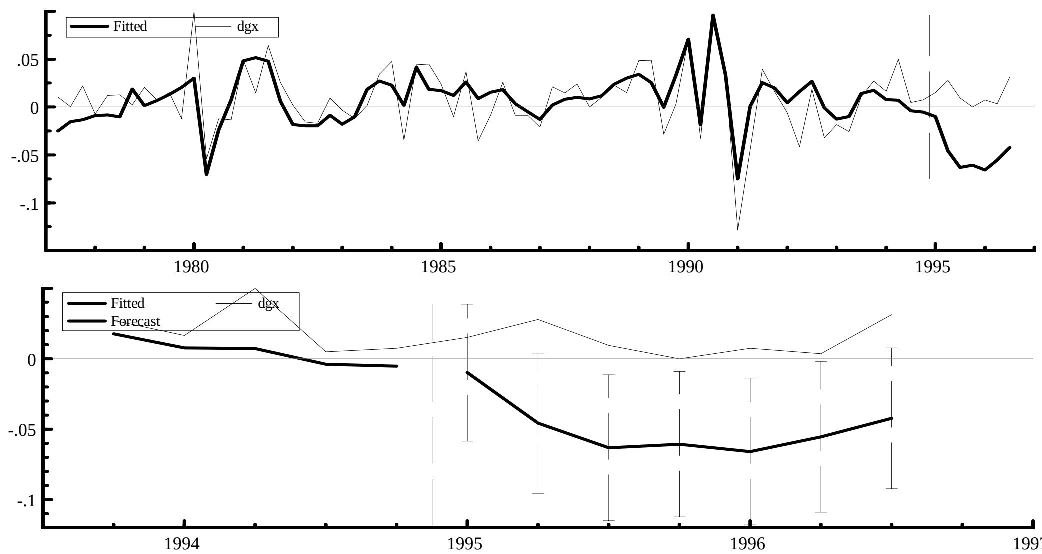


Figure 38: Predictive Accuracy-ECM-Exports with *reer*-Germany

Finally, both systems display evidence of parameter instability in terms of the Chow tests (figures 39 and 40) and in terms of the 95% confidence bands for the coefficient estimates (figures 41 and 42). Parameter instability is more pronounced for the system using *reer* than for the system using *rpx*.

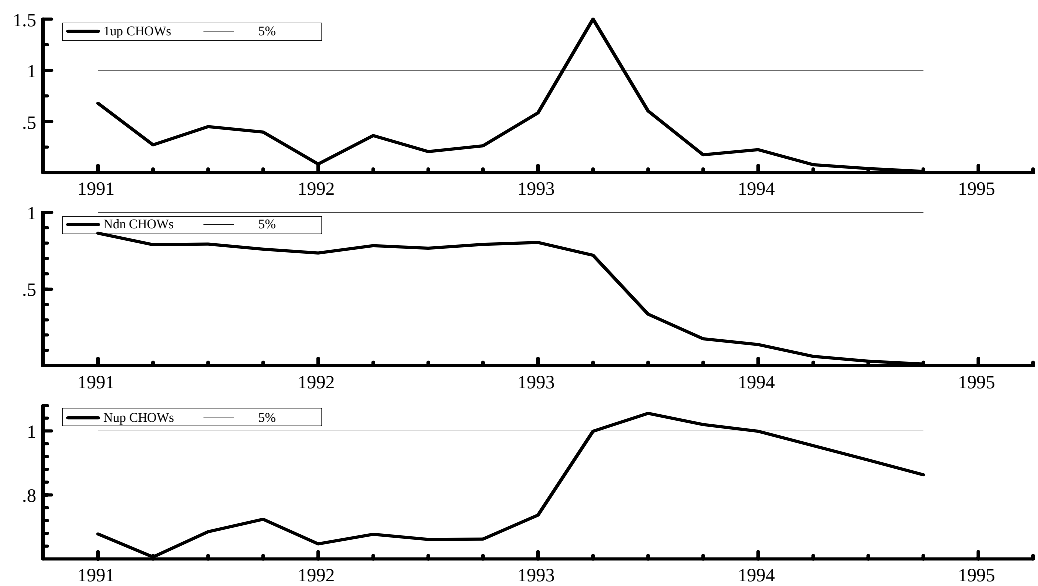


Figure 39: Chow Tests-ECM-Exports with *rpx*-Germany

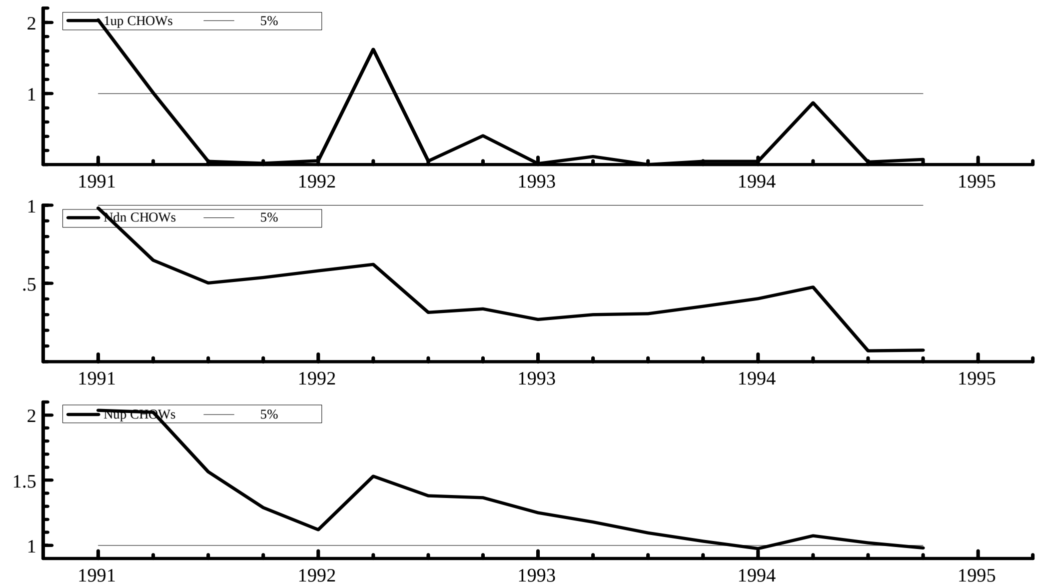


Figure 40: Chow Tests-ECM-Exports with *reer*-Germany

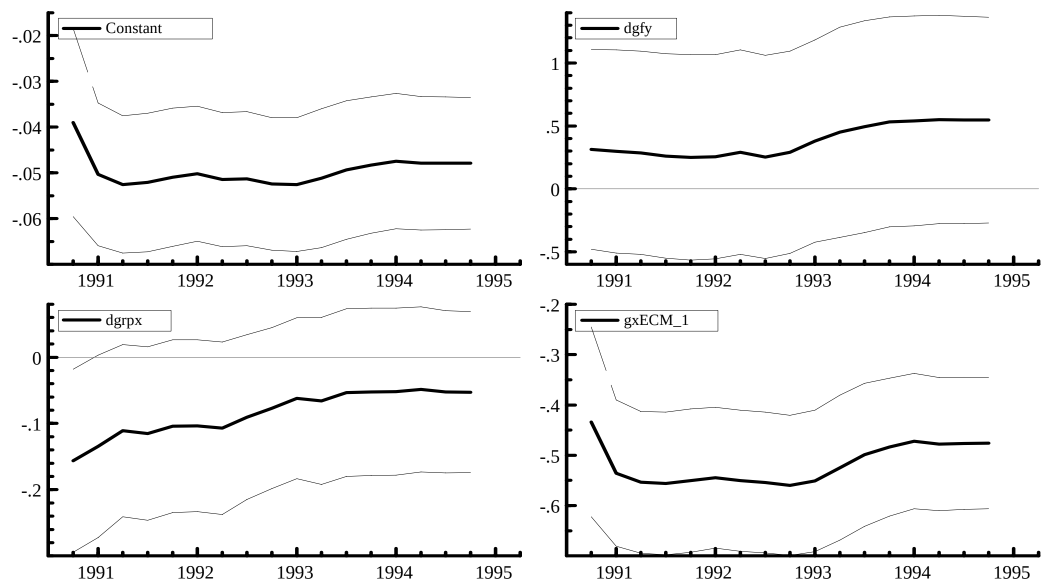


Figure 41: 95% Confidence Bands for ECM Coefficients-Exports with *rpx*-Germany

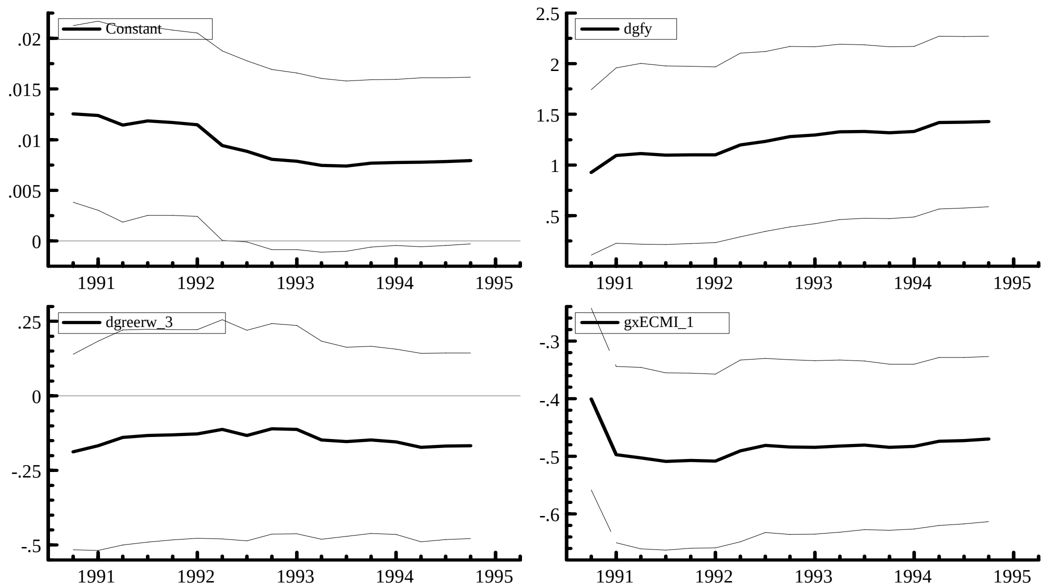


Figure 42: 95% Confidence Bands for ECM Coefficients-Exports with *reer*-Germany

German Imports

Long-run Forces Measuring prices as the import price relative to GDP deflator yields estimated elasticities for German imports that are not robust to the choice of lag. Indeed, systems using more than two lags either lack cointegration or have coefficient estimates with the wrong sign. The only case exhibiting cointegration and consistency with economic theory has lags of two quarters; the estimated price elasticity is -0.06 and the income elasticity is 1.5. The system exhibits, however, serial correlation. Measuring prices with the IMF's *reer* reveals three cases with cointegration but the estimated price elasticities have the wrong sign (they should be positive). Thus these formulations are not a satisfactory characterization of the behavior of German imports.

Import Cointegration Results with *rpm*-Germany

	Number of lags Included							
	9	8	7	6	5	4	3	2†
Cointegration Vectors	2	0	1	0	0	0	1	1
Price-Elasticity	ni	ni	0.25	ni	ni	ni	0.31	-0.06
Income Elasticity	ni	ni	1.65	ni	ni	ni	1.63	1.47
Loading Coefficient	ni	ni	-0.10	ni	ni	ni	0.07	0.15
System's R^2	0.98	0.98	0.98	0.98	0.98	0.98	0.98	0.98
Import's Serial Corr.	0.31	0.01*	0.18	0.05	0.31	0.02*	0.03*	0.05
System's Serial Corr.	0.74	0.31	0.29	0.27	0.31	0.18	0.17	0.00*

Import Cointegration Results with *reer*-Germany

	Number of lags Included							
	9	8	7	6	5	4	3	2
Cointegration Vectors	0	0	0	0	1	1	0	1
Price-Elasticity	ni	ni	ni	ni	-2.60	-2.11	ni	-0.66
Income Elasticity	ni	ni	ni	ni	3.08	2.78	ni	1.82
Loading Coefficient	ni	ni	ni	ni	0.09	0.12	ni	0.14
System's R^2	0.93	0.92	0.92	0.91	0.91	0.91	0.90	0.90
Import's Serial Corr.	0.78	0.34	0.47	0.24		0.78	0.08	0.13
System's Serial Corr.	0.98	0.54	0.39	0.71	-	0.82	0.53	0.29

ni: indicates that the elasticities are not identified.

The cointegration results using rpm are (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.125 & (0.02) \\ -0.039 & (0.01) \\ 0.031 & (0.02) \end{pmatrix} \begin{pmatrix} 1 & -1.468 & 0.064 \\ (na) & (0.125) & (0.228) \end{pmatrix} \begin{pmatrix} m \\ y \\ rpm \end{pmatrix}, \text{ 1968.3-1994.4.}$$

The estimated speed of adjustment is 0.1 and highly significant as is the income elasticity; the price elasticity is not significant however. The out-of-sample predictions from the cointegration system exhibit a reasonable degree of accuracy (figure 43). Chow tests reject, however, the hypothesis of parameter constancy (figure 44).

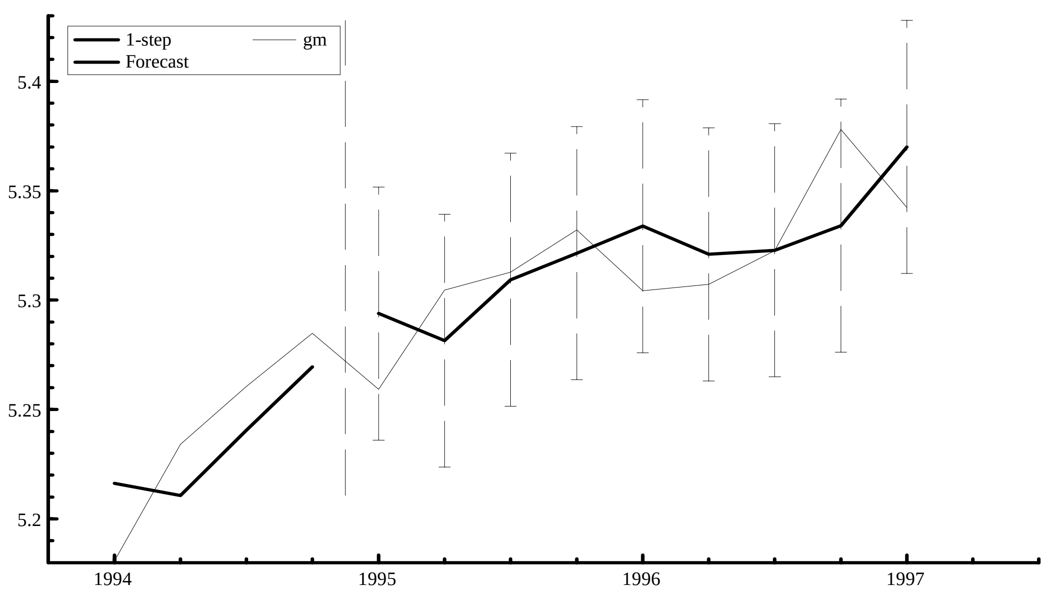


Figure 43: Predictive Accuracy-Cointegration-Imports-Germany

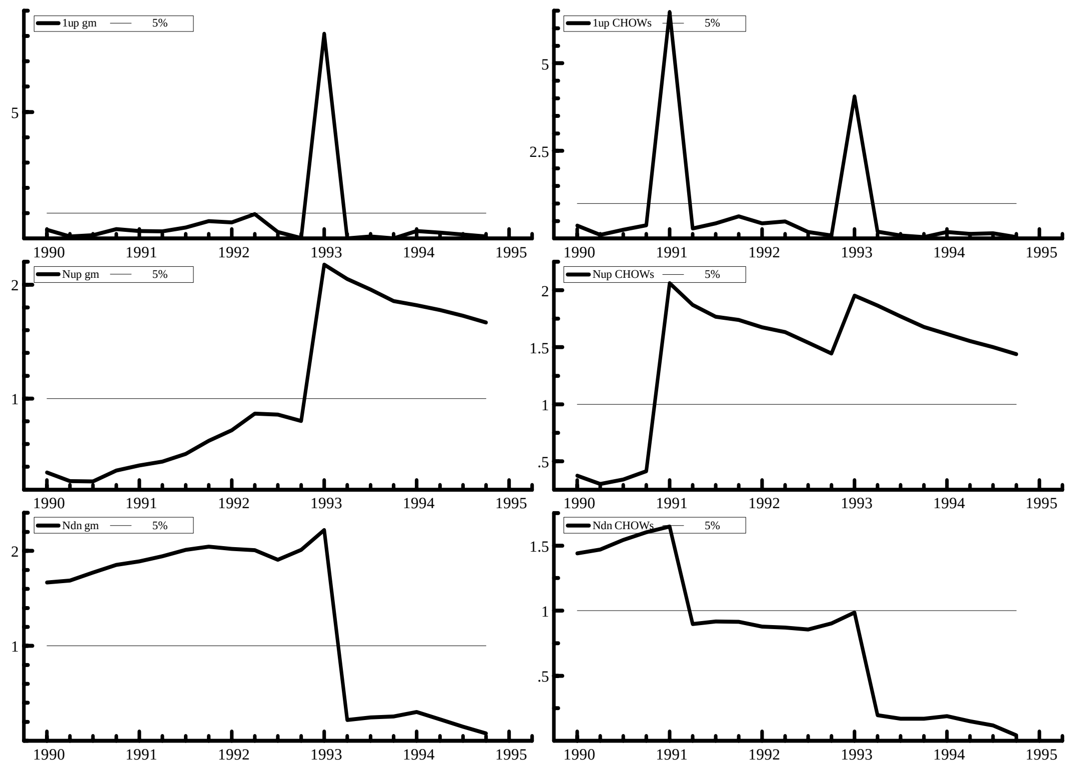


Figure 44: Chow Tests-Cointegration - Imports-Germany

Short-run Forces To explain short-run fluctuations, we estimate the parameters of an error-correction model including a dummy variable that takes a value of one in 1993.1. The estimation results are

$$\begin{array}{lccc} \Delta m_t = & +0.987\Delta y_t & -0.1698\Delta rpm_t & -0.1022ECM_{m,t-1} \\ (se) & (0.25) & (0.08) & (0.02) \end{array}$$

where $ECM_m = m - 1.4675 \cdot y + 0.0635 \cdot rpm$.

$R^2 = 0.49; SER = 2.32\%$	Null Hypothesis (p-value)	
Sample: 1968.3-1994.4	Serial-Independence (0.42)	Normality (0.96)
	Homoskedasticity (0.83)	Func. Form (0.02*)

The model explains about half of the variability of the growth rate of imports and the empirical distribution of the residuals satisfies the assumptions maintained for estimation except functional form. The prediction errors are not one-sided (figure 45), and both the Chow tests (figure 46) and the 95% confidence bands of the coefficients (figure 47) support the hypothesis of parameter constancy.

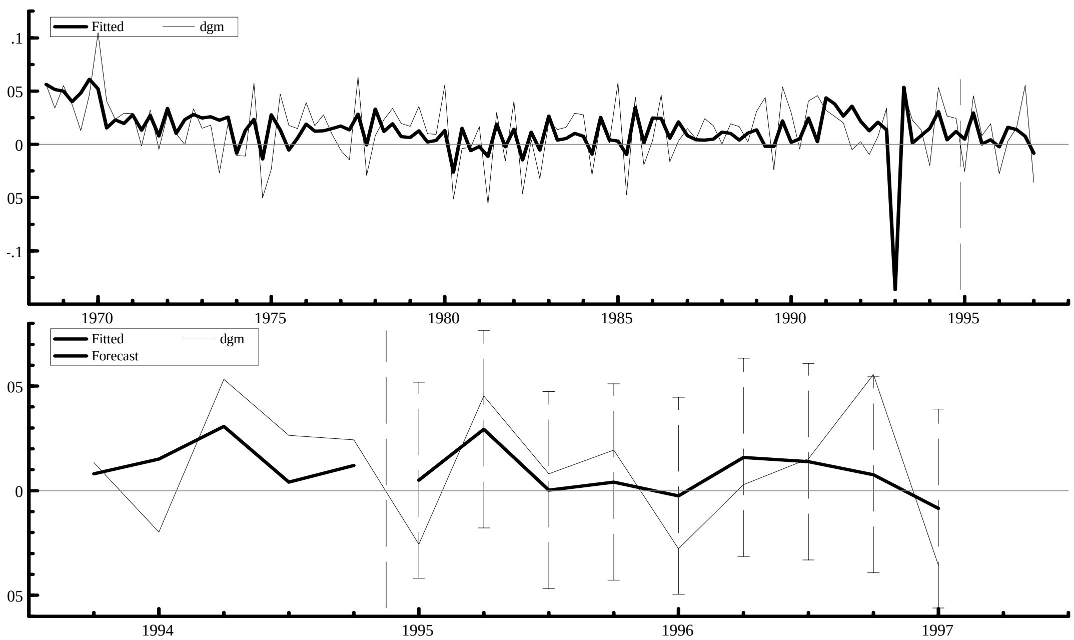


Figure 45: Predictive Accuracy - ECM - Imports-Germany

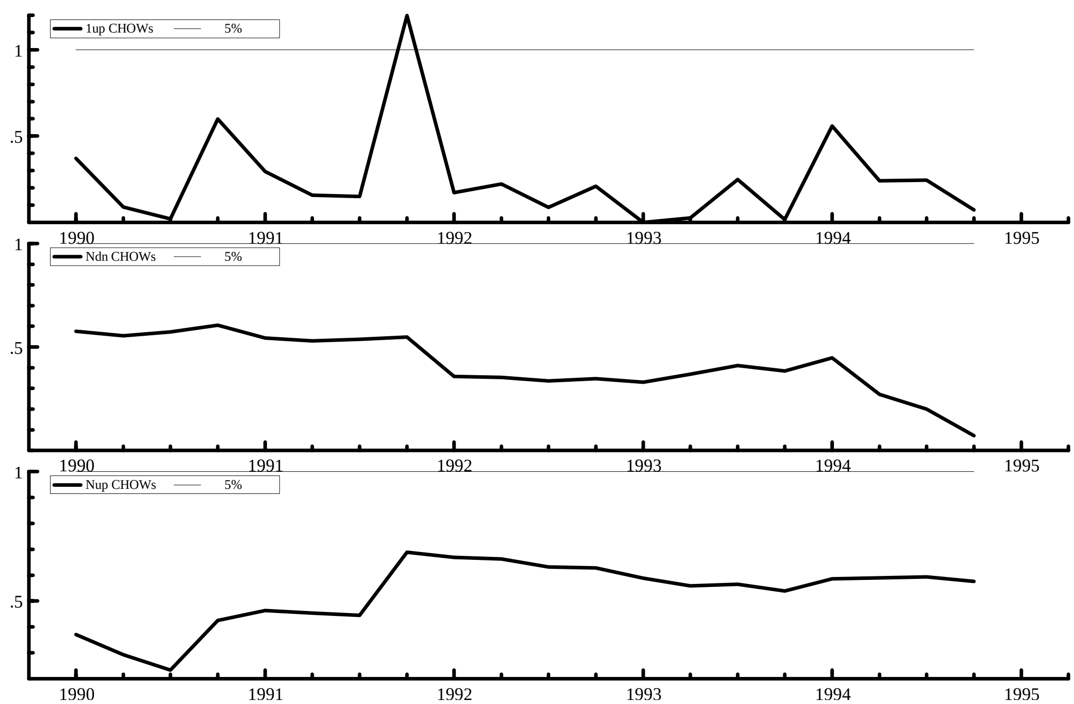


Figure 46: Chow Tests - ECM - Imports-Germany

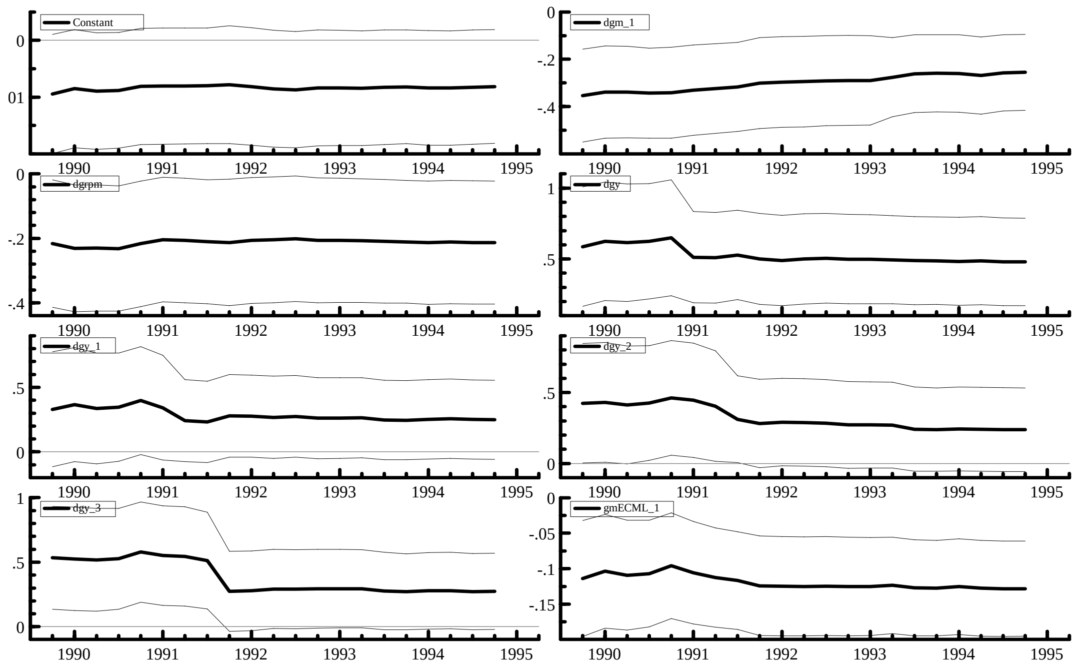


Figure 47: 95% Confidence Bands for - ECM Coefficients - Imports-Germany

H Italian Trade

The evolution of Italy's foreign trade and its proximate determinants are displayed in figure 48. Both export and import volumes grow over time along with foreign and domestic income. Relative prices for imports show a downward trend to a greater extent than exports' relative prices. In addition, the IMF's *reer* shows about the same volatility as relative prices based on deflators from the National Income Accounts.

