A DYNAMIC ANALYSIS OF FRANCE's EXTERNAL TRADE

Determinants of Merchandise Imports and Exports
and their Role in the Trade Surplus of the 1990s

by
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I. Introduction

One of the striking characteristics of France's economic performance since the beginning of the present decade has been the strength of the external accounts. In particular, on the basis of national accounts data, since 1989 the current account and the non-energy trade balance have recorded rising surpluses, which in 1996 amounted to 1.7 % of GDP and to close to 1.5 % of GDP, respectively. The service balance, on the other hand, has registered persistent surpluses which averaged between 1 and 1.4 percent of GDP in the 1990s. The largest component of merchandise trade is trade in manufactured goods[1]. The balance on manufacturing trade recorded surpluses throughout the period following the 1973 oil crisis to the late 1980s. The manufacturing trade balance slipped into deficit in the period 1987-1991 but subsequently moved into surplus which rose to almost 1 % of GDP in 1996. Graph 1 presents quarterly data on the evolution of these variables from the beginning of the 1970s to the end of 1996; data adjusted for inflation present a similar picture. It is clear that since movements in the current account are dominated by movements in the non-energy and, in particular, in the manufacturing trade balance throughout this period the sources of the current account improvement in the 1990s are likely to be those explaining the improvement of the manufacturing trade balance.

Awareness of the importance of developments in France's external transactions became more pronounced following the commitment to sustain a stable franc in the ERM. External accounts deficits were considered as leading to devaluations and, consequently, exchange rate stability required that the trade balance was not systematically in deficit, in order not to undermine exchange and the stability[2]. Since the dominant component of France's external transactions is trade in manufactures, the commitment to a stable franc inevitably implied the necessity to strengthen the manufacturing trade balance, principally through improvements in competitiveness. These took the form of cost and price restraint which has had a beneficial effect on export growth and on the manufacturing trade balance, and on supporting the external value of the franc[3].

Graph 1 also shows the importance of manufacturing in France's external and, more specifically, in non-energy trade. Over much of the period since the beginning of the 1970s peaks and troughs in the former coincide with peaks and troughs in the latter, while the level

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[1] Developments in the overall merchandise trade balance over the period since the early 1970s to 1987 have been the result of persistent deficits in energy trade largely offsetting surpluses in trade in manufactures and in agricultural and food products. Since 1985 the deficit in energy trade has been diminishing and by 1994 it was back to its value of the early 1970s of around F 75 billion (FOB-CIF basis). The principal contributor to the merchandise trade deficits of the period 1987-1992 was the deficit in trade in manufactures.

[2] See Bonnaz et Paquier (1993).

[3] The policy of strengthening competitiveness through price and cost restraint rather than through nominal exchange depreciation is associated with the notion of competitive disinflation; Muet (1992) and Blanchard and Muet (1992) discuss these issues in detail.

Graph 1
Non-energy, manufacturing, and current account balance
(percent of GDP)

Source: Commission services
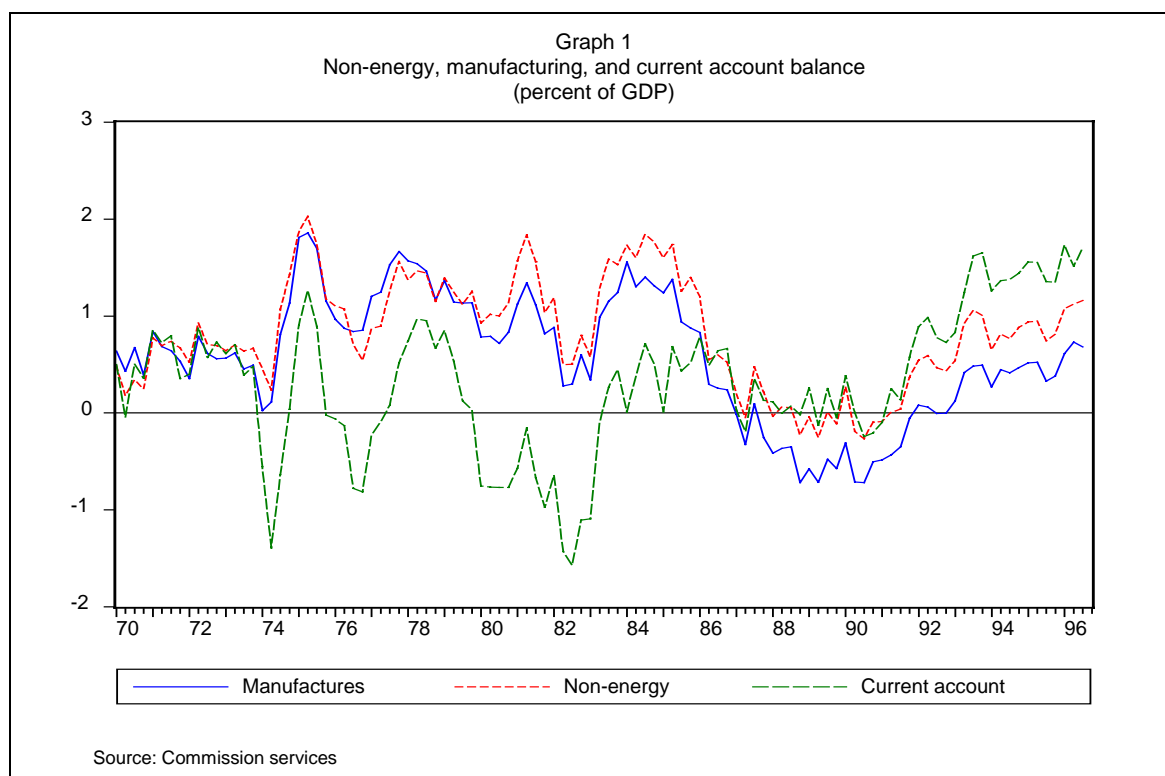
of the manufacturing trade balance accounts on average for virtually the level of the non-energy trade balance. This relationship has been particularly close in the period up to the second half of the 1980s. Since then, a systematic deviation has emerged where the level of the manufacturing trade balance has been lower than the level of the non-energy trade balance. Furthermore, since 1987 improvements in the non-energy trade balance have been larger than those in trade in manufactures where notably larger deficits have been recorded until late 1991. Clearly, marked improvements in non-manufacturing, non-energy trade (that is, trade in agricultural and food commodities) are at the background of these developments. By 1996, the non-energy trade balance had registered surpluses amounting to over 1 percent of GDP while the manufacturing trade balance had recovered from a peak deficit of 0.7 percent of GDP recorded at the end of the 1980s and in the beginning of the 1990s to a surplus of 0.7 percent of GDP in the first three quarters of 1996. This improvement has taken place against a background of turbulence in the ERM marked by the substantial nominal and real depreciations of the exchange rate for the Italian lira, the Spanish peseta, the Portuguese escudo, and the British and the Irish pound against the French franc[4].

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[4] The bilateral nominal French franc exchange rate for the lira and the peseta peaked at 37.9% and 26.2%, respectively, in April 1995 relative to January 1990; the exchange rate for the escudo peaked at 14.9% in December 1995; and the exchange rate for the Irish pound and for sterling peaked at 12.7% and 19.7%, respectively, in November 1995. The Swedish krona was not part of the ERM in the early 1990s when it also experienced substantial depreciations against the ERM currencies. Between January 1990 and September 1992 the krona exchange rate in the European currencies was stable, but from September 1992 to December 1993 the krona/French franc nominal exchange rate had appreciated by 31.4%, and there was another sharp appreciation in April 1995 when relative to September 1992 it had appreciated by 39.4%. Some of this has now unwound but the krona remains around 22% of its French franc value in September 1992. Bilateral real exchange rates, more on which is discussed in section II.b below, show similar behaviour.

There are three significant factors which have undoubtedly contributed to France's external performance in recent years. First, the different cyclical position of France relative to its main trading partners; secondly, relative price developments which have moved to France's advantage or disadvantage principally, but not exclusively, as a result of nominal and real exchange rate changes; and, third, supply improvements which have promoted export expansion and import substitution, principally through gains in cost and price competitiveness but also through improvements in non-price competitiveness associated with changing technology through new investment[5] in the trading sectors of the economy, increased export capacity, productivity-induced relative price changes etc. The impact of the first two factors has likely dominated, especially in short-term developments, the latter's influence on France's trade. Ultimately, however, many supply-side improvements have undoubtedly themselves taken the form of improvements in France's relative costs and relative prices.

The purpose of the present paper is to review the sources of France's trade surplus in recent years and to attribute trade balance movements strictly to those determinants of trade flows suggested by economic theory. These determinants are price and/or cost developments, and demand in France and in the rest of the world. Nominal exchange rate movements in the 1990s have been perceived as playing a significant role in France's trade performance, particularly during the depreciation episodes of 1992 and 1993, since they were considered to have imparted a competitive advantage to those trading partners whose currency had depreciated against the franc; ceteris paribus, and assuming that the nominal depreciation led to a real exchange depreciation, imports would increase and exports would decline, and the trade surplus in real terms would decline. This could have permanent effects on the trade balance since, according to some models of international trade, prolonged exchange rate appreciations, or depreciations can induce hysteresis phenomena (see Baldwin (1988), for example). At the same time, slow growth in France relative to the rest of the world in the 1990s would be expected to have led to a widening of the trade surplus. These two factors have a conflicting impact on trade balance movements, and since income elasticities are generally substantially larger than relative price elasticities it is possible to argue that the emergence of the trade surplus since the beginning of the 1990s is dominated by relative demand movements. A primary objective of the paper is to examine the empirical support for these propositions and to analyse the dynamics of adjustment of trade flows to changes in competitiveness and in relative demand. To do so, a cointegration/error-correction model is applied to flows of both imports and exports, and the estimates of key elasticities obtained are instrumental in shedding light on this question.

The paper, in addition to the introduction, is organized as follows: Section II reviews some salient characteristics of France's trade in recent years in terms of trade patterns, price and cost competitiveness developments, import penetration and export market performance, and in terms of demand developments in France and abroad; section III discusses an accounting decomposition of movements in the non-energy and in the manufacturing trade balance over the period 1990 to 1996 according to the state of price competitiveness and of relative demand; section IV presents the econometric methodology used in the empirical work; section V examines the error-correction model for non-energy and manufacturing imports;

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[5]   Francq et Lamoit (1990) attribute the better export performance of Germany compared to France to Germany's investment dynamism. More investment in the internationally trading sector of the economy makes possible, among other things, greater than otherwise response to demand shocks. Investment in Germany's manufacturing sector during the post-1985 recovery of world trade notably exceeded the corresponding investment in France - see Francq et Lamoit (1990).

section VI examines the possibility of hysteresis in imports through equation stability tests; section VII examines the corresponding model for non-energy and manufacturing exports and section VIII reviews the question of hysteresis in exports again through equation stability tests; section IX uses elasticity estimates from the econometric estimation to shed light on the relationship between the trade balance and demand and competitiveness developments; section X presents simulation results for the non-energy and the manufacturing trade balance since the beginning of the 1990s where the contribution of price competitiveness and of relative demand is evaluated; and, finally, section XI presents conclusions.

There are eight annexes complementing the paper. The sources and the time series properties of the data are presented in Annex A; Annexes B, C and D present additional cointegration results and further evidence on the stability of the import equations; and Annexes E, F, G, and H are devoted to reviewing further cointegration results and to examining the stability of the export functions.

II. Some salient features of France's external performance in recent years

II. a Trade patterns

Most of France's international trade is conducted with other European and, more specifically, European Union States, reflecting the progressive liberalization of the past thirty or so years undertaken in the context of European integration. According to customs data (which are not immediately comparable to the national accounts data shown in Graph 1 and used subsequently in the empirical work), the average share of intra-EU merchandise exports in the decade of the 1970s was 56.6%; as can be seen in Table 1, this share declined marginally in the 1980s, but advanced to 61.3% in the first half of the 1990s. Intra-EU exports of manufactured goods advanced even more markedly between the 1970s and the 1990s, rising by some 6.2 percentage points of the total to 58.6%. The origin of merchandise imports, on the other hand, has shown more diverse patterns. The share of intra-EU imports rose from 54.3% of

Table 1 The composition of non-energy trade and of trade in manufactures (annual averages as percent of total; balances in Ecu million, customs data)						
	Intra-EU			Extra-EU		
	1970-79	1980-89	1990-94	1970-79	1980-89	1990-94
Merchandise exports	56.6	55.9	61.3	43.4	44.1	38.7
of which: Non-energy exports	56.4	55.6	61.2	43.6	44.4	38.8
Exports of manufactures	52.4	52.2	58.6	47.6	47.8	41.4
Merchandise imports	54.3	59.2	63.9	45.7	40.8	36.1
of which: Non-energy imports	64.1	67.7	67.5	35.9	32.3	32.5
Imports of manufactures	73.0	71.4	67.2	27.0	28.6	32.8
Merchandise trade balance	-1004	-12432	-8735	-2541	-2332	2876
of which: Non-energy balance	-464	-8635	-7688	4497	15120	13664
Balance on manufactures	-2563	-12614	-13772	6573	13872	9744

Non-energy is defined as total merchandise exports (imports) net of exports (imports) of fuel products, SITC 3; manufactures are defined by the sum of SITC 6, 7 and 8.

Source: Calculated from original customs data from Eurostat: *External Trade*, various issues.

total merchandise imports in the 1970s to 63.9% in the first half of the 1990s, an increase of almost ten percentage points in twenty five years. However, opportunities to widen the sources of imports have also expanded during this period. Imports of manufactures, the most highly income-elastic component of merchandise imports, have increasingly been sourced

outside the Union, with the latter's share falling from an average of 73% of total imports of manufactures in the 1970s to an average of 67.2% in more recent years.

France's merchandise trade balance was in deficit in the decade of the 1970s, and the largest deficit was recorded with the non-EU world - see Table 1. In the 1980s, the intra-EU deficit widened sharply while the non-EU balance changed from a deficit to a large surplus, which has been sustained into the 1990s. The intra-EU non-energy balance has continued to record deficits during the past twenty five years. These, however, have been offset by large extra-EU surpluses. The balance on trade in manufactures has also been in deficit with the EU trading partners, which has increased over time, but in surplus with the rest of the world, the peak of which occurred in the 1980s. The extra-EU trade surplus in manufactures in recent years has been diminishing while the intra-EU deficit has increased over time, from ECU 2.5 billion in the 1970s to ECU 13.8 billion the first half of the 1990s.