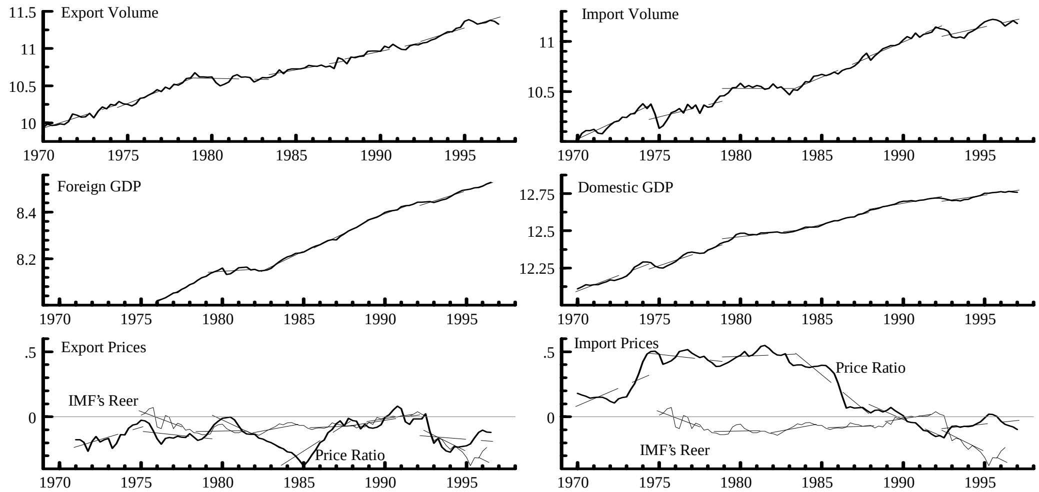


Figure 48: Trade, Income, and Prices - Italy

Italian Exports

Long-run Forces Measuring prices as the export price relative to foreign GDP deflator yields estimated elasticities for Italian exports that are not robust to the choice of lag. Systems with lag lengths greater than two do not exhibit cointegration. For the case exhibiting cointegration, the estimated price elasticity is -0.88 and the income elasticity is 1.6. Measuring prices with the IMF's real effective exchange rate yields one case with cointegration with a price elasticity of -0.4 and an income elasticity of 1.6.

Export Cointegration Results with *rpx*-Italy

	Number of lags Included							
	9	8	7	6	5	4	3	2
Cointegration Vectors	0	0	0	0	0	0	0	1
Price-Elasticity	ni	ni	ni	ni	ni	ni	ni	-0.88†
Income Elasticity	ni	ni	ni	ni	ni	ni	ni	1.62
Loading Coefficient	ni	ni	ni	ni	ni	ni	ni	-0.01
System's R^2	0.93	0.92	0.92	0.91	0.91	0.91	0.90	0.90
Export's Serial Corr.	0.09	0.04*	0.73	0.93	0.76	0.73	0.51	0.66
System's Serial Corr.	0.30	0.14	0.30	0.34	0.53	0.83	0.88	0.97

Export Cointegration Results with *reer*-Italy

	Number of lags Included							
	9	8	7	6	5	4	3	2
Cointegration Vectors	0	0	0	0	0	0	0	1
Price-Elasticity	ni	ni	ni	ni	ni	ni	ni	-0.51†
Income Elasticity	ni	ni	ni	ni	ni	ni	ni	1.65
Loading Coefficient	ni	ni	ni	ni	ni	ni	ni	-0.04
System's R^2	0.83	0.82	0.81	0.80	0.79	0.79	0.79	0.77
Export's Serial Corr.	0.35	0.22	0.47	0.57	0.70	0.83	0.89	0.74
System's Serial Corr.	0.07	0.14	0.58	0.74	0.93	0.94	0.96	0.99

ni: indicates that the elasticities are not identified.

The detailed cointegration results using rpx are (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.010 & (0.01) \\ -0.013 & (0.003) \\ -0.01 & (0.02) \end{pmatrix} \begin{pmatrix} 1 & -1.619 & 0.877 \\ (na) & (0.419) & (0.571) \end{pmatrix} \begin{pmatrix} x \\ fy \\ rpx \end{pmatrix}, \text{ 1976.3-1994.4}$$

whereas the cointegration results using the IMF's real effective exchange rate are

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.036 & (0.04) \\ 0.031 & (0.01) \\ -0.055 & (0.03) \end{pmatrix} \begin{pmatrix} 1 & -1.648 & 0.505 \\ (na) & (0.145) & (0.338) \end{pmatrix} \begin{pmatrix} x \\ fy \\ reer \end{pmatrix}, \text{ 1976.3-1994.4}$$

The results indicate that reliance on rpx gives comparable elasticity estimates to those obtained with $reer$ both in terms of magnitudes and statistical significance. The results also indicate, however, that the loading coefficients are not statistically significant which undermines the evidence on cointegration for both price measures.

Ex-post predictions for exports suggest comparable predictive performance for systems based on either rpx or $reer$ (figures 49 and 50). Figures 51 and 52 report the Chow tests for each measure of relative prices and the results indicate some parameter instability when relying on rpx and greater evidence of parameter constancy when using $reer$.

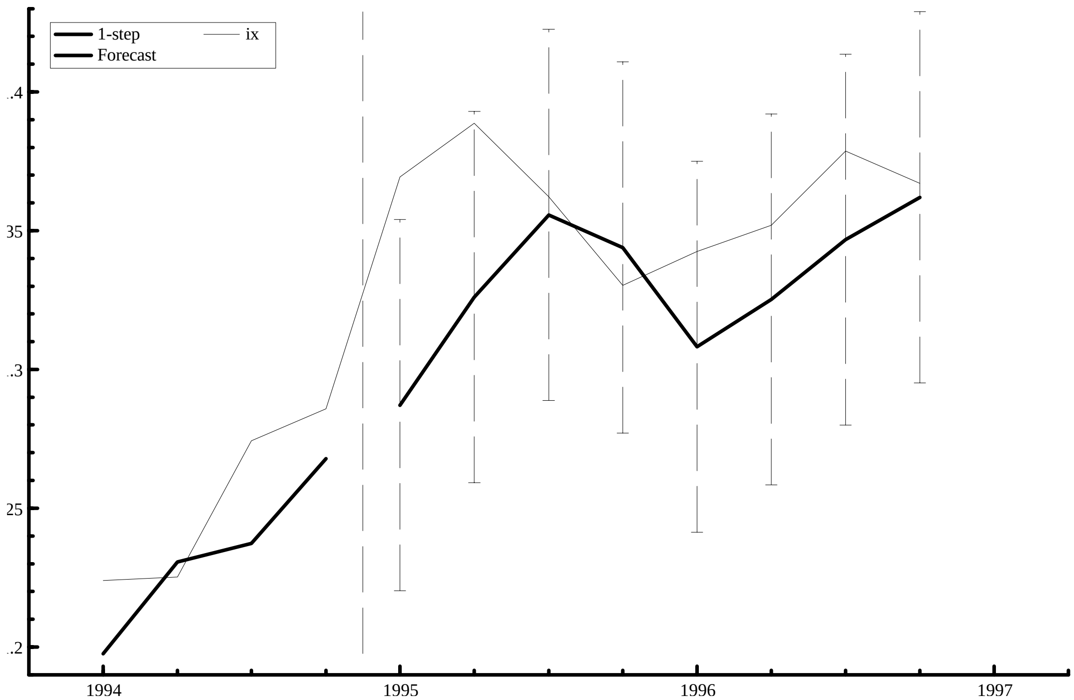


Figure 49: - Predictive Accuracy-Cointegration-Exports with *rpx*-Italy

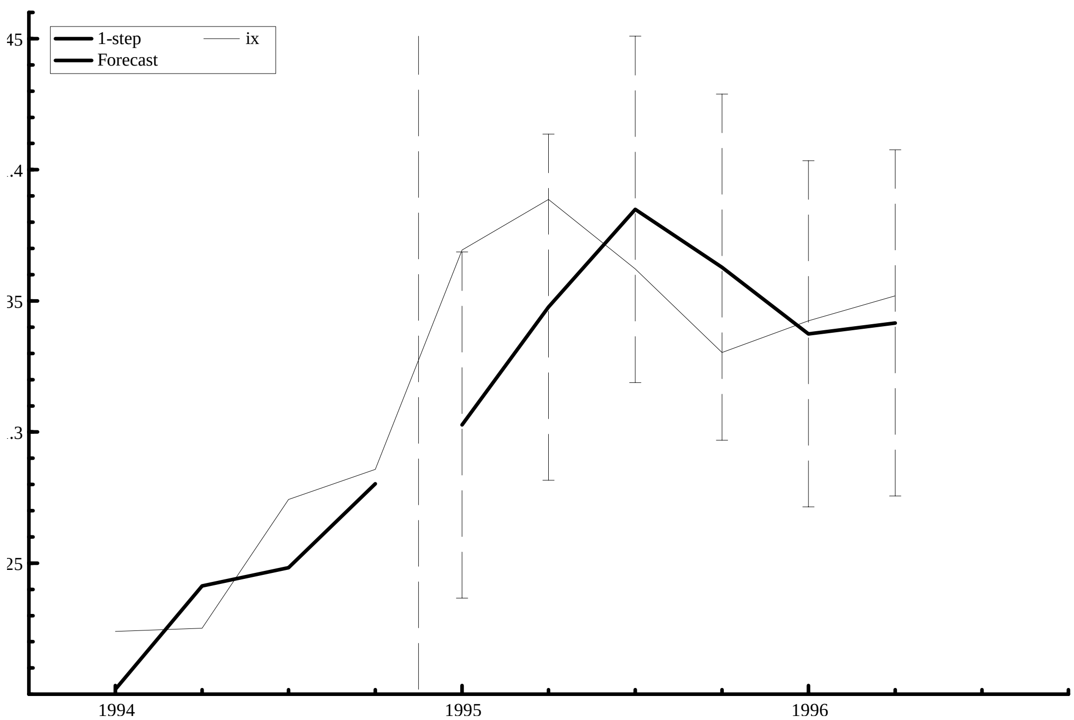


Figure 50: - Predictive Accuracy-Cointegration-Exports with *reer*-Italy

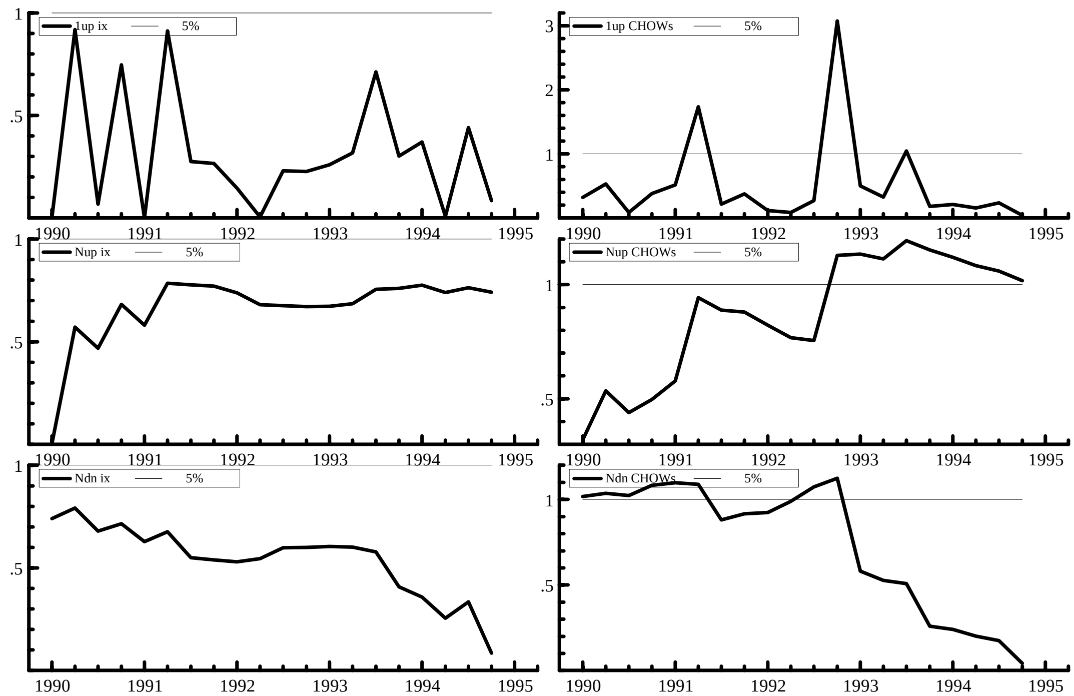


Figure 51: Chow Tests-Cointegration-Exports with *rpx*-Italy

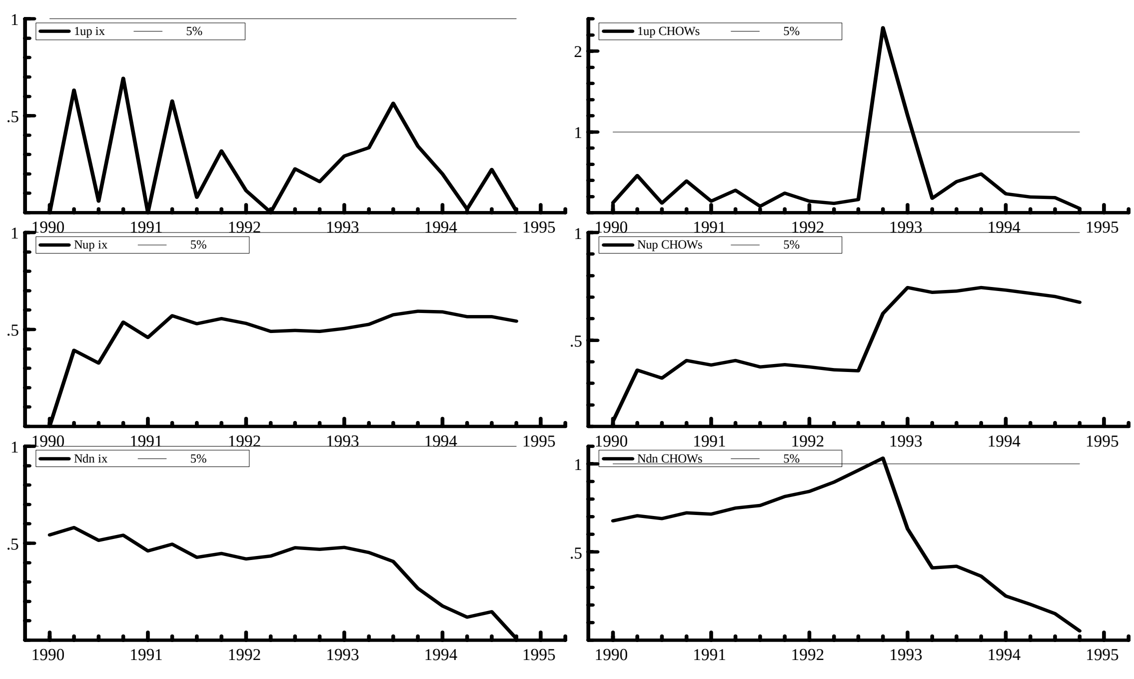


Figure 52: Chow Tests-Cointegration-Exports with *reer*-Italy

Short-run Forces To explain short-run fluctuations, we estimate the parameters of two error-correction models, one for each measure of relative prices. The estimation results using rpx are

$$\begin{array}{cccc} \Delta x_t = & +2.332\Delta fy_t & -0.325\Delta rpx_t & -0.029ECM_{x,t-1} \\ (se) & (0.60) & (0.13) & (0.04) \end{array},$$

where $ECM_x = x - 1.619 \cdot fy + 0.877 \cdot rpx + intercept$.

$R^2 = 0.33; SER = 3.34\%$	Null Hypothesis (p-value)	
Sample: 1976.3-1994.4	Serial-Independence (0.48)	Normality (0.23)
	Homoskedasticity (0.19)	Func. Form (0.76)

The estimation results using $reer$ are

$$\begin{array}{cccc} \Delta x_t = & +0.994\Delta fy_t & -0.224\Delta reer_t & -0.078ECM_{x,t-1} \\ (se) & (0.29) & (0.08) & (0.02) \end{array},$$

where $ECM_x = x - 1.648 \cdot fy + 0.505 \cdot reer + intercept$.

$R^2 = 0.33; SER = 3.28\%$	Null Hypothesis (p-value)	
Sample: 1976.3-1994.4	Serial-Independence (0.58)	Normality (0.18)
	Homoskedasticity (0.30)	Func. Form (0.49)

Both formulations explain one third of the variability of the growth rate of exports with the empirical distributions of the residuals satisfying the assumptions maintained for estimation. Notice, however, that the error-correction coefficient for the formulation using rpx is not significantly different from zero but it is for $reer$.

The predictions of the errors of the formulation using rpx are comparable to those of the system relying on $reer$ (figures 53 and 54). Similarly, support for the hypothesis of parameter constancy is strong and robust to the choice of measure of relative prices.

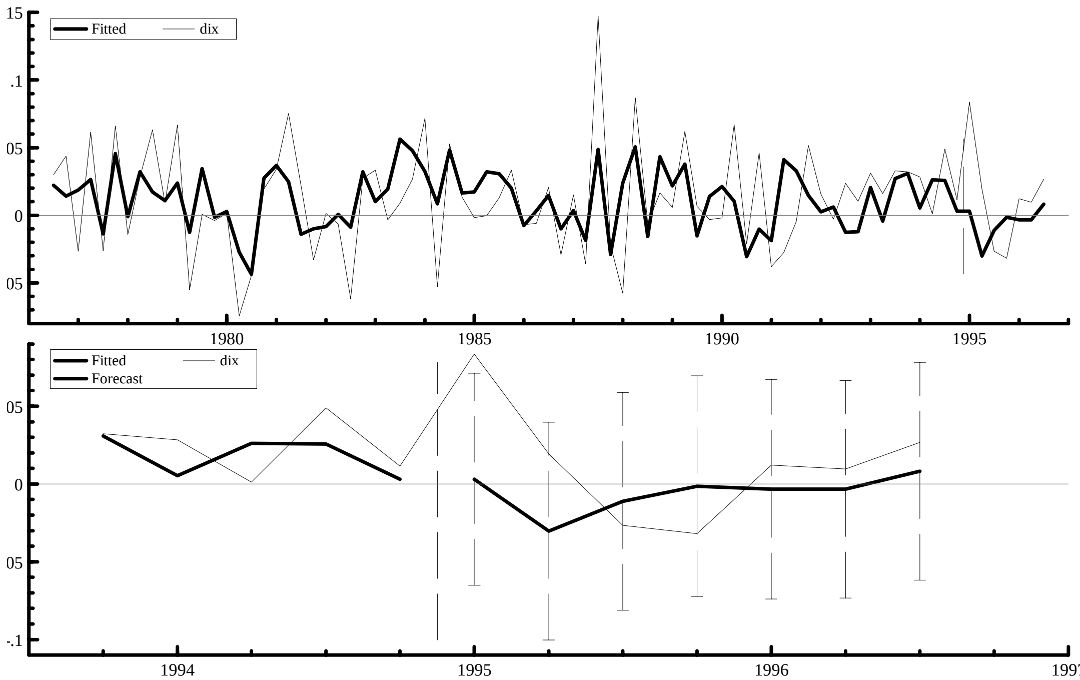


Figure 53: Predictive Accuracy-ECM- Exports with *rpx*-Italy

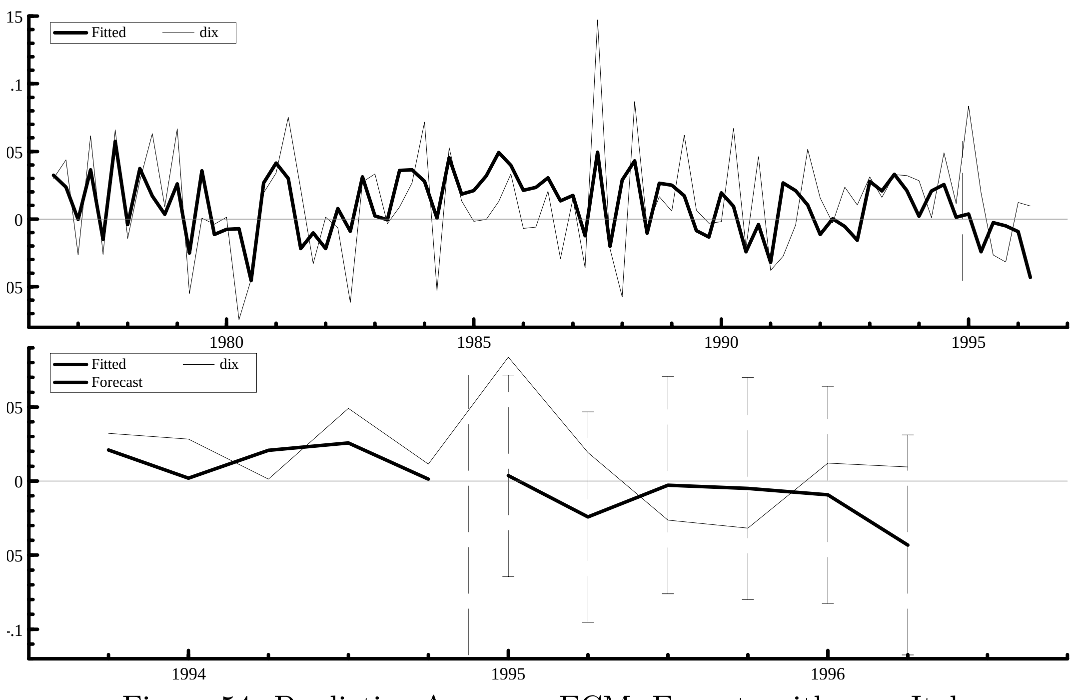


Figure 54: Predictive Accuracy-ECM- Exports with *reer*-Italy

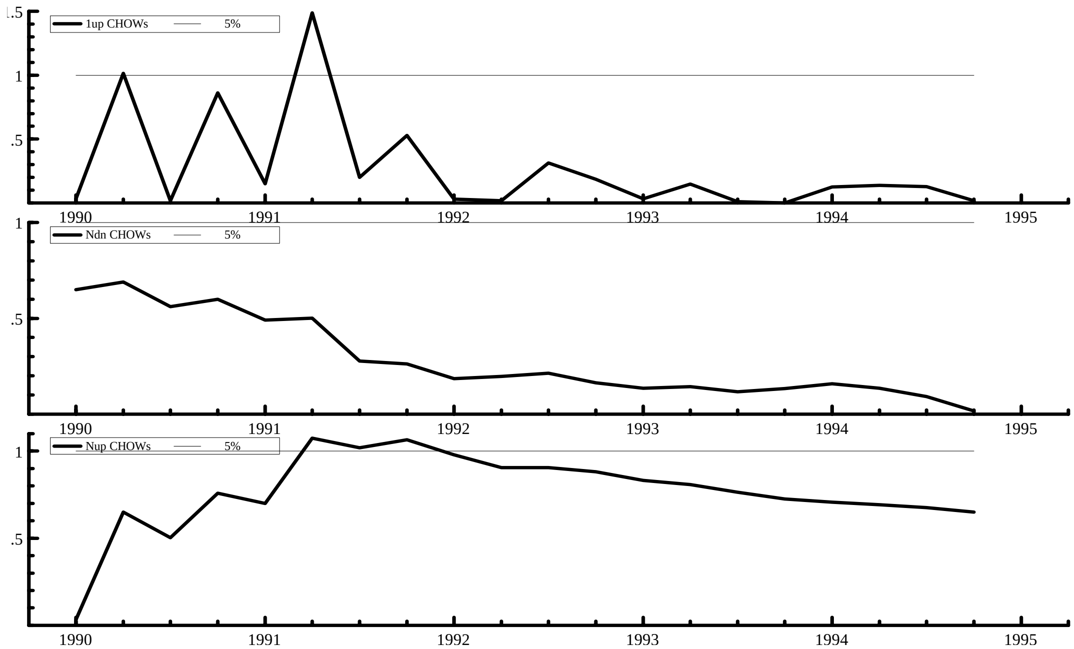


Figure 55: Chow Tests-ECM-Exports with rpx-Italy

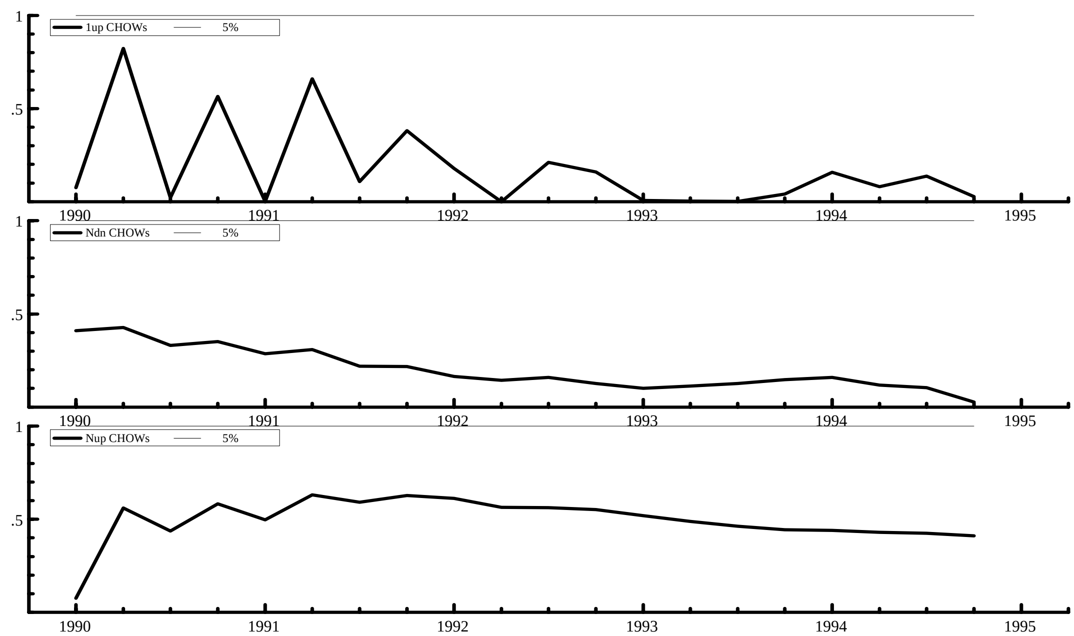


Figure 56: Chow Tests-ECM-Exports with $reer$-Italy

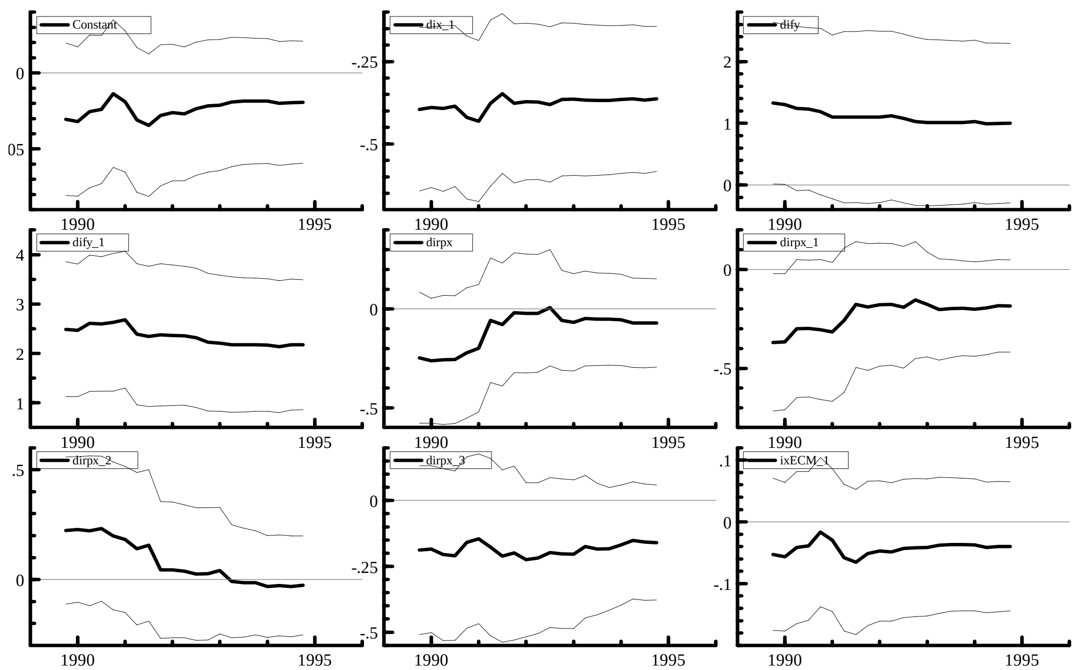


Figure 57: 95% Bands for Coefficients of ECM-Exports with rpx-Italy

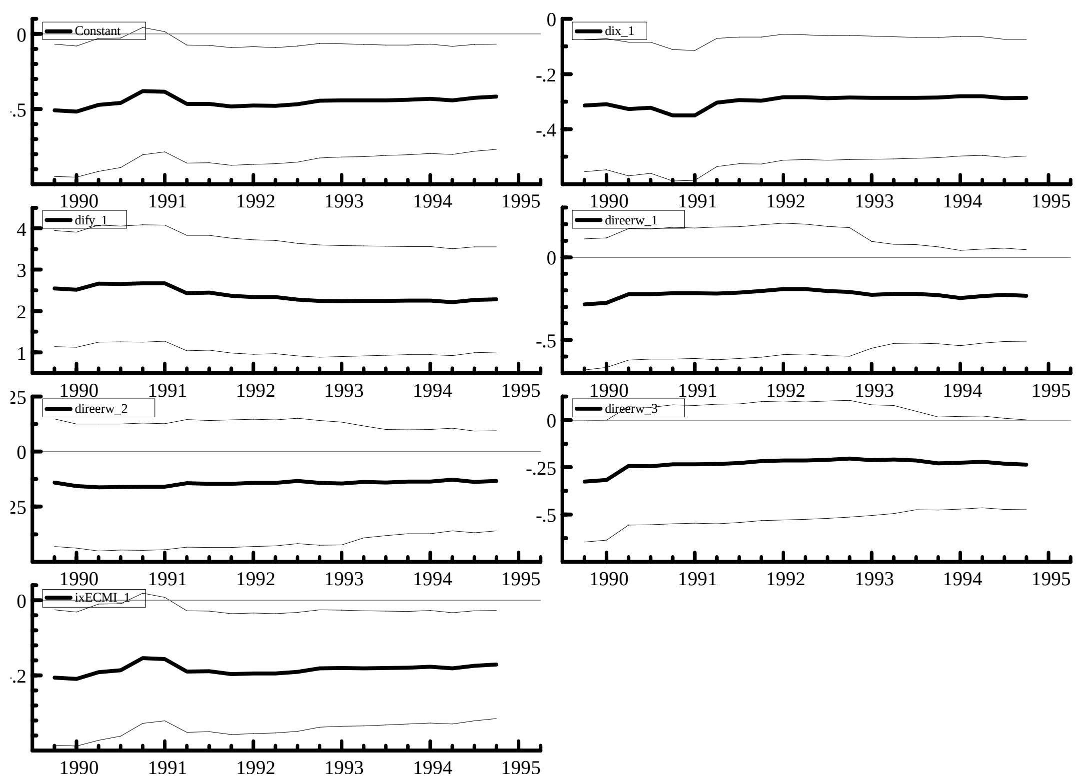


Figure 58: 95% Bands for Coefficients of ECM-Exports with $reer$-Italy

Italian Imports

Long-run Forces Measuring prices with *rpm* yields unique cointegration vectors for specifications with lag lengths fewer than six quarters. For these cases, the price elasticity ranges from -0.40 to -0.63 and the income elasticity ranges 0.8 to 1.4. Measuring prices with *reer* reveals no cointegration.

Import Cointegration-Results with *rpm*-Italy

	Number of lags Included							
	9	8	7	6	5	4†	3	2
Cointegration Vectors	2	2	2	2	1	1	1	1
Price-Elasticity	ni	ni	ni	ni	-0.63	-0.40	-0.39	-0.39
Income Elasticity	ni	ni	ni	ni	0.78	1.40	1.42	1.42
Loading Coefficient	ni	ni	ni	ni	-0.06	-0.50	-0.53	-0.48
System's R^2	0.87	0.87	0.85	0.85	0.84	0.83	0.83	0.82
Import's Serial Corr.	0.46	0.54	0.83	0.77	0.17	0.82	0.88	0.73
System's Serial Corr.	0.16	0.07	0.22	0.48	0.17	0.11	0.02*	0.02*

Import Cointegration Results with *reer*-Italy

	Number of lags Included							
	9	8	7	6	5	4	3	2
Cointegration Vectors	0	0	0	0	0	0	0	0
Price-Elasticity	ni	ni	ni	ni	ni	ni	ni	ni
Income Elasticity	ni	ni	ni	ni	ni	ni	ni	ni
Loading Coefficient	ni	ni	ni	ni	ni	ni	ni	ni
System's R^2	0.88	0.88	0.86	0.86	0.84	0.83	0.83	0.82
Import's Serial Corr.	0.20	0.42	0.07	0.24	0.64	0.69	0.75	0.83
System's Serial Corr.	0.40	0.67	0.36	0.65	0.01	0.41	0.47	0.37

ni: indicates that the elasticities are not identified.

Because the system with five lags has an statistically insignificant loading coefficient, we use *rpm* with four lags (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.497 & (0.12) \\ -0.0329 & (0.03) \\ -0.088 & (0.10) \end{pmatrix} \begin{pmatrix} 1 & -1.397 & 0.404 \\ (na) & (0.054) & (0.040 \end{pmatrix} \begin{pmatrix} m \\ y \\ rpm \end{pmatrix}, \text{ 1972.2-1994.4.}$$

Though the ex-post predictions are one-sided (figure 59), the Chow tests cannot reject the hypothesis of parameter constancy (figure 60).

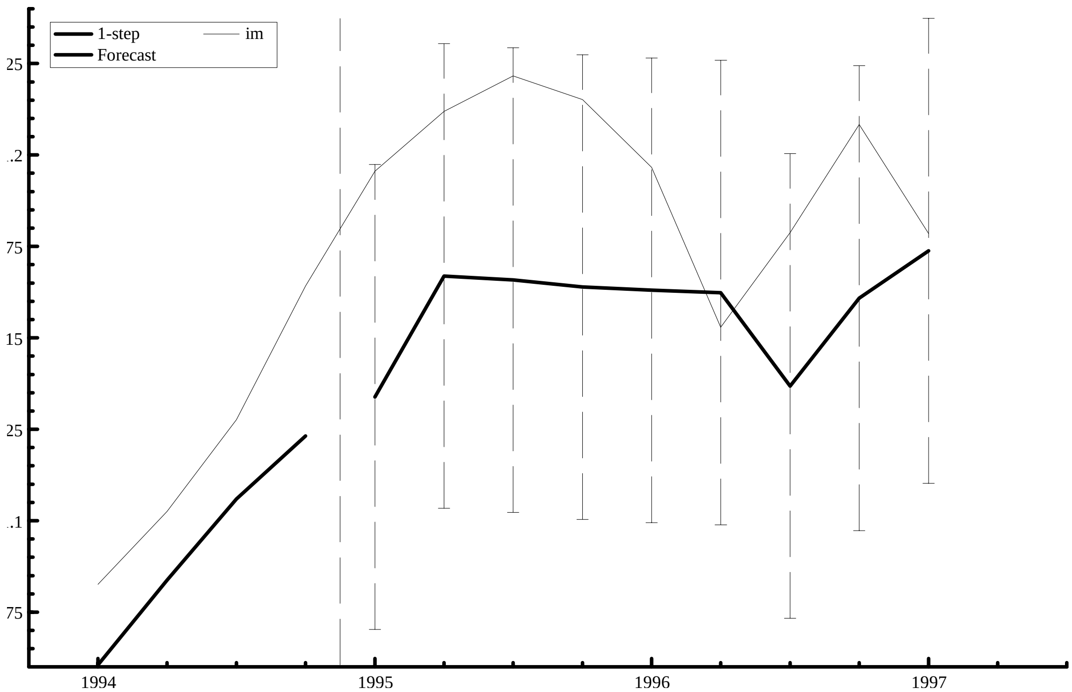


Figure 59: Predictive Accuracy-Cointegration-Imports-Italy

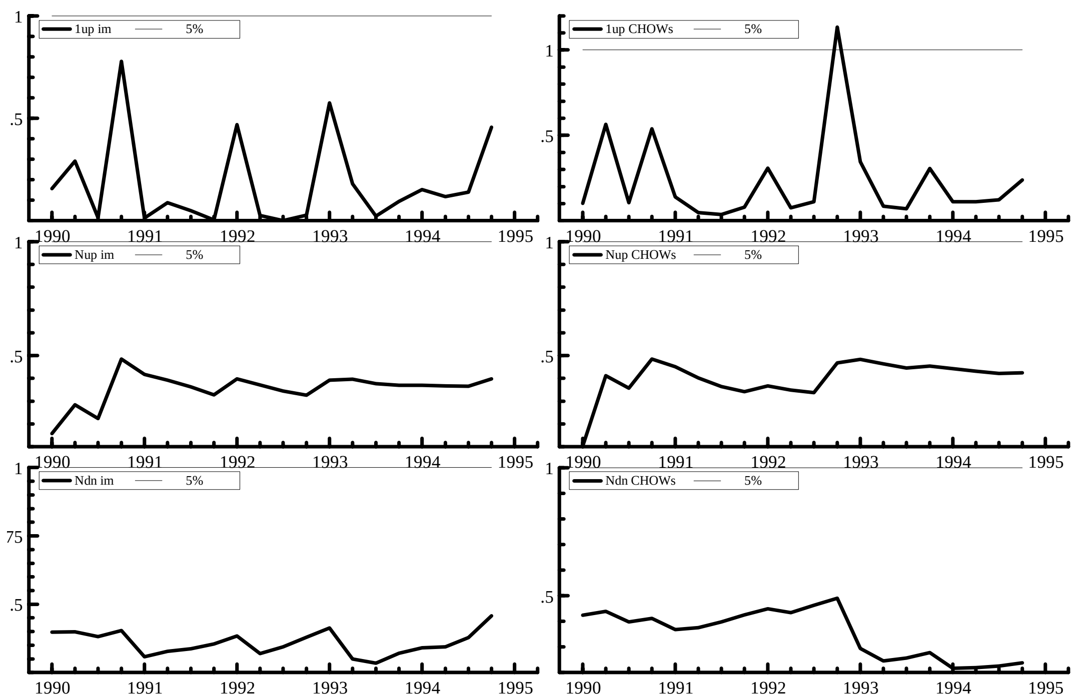


Figure 60: Chow Tests-Cointegration-Imports-Italy

Short-run Forces To explain short-run fluctuations, we estimate the parameters of an error-correction model; the estimation results are

$$\begin{array}{lll} \Delta m_t = & +1.012\Delta y_t & -0.3636 ECM_{m,t-1} \\ (se) & (0.37) & (0.06) \end{array}$$

where $ECM_m = m - 1.3968 \cdot y + 0.404 \cdot rpm$.

	Null Hypothesis (p-value)	
$R^2 = 0.42; SER = 3.02\%$		
Sample: 1970.3-1994.4	Serial-Independence (0.08)	Normality (0.34)
	Homoskedasticity (0.56)	Func. Form (0.04*)

The model explains less than half of the variability of the growth rate of imports and the empirical distribution of the residuals satisfies the assumptions maintained for estimation except the functional form. The prediction errors tend to be one-sided (figure 61). Finally, the Chow tests (figure 62) and the 95% confidence bands for the coefficient estimates (figure 63) suggest that one cannot reject the hypothesis of parameter constancy.

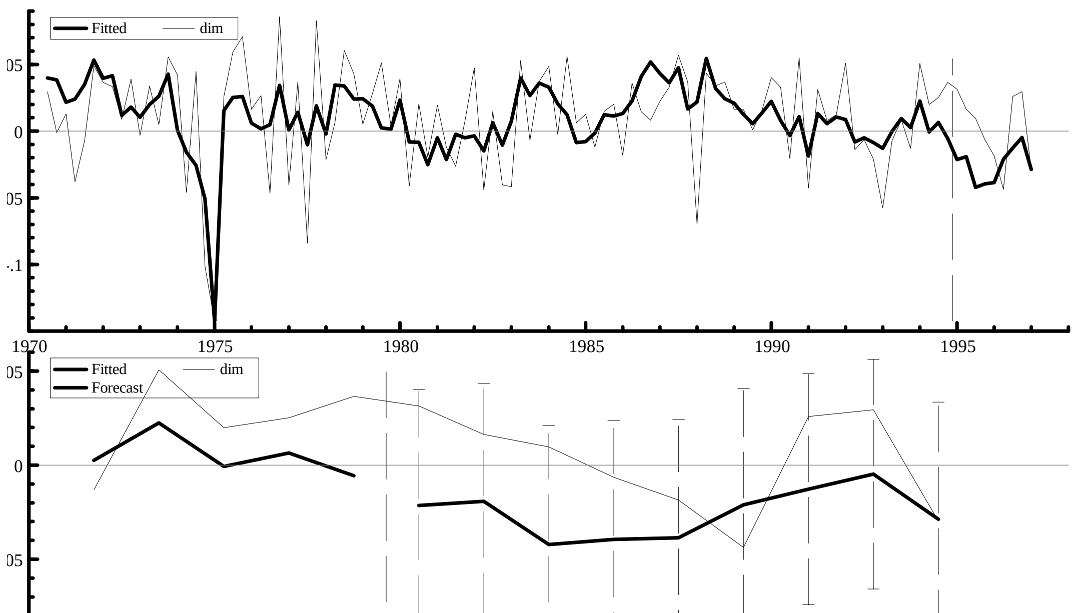


Figure 61: Predictive Accuracy-ECM- Imports-Italy

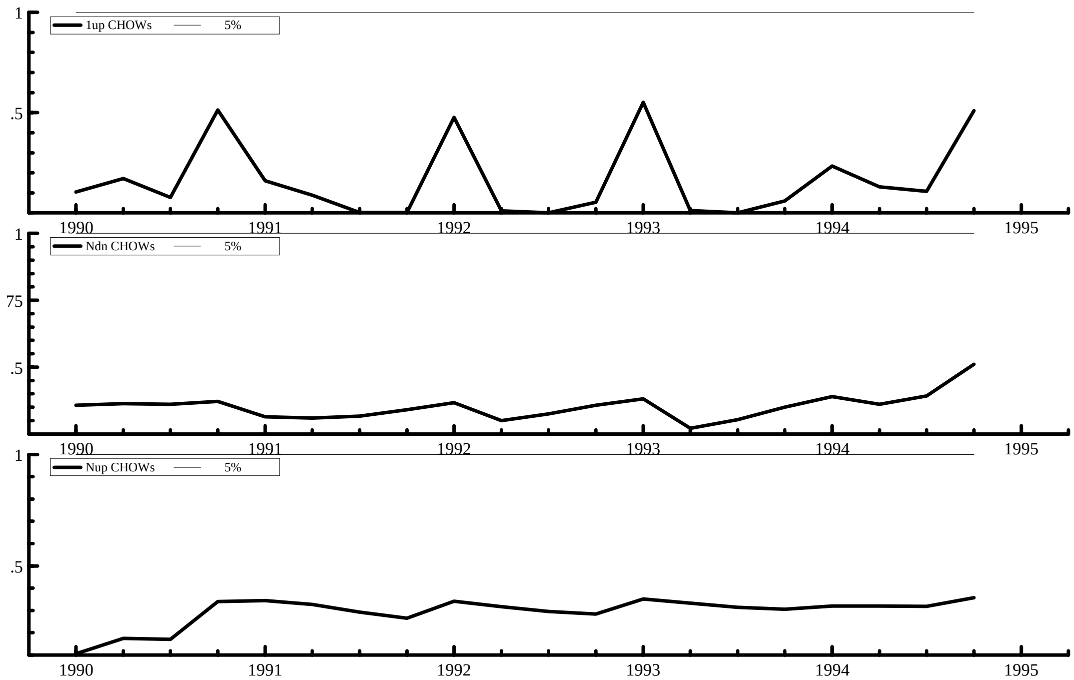


Figure 62: Chow Tests - ECM- Imports-Italy

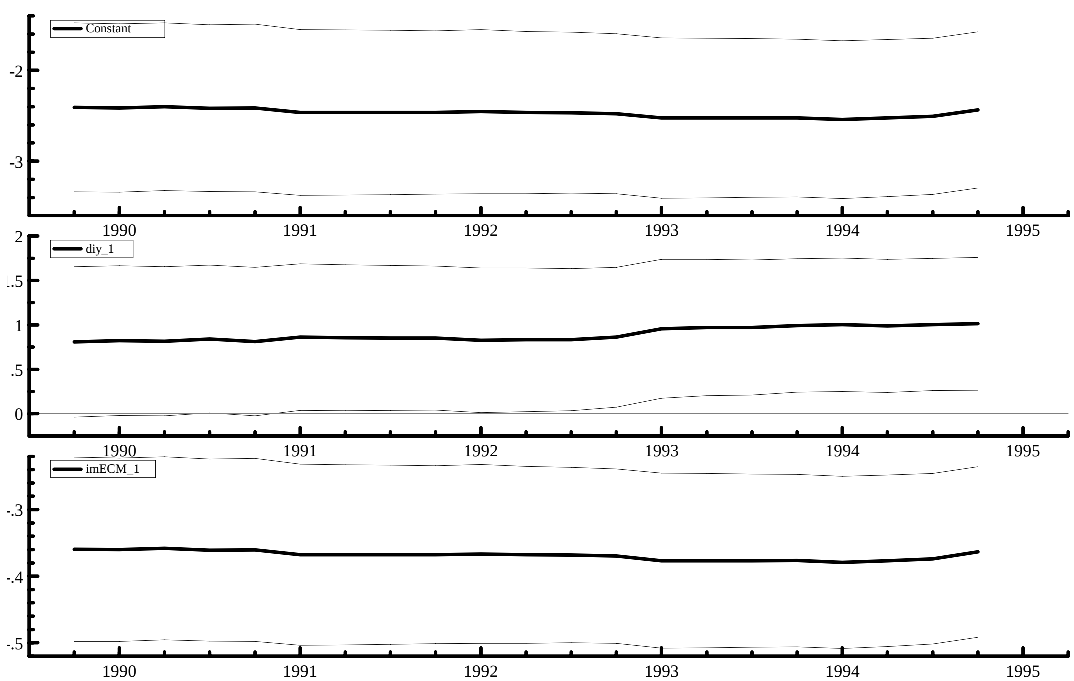


Figure 63: 95%Coefficient Bands for ECM Coefficients-Imports-Italy

I Japanese Trade

The evolution of Japan's foreign trade and its proximate determinants are displayed in figure 64. Both export and import volumes grow over time along with foreign and domestic income. Relative prices for imports show a downward trend to a greater extent than exports' relative price downward trend. In addition, the IMF's real effective exchange rate shows an upward trend which contrasts with the downward trend in the relative price of imports.

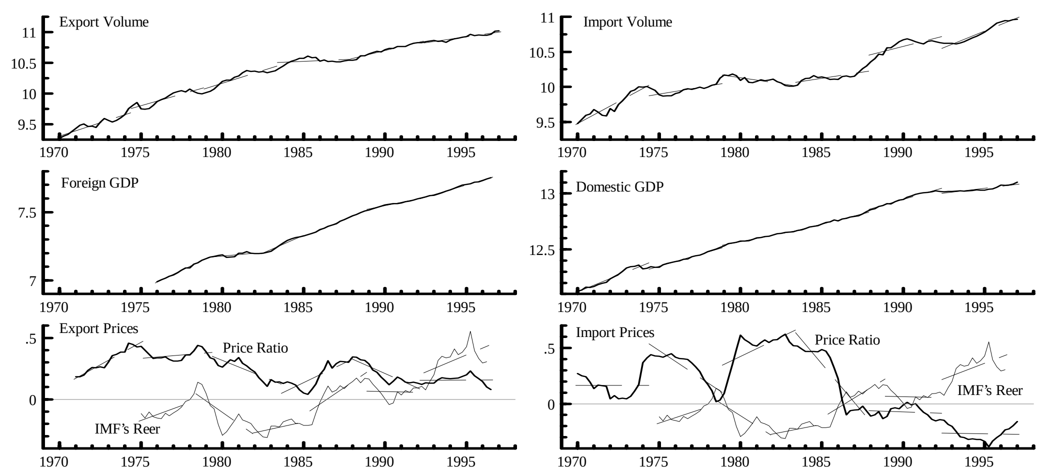


Figure 64: Trade, Income, and Prices-Japan

Japanese Exports

Long-run Forces For a given price measure, the choice of lag matters for the estimated elasticities; the sensitivity is greater for the price elasticity than for the income elasticity. For a given lag, the choice of price measure also noticeably affects the estimated elasticities.

Export Cointegration Results with *rpx*-Japan

	Number of lags Included							
	9	8	7	6	5†	4	3	2
Cointegration Vectors	1	0	0	0	1	1	1	0
Price-Elasticity	1.31	ni	ni	ni	-1.01	-1.51	-1.32	ni
Income Elasticity	1.21	ni	ni	ni	1.12	0.66	0.94	ni
Loading Coefficient	-0.04	ni	ni	ni	-0.13	-0.00	-0.02	ni
System's R^2	0.97	0.97	0.97	0.96	0.96	0.96	0.96	0.95
Export's Serial Corr.	0.00*	0.00*	0.02*	0.81	0.16	0.21	0.25	0.04*
System's Serial Corr.	0.01*	0.00*	0.00*	0.27	0.51	0.70	0.73	0.14

Export Cointegration Results with *reer*-Japan

	Number of lags Included							
	9	8	7	6	5	4	3	2
Cointegration Vectors	1	0	0	0	0	0	1	1
Price-Elasticity	-0.68	ni	ni	ni	ni	ni	-0.68	-1.80
Income Elasticity	1.75	ni	ni	ni	ni	ni	1.54	2.81
Loading Coefficient	0.10	ni	ni	ni	ni	ni	-0.02	0.01
System's R^2	0.93	0.92	0.91	0.91	0.91	0.90	0.90	0.89
Import's Serial Corr.	0.20	0.07	0.76	0.33	0.18	0.31	0.89	0.16
System's Serial Corr.	0.07	0.03*	0.02*	0.11	0.23	0.83	0.96	0.60

ni: indicates that the elasticities are not identified.