Table 2 shows the changing pattern of France's trade over a quarter century since 1970, depicted by the index of revealed comparative advantage[6]. The index, which is bounded between -1 and 1, is defined as $h = \{X_{ij} - M_{ij}\}/\{X_{ij} + M_{ij}\}$, where $X (M)$ is exports (imports) in nominal terms, i = commodity (non-energy and manufactures, respectively), and j = export destination (import origin). Positive values of the index correspond to a surplus in the commodity in question, negative to a deficit. The index shows that France in the 1970s had a comparative disadvantage with its EU trading partners in both non-energy trade and in trade in manufactures. However, this was largely offset by a comparative advantage against the rest of the world. This pattern has persisted throughout the period under review. France's comparative disadvantage against its EU partners deteriorated significantly in the 1980s, with the value of the index in the case of non-energy trade decreasing from -0.01 in the 1970s to -0.07 in the 1980s, and in the case of trade in manufactures from -0.08 to -0.15. During the same period, however, France improved its comparative advantage in its extra-EU trade, with the index in the case of non-energy trade rising from 0.15 in the 1970s to 0.19 in the 1980s, although in the case of trade in manufactures there was a decrease from 0.35 to 0.26. In its intra-EU trade in the 1990s France saw an improvement in both its non-energy trade and in its trade in manufactures, where in both cases the index of comparative advantage increased by around 50%. However, in its extra-EU trade France experienced a

Table 2 Indices of revealed comparative advantage: Non-energy trade and in trade in manufactures (annual averages)						
	Intra-EU			Extra-EU		
	1970-1979	1980-1989	1990-1994	1970-1979	1980-1989	1990-1994
Non-energy	-0.01	-0.07	-0.03	0.15	0.19	0.10
Manufactures	-0.08	-0.15	-0.08	0.35	0.26	0.10

The index is defined as h = {X_{ij} - M_{ij}}/{X_{ij} + M_{ij}}, where $X (M)$ is exports (imports) in nominal terms, i = non-energy (manufactures) and j = destination (origin); non-energy is defined as total merchandise exports (imports) net of exports (imports) of fuel products, SITC 3; manufactures is the sum of SITC 6, 7 and 8.
Source: Calculated from original data from Eurostat: *External Trade*, various issues.

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[6] The index was devised and first used in Balassa and Nolan (1989), among others. Since only observed data is used, the index reflects the revealed rather than the "true", structural, comparative advantage. The index here is calculated on one-digit SITC and it is perhaps less informative than an index calculated on a more detailed level. The index is ordinal in that the larger its value the greater the revealed comparative advantage is. Clearly, having a comparative advantage in a particular commodity is not sufficient to have a trade surplus in that commodity; for this to happen, demand for this commodity at home must be less than demand abroad.

deterioration, with the surpluses of the 1980s diminishing and, in the case of trade in manufactures the index declined from 0.35 in the decade of the 1970s to 0.10 in the first half of the 1990s. These trends stand in some contrast with the improvement in France's intra-EU trade in manufactures.

These developments indicate that France's non-energy trade surplus of the 1990s is singularly the result of improving trading opportunities outside the EU; within the EU France experienced persistent deficits. Trade in manufactures recorded diminishing surpluses with the rest of the world, but widening intra-EU deficits. We consider next developments in price and cost competitiveness, with particular emphasis on developments in the 1990s, which have undoubtedly played a role in these trade movements.

II. b Developments in price and cost competitiveness

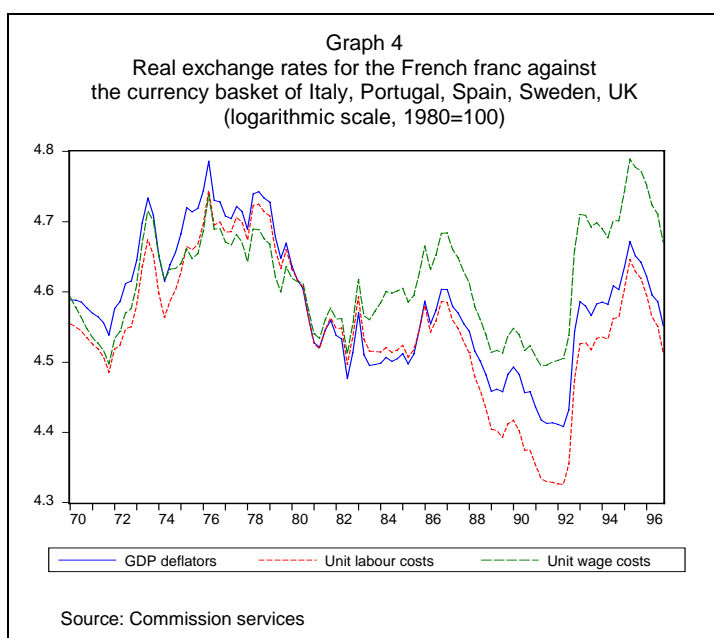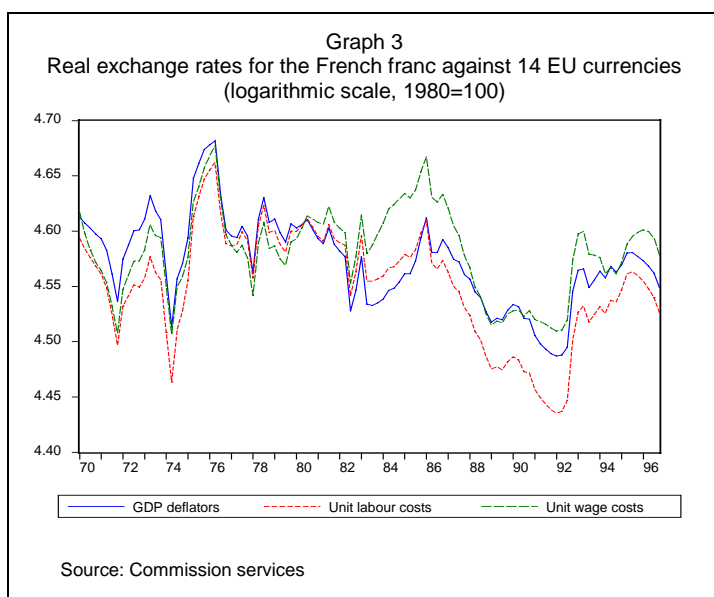Movements in France's price and cost competitiveness since 1970, measured by corresponding real exchange rates, have been erratic. The French franc has been particularly volatile against the basket of IC-23, where since 1986 significant deviations between the different measures of the real exchange rate (relative GDP deflators, unit labour costs, and unit wage costs in manufacturing) have emerged - see Graph 2. These most likely reflect the

Graph 2
Real exchange rates for the French franc against IC-23
(logarithmic scale, 1980=100)

Source: Commission services
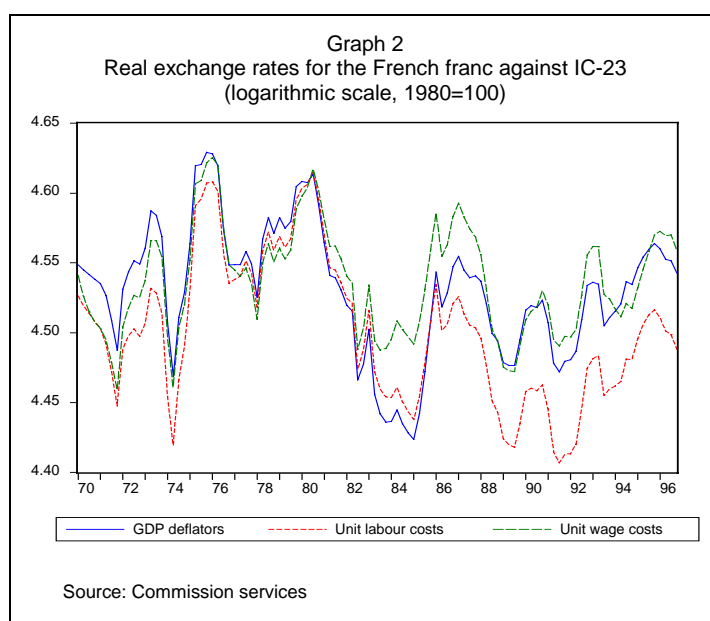
impact of improvements in France's productivity and wage moderation compared to these trading partners. These real exchange rate movements are also dominated by swings in the US dollar. The low point of the series and, correspondingly, the largest gain in competitiveness since 1980 took place between 1980:Q3 and 1989:Q3; During this period, the real exchange rate depreciated by between 13% and 17.9%, depending on the index. Two sub-periods within the overall period can be distinguished. First, the franc depreciated markedly in real terms from the beginning of the 1980s to the beginning of 1985, reflecting the large dollar appreciation of the early Reagan years. With the Plaza accords of the spring 1985 the initial real depreciation of the French franc was reversed. Secondly, there was a considerable loss of competitiveness during the period from the Plaza accords to the February 1987 Louvre accords, but it was also subsequently reversed. Since 1989 much of the early gains in competitiveness against the IC-23 were reversed, and by 1996 the real exchange rate had appreciated by between 7% and 10%.

Trends in price and cost competitiveness measured against the 14 EU partners are shown in Graph 3. Here, the swings in competitiveness have been less wide compared to IC-23, although the timing of peaks and troughs in the real exchange rates series coincide broadly with developments against the IC-23. Limited exchange rate flexibility within the ERM contributed to lessening the volatility of the real exchange rate for the French franc against

Graph 3
Real exchange rates for the French franc against 14 EU currencies
(logarithmic scale, 1980=100)

Legend: GDP deflators — Unit labour costs — Unit wage costs

Source: Commission services

Graph 4
Real exchange rates for the French franc against
the currency basket of Italy, Portugal, Spain, Sweden, UK
(logarithmic scale, 1980=100)

Legend: GDP deflators — Unit labour costs — Unit wage costs

Source: Commission services

the Union currencies. There was a marked deterioration in competitiveness in the early 1980s, but from 1986 onwards France's competitiveness against the 14 other Member States improved dramatically. Between the beginning of 1986 and the beginning of 1992 the real exchange rate for the French franc against the EU trading partners depreciated by 12% in terms of relative GDP deflators, by 16% in terms of relative unit labour costs in the total economy, and by 14.6% in terms of relative unit wage costs in manufacturing. These gains in competitiveness took place within the context of the ERM. Despite realignments during the early part of this period, since 1987 the hardening of the ERM virtually ruled out depreciations and there was none for the French franc[7]. Therefore, these gains in price and cost competitiveness reflected primarily favourable movements in France's prices and costs associated with the policy of competitive disinflation.

Competitiveness gains against a group of EU Member States which subsequently experienced substantial nominal depreciations are shown in Graph 4. The group in question is Italy, Portugal, Spain, Sweden and the UK Between the beginning of 1986 and the beginning of 1992 the real exchange rate for the French franc against this basket of currencies depreciated by 16.4% measured by relative GDP deflators, by 22.5% measured by relative unit labour costs in the total economy, and by 15% measured by relative unit wage costs in manufacturing. It is possible to see the subsequent reversal of these gains as a correction of previous competitiveness misalignments of the French franc relative to these currencies.

The French franc has appreciated markedly in the 1990s, both in nominal and in real terms, and this real exchange appreciation was particularly pronounced following the ERM turbulence of the summer of 1992, of the spring of 1993 and of the spring of 1995[8], thus

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[7]   In fact, between the summer of 1987 and September 1992 there were only a 8% depreciation of the Irish pound and a 3.7% depreciation of the Italian lira.

[8]   The beginning of the 1990s has been a particularly unstable period for the ERM. In September 1992 there was a generalized 3.5% devaluation of all 10 participating currencies against the DM; during the same

reversing a long period of real exchange depreciation which had started in the first quarter of 1986 and lasted until the beginning of 1992. Graphs 3 and 4 show that the real appreciation since the beginning of 1992 has especially taken place against the basket of currencies of Italy, Portugal, Spain, Sweden and the UK (heretofore denoted as EC-5), currencies which have devalued substantially in recent years. Between the first quarter of 1992 and the first quarter of 1996 the real exchange rate for the French franc appreciated by 23.8% measured by relative GDP deflators, by 30.9% measured by relative unit labour costs in the total economy, and by 28% measured by relative unit wage costs in manufacturing. These depreciations should in principle be expected to have adversely affected France's trade in the 1990s. In contrast, however, France's trade performance during this period, measured by the increasing external surpluses, has been very robust, suggesting that perhaps the real appreciation has had little effect on France's trade flows. As will be seen in section IX, this indeed has been the case.

Graph 5
Changes in cost competitiveness and in the import price of manufactures
(4-quarter moving averages of the French franc real exchange rate,
in terms of relative unit wage costs in manufacturing,
logarithmic quarter-to-quarter changes)

Import price of manufactures
Unit wage costs IC-23
Unit wage costs EC-14
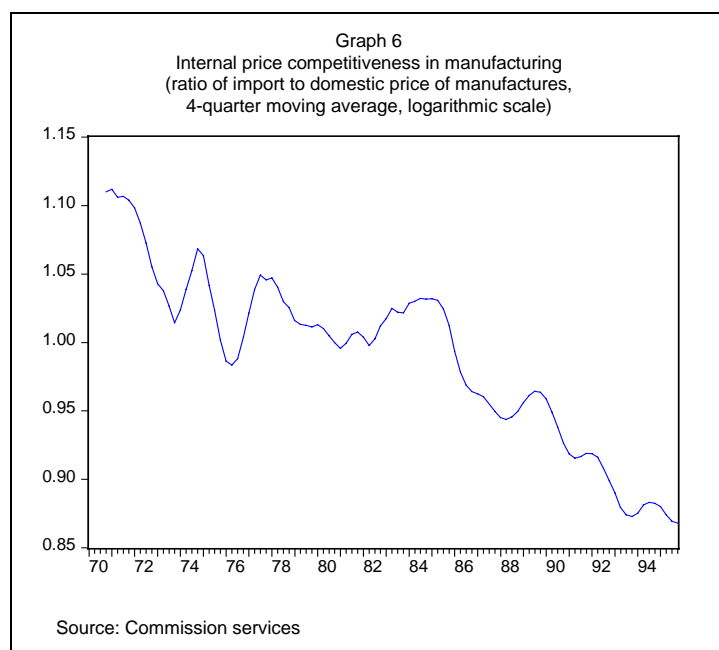Unit wage costs EC-5

Source: Commission services
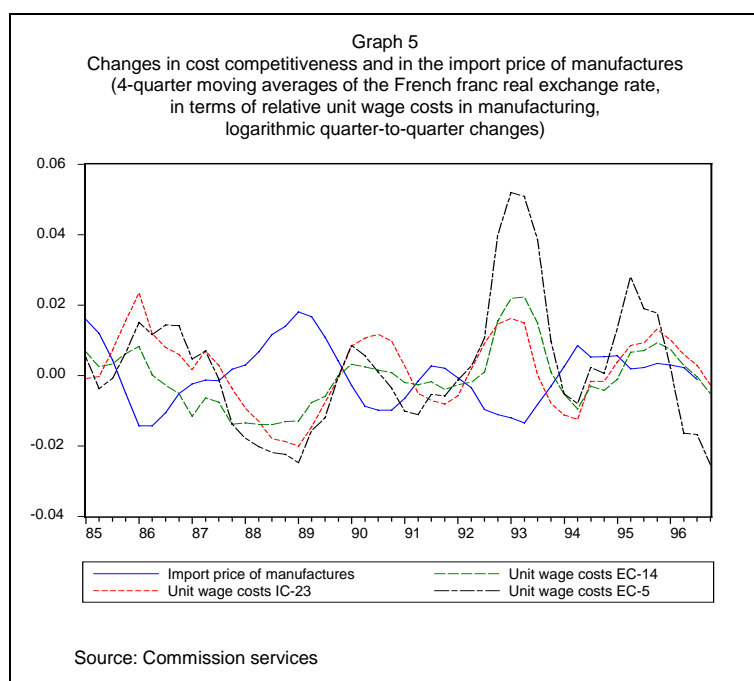
Graph 5 shows the implications of changes in cost competitiveness for inflation in the import price of manufactures over the period 1985:Q1-1996:Q3. The data used are the manufactures' import price, and the real exchange rate of the French franc against the IC-23, the 14 EU partners and the basket composed of the currencies of Italy, Portugal, Spain, Sweden and the UK; a 4-quarter moving average is also used to take account of lags in the effects of competitiveness changes on prices and to iron out the effect of differences in cyclical positions (the same moving average is used in Graphs 6, 7 and 8 below). Import prices appear to have changed less than competitiveness, especially in the post-1992 period and, in general changes in cost competitiveness have been partly offset by opposite changes in import prices. This was the case both during the depreciation of the French real exchange rate in the second half of the 1980s and in the appreciation of the most recent period. Where, however, these movements are particularly pronounced is in the case of the real appreciation against the basket of the five depreciating EU currencies. As can be seen in Graph 5, the large deterioration in cost competitiveness of France against Italy, Portugal, Spain, Sweden

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month sterling and the lira withdrew from the ERM and were devalued substantially, while the peseta was devalued by 5% and two months later by another 6% together with the escudo. In February 1993 the Irish pound was devalued by 10% and in May the peseta and the escudo were devalued by, respectively, 8% and 6.5%, against their central rates. Finally, in March 1995 the peseta lost another 7% and the escudo another 3.5%, against their central rates. The experience of the Swedish krona was more episodic. The krona was following a tight peg to the ECU until it was abandoned in November 1992 when the krona lost around 15% against the DM. By September 1993 the krona had devalued by a cumulative 31% vis-à-vis the DM, and by April 1995 it was down 41.1% compared to its October 1992 value. These devaluations affected very adversely the price and cost competitiveness of France.

and the UK was offset to some extent by reductions in import prices. Indeed, import price inflation in manufactures has been easing since 1989 with some acceleration in the post-1994 period. Nevertheless, it is clear that, responding to the real exchange appreciation, importers likely squeezed profit margins to offset the deterioration in cost competitiveness engendered by the ERM episodes of the 1990s, as they had symmetrically responded to opportunities to raise import prices and restore profit margins in the late 1980s when competitiveness was improving. As a result of this behavioural adaptation, the ratio of import to domestic prices of manufactures has been broadly stable although not invariant to real exchange rate changes. Moreover, manufactures' impor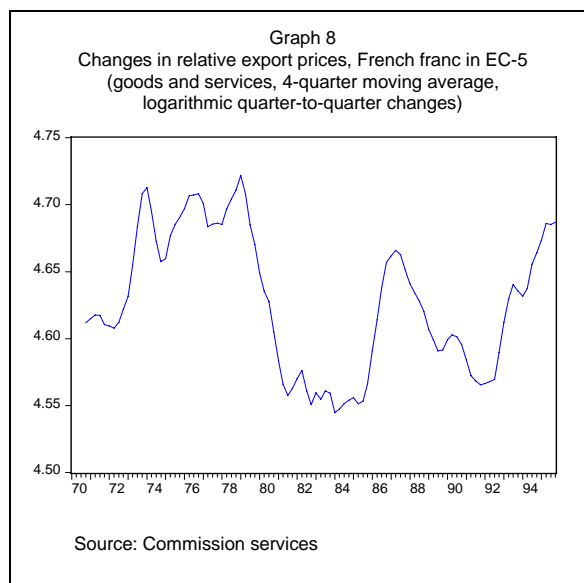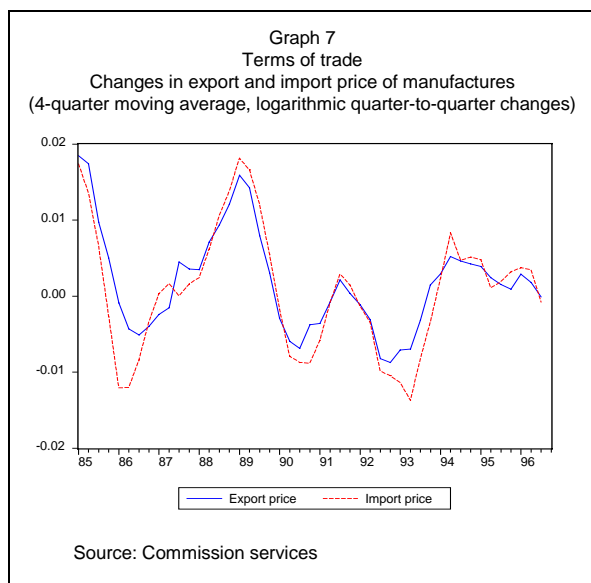t prices have been considerably more volatile than domestic manufactures' prices and clearly, more volatile than changes in competitiveness, as should be expected a priori, reflecting the greater volatility of the nominal exchange rate.

Graph 6
Internal price competitiveness in manufacturing
(ratio of import to domestic price of manufactures,
4-quarter moving average, logarithmic scale)

Source: Commission services

This deterioration in competitiveness is confirmed by another indicator, the ratio of import to domestic price of manufactures in French franc, shown in Graph 6. This indicator has been trending downwards, with some swings, throughout the period under consideration. In the post-1978 period to 1985 internal price competitiveness appeared to have stabilized, reflecting to a considerable extent the impact of the dollar appreciation on this ratio. However, from 1985 onwards the deterioration in internal price competitiveness has been uninterrupted, although some modest stabilization of these trends is taking place in recent quarters. Measured from its peak in 1985:Q1 to 1996:Q3, the import to domestic manufactures price ratio has declined by 14.7%. This should a priori be expected to have encouraged substitution of foreign for domestic goods and to have adversely affected trade performance in manufactures, although there is no doubt some adaptation of domestic producers to competitors' prices also taking place[9].

Real exchange rate movements have undoubtedly affected export prices and export performance. Graph 7 shows that, as far as the ratio of export to import prices is concerned, there has been very little movement in the terms of trade for manufactures; indeed, export prices in domestic currency have changed in an identical manner with import prices, with the exception of the sharper decline of import prices in the beginning of 1996 and in the 1992-93 period. However, the real exchange rate measured by the French export prices of goods and services relative to the respective export prices of EC-5, shown in Graph 8, confirms that

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[9]   Agénor (1995), using similar but more detailed indicators, notes that this indicator of competitiveness is less volatile than the cost and price measures of the real exchange rate; as he correctly points out, this may be a reflection of pricing to market strategies. This lesser volatility is confirmed in our data too. Agénor (1995) also notes that the deterioration of the internal price competitiveness index since 1985 is registered by all sectoral indices of manufactured goods with the exception of transport equipment where competitiveness gains have been registered since the late 1980s - see Agénor (1995), chart 5.

Graph 7
Terms of trade
Changes in export and import price of manufactures
(4-quarter moving average, logarithmic quarter-to-quarter changes)

Source: Commission services

Graph 8
Changes in relative export prices, French franc in EC-5
(goods and services, 4-quarter moving average,
logarithmic quarter-to-quarter changes)

Source: Commission services

there has been a substantial loss of competitiveness in recent years, mirroring the deterioration of competitiveness during the 1984-86 period against these same five Member States.

The behaviour of export prices suggests that French exporters responded to changes in competitiveness in a symmetrical way to foreign suppliers of manufactured goods to France. There is also a notable coincidence of the sharp real appreciation of the French franc in 1992 and the deceleration in the rise of export prices of manufactures. This likely suggests strategic behaviour on the part of French exporters in response to changes in cost competitiveness, or pricing to market, taking the form of a reduction of their export prices to partially or wholly offset the deterioration in competitiveness[10]. Agénor (1995) notes that French exporters trade off competitiveness and profitability, and he quotes evidence that French franc-denominated prices are very sensitive to exchange rate movement; thus, in the case of a franc appreciation, exporters appear to prefer to squeeze profits and to stabilize the local price in the destination market rather than increase export prices and endanger their market share.

The more restrained response of import and export prices to changes in price and cost competitiveness suggests that the competitiveness effects in trade flows may be smaller than suggested by the actual size of the real appreciation of the French franc in recent years. This suggests that the appropriate relative price variable in the empirical work ought to be defined not in terms of the nominal exchange rate and relative costs or prices but by the actual prices in common currency observed in the market. However, this was possible to do only in the import equations - see section V and VII below.

II. c   Import penetration, export/import ratios, and export market shares

The internationalization of the French economy over the past thirty or so years is revealed by the increase in the share of imports in domestic spending. EU and multilateral trade liberalization since the beginning of the 1970s have played an important role in import

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[10]   Clearly, French exporters may also have chosen to diversify towards high value-added products the demand for which is less price elastic; to test this hypothesis more detailed data than presently available would be required.

penetration. At the same time, however, competitiveness developments may also have been a factor in import penetration, but, given the wide swings in competitiveness over this period discussed in the previous section, it is likely that competitiveness has played a less systematic role[11].

Graph 9 presents quantity measures of import penetration for total imports of goods and services, imports of goods, non-energy imports and imports of manufactures over the period 1970 to 1995. The data are ratios of domestic demand in constant (1980) prices. The data show that the rise in the share of imports in domestic demand has been most rapid in the case of non-energy and manufactured goods, undoubtedly reflecting the high income elasticity of demand for such goods[12]. The ratio of imports of goods and services in domestic demand has risen from around 44% in 1970 to around 73% in 1996, an increase of some thirty percentage points; the ratio of imports of goods in domestic demand has risen over the same period by some 26 points to 65% in 1996; the share of non-energy imports in domestic demand has risen by around 31 points to over 55%; and the share of imported manufactured goods in domestic demand has risen by around 28 percentage points to around 48% in 1996. Amongst imports of manufactured goods, imports of household equipment goods is the largest component, followed by machinery and transport equipment while current consumption goods was a smaller category[13]. The data 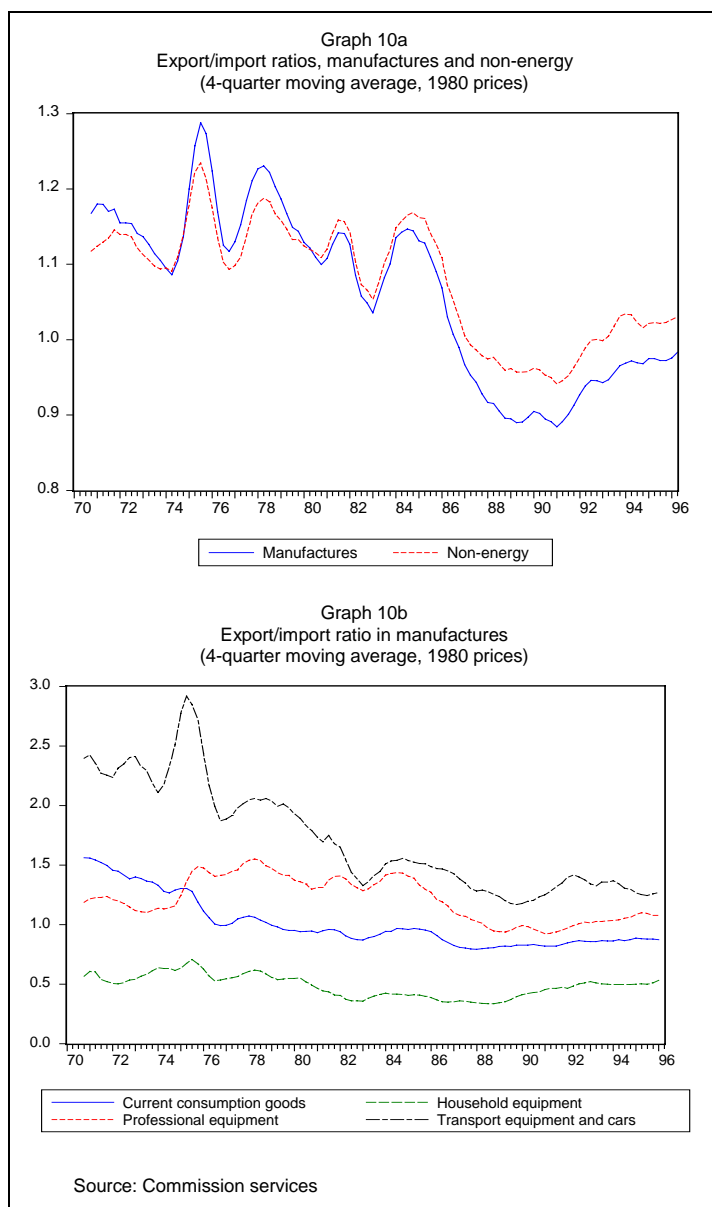in Graph 9 appear to suggest that, despite the increasing openness of the French economy over the past quarter century, the rate of growth of import penetration has slowed down in the 1990s, perhaps reflecting the impact of the slow growth and of the 1993 recession as well as the real appreciation of the French franc during this period on the demand for imports.
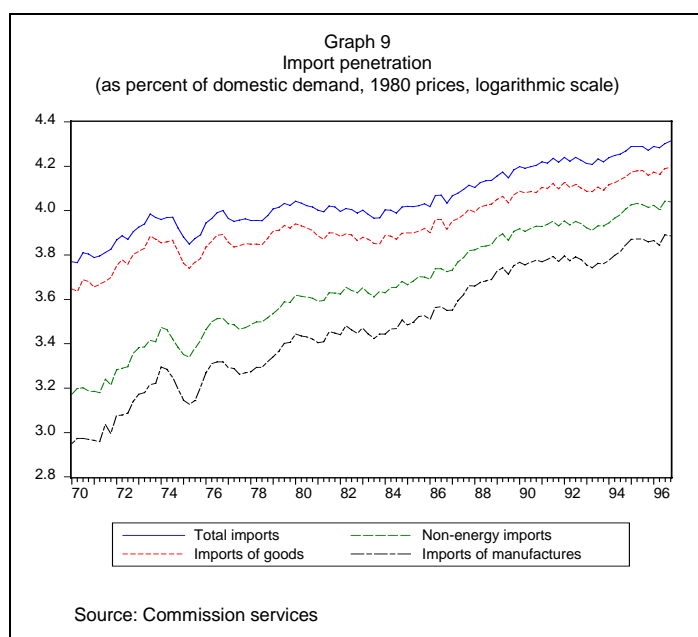
Graph 9
Import penetration
(as percent of domestic demand, 1980 prices, logarithmic scale)

Legend:
Total imports —— Non-energy imports - - - -
Imports of goods - - - - Imports of manufactures - - - -

Source: Commission services

The relative performance in exports and imports, another indicator of competitiveness, is presented in Graphs 10a and 10b. The data represent 4-quarter moving averages of export/import ratios for non-energy and for manufactured goods (Graph 10a) and for four categories of manufactured goods in constant prices (Graph 10b). There has been a smooth decline in the non-energy and manufactures ratios since their peak of 1975. These ratios,

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[11] It is possible, however, that competitiveness misalignments, which lead to trade hysteresis, have also been a significant factor in import penetration. In trade hysteresis the loss in competitiveness during the 1990s, for example, may have permanently raised import penetration, as the gains in competitiveness during the dollar appreciation of the early 1980s may have permanently lowered import penetration; see Baldwin (1988). Hysteresis in trade volumes is discussed in sections VI and VIII below.