Reliance on *reer* yields several candidate formulations, but the estimated loading coefficients in most of them are either small in magnitude or statistically insignificant. The detailed cointegration results using *rpx* are (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.129 & (0.08) \\ -0.026 & (0.02) \\ -0.214 & (0.01) \end{pmatrix} \begin{pmatrix} 1 & -1.1321 & 1.004 \\ (na) & (0.058) & (0.121 \end{pmatrix} \begin{pmatrix} x \\ fy \\ rpx \end{pmatrix}, \text{1977.2-1994.4}$$

The results indicate that reliance on the conventional measure of relative prices (rpx) gives significant estimates for income and prices, which are also not significantly different from one. The error-correction term is not, however, significant.

Figure 65 shows that the ex-post predictions are reasonably accurate. Figure 66 displays the Chow tests which suggest that one cannot reject the hypothesis of parameter constancy.

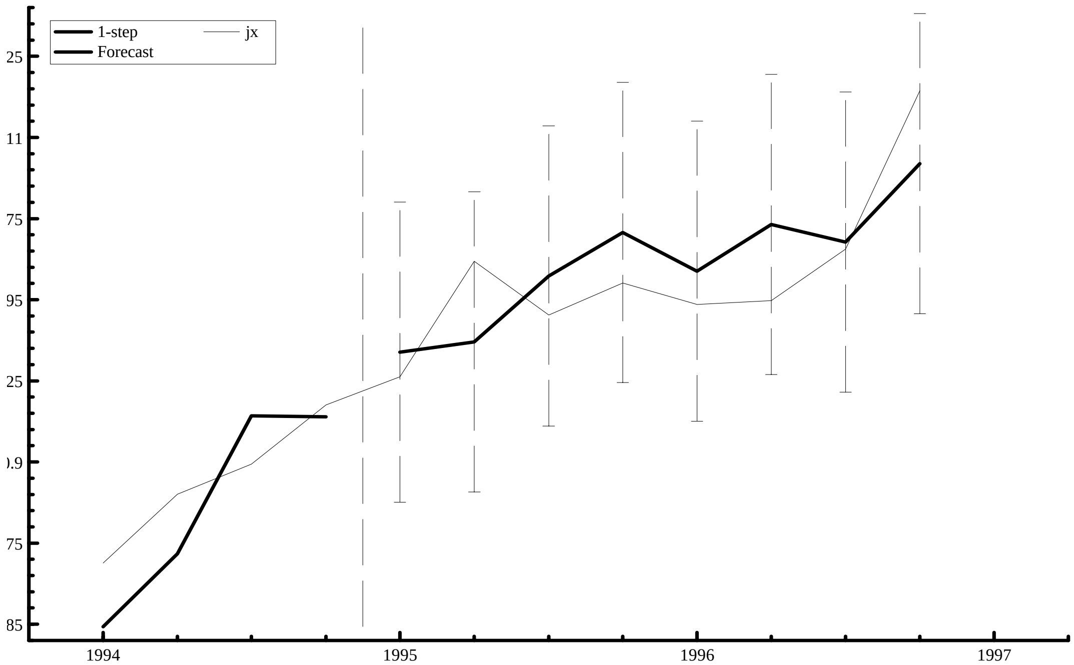


Figure 65: Predictive Accuracy-Cointegration-Exports-Japan

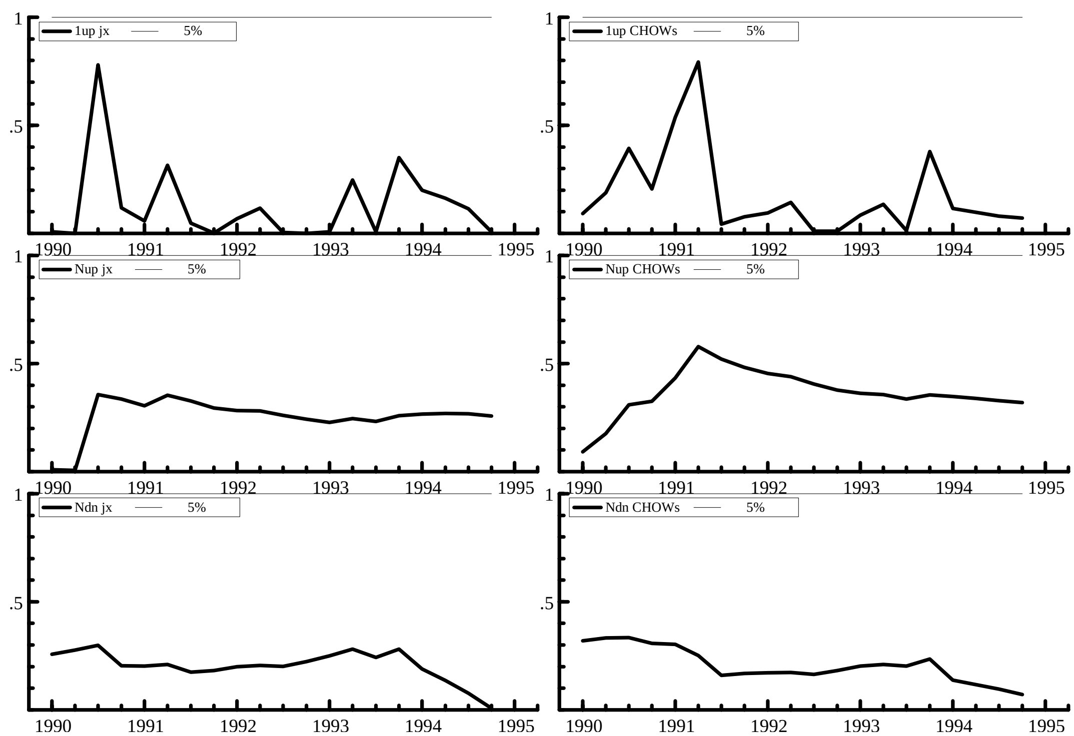


Figure 66: Chow Tests-Cointegration-Exports-Japan

Short-run Forces To explain short-run fluctuations, we estimate the parameters of an error-correction model using rpx :

$$\begin{array}{cccc} \Delta x_t = & +0.590\Delta fy_t & -0.453\Delta rpx_t & -0.169ECM_{x,t-1} \\ (se) & (0.34) & (0.12) & (0.06) \end{array},$$

where $ECM_x = x - 1.121 \cdot fy + 1.001 \cdot rpx + intercept.$

$R^2 = 0.49; SER = 2.04\%$	Null Hypothesis (p-value)	
Sample: 1977.3-1994.4	Serial-Independence (0.06)	Normality (0.93)
	Homoskedasticity (0.99)	Func. Form (0.87)

The formulation explains about half of the variability of the growth rate of exports and the empirical distribution of the residuals satisfies the assumptions maintained for estimation. Moreover, the error-correction coefficient is significantly different from zero which strengthens the evidence on cointegration among those variables.

The predictions of the error-correction model tend to be one-sided (figure 67). Finally, the Chow tests (figure 68) and the 95% confidence bands for the ECM coefficient estimates (figure 69) suggest that one cannot reject the hypothesis of parameter constancy.

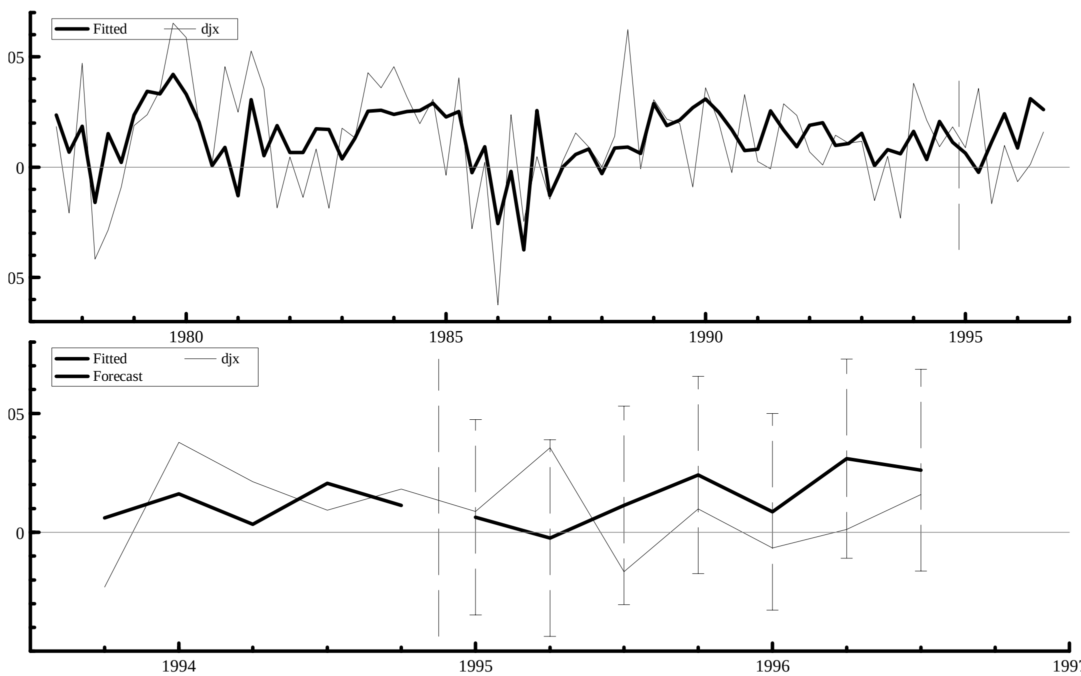


Figure 67: Predictive Accuracy-ECM-Exports-Japan

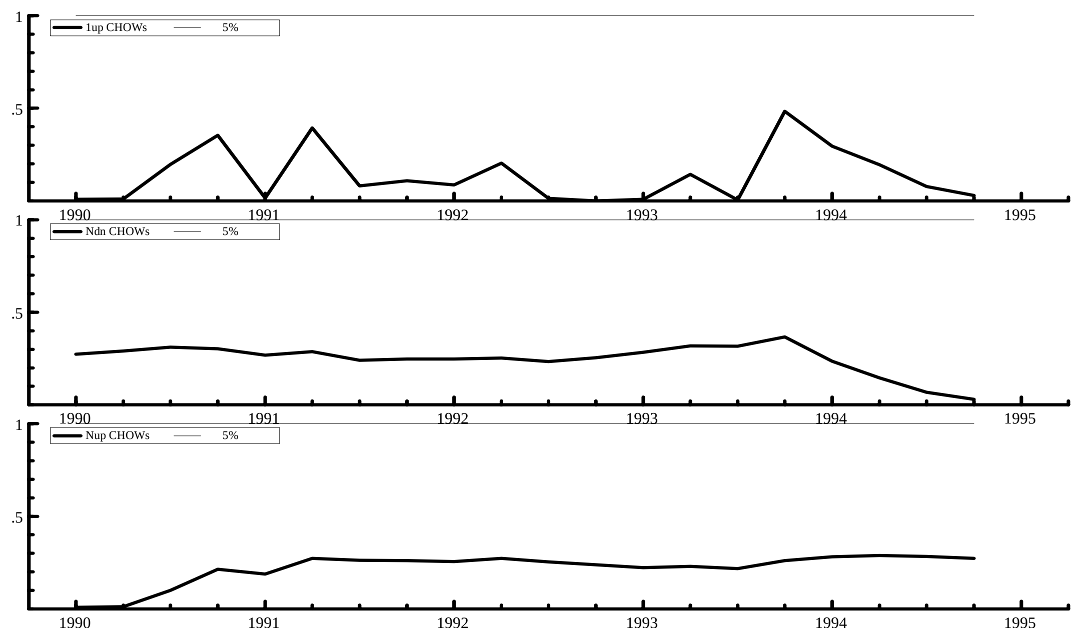


Figure 68: Chow Tests-ECM-Exports-Japan

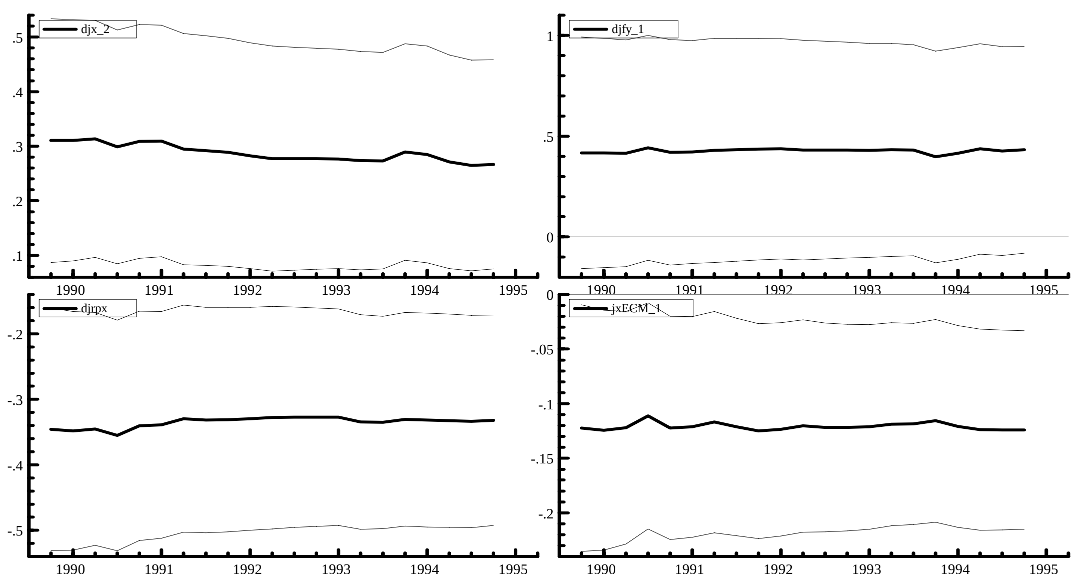


Figure 69: 95% Confidence Bands for ECM Coefficients-Exports-Japan

Japanese Imports

Long-run Forces For a given measure of prices, the number of cointegration vectors and the attending elasticity estimates are sensitive to the number of lags included in the specification. Using rpm gives multiple cointegration vectors except for the case of six lags; even in this case the systems's residuals lack serial independence.

Import Cointegration Results with *rpm*-Japan

	Number of lags Included							
	9	8	7	6†	5	4	3	2
Cointegration Vectors	2	2	2	1	2	2	2	2
Price-Elasticity	ni	ni	ni	-0.31	ni	ni	ni	ni
Income Elasticity	ni	ni	ni	0.96	ni	ni	ni	ni
Loading Coefficient	ni	ni	ni	-0.12	ni	ni	ni	ni
System's R^2	0.99	0.99	0.99	0.98	0.98	0.98	0.98	0.98
Import Serial Corr.	0.60	0.36	0.63	0.22	0.10	0.00*	0.00*	0.01*
System's Serial Corr.	0.22	0.12	0.27	0.03*	0.26	0.00*	0.00*	0.00*

Import Cointegration Results with *reer*-Japan

	Number of lags Included							
	9	8	7	6	5	4	3	2
Cointegration Vectors	1	1	1	0	0	0	1	2
Price-Elasticity	-11.03	-6.61	-1.16	ni	ni	ni	-0.95	ni
Income Elasticity	4.09	3.93	0.97	ni	ni	ni	1.47	ni
Loading Coefficient	–0.00	-0.00	-0.11	ni	ni	ni	-0.03	ni
System's R^2	0.96	0.96	0.96	0.95	0.95	0.95	0.95	0.94
Import's Serial Corr.	0.63	0.50	0.78	0.78	0.92	0.44	0.44	0.34
System's Serial Corr.	0.92	0.98	0.98	0.83	0.86	0.44	0.56	0.02*

ni: indicates that the elasticities are not identified.

The cointegration results using *rpm* are (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.025 & (0.016) \\ -0.023 & (0.01) \\ 0.046 & (0.02) \end{pmatrix} \begin{pmatrix} 1 & -0.917 & 0.326 \\ (na) & (0.119) & (0.209 \end{pmatrix} \begin{pmatrix} m \\ y \\ rpm \end{pmatrix}, \text{ 1956.4-1994.4.}$$

Figure 70 shows the ex-post predictions which are reasonably accurate. Figure 71 displays the Chow tests which suggest that one cannot reject the hypothesis of parameter constancy.

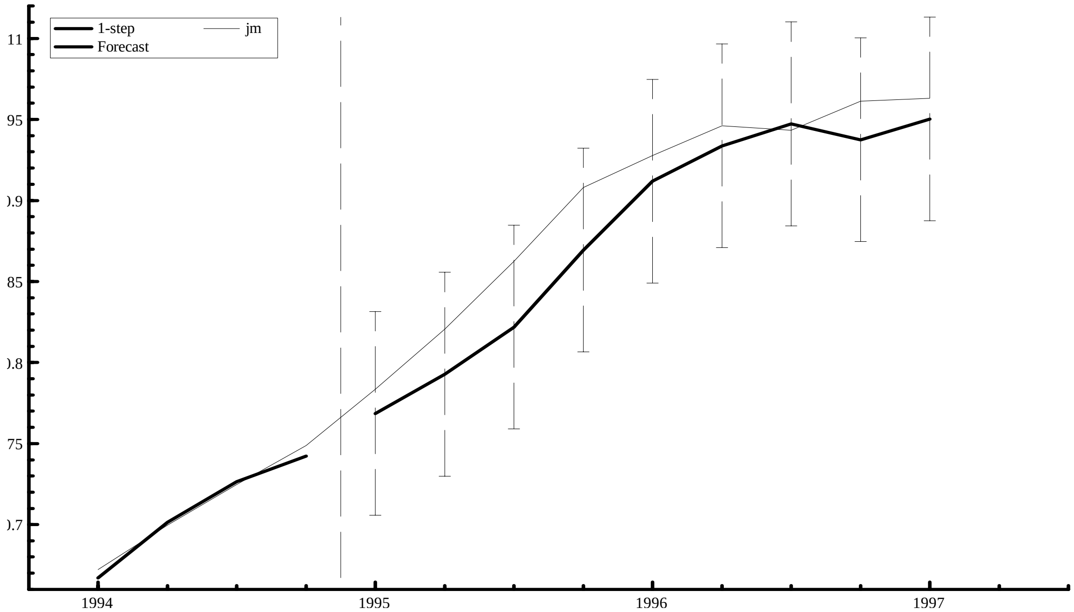
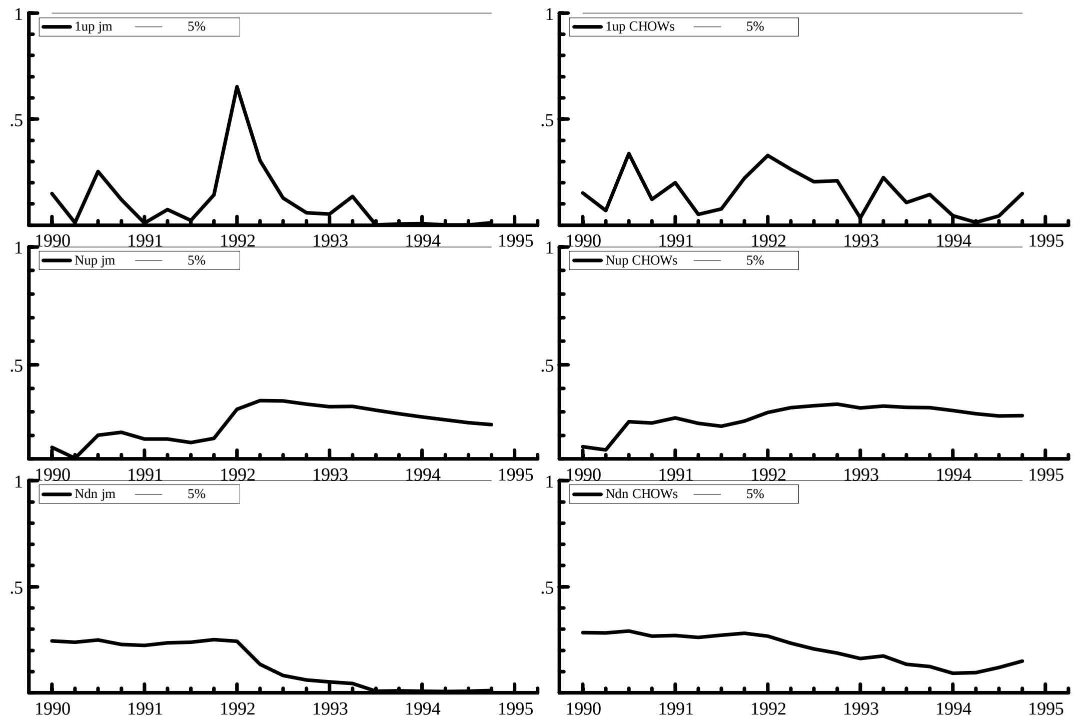


Figure 70: Predictive Accuracy-Cointegration-Imports-Japan



Figures 71: Chow Tests-Cointegration-Imports-Japan

Short-run Forces To explain short-run fluctuations, we estimate the parameters of an error-correction model including two dummy variables that take a value of one in 1972.1 and 1989.1. The estimation results are

$$\begin{array}{cccc} \Delta m_t = & +1.004\Delta y_t & -0.0456\Delta rpm_t & -0.03262 ECM_{m,t-1} \\ (se) & (0.37) & (0.09) & (0.017) \end{array}$$

where $ECM_m = m - 0.91713 \cdot y + 0.32626 \cdot rpm + intercept.$

$R^2 = 0.52; SER = 2.91\%$	Null Hypothesis (p-value)	
Sample: 1956.4-1994.4	Serial-Independence (0.28)	Normality (0.57)
	Homoskedasticity (0.21)	Func. Form (0.09)

The model explains about half of the variability of the growth rate of imports and the empirical distribution of the residuals satisfies the assumptions maintained for estimation. The predictions of the model are, however, one-sided (figure 72). Finally, the Chow tests (figure 73) and the 95% confidence bands for the ECM coefficient estimates (figure 74) suggest that one cannot reject the hypothesis of parameter constancy.

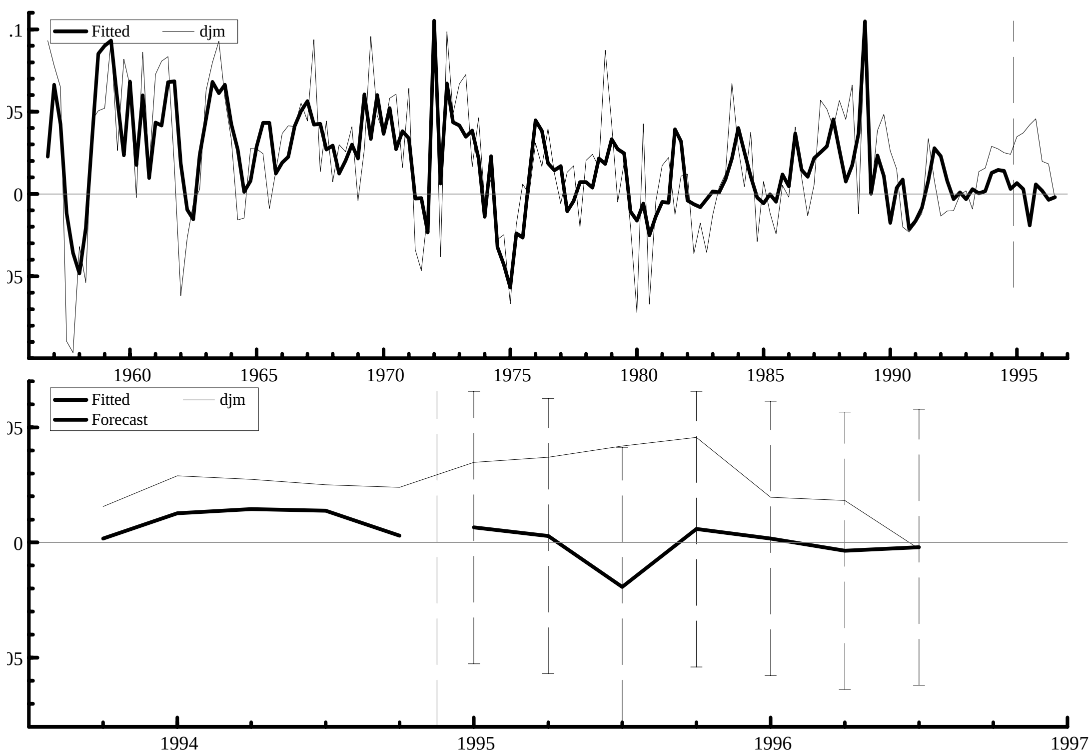


Figure 72: Predictive Accuracy-ECM-Imports-Japan

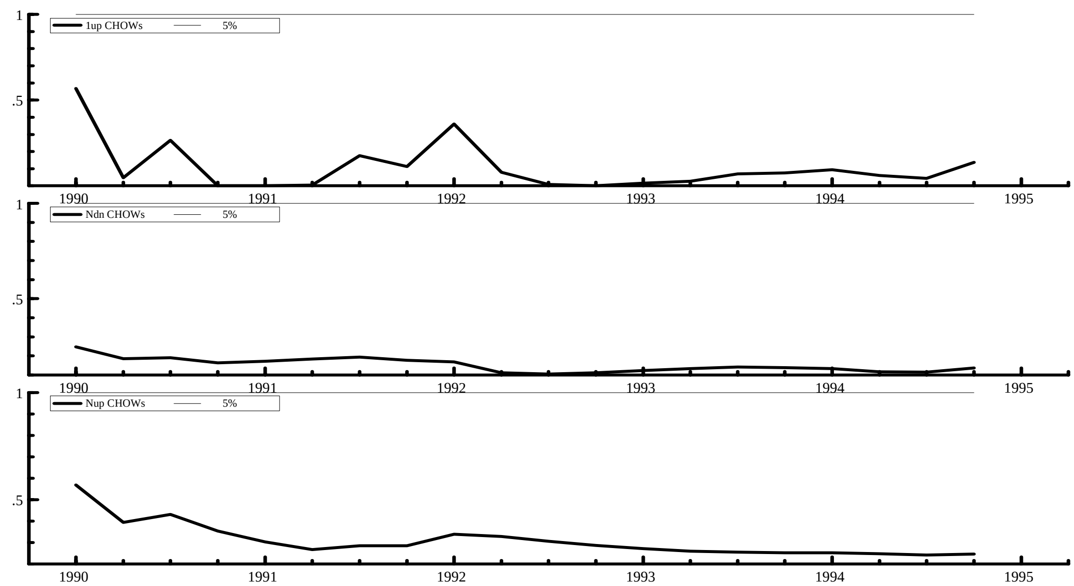


Figure 73: Chow Tests-ECM-Imports-Japan

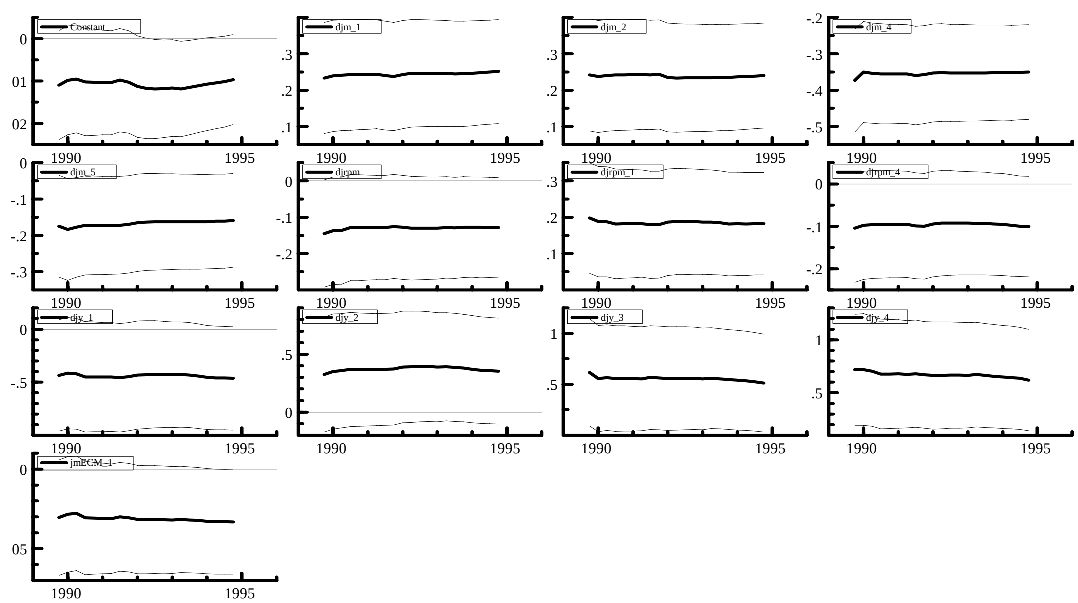


Figure 74: 95% Confidence Bands for ECM Coefficients-Imports-Japan

J U.K. Trade

The evolution of U.K's foreign trade and its proximate determinants are displayed in figure 75. Both export and import volumes grow over time along with foreign and domestic income. Relative prices for both exports and imports show a downward trend but the decline in the relative price of imports is more pronounced than that of exports' relative price. In addition, the IMF's real effective exchange rate shows smaller volatility than the conventional measure of relative prices based on deflators from the National Income Accounts.

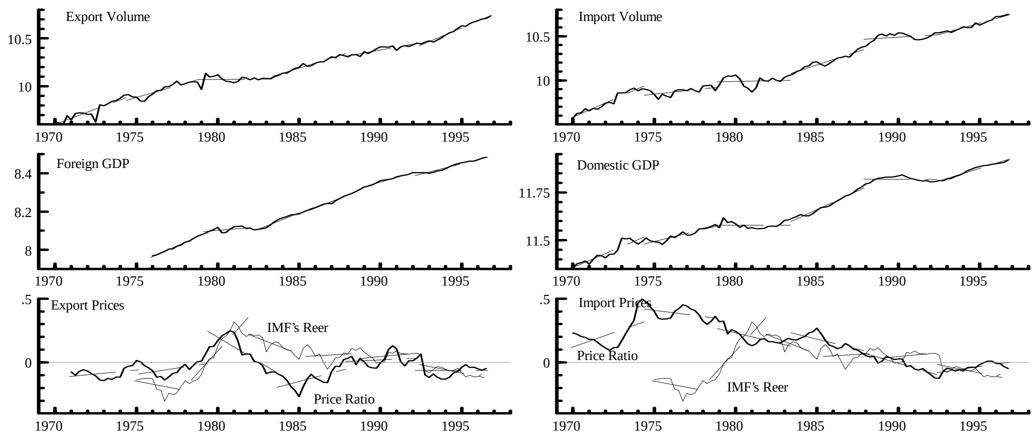


Figure 75: Trade, Income, and Prices-United Kingdom

U.K. Exports

Long-run Forces Measuring prices as the export price relative to foreign GDP deflator yields estimated elasticities for British exports that are not robust to the choice of lag. Though all of the systems exhibit cointegration, only the ones with fewer than five lags have negative price elasticities; for these cases, the vector of residuals exhibits serial correlation. Measuring prices with the IMF's effective real exchange rate yields comparable results; formulations without serial correlation exhibit relatively low price elasticities.

Export Cointegration Results with rpx-United Kingdom

	Number of lags Included							
	9	8	7	6	5	4†	3	2
Cointegration Vectors	1	1	1	1	1	1	1	1
Price-Elasticity	1.17	0.60	0.49	0.88	1.06	-1.55	-3.41	-1.02
Income Elasticity	0.96	1.11	1.18	1.16	1.16	1.10	1.18	2.28
Loading Coefficient	0.18	0.30	0.25	0.04	0.03	-0.09	-0.08	-0.06
System's R^2	0.86	0.85	0.82	0.79	0.79	0.78	0.76	0.76
Export's Serial Corr.	0.84	0.55	0.12	0.05	0.27	0.00*	0.18	0.37
System's Serial Corr.	0.90	0.74	0.50	0.16	0.17	0.00*	0.00*	0.00*

Export Cointegration Results with *reer*-United Kingdom

	Number of lags Included							
	9	8	7	6	5	4	3	2†
Cointegration Vectors	1	1	0	1	1	0	0	1
Price-Elasticity	-0.16	-0.15	ni	-0.07	-0.05	ni	ni	-1.11
Income Elasticity	1.26	1.26	ni	1.26	1.25	ni	ni	1.09
Loading Coefficient	0.10	0.06	ni	-0.08	-0.01	ni	ni	-0.03
System's R^2	0.82	0.81	0.80	0.78	0.77	0.74	0.72	0.72
Export's Serial Corr.	0.84	0.76	0.65	0.39	0.61	0.64	0.42	0.83
System's Serial Corr.	0.03*	0.25	0.86	0.18	0.06	0.00*	0.00*	0.00*

ni: indicates that the elasticities are not identified.

The detailed cointegration results using *rpx* are (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.023 & (0.016) \\ -0.018 & (0.004) \\ -0.010 & (0.027) \end{pmatrix} \begin{pmatrix} 1 & -1.105 & 1.550 \\ (na) & (0.323) & (0.505) \end{pmatrix} \begin{pmatrix} x \\ fy \\ rpx \end{pmatrix}, \text{ 1977.1-1994.4}$$

whereas the cointegration results using *reer* are

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.025 & (0.01) \\ -0.016 & (0.003) \\ -0.039 & (0.020) \end{pmatrix} \begin{pmatrix} 1 & -1.089 & 1.105 \\ (na) & (0.334) & (0.354) \end{pmatrix} \begin{pmatrix} x \\ fy \\ reer \end{pmatrix}, \text{ 1977.1-1994.4}$$

The results indicate that reliance on the conventional measure of relative prices (*rpx*) gives comparable elasticity estimates to those obtained by the IMF's measure of the real exchange rate (*reer*) both in terms of magnitudes and statistical significance. At the same time, the error-correction terms for *reer* has greater statistical significance than that for *rpx*.

Figures 76 and 77 show the ex-post predictions using the two measures of relative prices and a comparison of these data suggest that reliance on *reer* results in greater errors than reliance on *rpx*.

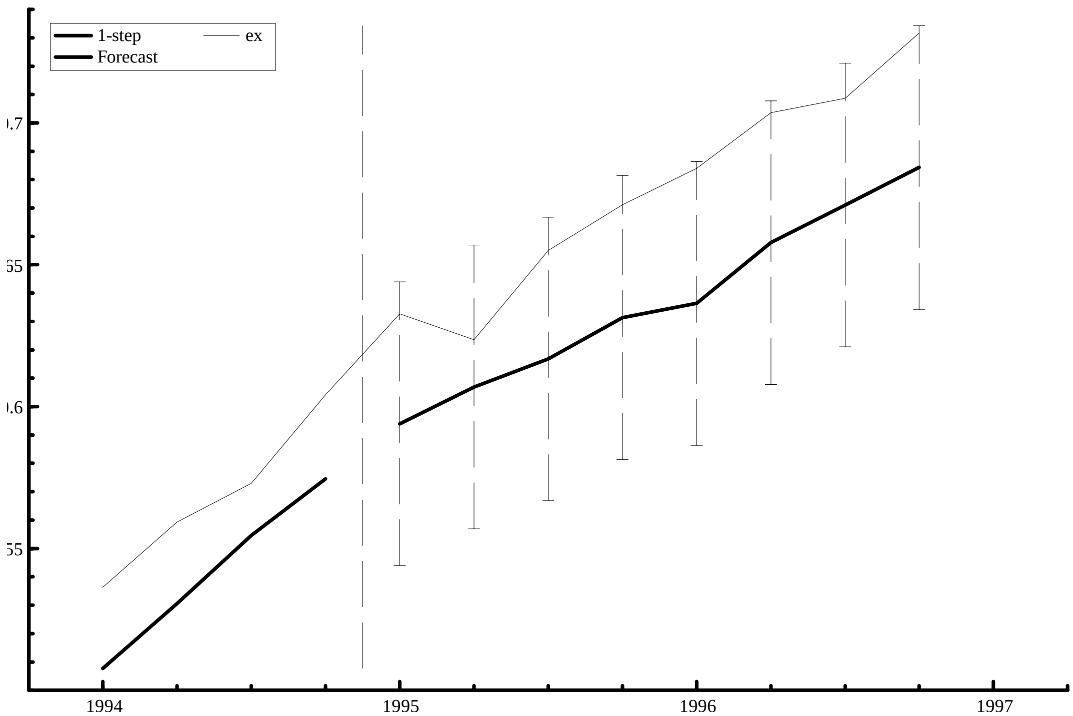


Figure 76: Predictive Accuracy-Cointegration-Exports with rpx-United Kingdom

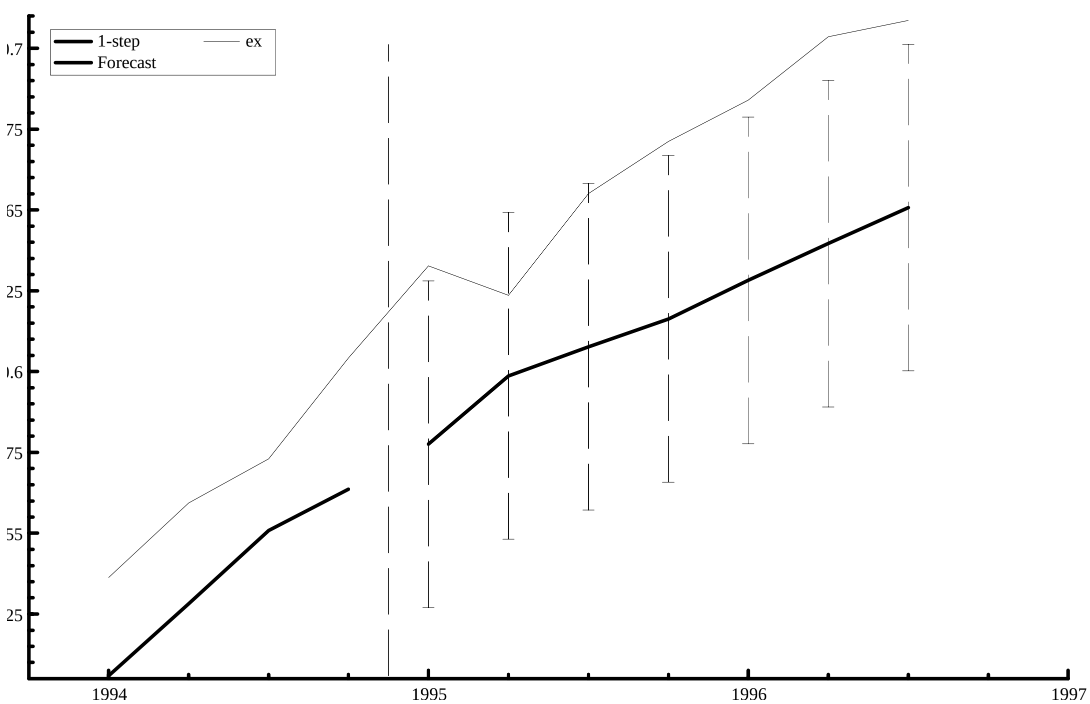


Figure 77: Predictive Accuracy-Cointegration-Exports with $reer$-United Kingdom

Figure 78 and 79 report the Chow tests associated with each measure of relative prices and the results indicate that, with a couple of exceptions, we cannot reject the hypothesis of parameter constancy for either measure of relative prices.

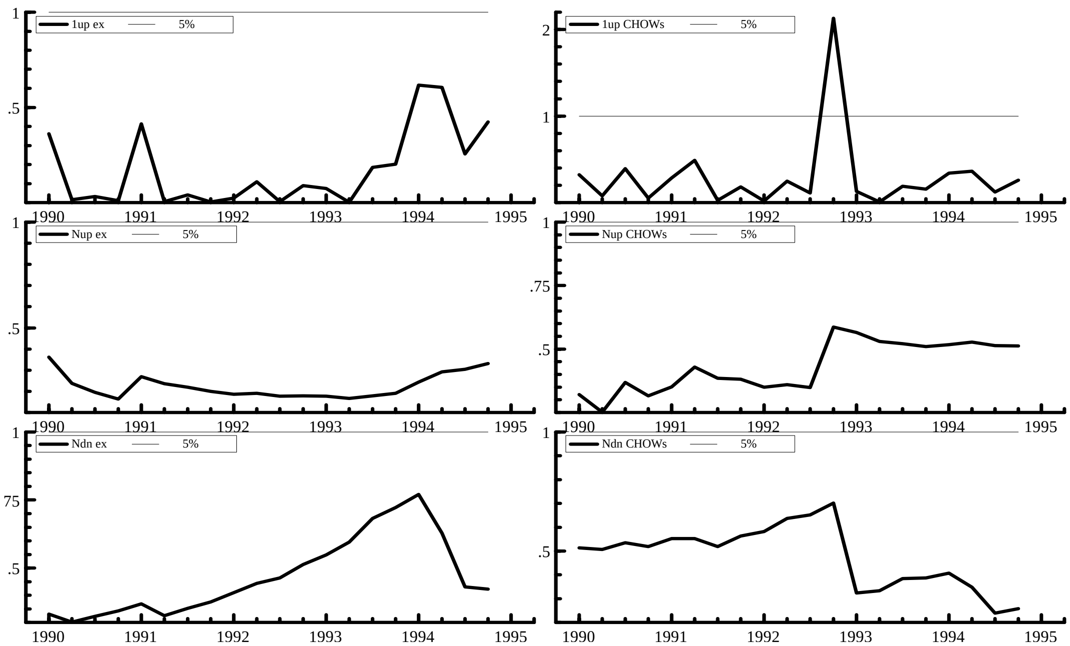


Figure 78: Chow Tests-Cointegration-Exports with *rpx*-United Kingdom

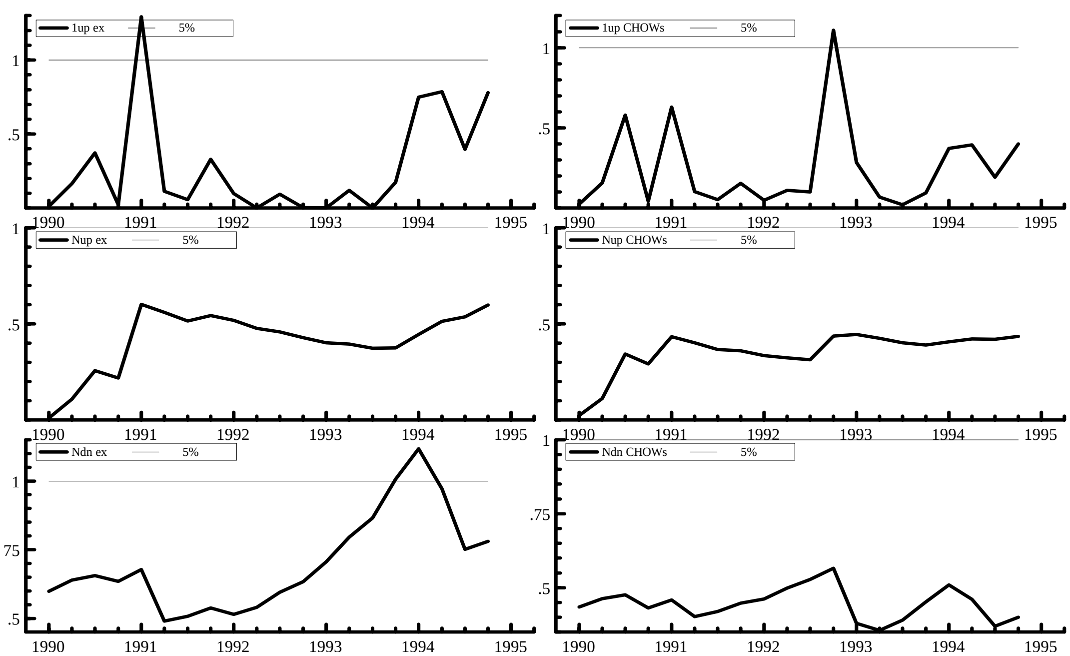


Figure 79: Chow Tests-Cointegration-Exports with *reer*-United Kingdom

Short-run Forces To explain short-run fluctuations, we estimate the parameters of two error-correction models, one for each measure of relative prices; we include two dummy variables for 1979.1 and 1979.2. The estimation results using rpx are

$$\underset{(se)}{\Delta x_t =} \quad \underset{(0.65)}{+1.092\Delta fy_t} \quad \underset{(0.10)}{-0.242\Delta rpx_t} \quad \underset{(0.015)}{-0.025 ECM_{x,t-1}},$$

where $ECM_x = x - 1.105 \cdot fy + 1.550 \cdot rpx + intercept$.

$R^2 = 0.74; SER = 1.178\%$	Null Hypothesis (p-value)	
Sample: 1977.3-1994.4	Serial-Independence (0.06)	Normality (0.41)
	Homoskedasticity (0.69)	Func. Form (0.87)

The estimation results using $reer$ are

$$\underset{(se)}{\Delta x_t =} \quad \underset{(0.60)}{+0.800\Delta fy_t} \quad \underset{(0.09)}{-0.289\Delta reer_t} \quad \underset{(0.02)}{-0.015 ECM_{x,t-1}},$$

where $ECM_x = x - 1.089 \cdot fy + 1.105 \cdot reer + intercept$.

$R^2 = 0.72; SER = 1.85\%$	Null Hypothesis (p-value)	
Sample: 1977.3-1994.4	Serial-Independence (0.08)	Normality (0.21)
	Homoskedasticity (0.95)	Func. Form (0.59)

Both formulations explain nearly three quarters of the variability of the growth rate of exports and the empirical distributions of the residuals satisfy the assumptions maintained for estimation. Notice, however, that the error-correction coefficients are not significantly different from zero. This result weakens the evidence on cointegration among those variables.

The prediction errors tend to be one-sided (figures 80 and 81) with the errors of formulation using rpx being worse than those of $reer$.

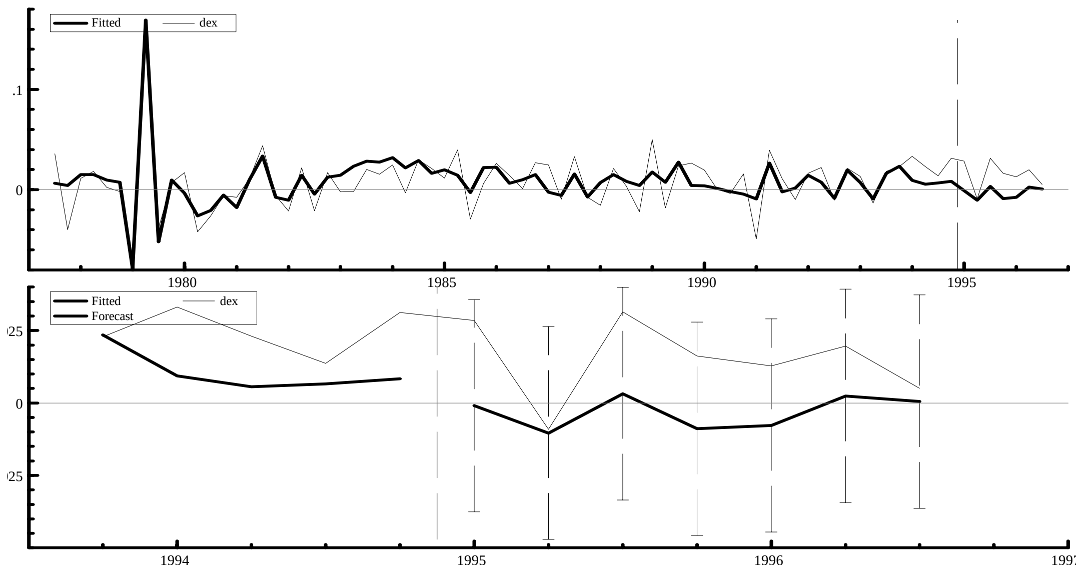


Figure 80: Predictive Accuracy-ECM-Exports with rpx-United Kingdom

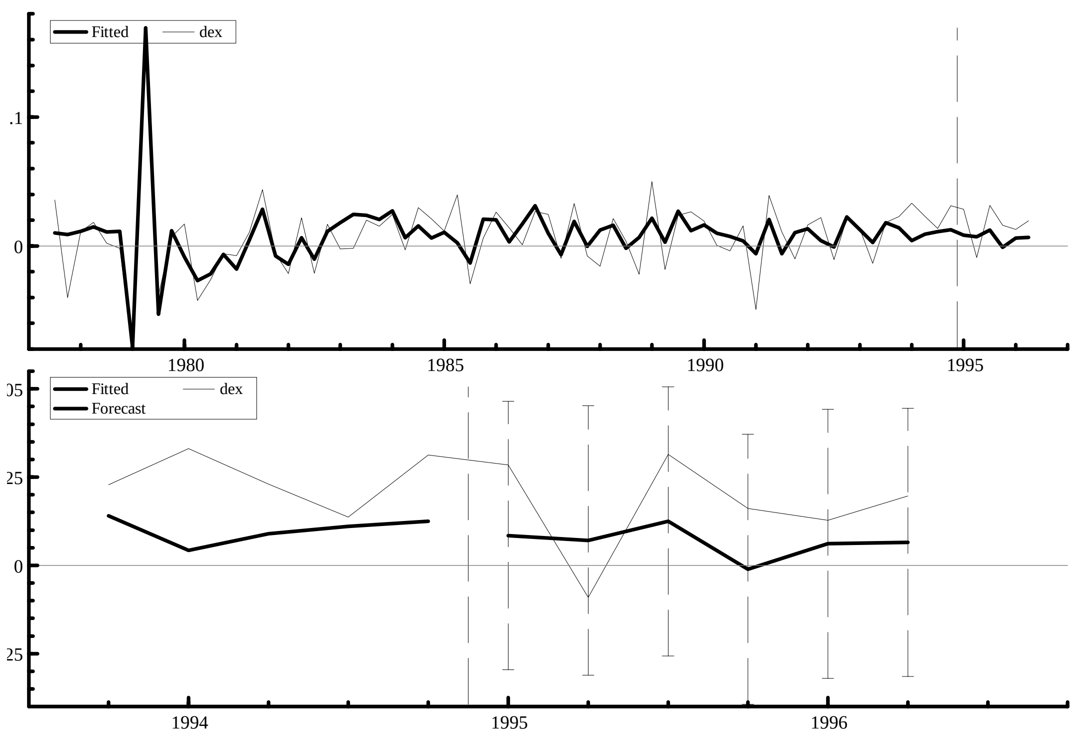


Figure 81: Predictive Accuracy-ECM-Exports with $reer$-United Kingdom

Figures 82 and 83 report the results from the Chow tests which cannot reject the hypothesis of parameter constancy. This evidence is corroborated by the evolution of the 95% confidence bands for the recursive estimates (figures 84 and 85).