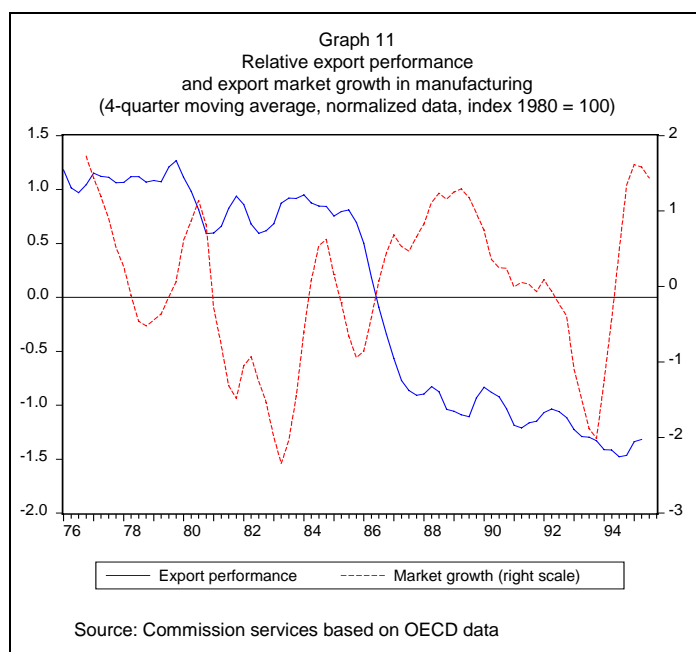[12] Smoothing the data in the sample 1970:Q1-1996:Q4 yields a trend value of 0.005 for total imports, 0.004 for imports of goods, 0.009 for non-energy imports and 0.01 for imports of manufactures.

[13] According to data in Agénor (1995), chart 6, household equipment goods represented over 60% of domestic demand in 1994, machinery and transport equipment was around 50%, and current consumption goods were around 30%.

Graph 10a
Export/import ratios, manufactures and non-energy
(4-quarter moving average, 1980 prices)

— Manufactures   ----- Non-energy

Graph 10b
Export/import ratio in manufactures
(4-quarter moving average, 1980 prices)

— Current consumption goods   ----- Household equipment
----- Professional equipment   —·— Transport equipment and cars

Source: Commission services

while fluctuating around a downward trend, reached another peak in the period of the dollar appreciation during the early 1980s. Subsequently, however, coinciding with the dollar depreciation, there was a sharp deterioration in the export/import ratio which lasted until the beginning of the 1990s. Between 1988 and 1991 export performance relative to imports stabilized, and since 1991 there has been some continuous and considerable recovery in export performance compared to imports, although the export/import ratio remains substantially below its values in the pre-1985 period. In a manner parallel to the 1985-88 period, the franc appreciation of recent years could be expected to have affected negatively the export/import ratio by reducing exports and encouraging imports. The data suggest that this has not been the case. Since the early 1990s export performance has outstripped import growth with the result that there has been a marked recovery in the export/import ratios.

Graph 10b shows that the deterioration of the export/import ratio for manufactures masks important differences at a more disaggregated level. Of the four categories of imports shown, the most marked deterioration is in the case of transport equipment and automobiles. While there has been little change in the ratio for household equipment up to the late 1980s, it has subsequently shown a modest increase. There was a notable increase in the export/import ratio for professional equipment starting in the aftermath of the first oil shock and lasting until the Plaza accords; it subsequently decreased but again since the late 1980s there has been some modest recovery. The ratio for current consumption goods has also declined in the early part of the sample, but it has stabilized since the beginning of the 1980s and has been virtually flat since then. The data in Graph 10b confirm the recovery in the relative export/import performance since the beginning of the present decade presented in the aggregate data of Graph 10a.

Graph 11
Relative export performance
and export market growth in manufacturing
(4-quarter moving average, normalized data, index 1980 = 100)

Export performance ----- Market growth (right scale)

Source: Commission services based on OECD data

Graph 11 presents alternative indicators of export performance for the French manufacturing sector. The data, which is from the OECD[14], shows France's export market growth and her relative export performance in manufactures. Export market growth is a weighted average of all imports, in volume terms, by all trading partners to which France exports; relative export performance measures the ratio of France's export volumes to the export market. As can be seen in the Graph, market growth fluctuates widely, consistent with cyclical developments in the trading partners' economies.
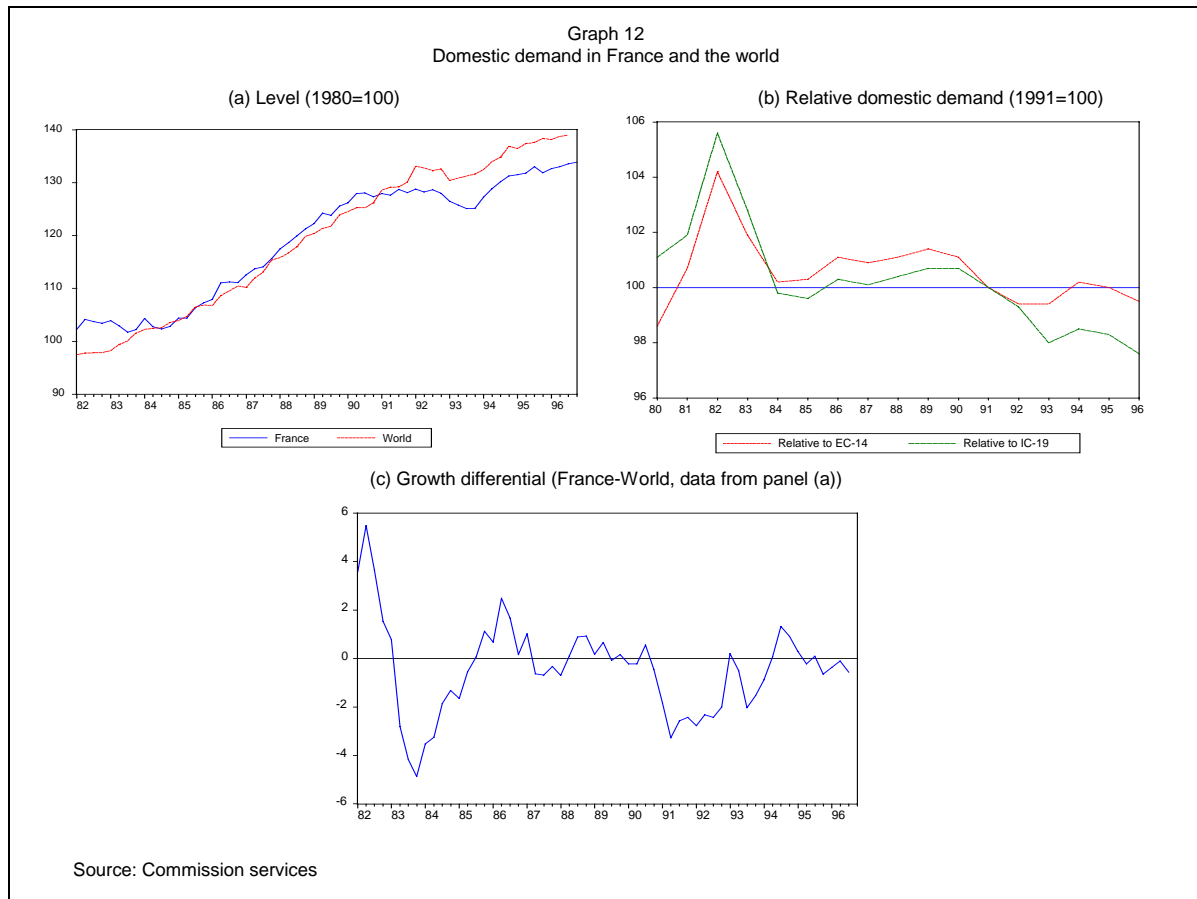
France's market share declined sharply in the beginning of the 1980s, then it gained ground until the late 1980s when there was another sharp loss of market share in the beginning of the 1990s. Since 1993, however, market growth has rebounded significantly. Export performance, on the other hand, which contrasts potential to actual exports, shows that export performance has deteriorated markedly in the period since the Plaza accords. Notwithstanding the sharp increase in market growth in the post-1983 to the late 1980s period, France has experienced a trend decline in its export performance in the sense that actual exports have fallen short of the export market growth. The deterioration was particularly pronounced until 1989 and it has continued, albeit at a slower pace, beyond 1994.

Developments in market share and in export performance in manufactures may not be correlated closely with developments in price and cost competitiveness discussed earlier. First, there are long and possibly variable lags between changes in apparent competitiveness and changes in trade flows; secondly, changes in cyclical positions most certainly dominate price and cost competitiveness in trade flows adjustments; and, thirdly, as noted previously, exchange rate and cost indices of competitiveness disregard the fact that manufactures are imperfect substitutes with other internationally traded and also domestically produced manufactures and, as a result, they fail to acknowledge pricing-to-market behaviour. As a result, trade flows adjustments in response to changes in competitiveness and in relative cyclical positions can only be approximated by a dynamic model of the kind used in sections V and VII.

II. d   Demand developments in France and abroad

The cyclical position of France relative to the rest of the world has been an important determinant of movements in the external accounts; it is also likely that this factor has played a dominant role in the surplus of the 1990s since empirical evidence suggests that the

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[14]   OECD (1996); for a review of the construction of these data and other technical documentation see Durand et al. (1992).

Graph 12
Domestic demand in France and the world

(a) Level (1980=100)

(b) Relative domestic demand (1991=100)

(c) Growth differential (France-World, data from panel (a))

Source: Commission services

demand elasticity of trade flows is substantially larger than the relative price or competitiveness elasticity. Clearly, small movements in France's cyclical position against the rest of the world can have large effects on the trade balance, such that only unreasonably large movements in competitiveness would generate; as will be seen later, the evidence collected in section IX supports well this hypothesis.

Much of the 1970s to the beginning of the 1980s domestic demand in France and abroad developed in step. Deviations from these trends emerged in the 1980s and became particularly pronounced in the 1990s. Graph 12 presents developments in domestic demand in France and in the rest of the world (the exact composition of the world demand variable is discussed in section VII, footnote 58), as well as on Commission data for France and the IC-19 and the EC-14 (index of double export weighted national final uses including stocks in 1990 prices, annual data), over the period 1980:Q1 - 1996:Q3; the level of the indices is shown in panels (a) and (b), respectively, and the growth differential (4th/4th) between France and the rest of the world based on the data of panel (a) is shown in panel (c). In the second half of the 1980s the level of demand in France grew in parallel with demand abroad and the growth differential fluctuated around zero. However, since 1990 the index for domestic demand in France has started to deviate significantly from its past trend and to fall below the corresponding index for world demand, which has developed broadly in line with its trend of the past decade. The Commission data confirm that while the drop in domestic demand in France was not particularly marked relative to her EC trading partners, it was notably greater when measured against the IC-19. The growth differential became

particularly large from the beginning of 1991 onwards, peaking at 3.3% in 1992:Q2[15]. This negative differential persisted for about 13 quarters until the second quarter of 1994; subsequently, however, growth in world domestic demand outstripped growth in France over much of the period between the first quarter of 1995 and the third quarter of 1996. This is the period when France's external surplus also developed.

Movements in the external balance are highly correlated with differences in domestic demand growth between France and the rest of the world. Over the period 1970:Q1 - 1996:Q3 the simple correlation coefficient for the non-energy trade balance is -0.61, for the manufacturing trade balance -0.63, and for the current account balance -0.62. These coefficients are suggestive of the substantial dependence of the external accounts on cyclical developments in France and abroad. Large and sustained deviations of France's demand growth from the rest of the world would tend to reduce (increase) imports and raise (lower) exports, thus contributing to an incipient increase in the external imbalances.

III. Factors underlying the trade surplus of the 1990s

The current account, merchandise trade, and manufacturing trade surpluses of the 1990s have been principally the result of trade volume changes rather than price changes. To decompose the contribution of volume and price effects in movements in these balances, the following identity is used:

$$B_t = Px_t * X_t - Pm_t * M_t \tag{1}$$

where B is the balance in question, Px and X are the export price deflator and the volume of exports, and Pm and M is the import price deflator and the volume of imports, respectively. Totally differentiating (1) and collecting terms, we have the following decomposition:

$$dB_t = [dX_t * Px_t - dM_t * Pm_t] + [dPx_t * X_t - dPm_t * M_t]$$

$$+ [dPx_t * dX_t] - [dPm_t * dM_t] \tag{2}$$

where the first difference operator is $d = Z_t - Z_{t-1}$. The first term in brackets is the export and import volume contribution, respectively, to adjustments in the balance; the second bracket provides the contribution of the terms of trade changes; and the third and fourth terms are cross-terms of the interaction of volume and price changes. Applying this decomposition to the French, national accounts, trade data for the period 1990-95 we obtain the results reported in Table 3[16].

Over the period 1990-95 the non-energy trade balance registered a cumulative surplus of FF 115.5 billion, and the manufacturing trade balance a cumulative surplus of FF 111.5 billion; thus, the dominant component of the surplus in recent years derives from the good performance in trade in manufactures. This is clearly confirmed by the cumulative change in exports of manufactures during this period which, at FF 284.6 billion, were equivalent to

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[15] Differentials of comparable size have coincided with unstable macroeconomic policies; large demand growth differentials in France's favour were recorded during the period 1981:Q4-1982:Q4 of the Mitterrand socialist expansion; the Delors reversal of this expansion led to correspondingly large negative growth differentials in the period 1983:Q2-1985:Q1.

[16] See Mastropasqua and Vona (1988) for a similar decomposition of the US trade balance movements in the 1980s.