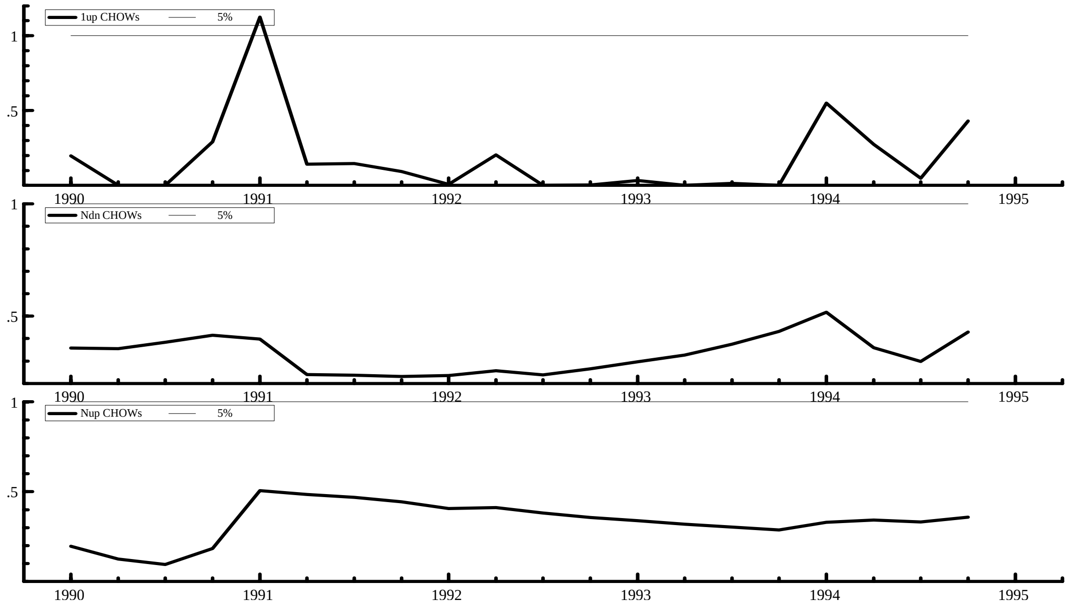


Figure 82: Chow Tests-ECM-Exports with *rpx*-United Kingdom

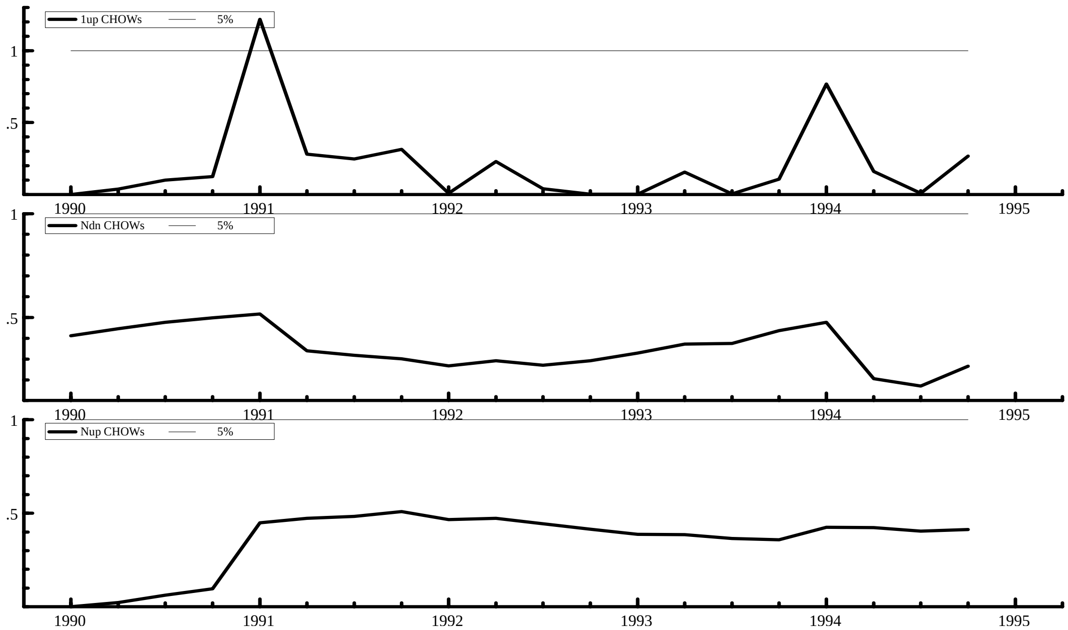


Figure 83: Chow Tests-ECM-Exports with *reer*-United Kingdom

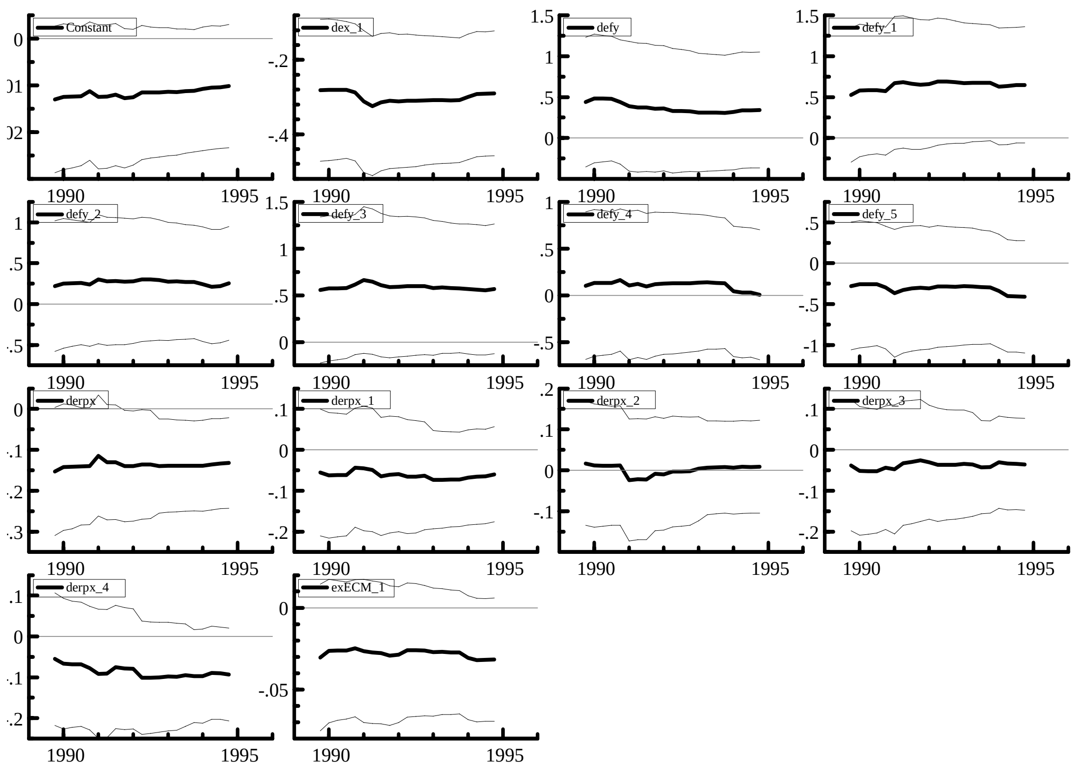


Figure 84: 95% Bands for ECM Coefficients-Exports with rpx-United Kingdom

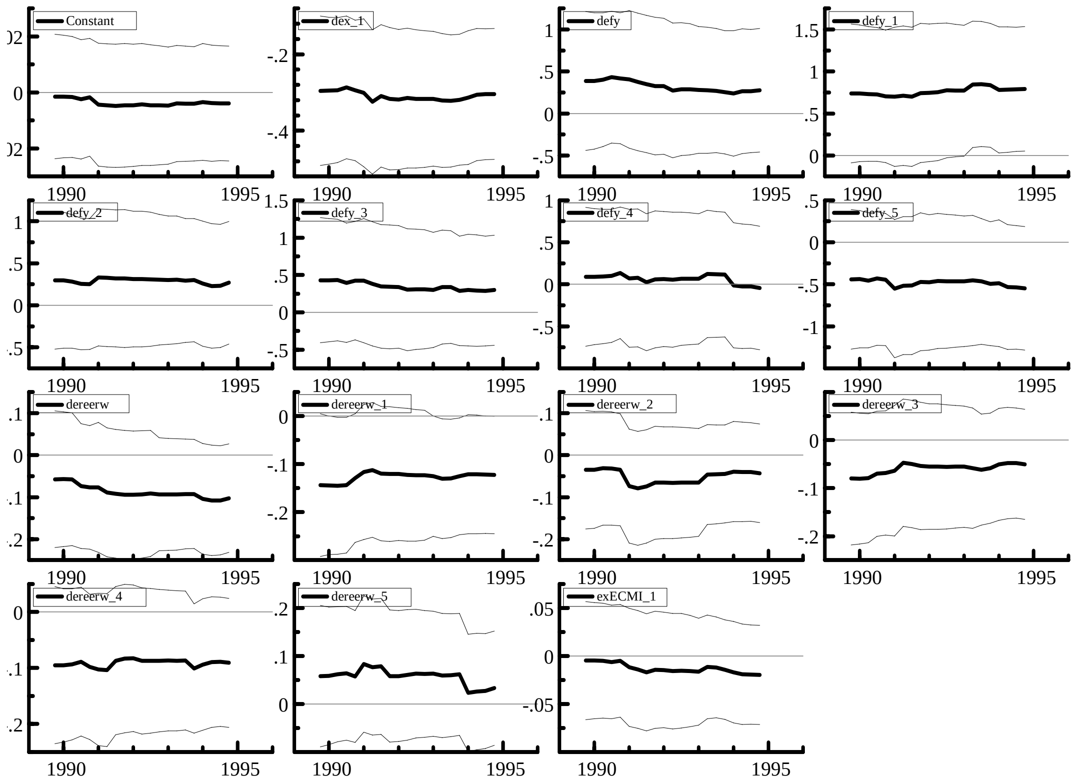


Figure 85: 95% Bands for ECM Coefficients-Exports with *reer*-United Kingdom

U.K. Imports

Long-run Forces Measuring prices as the import price relative to the GDP deflator yields estimates that are not robust to the choice of lag. Only the specification with five lags exhibits cointegration with an estimated price elasticity of -0.5 and an estimated income elasticity of 2.2. Measuring prices with the IMF's real effective exchange rate yields worse results: the only formulation exhibiting cointegration has lags of two quarters, a price elasticity of 0.2 and an income elasticity of 2.

Import Cointegration Results with *rpm*-United Kingdom

	Number of lags Included							
	9	8	7	6	5†	4	3	2
Cointegration Vectors	0	0	0	0	1	2	2	2
Price-Elasticity	ni	ni	ni	ni	-0.58	ni	ni	ni
Income Elasticity	ni	ni	ni	ni	2.21	ni	ni	ni
Loading Coefficient	ni	ni	ni	ni	0.01	ni	ni	ni
System's R^2	0.97	0.97	0.97	0.97	0.97	0.97	0.97	0.97
Import's Serial Corr.	0.31	0.63	0.72	0.20	0.33	0.44	0.67	0.90
System's Serial Corr.	0.41	0.16	0.41	0.05	0.13	0.06	0.46	0.35

Import Cointegration Results with *reer*-United Kingdom

	Number of lags Included							
	9	8	7	6	5	4	3	2
Cointegration Vectors	0	0	0	0	0	0	0	1
Price-Elasticity	ni	ni	ni	ni	ni	ni	ni	-0.17
Income Elasticity	ni	ni	ni	ni	ni	ni	ni	2.04
Loading Coefficient	ni	ni	ni	ni	ni	ni	ni	-0.23
System's R^2	0.97	0.96	0.96	0.96	0.96	0.96	0.96	0.96
Import's Serial Corr.	0.58	0.59	0.16	0.28	0.83	0.66	0.14	0.57
System's Serial Corr.	0.20	0.55	0.12	0.50	0.66	0.15	0.10	0.08

ni: indicates that the elasticities are not identified.

The results using *rpx* are (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} 0.005 & (0.004) \\ 0.006 & (0.002) \\ -0.006 & (0.03) \end{pmatrix} \begin{pmatrix} 1 & -2.209 & 0.580 \\ (na) & (0.817) & (1.511) \end{pmatrix} \begin{pmatrix} m \\ y \\ rpm \end{pmatrix}, \text{1956.2-1994.4.}$$

The evidence in favor of cointegration for U.K. imports is weak, as evidenced by a loading coefficient that is not significantly different from zero and a price elasticity that, again, is not statistically significant. Given the weak character of the evidence, the results should be treated with greater caution than usually. The out-of-sample prediction errors tend to be one-sided though not significantly different from zero (figure 86). Chow tests cannot reject the hypothesis of parameter constancy (figure 87).

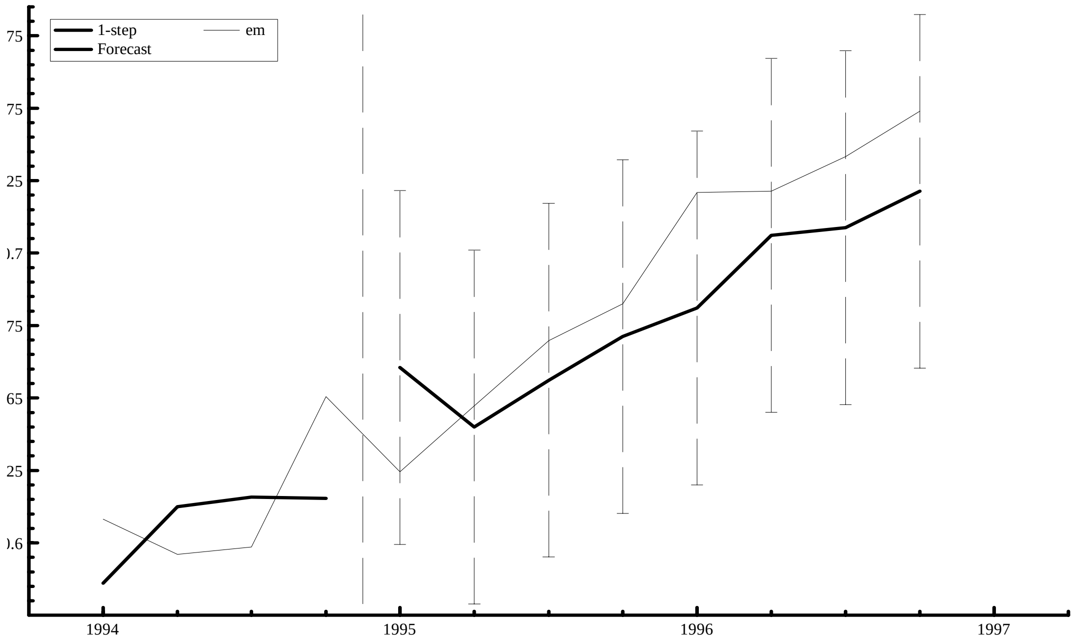


Figure 86: Predictive Accuracy-Cointegration-Imports-United Kingdom

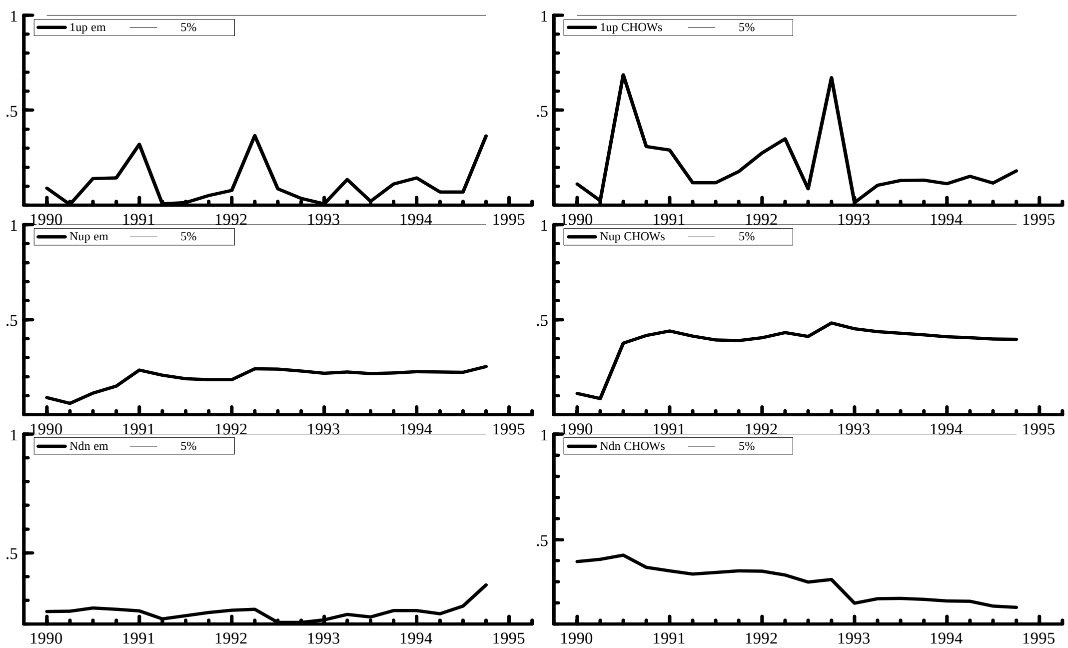


Figure 87: Chow Tests - Cointegration-Imports-United Kingdom

Short-run Forces To explain short-run fluctuations, we estimate the parameters of an error-correction model including three dummy variables that take a value of one in 1972.4, 1975.2, 1979.4. The estimation results are

$$\begin{array}{llll} \Delta m_t = & +1.008\Delta y_t & -0.0192 ECM_{m,t-1} \\ (se) & (0.30) & (0.010) \end{array}$$

where $ECM_m = m - 2.209 \cdot y + 0.580 \cdot rpm + intercept.$

$R^2 = 0.23; SER = 2.91\%$	Null Hypothesis (p-value)	
Sample: 1955.4-1994.4	Serial-Independence (0.05)	Normality (0.00*)
	Homoskedasticity (0.02*)	Func. Form (0.78)

The model barely explains one quarter of the variability of the growth rate of imports and the empirical distribution of the residuals does not satisfy all of the assumptions maintained for estimation, but the prediction errors are not one-sided (figure 88). Despite these limitations, the Chow tests (figure 89) and coefficient estimates (figure 90) support the hypothesis of parameter constancy.

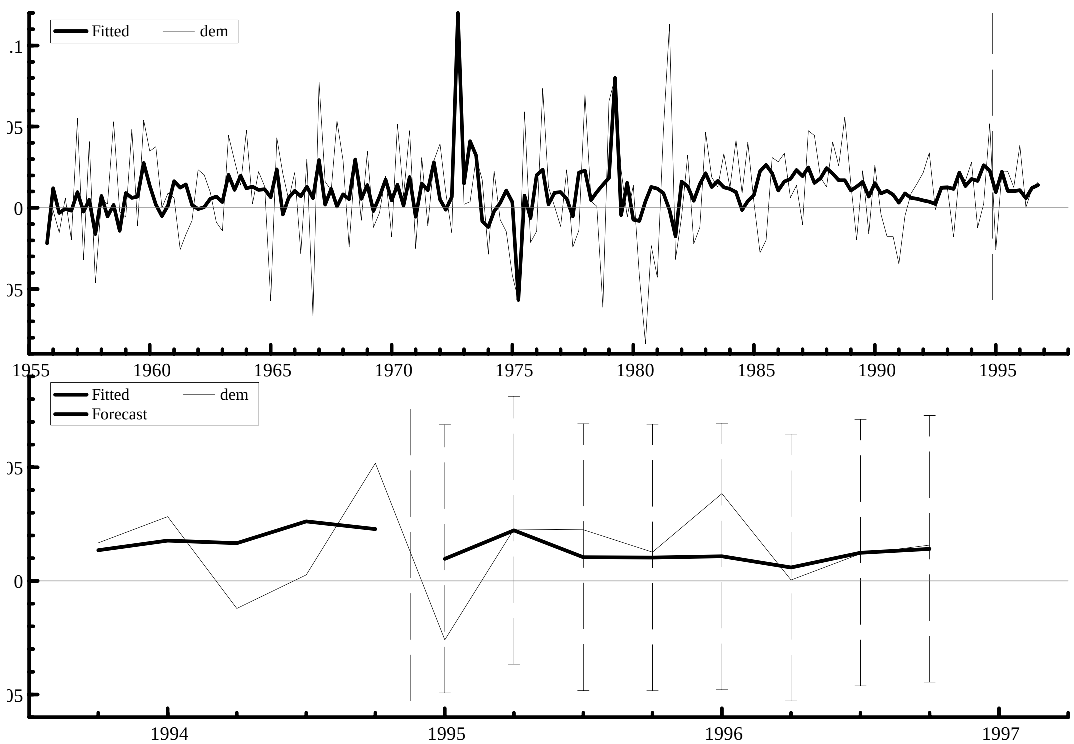


Figure 88: Predictive Accuracy-ECM-Imports-United Kigndom

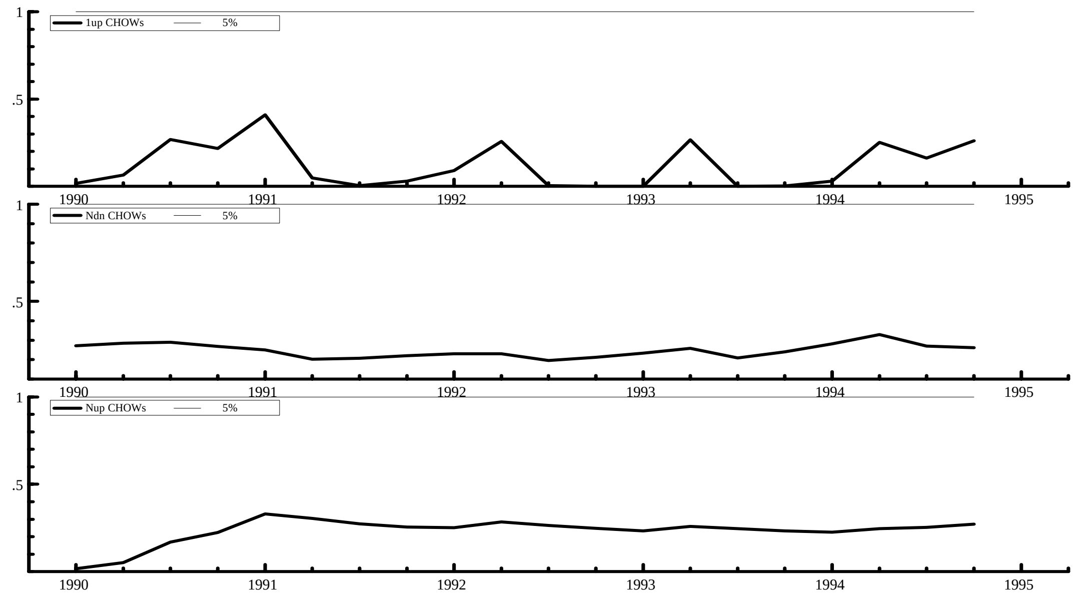


Figure 89: Chow Tests-ECM-Imports-United Kingdom

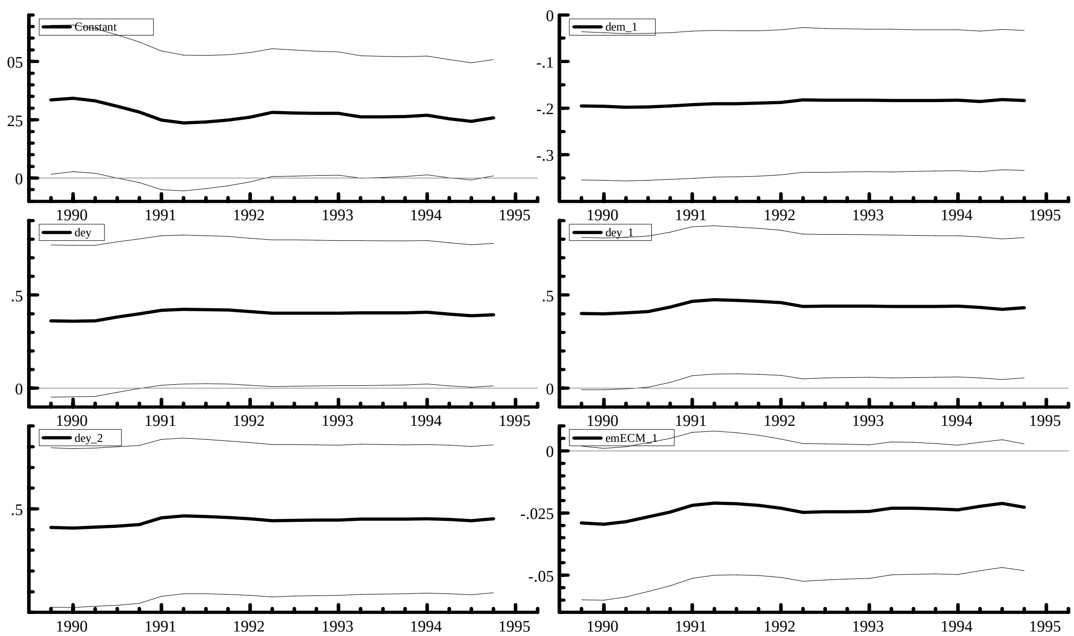


Figure 90: 95% Bands for ECM Coefficients - Imports-United Kingdom

K U.S. Trade

The evolution of U.S. foreign trade and its proximate determinants are displayed in figure 91. Both export and import volumes grow over time along with foreign and domestic income. Relative prices for both trade flows show downward trends, though the downward trend in the relative price of imports is considerably less accentuated than that of export prices. The IMF's *reer* shows less volatility than the conventional measure of relative prices based on deflators from the National Income Accounts.