87.3% of the cumulative change in non-energy exports. The cumulative increase in non-energy imports is also dominated by imports of manufactures which, in the period 1990-95 represented approximately 84% of the cumulative change in non-energy imports.

The data in Table 3 show that the largest rise in the surplus occurred in 1992 and also in 1993, the years of slow growth and of recession. While there has been some positive contribution from terms of trade effects throughout the period, these were also particularly prominent in 1993. It is clear that during 1992-93 export prices were more resilient to the decline in demand, evidenced by the fall in export volumes, than import prices which likely responded more vigorously to the fall in French imports (by FF 50 billion in the case of non-energy imports and by FF 57 billion in the case of imports of manufactures).

These beneficial effects of the recession on the trade balances were reversed, but on a modest scale, with the recovery of the post-1993 period. The non-energy trade balance deteriorated by some FF 8 billion in 1994, notably as a result of the sharp recovery in the demand for imports which outstripped the rise in exports. In manufacturing trade, however, the rise in

Table 3 Decomposition of non-energy and manufacturing trade balance (in billion of francs)							
	1990	1991	1992	1993	1994	1995	Cumulative
Non-energy trade							
Change in trade balance	2.3	16.6	47.8	43.0	-7.8	13.6	115.5
due to: *Export volume*	56.7	45.8	57.5	-15.8	80.3	101.4	325.9
Import volume	67.0	29.2	11.6	-49.7	90.9	85.3	234.4
Terms of trade	11.7	0.1	3.1	10.3	3.0	2.4	30.6
Other	0.9	-0.2	-1.5	2.2	-1.1	-1.6	-1.3
Residual	0.0	-0.1	-0.3	-3.4	-0.9	-0.9	-2.5
Trade in manufactures							
Change in trade balance	-0.8	23.0	39.3	39.3	4.2	6.5	111.5
due to: *Export volume*	50.2	40.1	44.7	-31.5	87.9	93.2	284.6
Import volume	60.3	18.1	9.4	-56.7	81.1	84.6	196.8
Terms of trade	8.6	0.5	4.7	19.4	-2.5	-2.0	28.7
Other	0.7	0.3	0.7	2.8	-0.3	-0.1	-4.1
Residual	0.0	-0.2	-1.4	-2.5	-0.2	0.0	-4.3
The entries in the table are based on the decomposition, according to equation (2), discussed in the text; "Other" refers to the cross-terms in equation (2) of the text. Source: Calculations based on original INSEE data; see Annex A.							

exports was significantly larger than the rise in imports with the result that the manufacturing trade balance has continued to record improvements albeit substantially smaller than those of the 1991-93 period.

A notable feature of the data is that the deterioration in price and cost competitiveness during the 1990s does not appear to have contributed adversely to trade balance movements. With the exception of 1994 and 1995 in the case of manufactures, the terms of trade effects have made a positive but relatively small contribution to changes in the trade balance in the 1990s. It is possible to note that should competitiveness have remained at the same level as in the late 1980s the potential improvement in France's trade balance, ceteris paribus, would not have been affected significantly. On the other hand, had domestic demand remained at its trend value during the second half of the 1980s, there likely would have been, ceteris paribus, no improvement in France's external performance. Finally, together with the evidence suggested in Graph 11, it is likely that had export performance kept up with France's international market growth, certainly in the post-Plaza accords period as a whole and more specifically during the post-1989 period, France's external performance would have been markedly better.

The decomposition of Table 3 is only indicative of the behavioural adjustments in export and import functions in response to the competitiveness and output shocks which have occurred in recent years. An exact quantitative evaluation requires estimates of the export and import functions with respect to the determinants postulated by economic theory. Moreover, as the impact of these shocks on trade flows is likely not instantaneous, it is essential to obtain estimates of the lags which typically characterize adjustments in trade flows[17]. Finally, one method of uncovering the sources of the external surplus of recent years would be through the use of simulations. For these reasons, a more thorough econometric estimation approach is undertaken in subsequent sections.

IV. Econometric methodology and modelling strategy

This section reviews the main methodological issues that impinge upon the specification and econometric estimation of models of trade flows.

IV.a Specification of trade functions

Conventionally, empirical analysis of trade flows has been carried out through a partial-equilibrium model based on the hypothesis of imperfect substitution between foreign and domestic goods. The model assumes that, in a simplified two-country world, each country produces a single tradable good that is an imperfect substitute for the good produced in the other country. The simplest and most widely used procedure for estimating aggregate export and import demand functions in this framework[18] is based on the Marshallian demand function.

The general function for aggregate imports has the following theoretical form:

$$M_d = F(Y, P_m, P_d) \tag{3}$$
$$f_1 > 0, \ f_2 < 0, \ f_3 > 0$$

where M_d = volume of imports demanded, P_m = price of imports, P_d = domestic price of import-competing goods, and Y = real domestic economic activity; f_i are the expected partial derivatives with respect to the ith argument. Equation (3) suggests that the demand for imports is determined by the level of expenditure or income (or another scale variable that captures domestic demand conditions) and by the price of imports and domestic substitutes measured in the same currency[19]. With homogeneity of degree zero in prices and nominal income, consistent with absence of money illusion, equation (3) can be re-written with a relative price term:

$$M_d = F(Y, RP) \tag{4}$$

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[17]   One important effect, for example, is associated with the lagged response of volumes and the instantaneous response of prices in the event of competitiveness shocks, giving rise to J-curve movements in external balances; however, examination of such effects is beyond the scope of the present study.

[18]   For an excellent but somewhat dated survey of the literature on imperfect substitutes and perfect substitutes models in international trade, the latter implying the law of one price, see Goldstein and Khan (1985).

[19]   This is in accordance with conventional demand theory that assumes the consumer to maximize utility subject to a budget constraint and, for the case in which the importer is a producer, to maximize production subject to a cost constraint; see Goldstein and Khan (1985).

$f1 > 0, \; f2 < 0$

Correspondingly, the general function for aggregate exports has this theoretical form:

$$Xd = F(Yw, Px, E*Pw) \tag{5}$$

$$f_1 > 0, \; f_2 < 0, \; f_3 > 0$$

where X_d = volume of exports demanded by foreigners, Y_w = world economic activity in constant prices, P_x = price of exports, P_w = foreign competitors' prices in the country's export markets and E = nominal exchange rate in units of foreign currency per unit of home currency; f_i are the expected partial derivatives of the export function with respect to the ith argument. According to equation (5), the foreign country's demand for exports is a function of its real income or expenditure, of the price of the domestically produced substitute good and of the price of the foreign competing good. As in equation (4), equation (5) can be re-written with a relative price term $(P_x/E*P_w)$ which can be viewed, alternatively, as the terms of trade or the real exchange rate. The use of this variable imposes the constraint of homogeneity in prices which is not testable.

$$Xd = F(Yw, Px/E*Pw) \tag{6}$$

$$f_1 > 0, \; f_2 < 0$$

The supply side function of the imperfect substitute model postulates that the quantity supplied is a positive function of the own price and a negative function of the price of domestic goods in the exporting or importing country. The price elasticity of supply is often assumed to be infinity or at least largely independent of the quantity of trade. This allows estimation of a single equation, reduced form, model by ordinary least squares since export and import prices can be considered as exogenous. As stressed by Goldstein and Kahn (1985), this assumption appears more plausible for the case of imports rather than for the case of exports since for a single country it is more difficult to increase its supply of exports without an increase in costs and thus in prices compared to the rest of the world .

Another basic assumption of the model is that exporters and importers are always on their demand schedules so that demand always equal the actual level of trade flows. However, it is generally recognised that imports and exports do not immediately adjust to their long-run equilibrium level, following a change in any of their determinants. To take into account the slow reaction of the economic agents to changes in the explanatory variables due to adjustment costs, inertia, habit or lags in perceiving changes, a simple first-order partial-adjustment mechanism has often been included in the equations, permitting a short-run adjustment towards the long-run equilibrium. More generally, the relaxation of the equilibrium hypothesis in the general function has usually been achieved by specifying a dynamic model within the framework of a general distributed model with geometrically declining weights (the Koyck model) or, following Almon, by assuming that distributed lag coefficients lie on an exact polynomial curve[20].

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[20] One of the drawbacks of the partial-adjustment model is that it imposes the same geometrically weighted lag distribution on both the real income and relative price variables, while trade flows are known to

IV.b Econometric methodology

Most of the empirical work on the OLS estimates of aggregate and disaggregated trade flows has used variables in (log) levels rather than first-differenced or otherwise filtered data. Yet, in the case of non-stationarity[21] of the time series involved in the analysis, estimated coefficients are likely to be inconsistent. In addition, the assumption of conventional asymptotic econometric theory does not hold and the use of standard statistic tests is not valid. Most often, to cope with the presence of autocorrelation in the residuals (arbitrarily interpreted as indicative of autoregressive errors) that often arises when quarterly data of variables in level are used, a re-estimation with Cochrane-Orcutt or other serial-correlation correction methods has been carried out. Yet, this procedure is justified and valid only in presence of a "common factor", otherwise it will give rise to inconsistent estimates[22].

In some cases, to avoid obtaining inference from spurious regressions[23] and problems in the error process, only first-differenced time series have been employed in the estimation. However, such a modelling is valid only if the time series are not cointegrated. If the series are cointegrated, a model with only differenced variables will be misspecified and the estimates will be biased and inconsistent. Furthermore, such models do not include information about the long-run equilibrium relationship between the variables.

The parametric error-correction methodology avoids the shortcomings of the conventional least squares regression. We have examined aggregate trade demand equations using the following two-stage procedure. First, we have tested and estimated the cointegrating relationship, by applying the Johansen maximum likelihood methodology to the aggregate non-energy and manufacturing import and export functions derived from the imperfect substitution model. Then we have imposed this long-run result as an error correction

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respond more rapidly to income changes than to price changes. Moreover, this model arbitrarily imposes rigid conditions, such as a specific pattern to lag structure estimates.

[21] We adopt, as usual, the notation that a non-stationary variable X_t is integrated of order d (i.e. *I[d]*) if it requires to be differenced d times to become stationary *(I[0])*, i.e. to remove the non-deterministic trend. Alternatively, if adding a time trend is sufficient to induce stationarity, the series is defined to be trend stationary. If $X_t \sim I[1]$, then $\Delta X_t \sim I[0]$, and when the data are in logarithms these changes (i.e. $\Delta logX_t = logX_t - logX_{t-1} = log(X_t/X_{t-1})$) are growth rates.

[22] More formally, a linear regression equation relating a variable Y_t to its own lagged value and to the current and lagged values of a regressor variable, X_t: $Y_t = \alpha Y_{t-1} + \beta X_t + \delta X_{t-1} + \mu_t$, where $|\alpha|<1$ and $\mu_t \sim (0,\sigma^2)$, can be transformed into: $Y_t = \alpha Ly_t + \beta X_t + \delta X_{t-1} + \mu_t$ using the lag operator L (where $LY_t = Y_{t-1}$), or $(1-\alpha L)Y_t = (\beta+\delta L)X_t + \mu_t$. If the parameter $\delta = -\alpha\beta$, then the last equation becomes: $(1-\alpha L)Y_t = \beta(1-\alpha L)X_t + \mu_t$ where $(1-\alpha L)$ is the common factor of the terms involving Y_t and X_t, and implies an autoregressive error process. In fact, dividing both sides of the last equation by the common factor, we obtain: $Y_t = \beta X_t + \mu_t/(1-\alpha L) = \beta X_t + \varepsilon_t$ where $\varepsilon_t = \mu_t/(1-\alpha L)$, or $\varepsilon_t = \beta\varepsilon_{t-1} + \mu_t$, which represents the first order autoregressive process generating the error term. Only in this case (that is when the condition $\delta = -\alpha\beta$ holds) the use of Cochrane-Orcutt transformation (which, indeed, imposes a common-factor restriction assuming an *AR* error process) for removing a first order autocorrelation in the residuals may be optimal, and serial correlation can be considered a convenience rather than a nuisance, because it allows estimates of only three parameters $(\alpha, \beta$ and $\sigma^2)$, instead of the four parameters $(\alpha, \beta, \delta$ and $\sigma^2)$ of the first model; see Hendry and Mizon (1978).

[23] Spurious regressions refer to relationships between the variables in an equation only due to correlated stochastic time trends in the data.

mechanism to a single-equation-error-correction-model *(SEECM)*[24] and we have estimated the short-run dynamics.

The explicit *SEECM* specification of the trade equations is:

$$\Delta m_t = A(L)\Delta m_{t-1} + B(L)\Delta y_t + C(L)\Delta rp_t + \theta[m_{t-1} - m^*_{t-1}] + \mu_t \quad (7)$$

$$\Delta x_t = A(L)\Delta x_{t-1} + B(L)\Delta yw_t + C(L)\Delta reer_t + \varphi[x_{t-1} - x^*_{t-1}] + \varepsilon_t \quad (8)$$

where *A(L), B(L)* and *C(L)* are finite polynomials and Δ is the first difference operator. The term in squared brackets is the error-correction term, that is the deviation of actual import and export demand, m_t and x_t, from the long-run equilibrium or cointegrating relation, m^*_t and x^*_t, respectively. The equations have been estimates in log functional form, so that the coefficients are elasticities.

IV.b.1 Unit root analysis

Before applying the cointegration and the error-correction methodology mentioned above it is necessary to determine the time-series properties (i.e. the order of integration) of each variable, by testing whether they are stationary or they include a stochastic trend[25]. We employed the most commonly used Dickey-Fuller (DF) and augmented Dickey-Fuller (ADF) univariate tests of the null hypothesis of a unit root in our observed time-series, against the alternative that the process is stationary[26]. In computing the ADF test we have employed up to four lags to remove residual autocorrelation in the quarterly data used here.

The results of the two tests are reported in Annex A and show that at the 5% significance level the null hypothesis of a unit root in the series under consideration cannot be rejected at least in the sample period under examination (1970:Q1-1996:Q4)[27]. The only exception is the real effective exchange rate defined in terms of unit wage costs in manufacturing against 23 industrial countries and 20 industrial countries which appear to be *I(0)*.

To examine the data for the presence of a second unit root, the DF and ADF tests were applied to the first differences in logs of the time series. The results indicate that the presence of a second unit root is easily rejected; therefore, the first difference of all the series under consideration is stationary, thus confirming that the series are likely to be *I[1]* in (log)level. Hence, with usual caveats arising from the low power (the tendency to under-reject the null of non-stationarity when it is false) and poor size (over-rejecting the null hypothesis when it is true) of the tests in finite sample, and the problem of "near

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[24] The *SEECM* can be considered as a generalization of conventional partial adjustment model widely used in the specification of import and export demand functions, and is consistent with optimizing behaviour of economic agents in a dynamic environment.