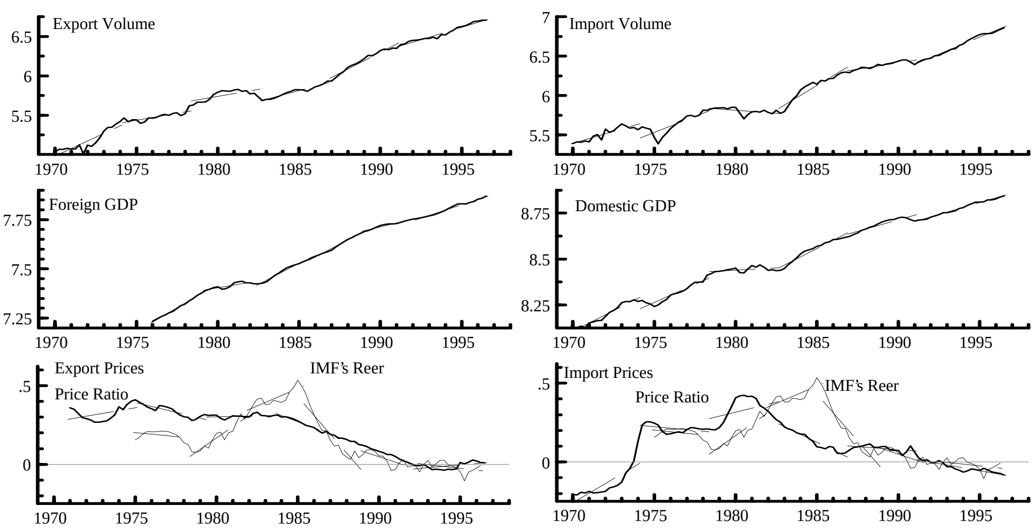


Figure 91: Trade, Income, and Prices-United States

U.S. Exports

Long-run Forces For a given measure of prices, elasticities for U.S. exports appear robust to the choice of lag-length. For a given lag, however, the estimates are not robust to the choice of price measure.

Export Cointegration with Results with *rpx*-United States

	Number of lags Included							
	9	8	7	6	5	4	3	2†
Cointegration Vectors	2	2	1	1	1	0	0	1
Price-Elasticity	ni	ni	-1.38	-1.44	-1.51	ni	ni	-1.47
Income Elasticity	ni	ni	0.89	0.73	0.74	ni	ni	0.83
Loading Coefficient	ni	ni	-0.15	0.03	-0.05	ni	ni	-0.06
System's R^2	0.96	0.95	0.93	0.92	0.92	0.92	0.92	0.91
Export's Serial Corr.	0.50	0.28	0.09	0.14	0.06	0.64	0.97	0.81
System's Serial Corr.	0.27	0.01*	0.00*	0.00*	0.01*	0.05	0.84	0.71

Export Cointegration Results with *reer*-United States

	Number of lags Included							
	9	8	7	6	5	4	3	2†
Cointegration Vectors	2	1	1	1	1	2	1	1
Price-Elasticity	ni	-0.72	-0.52	-0.66	-0.62	ni	-0.57	-0.56
Income Elasticity	ni	1.68	1.73	1.81	1.75	ni	1.79	1.79
Loading Coefficient	ni	-0.25	-0.35	-0.24	-0.30	ni	-0.20	-0.22
System's R^2	0.97	0.97	0.96	0.96	0.95	0.95	0.95	0.95
Export's Serial Corr.	0.02*	0.04*	0.71	0.20	0.20	0.64	0.66	0.64
System's Serial Corr.	0.03*	0.00*	0.07	0.00*	0.01*	0.36	0.89	0.96

ni: indicates that the elasticities are not identified.

The detailed cointegration results using *rpx* are (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.056 & (0.048) \\ -0.041 & (0.010) \\ 0.037 & (0.023) \end{pmatrix} \begin{pmatrix} 1 & -0.828 & 1.469 \\ (na) & (0.191) & (0.244) \end{pmatrix} \begin{pmatrix} x \\ fy \\ rpx \end{pmatrix}, \text{ 1976.3-1994.4}$$

whereas the cointegration results using *reer* are

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.204 & (0.04) \\ -0.006 & (0.010) \\ 0.101 & (0.073) \end{pmatrix} \begin{pmatrix} 1 & -1.785 & 0.561 \\ (na) & (0.079) & (0.084) \end{pmatrix} \begin{pmatrix} x \\ fy \\ reer \end{pmatrix}, \text{ 1976.4-1994.4}$$

The results indicate that reliance on *rpx* gives rather different elasticity estimates to those obtained with *reer* both in terms of magnitudes and statistical significance. Specifically, reliance on *reer* doubles the estimated income elasticity and halves the price elasticity relative to the model using *rpx*. Moreover, the error-correction term for *reer* is larger and it has greater statistical significance than the one for *rpx*.

Figures 92 and 93 below compare the ex-post predictions associated with the two measures of relative prices and the evidence indicates smaller forecast errors using *reer*.

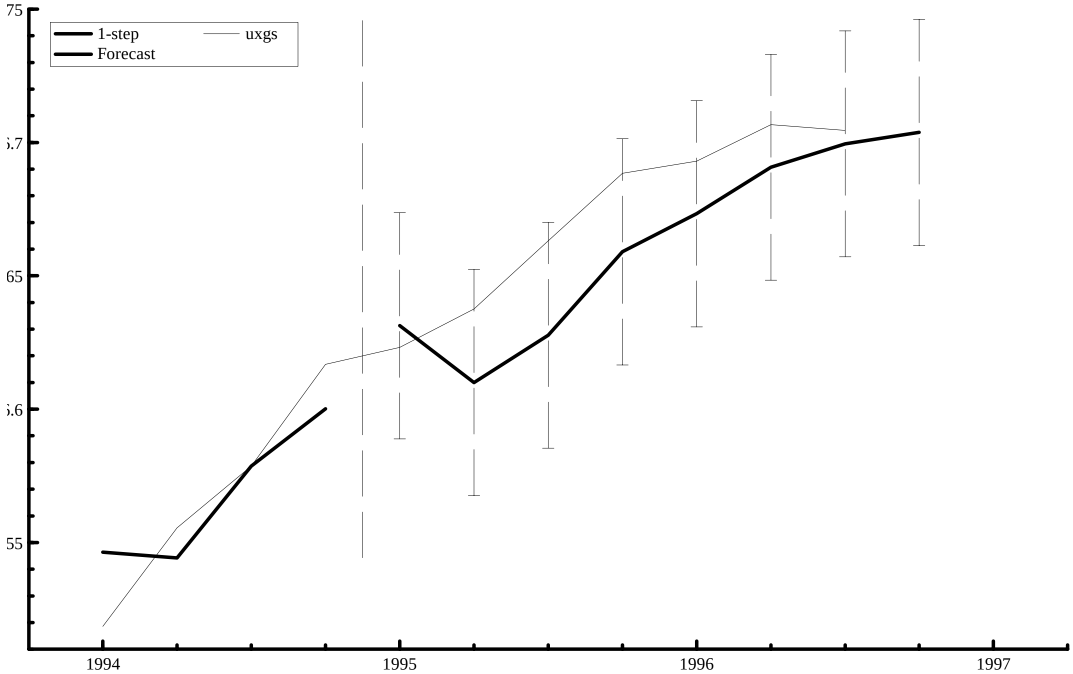


Figure 92: Predictive Accuracy-Cointegration-Exports with rpx-United States

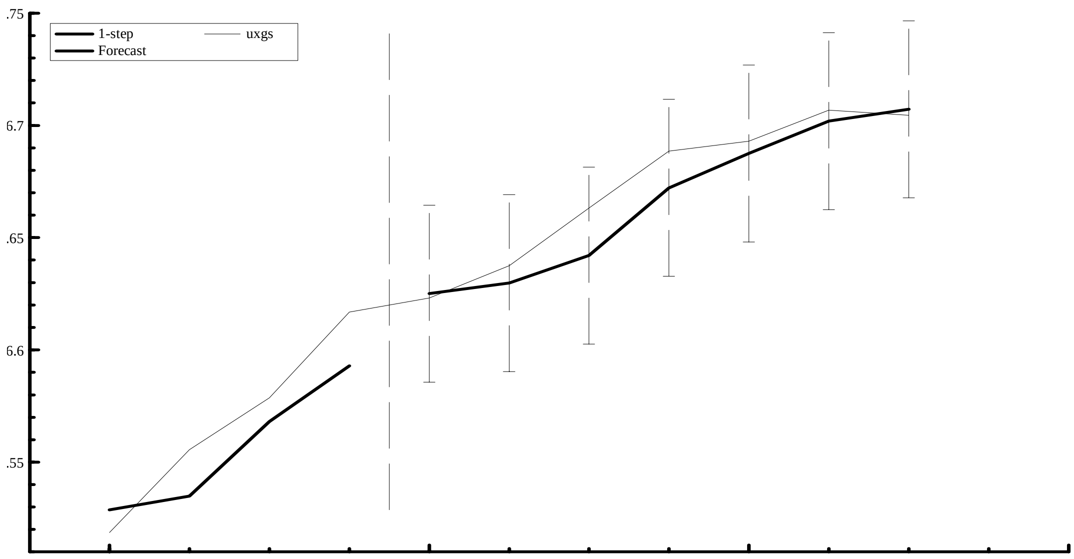


Figure 93: Predictive Accuracy-Cointegration-Exports with $reer$-United States

Figures 94 and 95 report the Chow tests for each measure of relative prices and the evidence, once again, supports the hypothesis of parameter constancy but the use of $reer$ has fewer failures of the Chow test.

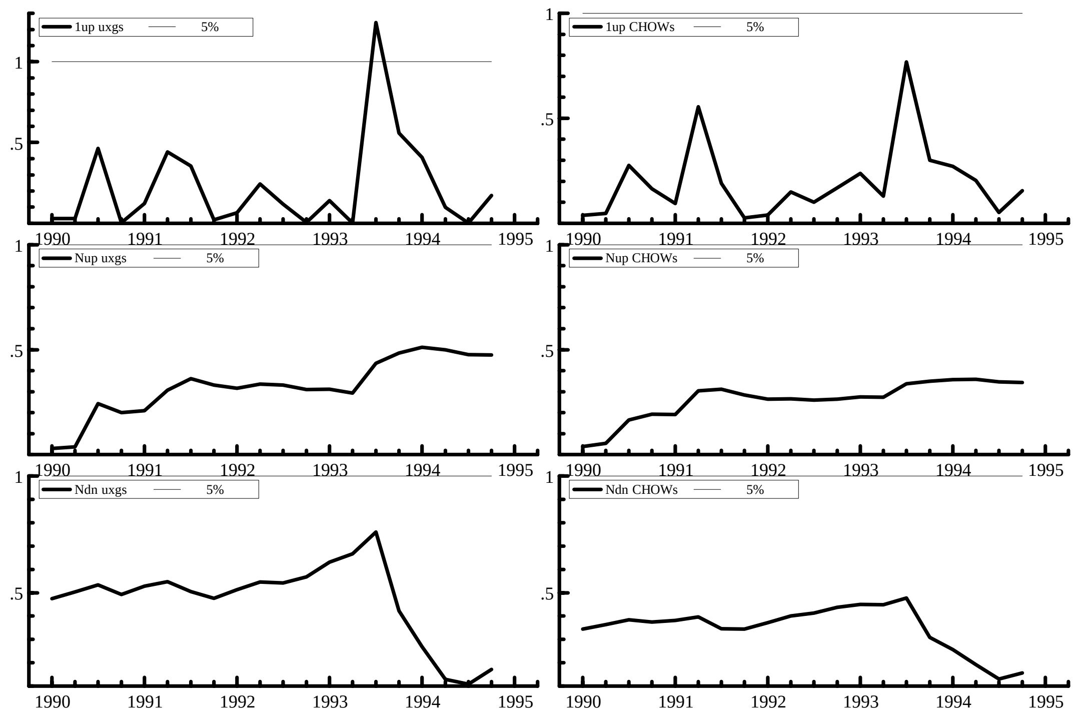


Figure 94: Chow Tests-Cointegration-Exports with rpx-United States

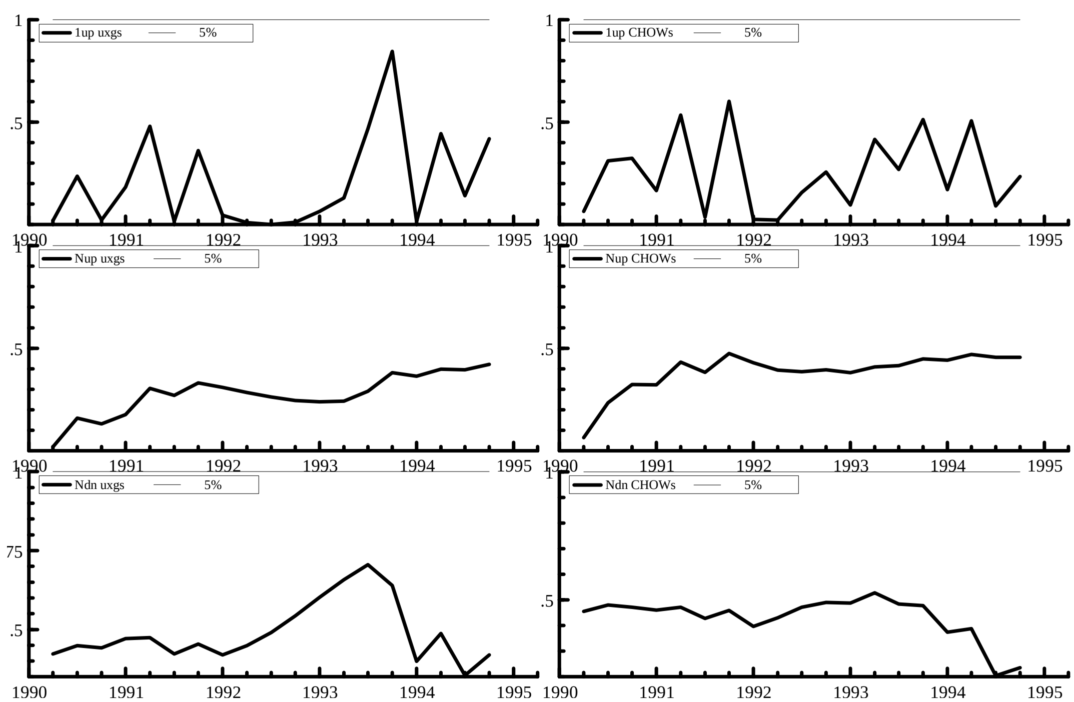


Figure 95: Chow Tests-Cointegration-Exports with $reer$-United States

Short-run Forces To explain short-run fluctuations, we estimate the parameters of two error-correction models, one for each measure of relative prices, and with two dummy variables taking a value of one in 1978.2 and 1977.4. The estimation results using rpx are

$$\underset{(se)}{\Delta x_t =} \quad \underset{(0.48)}{+1.826\Delta fy_t} \quad \underset{(0.23)}{-0.533\Delta rpx_t} \quad \underset{(0.05)}{-0.0519 ECM_{x,t-1}},$$

where $ECM_x = x - 0.8276 \cdot fy + 1.469 \cdot rpx + intercept$

$R^2 = 0.49; SER = 1.79\%$	Null Hypothesis (p-value)	
Sample: 1976.3-1994.4	Serial-Independence (0.18)	Normality (0.49)
	Homoskedasticity (0.50)	Func. Form (0.06)

The estimation results using $reer$ are

$$\underset{(se)}{\Delta x_t =} \quad \underset{(0.29)}{+0.907\Delta fy_t} \quad \underset{(0.05)}{-0.001\Delta reer_t} \quad \underset{(0.03)}{-0.1507 ECM_{x,t-1}},$$

where $ECM_x = x - 1.785 \cdot fy + 0.561 \cdot reer + intercept$

$R^2 = 0.69; SER = 1.65\%$	Null Hypothesis (p-value)	
Sample: 1976.3-1994.4	Serial-Independence (0.62)	Normality (0.90)
	Homoskedasticity (0.22)	Func. Form (0.25)

The formulation using rpx explains about half of the variability of the growth rate of exports and the empirical distribution of the residuals satisfies the assumptions maintained for estimation. This formulation has, however, an error-correction coefficient that is not significantly different from zero which weakens the evidence on cointegration among those variables. The formulation using $reer$ explains two thirds of the variability of growth rate and its residuals are also consistent with the maintained assumptions.

The predictions of both error-correction models are one-sided but the out-of-sample predictions of the formulation using rpx are worse than those of $reer$ (figures 96 and 97).

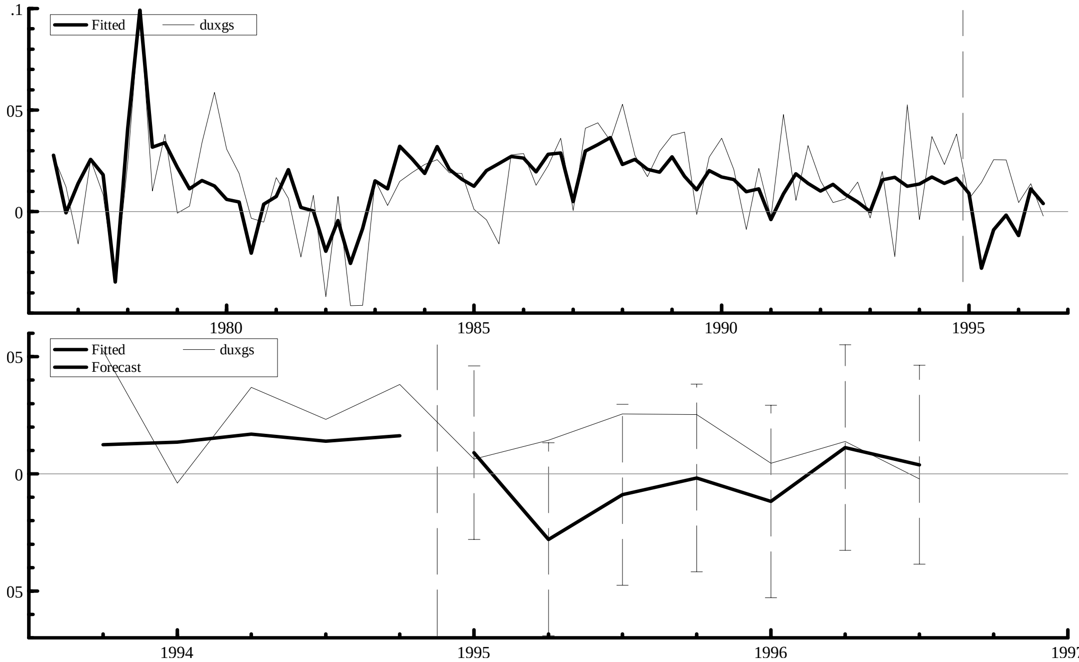


Figure 96: Predictive Accuracy-ECM-Exports with *rpx*-United States

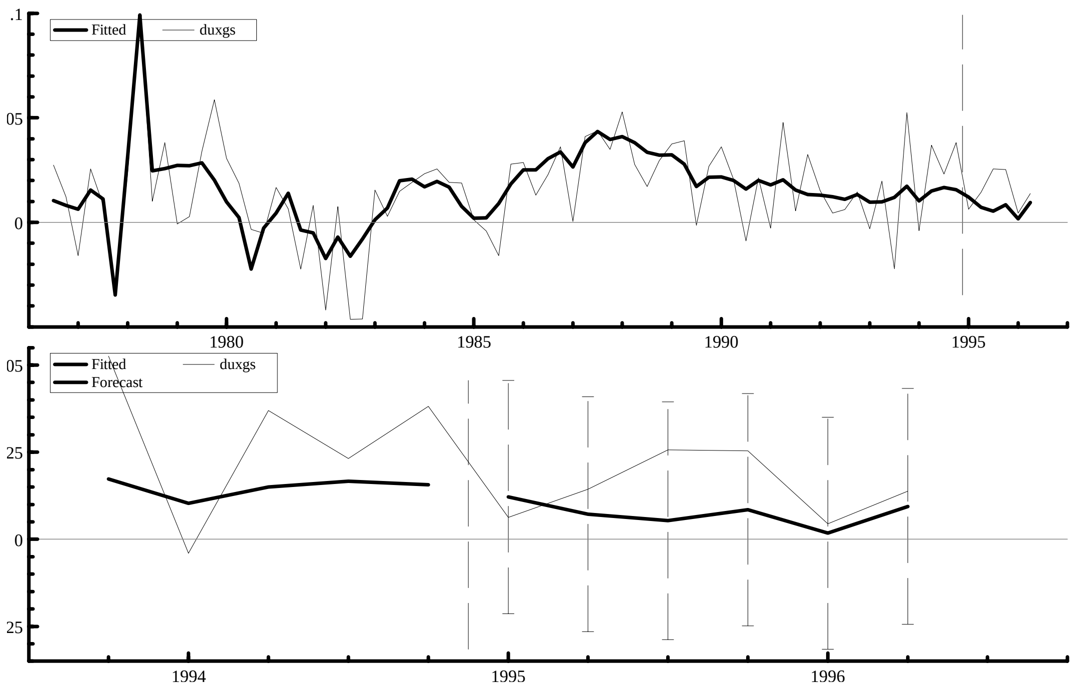


Figure 97: Predictive Accuracy-ECM-Exports with *reer*-United States

In terms of parameter constancy, both error-correction models offer comparable reliability as reflected in the Chow tests (figures 98 and 99) and the 95% confidence bands of the recursive coefficients (figures 100 and 101).

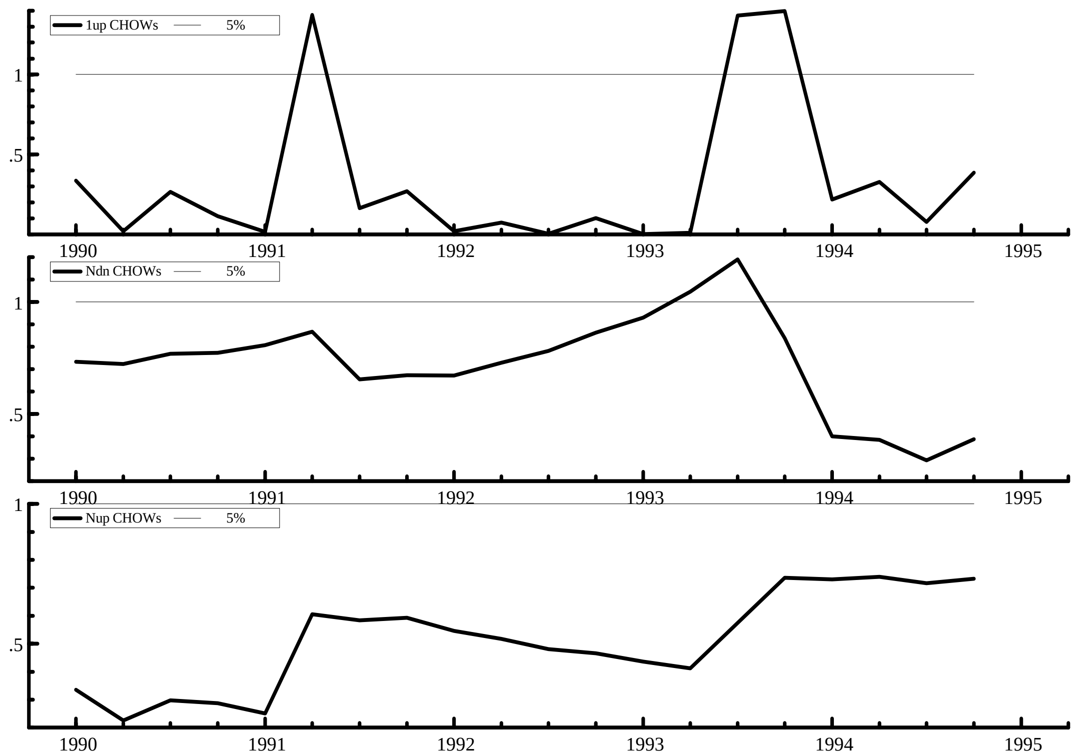


Figure 98: Chow Tests-ECM-Exports with rpx-United States

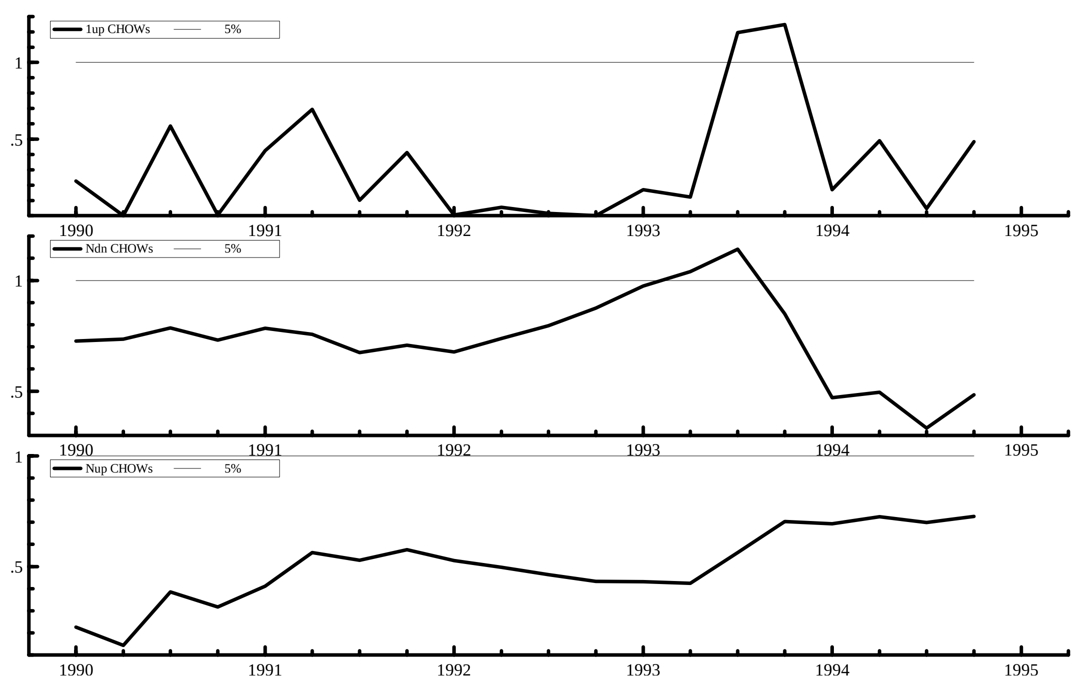


Figure 99: Chow Tests-ECM-Exports with $reer$-United States

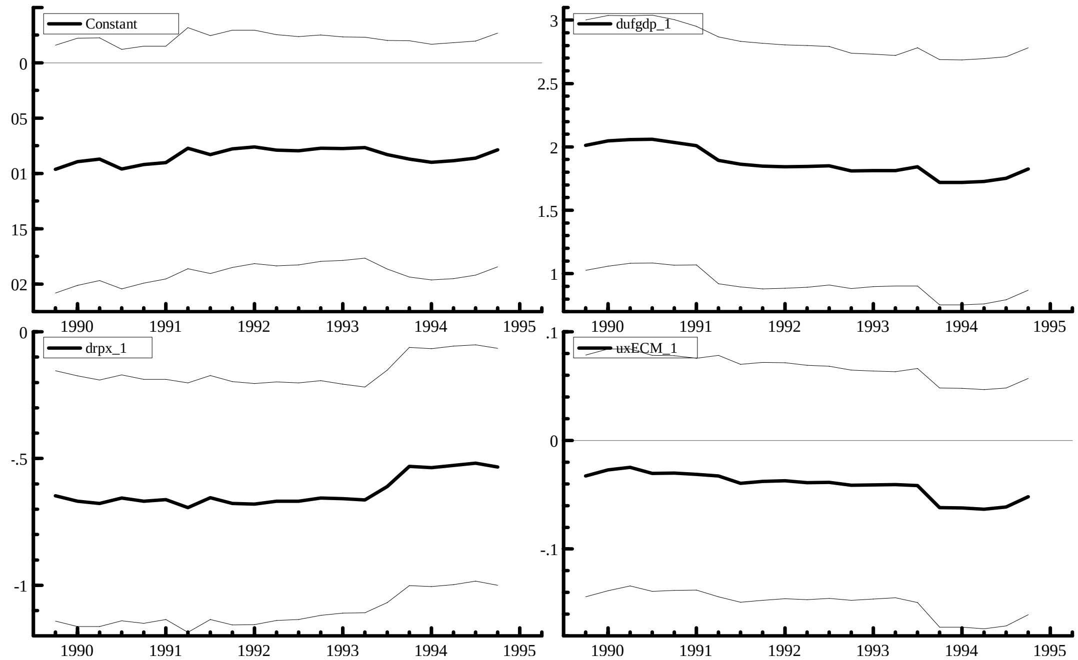


Figure 100: 95% Bands for ECM Coefficients-Exports with rpx-United States

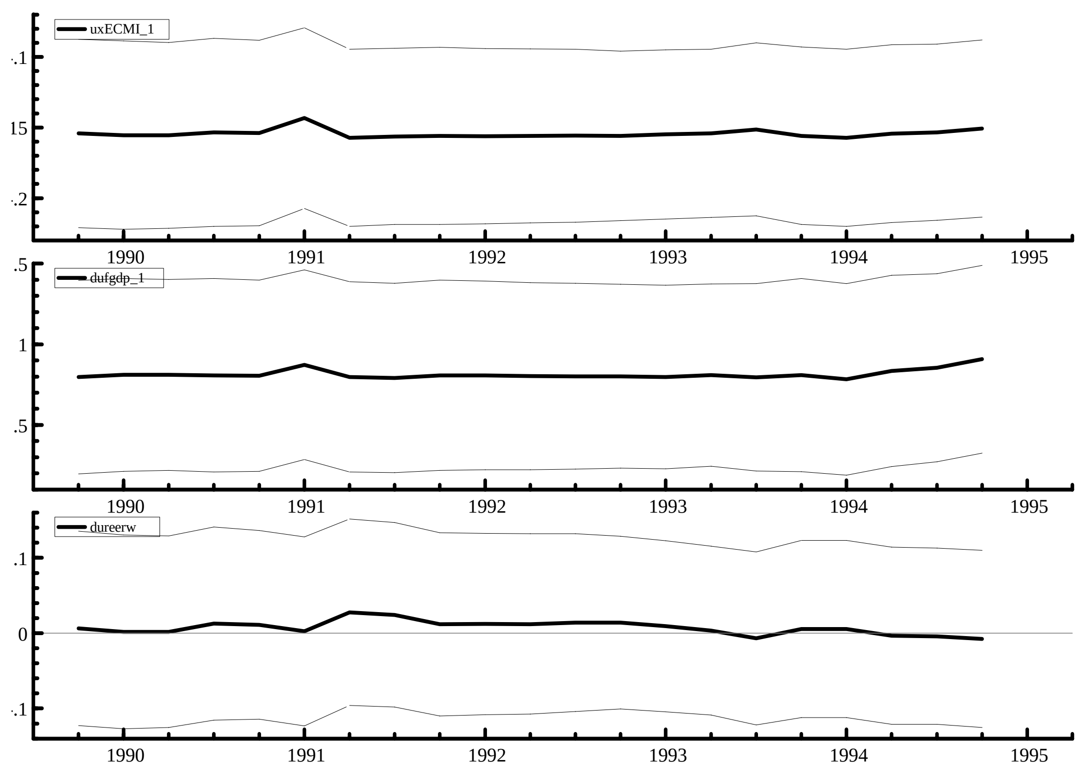


Figure 101: 95% Bands for ECM Coefficients-Exports with $reer$-United States

U.S. Imports

Long-run Forces Income elasticities for imports appear robust to both the choice of lag-length and price measure. For a given measure of prices, price elasticities are also robust to the choice of lag-length; the estimates are not robust to the choice of price measure and take the wrong sign for the IMF's measure of real exchange rates.

Import Cointegration Results with *rpm*-United States

	Number of lags Included							
	9†	8	7	6	5	4	3	2
Cointegration Vectors	1	1	1	1	1	1	2	2
Price-Elasticity	-0.31	-0.21	-0.37	-0.36	-0.41	-0.43	ni	ni
Income Elasticity	1.79	1.72	2.76	2.12	2.10	2.10	ni	ni
Loading Coefficient	-0.10	-0.08	0.01	0.01	-0.03	-0.03	ni	ni
System's R^2	0.99	0.99	0.99	0.99	0.99	0.99	0.99	0.99
Import's Serial Corr.	0.12	0.37	0.11	0.03*	0.31	0.37	0.06	0.81
System's Serial Corr.	0.26	0.12	0.02*	0.02*	0.03*	0.02*	0.00*	0.00*

Import Cointegration Results with *reer*-United States

	Number of lags Included							
	9	8	7	6	5	4	3	2
Cointegration Vectors	2	2	1	1	1	0	0	1
Price-Elasticity	ni	ni	-0.13	-0.13	-0.12	ni	ni	-0.11
Income Elasticity	ni	ni	2.11	2.13	2.14	ni	ni	2.22
Loading Coefficient	ni	ni	-0.10	-0.02	-0.06	ni	ni	0.04
System's R^2	0.98	0.98	0.98	0.98	0.98	0.98	0.97	0.97
Import's Serial Corr.	0.01*	0.05	0.00*	0.03*	0.64	0.28	0.25	0.03*
System's Serial Corr.	0.01*	0.08	0.44	0.08	0.55	0.20	0.21	0.28

ni: indicates that the elasticities are not identified.

The cointegration results using rpm are (standard errors in parentheses)

$$\widehat{\alpha\beta' z} = \begin{pmatrix} -0.100 & (0.03) \\ -0.029 & (0.01) \\ 0.019 & (0.01) \end{pmatrix} \begin{pmatrix} 1 & -1.786 & 0.309 \\ (na) & (0.094) & (0.146) \end{pmatrix} \begin{pmatrix} m \\ y \\ rpm \end{pmatrix}, \text{ 1961.4-1994.4.}$$

which shows significant long-run elasticities for income and prices and a significant error-correction term.

The out-sample predictions based on these estimates are one sided (figure 102) and the recursive Chow tests for parameter-constancy tests suggest somewhat stable elasticities through 1994 (figure 103).

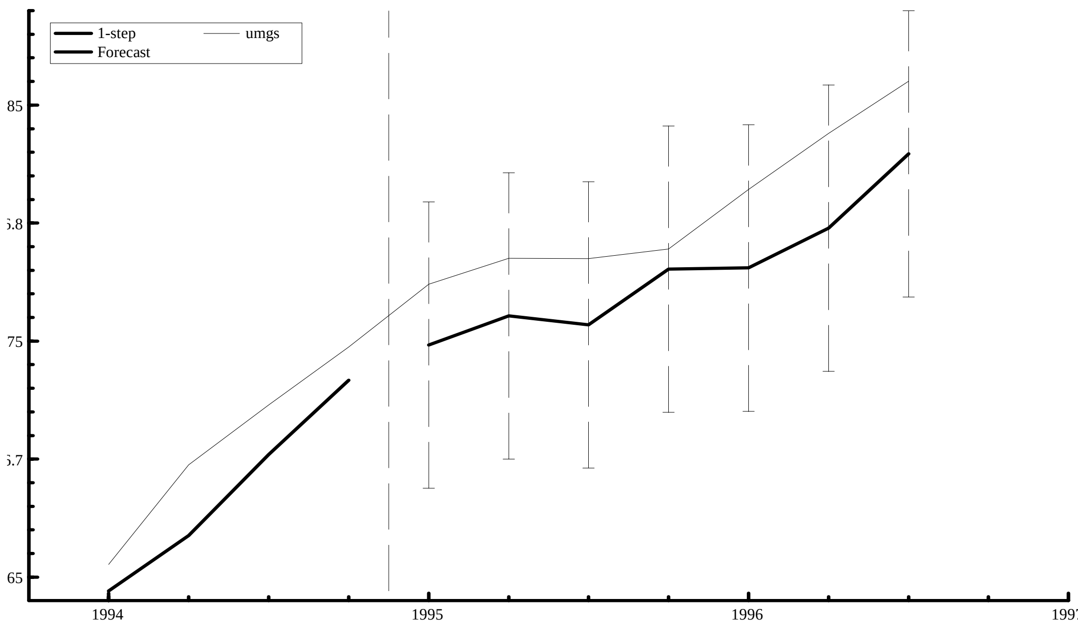


Figure 102: Predictive Accuracy-Cointegration-Imports-United States

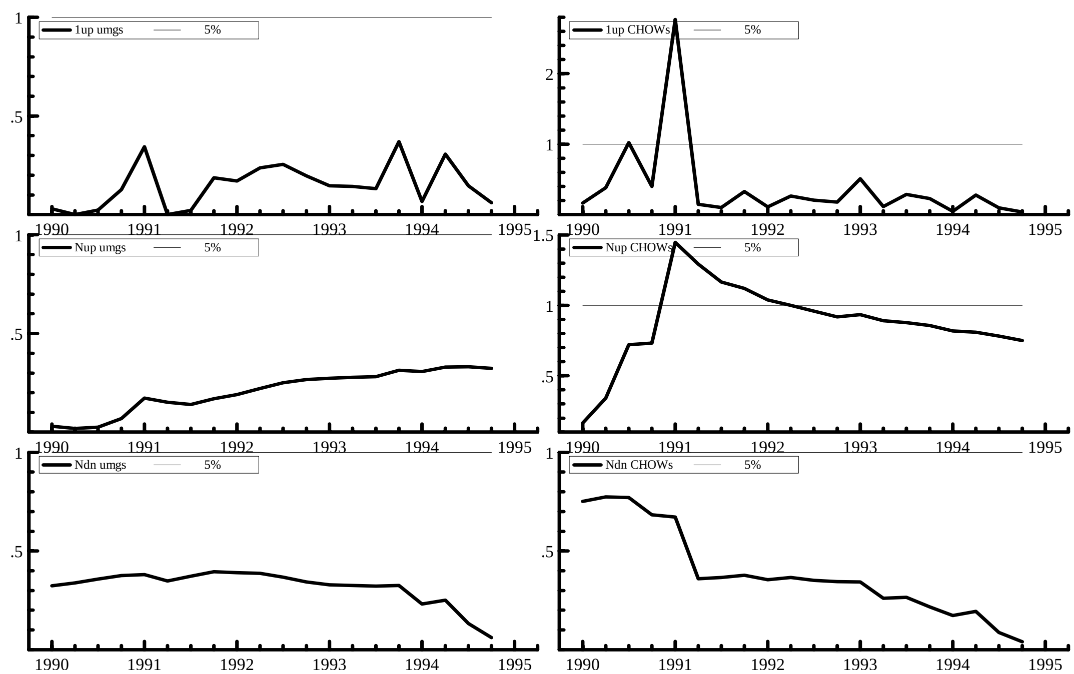


Figure 103: Chow Tests-Cointegration-Imports-United States

Short-run Forces To explain short-run fluctuations, we estimate the parameters of an error-correction model including five dummy variables that take a value of one in 1969.1, 1969.2, 1972.1, 1972.2 and 1974.2. The estimation results are

$$
\begin{array}{lccc}
\Delta m_t = & +1.004\Delta y_t & -0.0456\Delta rpm_t & -0.03262ECM_{m,t-1} \\
(se) & (0.37) & (0.09) & (0.017)
\end{array}
$$

where $ECM_m = m - 1.7855 \cdot y + 0.3086 \cdot rpm + intercept$.

$R^2 = 0.63; SER = 2.48\%$	Null Hypothesis (p-value)	
Sample: 1960.3-1994.4	Serial-Independence (0.39)	Normality (0.06)
	Homoskedasticity (0.20)	Func. Form (0.60)

The model explains about two-thirds of the variability of the growth rate of imports (figure 104) and the empirical distribution of the residuals satisfies the assumptions maintained for estimation. Unlike the predictions of the cointegration system, the predictions of the error-correction model are not one-sided though the variable being predicted is the growth rate. Finally, the Chow tests (figure 105) and coefficient estimates (figure 106) support parameter constancy.

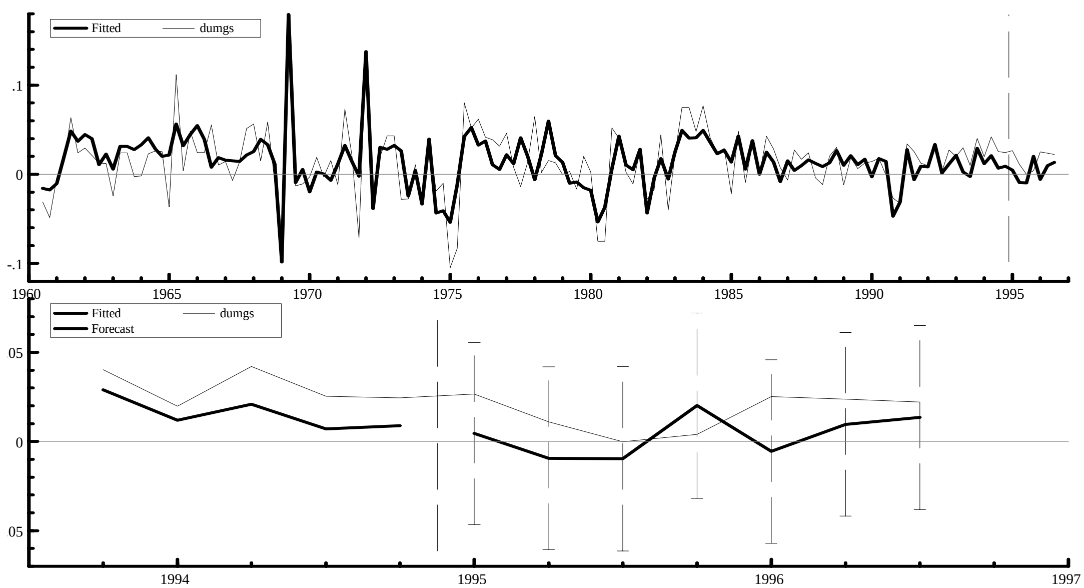


Figure 104: Predictive Accuracy-ECM-Imports-United States

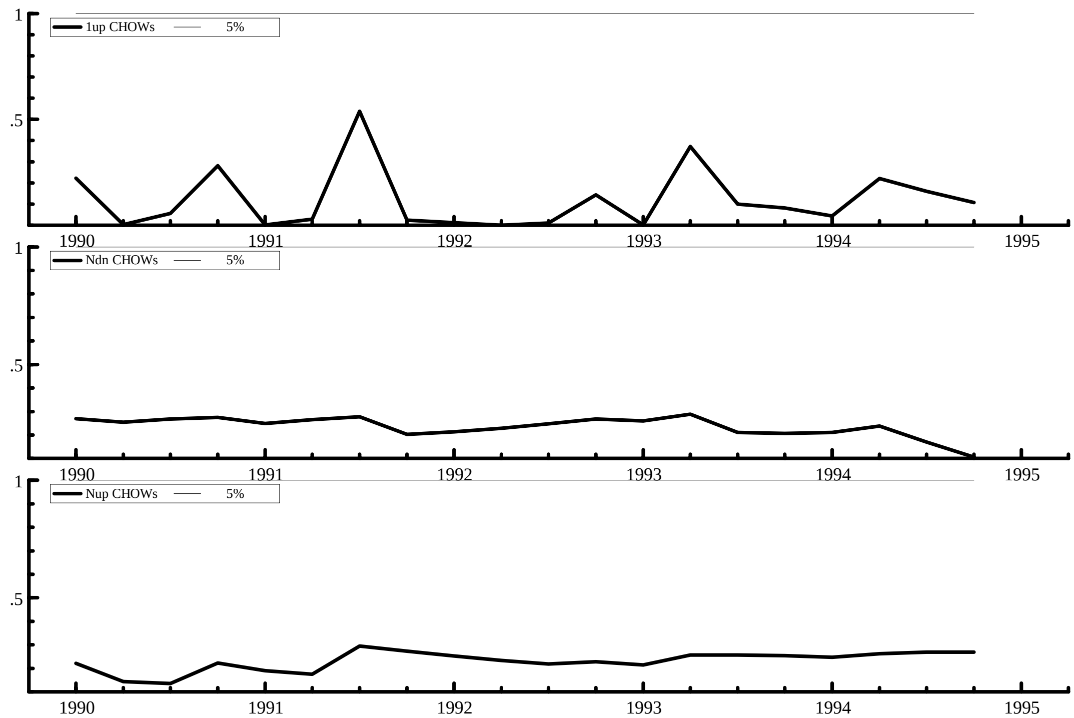


Figure 105: Chow Tests - ECM-Imports-United States

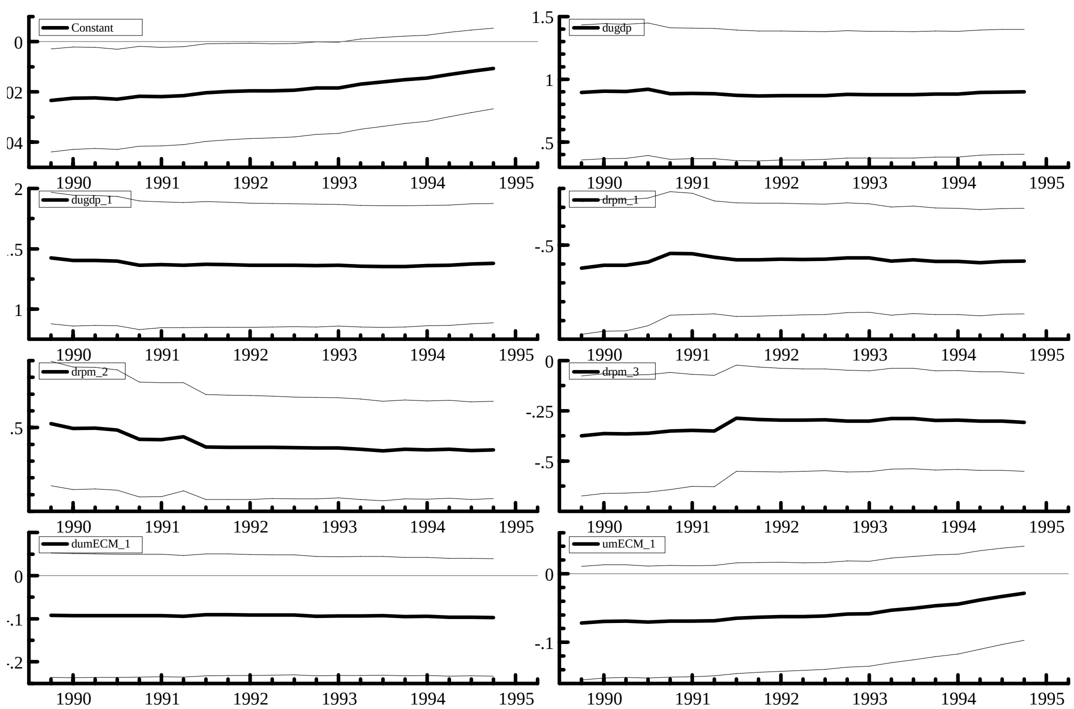


Figure 106: 95% Bands for ECM Coefficients-Imports-United States

L Bilateral Trade Equations

Because our initial estimates of German and French import and export volume equations showed price elasticities to be extremely low, we have undertaken further tests to try to determine the robustness of this result by estimating equations for German and French trade with non-EU countries, along with several bilateral equations (involving German trade with France and with the United States). The price elasticity estimates obtained in these additional tests are presented in the table below. The new results can be summarized as follows.

On the whole, these additional tests failed to overturn our earlier finding with respect to price elasticities in the German and French import equations. With one exception, price elasticities in the import equations either had the wrong sign (positive) or were not statistically significant. The exception was in the case of the single-equation (OLS) estimation of German bilateral imports from France, which yielded a statistically significant elasticity of -0.6. It may be worth noting that elasticity estimates in excess of -1.0 were obtained in the bilateral equation for German exports to France. However, the estimate is insignificantly different from zero. More noteworthy is the fact that the estimated price elasticities of German and French imports from non-EU countries appear to be about the same (zero or close to it) as those of the total imports of these two countries.

On the export side, in contrast to the results for imports, we did turn up evidence of price-elasticity. The conventional (in this case bilateral) relative price term yielded an elasticity of -1.2 in single-equation estimation for German exports to the United States. However, the conventional relative price terms proved to be only moderately higher for German and French exports to non-EU countries (-0.4 and -0.3, respectively) than they had been for German and French exports to all countries (between -0.1 and -0.2).

On balance, these results tend to support the view that German and French imports are relatively price inelastic. At the same time, they also suggest that German and French exports are more price elastic than our initial estimates indicated. Finally, they lend at most only weak support to the view that intra-European trade is less price elastic than extra-European trade.

Bilateral Elasticities for German Trade (OLS)

		Price	Income	Relative Price	Stability
Imports	Total	0.1	1.7	*rpm*	fail
	From Non-EU	0.0	1.5	*rpm*	fail
		-0.5	1.2	*reer*	pass
	From France	-0.6	1.6	*rpm*	pass
	From U.S.	-0.0	1.3	*rpm*	pass
Exports	Total	-0.2	1.5	*rpx*	fail
	To Non-EU	-0.4	0.8	*rpx*	fail
	To France	-2.1	1.0	*rpx*	pass
	To U.S.	-1.2	1.6	*rpx*	pass

Bilateral Elasticities for French Trade (OLS)

		Price	Income	Relative Price	Stability
Imports	Total	-0.3	1.7	*rpm*	pass
	From Non-EU	-0.2	1.3	*rpm*	pass
		-1.6	1.7	*reer*	pass
Exports	Total	-0.1	1.6	*rpx*	fail
	To Non-EU	-0.3	0.6	*rpx*	pass