[25] A stochastic variable Y_t generated by a first-order autoregressive process $Y_t = \rho Y_{t-1} + \mu_t$ is stationary if $|\rho| \langle 1$. A stationary time-series tends to return to its mean value and fluctuate around it with a finite variance.

[26] For a survey on new extension of the Dickey-Fuller procedure and other alternative non-parametric tests, such as Phillips and Perron test, see the introduction of Banerjee and Hendry (1992) and the following three articles in the special issue of the Oxford Bulletin of Economics and Statistics.

[27] In fact, the order of integration of a time-series is sensitive to the period over which the tests are performed. The degree of integration is not necessarily an intrinsic property and could change over historical periods.

observational equivalence"[28], it seems reasonable to carry out the analysis assuming that all the variables in our information set, with the exception of the real exchange rates mentioned previously, are non-stationary, i.e. they contain a stochastic trend over the 1970:Q1-1996:Q4 sample. Accordingly, standard asymptotic theory can not be applied estimating equations containing variables in level because the properties of the series do not satisfy the classical assumptions of constant mean and variance which, on the contrary, evolve with time; standard errors of the parameters will be biased and could give rise to "spurious" regressions in the non-cointegrated case[29].

The same problem could arise if the levels of non-stationary variables were introduced in regressions that are formulated in differences (as it is done modelling an unrestricted error correction model). In fact, if the regression is to have stationary residuals in order to avoid spurious regression when the variables are integrated, there must exist at least one linear combination of the levels of variables which is stationary[30]. If it does exist, the variables are said to be cointegrated and the linear combination can be interpreted as a long-run equilibrium relationship which holds apart from a stationary stochastic error representing short-run deviations. Therefore, we have first tested for the presence of cointegration among the variables of interest and then formulated the error-correction dynamic model.

IV.b.2 Cointegration analysis

The cointegration methodology[31], first proposed by Granger (1981), Engle-Granger (1987) and extended by Johansen (1988), is now commonly used in the construction of single-equation dynamic models. In our case, cointegration has been used to evaluate the long-run stationary steady-state between the level of imports and exports and their theoretically most important determinants.

When there are more than two *I[1]* variables under consideration, as in our case, the most common cointegration analysis proposed by Engle and Granger (EG), based on a OLS static regression and on the DF and ADF cointegration tests of the residuals (using correct critical values), has proved to be inefficient[32]. In finite samples the EG method is sensitive to the so-called direction normalisation rule, that is to the specific choice of the endogenous variable to put on the left-hand side of the equation. The method also ignores the possibility of more than one cointegrating vectors when more than two variables are included in the analysis. In addition, the ADF test for cointegration imposes an implicit common factor restriction on the

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[28]   In finite samples, any trend-stationary process could be approximated arbitrarily well by a unit root process and vice versa; in addiction, the usual tests are not able to reject the null hypothesis if the deterministic trend of the series has a single break. For this and other criticisms see Campbell and Perron (1991).

[29]   On this issue see Granger and Newbold (1974), Dickey and Fuller (1979) and Engle and Granger (1987).

[30]   In a bivariate case, two variables *(X, Y)* are said to be cointegrated of order one, i.e. *CI[1, 1]*, if they are individually *I[1]*, but some linear combination (i.e. $X+\alpha Y$) of the two is *I[0]*; see Engle and Granger (1987).

[31]   Cointegration is a statistical property that describes the long-run behaviour of economic variables and provides a formal underpinning to the use of the error-correction model (ECM) as it has been demonstrated in the Granger Representation Theorem. A treatment of this topic is in Banerjee et al. (1993).

[32]   Stock (1987) has shown that if each variable is *I[1]* and if there is a cointegrating vector such that the linear combination is *I[0]*, then the OLS estimators of this vector are consistent and converge in probability faster than in ordinary case, that is they are "super consistent". On the contrary, Banerjee et al. (1986) found with Monte Carlo simulations and asymptotic approximations that in finite samples the OLS procedure can lead to some bias which decreases only slowly as the sample size increases.

long-run model; if the restriction is not valid, the test loses power[33]. In view of these considerations, we presently test the cointegration hypothesis using the more powerful Johansen FIML (Full-Information Maximum Likelihood) approach in the multivariate framework.

The Johansen procedure is based on maximum likelihood estimation of a vector autoregressive (VAR) system. It is a method for estimating both the distinct cointegrating vectors which may exist within a set of variables and for carrying out a range of statistical tests. Given a *Nx1* vector of non-stationary *I(1)* variables X_t and considering a VAR model of order p, with Gaussian errors, we obtain[34]:

$$X_t = C_t + \Pi_1 X_{t-1} + \ldots + \Pi_p X_{t-p} + \varepsilon_t \tag{9}$$

where C_t is an $N \times 1$ vector of deterministic components (such as constants, dummies or drift terms), Π_i $i = 1..$p are $N \times N$ parameter matrices, ε_t is a vector of white noise errors with covariance matrix >0 and $t = 1...T$.

By re-parametrizing[35], the VAR can be transformed into a reduced-form error-correction model or vector error-correction model (*VECM*), in which we can directly distinguish between the effects related to the short-and long-run variation in the data:

$$\Delta X_t = C_t + \sum \Gamma_i \Delta X_{t-i} - \Pi X_{t-1} + \varepsilon_t \tag{10}$$

where

$$\Gamma_i = -(\Pi_{i+1}... + .. + \Pi_p), \quad \Pi = (\Pi_i... + ...+ \Pi_p) - I, \quad i = 1,...p-1.$$

When there are more than two variables in the VAR, it is not necessary for all of them to have the same order of integration. However, for every stationary variable included in the VAR, the number of cointegrating relationship will automatically increase correspondingly because each *I(0)* variable forms a cointegration relation by itself[36]. What really matters is that the variables on both side of the *VECM* were jointly "balanced" in order to preserve the assumption of a stationary (zero mean) error term ε_t.

The matrices of parameters (Γ, C and Π) are estimated using maximum likelihood method. The choice of deterministic components (constant, trend) to introduce in the *VECM*, restricted to enter the cointegration space or unrestricted, influences the distribution of the cointegrating tests. In the formulation of the system the lag-length of the ΔX_{t-i} should be high enough to assure a white noise disturbance vector ε_t. In the reduced-form model, which is likely to be heavily overparametrized, the short run dynamics is given by the elements of matrix Γ_i. The estimates of these matrices is intended to correct the variation in X_t related to the short run. The cointegration test involves determining the rank of the matrix Π, which

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[33]   See Banerjee et al. (1993).

[34]   Contrary to the approach of Engle and Granger, in this multivariate approach all the variables are explicitly endogenous so it is not necessary to make any preliminary and arbitrary normalization.

[35]   Adding and subtracting various lags or simply using the first-difference operator Δ, defined as $\Delta X_t = X_t - X_{t-1}$. This re-parametrization leaves all the properties of the residual unchanged.

[36]   As stressed by Harris (1995) "stationary variables might play a key role in establishing a sensible long-run relationship between non-stationary variables, especially if theory a priori suggests that such variables should be included", p. 80.

corresponds to the number of its non-zero eigenvalues, which contain information on the long-run properties of the X_t variables. Therefore, the magnitude of the eigenvalues (λ_i) provides information on the presence of a cointegrating relation. For the eigenvectors corresponding to the non-stationary part of the model (or "common trends"), $\lambda_i \cong 0$ for $i = r +1,...n$. In the situation where the components are non-stationary and integrated of order 1, the rank of the Π matrix determines the cointegration properties of X_t, that is the number of cointegrating vectors in the system.

There can be three different cases for the rank of Π which are of particular importance[37]:

1) If $r = n$, that is the matrix Π has full rank, this implies that all the n variables in X are stationary.

2) If $r = 0$, that is Π is a null matrix, this implies that no cointegrating vector exists because all linear combination of X are $I[1]$. In this case the reduced-form becomes a model of only differenced variables.

3) If $0 < r < n$, that is the matrix Π has a reduced rank, then there are $(n-r)$ linear combinations of X which act as a common stochastic trend, and r cointegrated linear combinations.

Summing up, the hypothesis of cointegration is formulated as the hypothesis of reduced rank of the coefficients matrix Π. If this holds, the latter can be decomposed as: $\Pi=\alpha\beta'$ under the Johansen ML procedure, where β is the $n \times r$ matrix of cointegrating vectors (each row β_i is a cointegrating vector) and α is the $n \times r$ matrix of "weighting elements". The matrix α contains the adjustment coefficients which represent the feedback effects from disequilibrium to the dependent variables. In case of only one cointegrating vector, if a given column of α is not different from zero, except for the first entry, single-equation estimation of the relation will not lead to loss of information on cointegration and dynamic behaviour. This is because the right-hand variables can be considered weakly exogenous for the variable of interests[38].

The stationary relations $\beta'X_{t-1}$ are referred to as the cointegrating relations. The estimate of ß' is obtained by solving an eigenvalue problem, whose solution is represented by the eigenvectors β_i and the eigenvalues λ_i. When there is more than one cointegrating vector the multivariate model determines the cointegration space, instead of the individual vectors, and this makes the analysis more complicated due to the difficulties of interpreting the cointegration space[39].

For testing cointegration, the Johansen approach is based on sequential likelihood ratio test of the null hypothesis of $n-r$ unit roots against the alternative of $n-r-1$ unit roots. Two different tests have been developed[40]:

1) The trace statistic test: $\lambda_{trace} = - T \sum_{i=r+1} ln(1-\lambda_i)$, r=0, 1,..n-1 were λ_i are the ordered largest eigenvalues.

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[37]   See Johansen and Juselius (1990) for a detailed exposition.

[38]   See Banerjee et al. (1993).

[39]   See Juselius (1992).

[40]   The test strategy is the multivariate analogue of the Dickey-Fuller (DF) test. Critical value for these tests have been tabulated by Johansen (1988), Johansen and Juselius (1990), Osterwald-Lenum (1992).

2)   The max statistic test: $\lambda_{max} = -Tln(1-\lambda_{r+1})$ which is based on the largest eigenvalue. In this case, as noted before, a small value of λ_i implies that the unit root hypothesis cannot be rejected at the *rth* level of significance, as in the case of scalar unit root tests[41].

The trace test has been found to show more robustness to both skewness and excess kurtosis in the residuals than the λ_{max} test[42]. As usual, if λ_{trace} or λ_{max} exceeds the critical value, the null can be rejected in favour of the alternative. It should also be stressed that testing for cointegration by investigating for the number of *r* linearly independent columns in Π is equivalent to testing that the last *(n-r)* columns of α are insignificantly small.

In this multivariate framework, after testing hypotheses on the number of cointegrating relationships, we can also test certain linear restrictions on β. This involves testing restrictions of the form $\beta = H\phi$ where H is the restriction matrix and ϕ is the parameter matrix. To carry out this test the statistics: $n \sum_{i=1} ln [(1- \lambda^*)/(1- \lambda)]$ can be used; here λ^* refers to eigenvalues derived under the parameter constrained and λ refers to unconstrained eigenvalues. This is a likelihood ratio and the statistic is asymptotically distributed as χ^2 under the null.

V. Determinants of imports: Specification and estimation results

The explicit restricted form of equation (4), the long-run demand for imports, has the following form:

$$M_t = \alpha + \beta*DDO2_t + \gamma*RP_t \tag{11}$$

where all data are in logarithms and M = volume of imports, $DDO2$ = domestic demand including stocks, and RP is the relative price term, $RP = PM/PD,$ where PD = domestic price of import-competing goods, and PM = price of imports. The restricted equation assumes that the individual price terms have identical but opposite in sign coefficients, a restriction generally supported by the data.

Equation (11) is assumed to represent the cointegrating, long-run, relationship between imports and their determinants. The short-term dynamics of adjustment of actual imports to the level consistent with the long-term determinants is represented by an error-correction process, an explicit form of equation (7), of the following form:

$$\Delta M_t = a_0 + \Sigma a_i W_{t-j} + \theta[M_{t-1} - (\alpha + \beta*DDO2_{t-1} + g*RP_{t-1})]$$

$$= a_0 + \Sigma a_i W_{t-j} + \theta[ER_{t-1}] \tag{12}$$

where W is a vector of independent variables in first-difference form lagged j quarters which includes also values of the dependent variable starting from t-1 and lagged j quarters, θ is the error-correction coefficient, and ER is the error term from the cointegration vector which, for

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[41]   In the trace test, the null hypothesis is that the number of cointegrating vectors (*r*) is less than or equal to *n* (where *n = 0, 1,...N)* against a general alternative that *r > n*; in the maximum eigenvalue test the alternative hypothesis is explicit *(r = n+1)*.

[42]   See Cheung and Lai (1993).

cointegration, must be stationary. θ must be negative in order that values of imports larger (smaller) than the prediction of the cointegrating equation lead to reduction (increase) in current imports; thus, a negative value of θ ensures stability. The model is deliberately parsimonious in the choice of variables in order to test strictly the predictions of economic theory and the value of the error-correction approach to modelling the demand for imports.

The estimation strategy follows the two-step procedure. First, the cointegration vector (11) is obtained with the Johansen procedure through a search process based on the Pantula principle[43]; as a result, only a constant term is included in the cointegration equation as well as in the dynamic model; secondly, the error term of the cointegration vector is used as an independent variable in the dynamic regressions (12). The data are all in logarithms, and the mnemonics used in the regressions reported in Table 5 below are as follows: *DDO2* = domestic demand including stocks (1980 prices); *PMPGDP* = ratio of import to GDP deflator; *PMPDD2* = ratio of import to domestic demand deflator; I*MPMQ* = imports of manufactures (1980 prices); and *PMDMANU* = ratio of import price to domestic price of manufactures.

Two sets of regressions are reported, equation (I) where the dependent variable is non-energy imports and equation (III) where the dependent variable is imports of manufactures.

Furthermore, equation (I) is estimated with the ratio of the import to the GDP deflator as the relative price term, and equation (II) with the ratio of import to the domestic demand deflator as the relative price term.

Table 4 Imports, domestic demand and relative prices Johansen likelihood ratio tests for cointegration, trace test			
	No cointegration	At most one vector	At most two vectors
Sample 1970:Q1-1996:Q3			
Eigenvalue	*0.190*	*0.062*	*0.023*
Non-energy imports, equation (I)	31.07*	9.10	2.46
Eigenvalue	*0.215*	*0.079*	*0.034*
Non-energy imports, equation (II)	37.45**	12.14	3.62
Eigenvalue	*0.232*	*0.078*	*0.036*
Imports of manufactures, equation (III)	39.59**	12.16	3.76*
The null hypotheses are that there is no cointegration, that there is at most one cointegration vector, and that there is at most two cointegration vectors; the 5% critical values of the likelihood ratio statistic at the three largest eigenvalues reported for each equation in the Table are 29.68, 15.41 and 3.76; the 1% critical values are, correspondingly, 35.65, 20.04 and 6.65; * (**) rejects the null at the 5% (1%) level of significance; equations (I), (II), and (III) refer to those in Table 5.			

The results of cointegration tests, based on the Johansen method with a lag order of three in *VECM*, are reported in Table 4. The test statistic is λ_{trace} discussed previously. The Pantula search suggested that the long-run equations had intercepts but there was no trend in the relationship between import demand and its determinants; this specification not only described the data best but it also likely yielded Gaussian errors[44]. The results confirm that the variables in question are cointegrated, and in all cases the cointegration vector is unique. The null of no cointegration is rejected at the 99% level of significance in the case of equations (II) and (III), and at the 95% level of significance in the case of equation (I), at the largest eigenvalue. The somewhat lower level of significance at which cointegration is supported by the data in the case of the non-energy equation defined over the relative price of the ratio of import to the GDP deflator points to the presence of some instability, an issue to which we will return once more in section VI. Nevertheless, the results support the
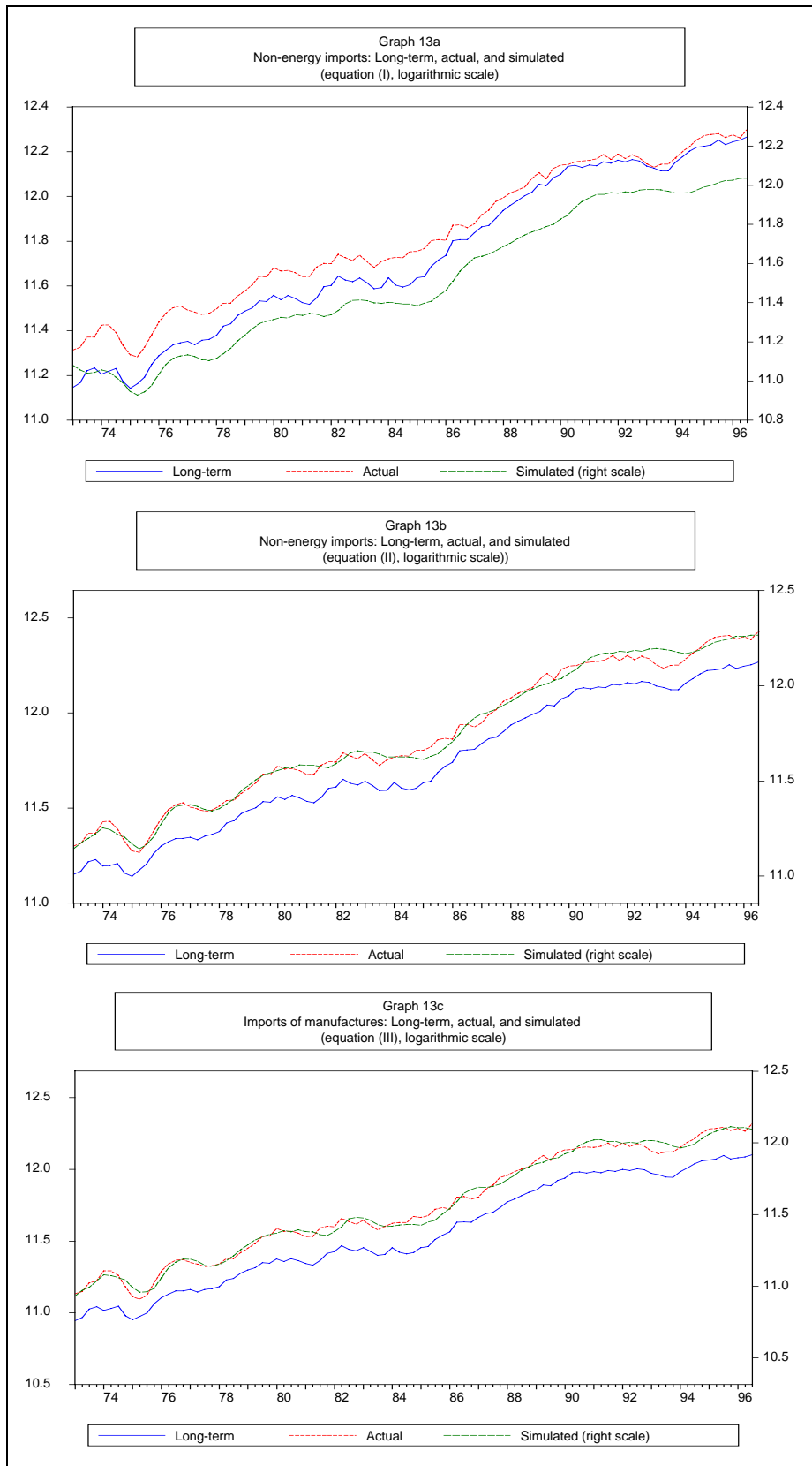
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[43]   See Harris (1995), p. 97, on this.

[44]   An alternative specification, where it was postulated that the long-run equation was defined over no intercept or trend, yielded essentially robust results but unreasonably large elasticity estimates.

hypothesis that there is a long-term relationship between imports, domestic demand and relative prices, as postulated by economic theory, and that this relationship is unique. This cointegration vector can, therefore, be interpreted as representing the long-term demand for imports[45].

The estimated long-run demand for imports is shown in the upper panel of Table 5. The first adjustment weight suggests that a 1% deviation of actual imports from their predicted value causes an adjustment of non-energy imports equal to around 55% each quarter; in the case of imports of manufactures, this adjustment is even faster, around 70% of the deviation in each

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[45]   Cointegration is confirmed by unit root tests for the residuals of the estimated vectors. The *DF (ADF (4))* statistic for the error of equation (I) is -5.05 (-4.51); for the error of equation (II) it is -4.72 (-4.74); and for the error of equation (III) it is -5.12 (-4.93). The critical value of the test statistic at the 99% (95%) level of significance is -3.50 (-2.89).

Table 5
Demand for imports
Restricted cointegration/error correction model for imports, domestic demand, and relative import/domestic price, 1970:Q1-1996:Q3
(equations (I)-(III), standard errors in parentheses)

Non-energy imports			Imports of manufactures	
	(I)	(II)		(III)
Cointegrating vector				
DDO2	2.253 (0.051)	2.142 (0.070)	DDO2	2.357 (0.055)
PMPGDP	-0.201 (0.066)	-	PMDMANU	-0.404 (0.142)
PMPDD2	-	-0.358 (0.095)		-
Constant	18.806	16.949	Constant	19.988
$③_1$	-0.536 (0.155)	-0.556 (0.142)	$③_1$	-0.691 (0.145)
$③_2$	-0.034 (0.057)	-0.038 (0.053)	$③_2$	-0.089 (0.048)
$③_3$	0.057 (0.122)	-0.018 (0.108)	$③_3$	0.068 (0.084)
Short-term adjustment model				
θ	-0.272 (0.126)	-0.308 (0.119)	θ	-0.519 (0.120)
d(DDO2(-1))	-	0.187 (0.260)	d(IMPMQ(-1))	0.120 (0.091)
d(DDO2(-2))	0.275 (0.251)	0.250 (0.253)	d(IMPMQ(-2))	0.137 (0.090)
d(DDO2(-3))	0.380 (0.247)	0.392 (0.247)	d(DDO2(-3))	0.456 (0.285)
d(PMPGDP(-2))	-0.227 (0.132)	-	d(PMDMANU(-2))	-0.209 (0.181)
d(PMPGDP(-3))	-0.132 (0.129)	-	d(PMDMANU(-3))	-0.148 (0.167)
d(PMPGDP(-4))	-0.393 (0.122)	-	d(PMDMANU(-4))	-0.490 (0.176)
d(PMPDD2(-2))	-	-0.281 (0.135)		-
d(PMPDD2(-4))	-	-0.400 (0.125)		-
Constant	-10.213 (4.606)	-10.425 (4.031)	Constant	-20.729 (4.815)
R^2	0.284	0.304	R^2	0.311
SER	0.024	0.024	SER	0.027
DW	1.839	1.966	DW	1.922
Jarque-Bera χ^2 (2)	2.369	2.240	Jarque-Bera χ^2 (2)	1.279
F(ARCH(4))	0.791 (0.50)	1.001 (0.41)	F(ARCH(4))	1.200 (0.32)
F(RESET)	0.813 (0.54)	1.119 (0.36)	F(RESET)	1.277 (0.28)
F(HET1)	1.174 (0.32)	0.463 (0.92)	F(HET1)	1.086 (0.38)
F(HET2)	1.032 (0.44)	0.367 (1.00)	F(HET2)	0.909 (0.61)
F(AR(4))	1.705 (0.16)	0.954 (0.44)	F(AR(4))	1.674 (0.16)

$\tilde{a}_i$ refers to the value of the adjustment coefficient in the import function entering the first, the second (DDO2) and third (relative price) equation, respectively, in the VAR; θ is the error correction term; d(Z) is the first-difference operator for Z; R^2 is the adjusted R^2; SER is the standard error of the regression; DW is the Durbin-Watson statistic; the Jarque-Bera statistic for the normality of the residuals, distributed as χ^2 with two degrees of freedom, has a 95% critical value of 5.99; F(ARCH(4)) of lag order 4 is the F test for autoregressive conditional heteroscedasticity (probability in parentheses); F(RESET) is Ramsey's F statistic for specification error run on five fitted terms; F(HET1) is White's F test for heteroscedasticity, no cross terms; F(HET2) is White's F test for heteroscedasticity, with cross terms; F(AR(4)) is the LM test for serial correlation of order 4; sample 1970:1-1996:4.

Graph 13a
Non-energy imports: Long-term, actual, and simulated
(equation (I), logarithmic scale)

| Long-term | Actual | Simulated (right scale) |

Graph 13b
Non-energy imports: Long-term, actual, and simulated
(equation (II), logarithmic scale))

| Long-term | Actual | Simulated (right scale) |

Graph 13c
Imports of manufactures: Long-term, actual, and simulated
(equation (III), logarithmic scale)

| Long-term | Actual | Simulated (right scale) |

quarter. The remaining adjustment weights are of negligible value, implying that disequilibrium in the import demand equation causes adjustment only in that equation alone, supporting the hypothesis that domestic demand and the real exchange rate are weakly exogenous for the parameters of the conditional import equation and also supporting the single-equation estimation procedure followed here[46].

The estimated long-run import equations, where the cointegration space is restricted to one vector, have the following form:

$$IMPNONE = 18.81 + 2.25*DDO2 - 0.20*PMPGDP \qquad (13)$$

$$IMPNONE = 16.95 + 2.14*DDO2 - 0.36*PMPDD2 \qquad (14)$$

$$IMPMQ = 19.99 + 2.36*DDO2 - 0.40*PMDMANU \qquad (15)$$

where *IMPNONE* = non-energy imports and *IMPMQ* = imports of manufactures. The results suggest that imports of non-energy goods and imports of manufactures are almost equally, and highly, elastic with respect to domestic demand, with the elasticity ranging from around 2.2 in the case of non-energy imports to close to 2.4 in the case of imports of manufactured goods. This high elasticity explains the rapid rise of imports in domestic demand, shown by the penetration ratios of Graph 9[47]; this is a property of the sample period characterized by trade liberalization and by rapid growth in intra-European and world trade. The estimates reported presently compare well with those of Germany (1.86) and that of the US (2.04) obtained in Carone (1995) for total imports of goods. In contrast, estimates reported in the Banque de France (1996) for imports of manufactures are substantially lower, and they are either restricted to or estimated at around unity, but the equation includes an import penetration term which undoubtedly biases the estimate of the elasticity with respect to demand downwards. This is clearly a model-specific matter, reflecting different modelling strategies, choice of variables and sample size, as is also the case in Bonnaz et Paquier (1993) who estimate an elasticity for imports of manufactures with respect to domestic demand equal to unity and a long-term price competitiveness elasticity of -0.7. Capet et Gudin de Vallerin (1993), from the sample 1970:Q1-1988:Q4, obtain a domestic demand elasticity of imports of manufactures also of unity and a price competitiveness coefficient of -0.81; their equation includes a relative capacity utilization term and a time trend too[48].

Similarly, the long-term demand elasticity estimates reported in Goldstein and Khan (1985) are obtained from earlier samples ending at the latest in 1980 and are undoubtedly contaminated

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[46] The exogeneity status of variables is not presently investigated further. For an overview of the three main concepts of exogeneity (weak, strong and super) and their importance for conducting valid inference as well as forecasting and policy analysis in a conditional econometric model see Engle, Hendry and Richard (1983), and Ericsson and Irons (1994).

[47] Note also in this respect that a long-run elasticity of imports with respect to domestic demand exceeding unity is not admissible since it would imply that imports would ultimately dwarf domestic demand and output. However, it is quite conceivable, as the empirical evidence shows, that specific samples such as the one used here may be characterized by high values of this elasticity consistent with the rapid rise in import penetration.

[48] Capet et Gudin de Vallerin (1993) find two cointegration vectors for imports; although both vectors are economically meaningful, the presence of cyclical variables in the long-term relationship is problematic and difficult to defend - see also footnote 51. The estimated demand elasticity in the present parsimonious model may be a mixture of demand and of other factors which are not explicitly articulated here; as a possible alternative specification see the estimates in Temple and Urga (1997) where, on UK manufacturing data, the demand elasticity is unity but the elasticity with respect to an OECD exports/domestic production variable is 1.55.

by, and reflect the importance of, restrictions in France's trade during that time. Most estimates surveyed in Goldstein and Khan (1985) range from 1.07 to 1.57 (see Table 4.3) and they apply to total imports. Similar observations apply to the estimated price elasticity of the import function. The long-run estimate for imports of manufactures reported in the Banque de France (1996) range from -0.56 to -0.94; Goldstein and Khan (1985) report estimates for total imports ranging from -0.33 to -1.31 on samples ending at the latest again in 1980; finally, Carone (1995) finds relative price elasticities for Germany (-0.58), Japan (-0.59) and the US (-0.92) for imports of goods which are clearly higher than those obtained here. As noted previously, model specification, choice of variables and the data sample account for much of these differences.

Variance decomposition analysis attributes, among innovations in all variables, the dominant contribution to forecast variance of the dependent variable to shocks in domestic demand[49]. In equation (I) this was around 81% after five quarters but declined to around 75% after twenty quarters and to around 74% after 35 quarters; in equation (II) it was also around 75% after five quarters but also declined to around 62% after 35 quarters; and in equation (III) it was around 80% in after 5 quarters but fell to around 71% after 35 quarters. The contribution of the price term was less 2% in the beginning of the period in equation (I) but rose quickly to almost 10% after 6 quarters and to over 21% after 35 quarters; in equation (II), its contribution rose to over 18% after 5 quarters and to around 35% after 35 quarters, and it continued to increase beyond this horizon; finally, in equation (III) the contribution of the relative price term rose from over 8% in the fifth quarter to over 14% in the twentieth and to over 15% in the thirty-fifth quarter. These results suggest that while demand shocks play a predominant role in import determination in the short-term, over longer horizons their contribution, while remaining important, diminishes. The opposite pattern is observed in competitiveness shocks[50].

Results from the estimation of the short-term adjustment model are shown in the lower panel of Table 5. The equations were estimated in an unrestricted form and selectively lagged variables with insignificant coefficients were dropped from the equation; the results reported in the Table correspond to the final equations. The equations are well determined, as can be seen from the results of specification tests reported in the lower part of the Table. Moreover, the equations reproduce the actual data quite well in dynamic simulations (see Graphs 13a-13c for within-sample simulations; out-of-sample simulations are discussed in section VI below). The lags in the final equations go as far back as four quarters in some cases.

The estimate for the error-correction coefficient, θ, is negative and statistically significant in each of the three equations, supporting the hypothesis of stability and of cointegration. The value of θ rises from -0.27 in the case of the first equation to -0.52 in the case of imports of manufactures. These estimates suggest that, depending on the equation, between around 30% and 52% of last period's deviations from equilibrium relationship is corrected in the current

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[49] The cointegration-error correction framework ensures that the dynamic responses of the variables in the system, resulting from variance decomposition exercises, where all variables are endogenous, are mutually consistent. The ordering used in these variance decomposition exercises was domestic demand, relative price term, and imports; this suggests that imports is the contemporaneously "most" endogenous variable, the last equation in the three-equation VAR implicit in the error-correction model which depends on the other two equations. Alternative orderings produced virtually identical results. Therefore, the results discussed in the text are robust across alternative models.

[50] In contrast to these results, and from an equation defining the dependent variable as the export/import ratio of manufactures, Agénor (1996) finds that shocks to price and non-price competitiveness account for about 40% of the variation in the trade ratio, while shocks to domestic and foreign output account for around 30%.

period and, therefore, within less than a year the adjustment is complete[51]. Changes in domestic demand lead to changes in imports in the same direction but the effect is not instantaneous; the one-quarter lagged value of the domestic demand variable is not significant in equation (I) but two- and three-quarter lags dominate. A similar pattern holds in equation (II) but, despite the high standard error, the one-quarter lag of this variable improved the fit. A three-quarter lag in this variable in imports of manufactures is also found to have a sizeable effect on current import adjustment. Somewhat surprisingly, changes in relative prices, defined over two- to four-quarter lags, exert a sizeable negative influence on import adjustment, a result which is particularly pronounced in the case of equation (III), imports of manufactures. The relative price elasticity of import demand suggests that exchange rate changes, to the extent they are passed through into domestic prices, can have significant effects on trade balance movements in the short-run. Furthermore, to the extent that imported manufactures are imperfect substitutes with domestically produced manufactures, there are opportunities for strategic price behaviour on the part of both importers and domestic producers in the event of real exchange rate shocks. Finally, imports of manufactures are dependent on their own value lagged one and two quarters.

Graphs 13a-13c present actual, simulated, and long-term import demand data. The long-term demand is defined by the cointegration vectors (13) - (15). The equations reproduce the data generally very well, especially in the case of imports of manufactures. The shifts in the long-term demand for imports in the three equations in 1973 reflect the impact of the first oil shock; the increase in imports in the period 1985-87 is likely associated with the rise in domestic demand; the steep rise in imports in the post-1989 period likely reflects the impact of both demand developments and the emerging real franc appreciation; finally, since 1993 imports according to the long-term relationship have flattened out, again possibly reflecting slow demand growth.

Equation (I) underpredicts import demand throughout the simulation period, while actual imports appear to be close to their long-term value over the sample. Equation (II) explains the data very well and the long-term and actual data are again very close to each other, even though the equation fails to predict accurately the downturn in import demand in the 1992-94 period. Equation (III), imports of manufactures, also behaves very well and has a smaller deviation from the actual during this period. Predictions for the long-run demand for imports appear to be largely invariant to the relative price term used in the cointegration vector. Thus, the equation with the ratio of non-energy import prices to the GDP deflator as the real exchange rate (Graph 13a) predicts long-run demand for imports as well as when the relative price term is approximated by the ratio of the non-energy import price deflator to the deflator of domestic demand (Graph 13b). It is likely that the rather high value of the error-correction coefficient of the each adjustment equation discussed previously contributes notably to the very close approximation of the actual data in the simulations.

VI. Hysteresis in imports and stability of the estimates

An important consideration in the analysis of trade flows is whether real exchange rate shocks have permanent effects on the behaviour of imports and exports. Conventional theory views trade flows as elastic with respect to the real exchange rate, and competitiveness shocks force movements along the import and export functions; once the shock is reversed the quantity of

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[51] Capet et Gudin de Vallerin (1993) estimate an error-correction coefficient of -0.21 in the imports of manufactures equation. Their cointegration equation includes, apart from lagged domestic demand, price competitiveness and imports, a relative capacity utilization term, intended to capture supply conditions in the domestic and the international market; see also footnote 48.

imports and exports reverts to its original equilibrium. However, when real exchange rate shocks affect permanently trade flows, then imports and exports are characterized by hysteresis[52]. The empirical counterpart of hysteresis is the presence of structural breaks in trade flow equations associated with large competitiveness shocks[53].

According to Baldwin's (1988) version of the hysteresis model, large real exchange rate misalignments are likely to induce entry and exit of firms which affect pricing behaviour. In empirical studies, the data should, as a result, indicate the presence of a structural break in pass-through equations; moreover, the model predicts that import equations should exhibit an increase in the absolute price elasticity synchronous with the structural break. In the present section, the import equations estimated previously are reviewed to examine the presence of hysteresis characteristics taking the form of instability. While no direct test for hysteresis is performed, the results are suggestive rather than conclusive concerning hysteresis[54]. The point of reference is the franc real appreciation since the beginning of 1992; evidence discussed later indicates that 1992:Q2 represents a possible break point.

Table 6 Stability of long-term import equations Johansen likelihood ratio tests for cointegration	No cointegration	At most one vector	At most two vectors
Sample 1970:Q1-1992:Q2			
Eigenvalue	0.204	0.148	0.017
Non-energy imports, equation (i)	35.23*	15.36	1.45
Eigenvalue	0.243	0.172	0.023
Non-energy imports, equation (II)	42.62**	18.42*	2.03
Eigenvalue	0.253	0.113	0.020
Imports of manufactures, equation (III)	37.54**	12.17	1.78
Sample 1970:Q1-1996:Q3			
Eigenvalue	0.190	0.062	0.023
Non-energy imports, equation (i)	31.07*	9.10	2.46
Eigenvalue	0.215	0.079	0.034
Non-energy imports, equation (II)	37.45**	12.14	3.62
Eigenvalue	0.232	0.078	0.036
Imports of manufactures, equation (III)	39.59**	12.16	3.76*

The null hypotheses are that there is no cointegration, that there is at most one cointegration vector, and that there is at most two cointegration vectors; the 5% critical values of the likelihood ratio statistic at the three largest eigenvalues reported for each equation in the Table are 29.68, 15.41 and 3.76; the 1% critical values are, correspondingly, 35.65, 20.04 and 6.65; * (**) rejects the null at the 5% (1%) level of significance.

We first examine the stability of the long-term relationship. If there is hysteresis in import demand, then the data should indicate that there is a structural break in the cointegrating relationship taking the specific form of lack of cointegration in the post-break sample. Table 6

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[52] See Baldwin (1988), Mastropasqua and Vona (1989), Harris (1993) and Amano et. al. (1993), and references therein. Amano et al. (1993) provide a very good review of the analytical and empirical issues related to hysteresis in international trade.

[53] Amano et. al. (1993) represent this succinctly with the following dynamic process: $X_t = a*X_{t-1} + d*Z_t$ where Z is a vector of exogenous variables. Hysteresis occurs when the process has a unit root or when $a = 1$; in this case the steady state solution for X is not unique but depends on the initial value of X and the path of Z; shocks to Z alter permanently the steady state. An alternative representation of hysteresis involves changes in parameter d of the process in response to large shocks in Z. In trade flow equations, real exchange rate shocks in the Baldwin (1988) model correspond to large changes in Z and affect, among other predictions, the price elasticity of the import function.

[54] Clearly, a conclusive test of hysteresis would require a structural model since hysteresis is in fact a structural break in conventional trade equations. This cannot be undertaken in the present paper. Moreover, the real appreciation of the French franc has occurred only in recent quarters and it would perhaps take data from a complete cycle of appreciation and depreciation to establish conclusively whether there has been hysteresis in France's trade flows or not.

presents results of cointegration tests[55], based on the Johansen maximum likelihood method for imports, domestic demand and relative prices, defined by the long-term equations (14)-(16) reported previously and over two samples, 1970:Q1-1992:Q2, and 1970:Q1-1996:Q3. The rejection of the null of no cointegration in the extended sample beyond 1992:Q2 would indicate the absence of structural changes in the long-term relationship and, consequently, the absence of hysteresis; conversely, if the data indicate that there is no cointegration in the sample beyond the point of the presumed structural break, this would constitute evidence of changes in the relationship between imports and their postulated determinants[56].

The results of cointegration tests suggest that the cointegration found in the shorter sample holds well over the longer sample. At the largest eigenvalue the null of no cointegration is rejected at the 95% level of significance for equation (I), non-energy imports with the relative price term defined as the ratio of the import to the GDP deflator, and at the 99% level of significance in the case of equations (II) and (III). Moreover, the uniqueness of the vector cannot be generally rejected at stringent level of significance. Therefore, there is no evidence that the long-run demand for imports of non-energy goods and for manufactures has shifted between the two samples.

It is possible that the dynamic part of the import equations has been subject to breaks and this possibility is investigated presently. The overall evidence supports rejection of structural breaks in the dynamic equation. Results from CUSUM and CUSUM of squares tests, reported in Annex D, confirm the absence of instability in each of the import equations (I)-(III). Furthermore, detailed examination of coefficient stability in these equations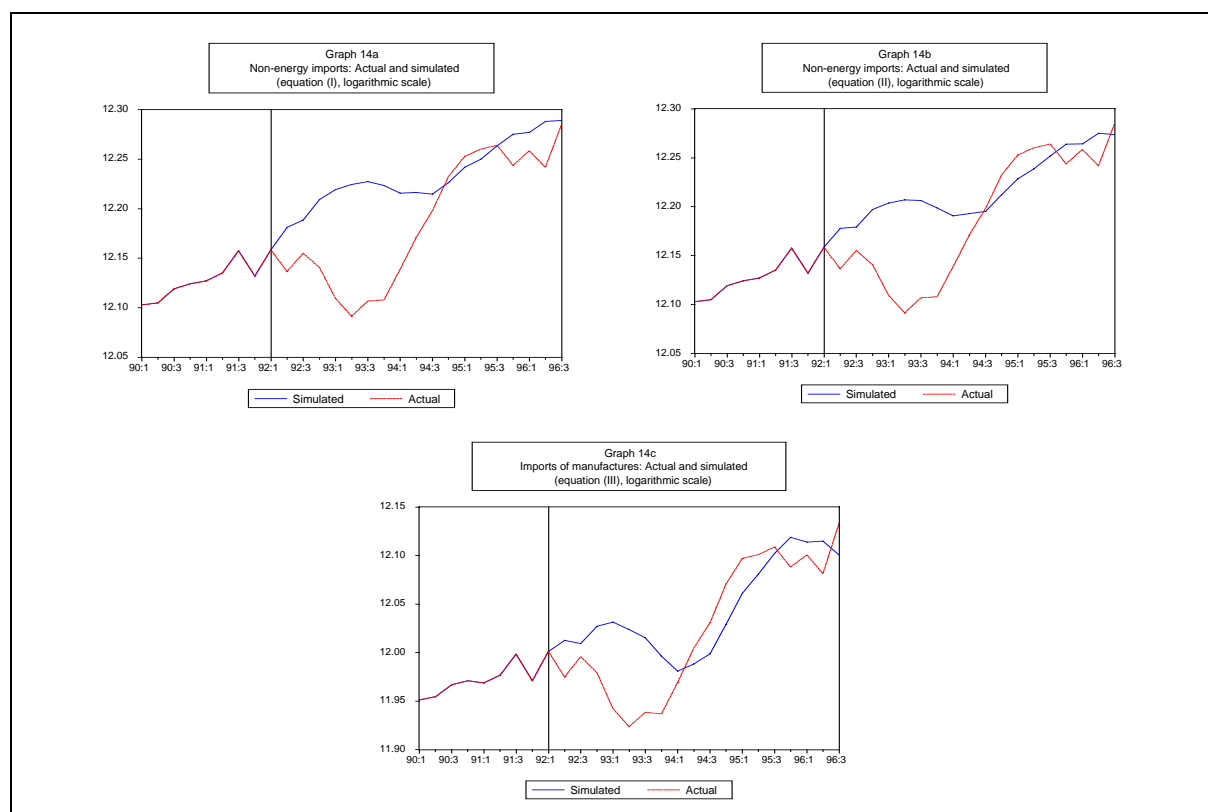, results of which are reported in Annex E, also confirm that no coefficient has been subject to instability. Nevertheless, it is possible that instability has been a temporary event characterizing only a few observations in the sample, and that stability holds both prior to and after these observations. To review this possibility Chow tests were employed. Since the franc appreciation has been particularly pronounced since 1992:Q2 a search for structural break at that quarter revealed that the data would support this hypothesis in some cases. In order to implement the tests, the coefficient of the dynamic relative price terms in the short-term adjustment equations were multiplied by a dummy variable with a value of one for the period 1992:Q2-1996:Q4, and zero otherwise.

The results are presented in Table 7. The data appear to support the hypothesis that there was a structural break on the price term from the second quarter of 1992 onwards in the case of the two equations for non-energy imports; the probability value for the calculated F is

Table 7 Stability of import equations: Tests for structural break in 1992:Q2 (complete sample 1970:Q1-1996:Q4)				
	F	Prob	LR (χ^2)	Prob
Non-energy imports: Equation (I)				
Chow breakpoint test	3.135	0.008	-	-
Chow forecast test	-	-	27.884	0.086
Non-energy imports: Equation (II)				
Chow breakpoint test	2.663	0.015	-	-
Chow forecast test	-	-	25.795	0.136
Imports of manufactures: Equation (III)				
Chow breakpoint test	1.770	0.094	-	-
Chow forecast test	-	-	23.561	0.214
Prob is the probability of drawing an F or a χ^2 statistic of the value shown in the Table from the respective distribution; the null hypothesis is that there in no structural break, and probability values of 5% or less can be taken to indicate rejection of the null.				

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[55] The cointegration vector has an intercept and it was assumed that there was no trend in the data-generating process; the lag order used in the tests was three.

[56] Amano et al. (1993) also follow this procedure in their test of hysteresis on Canadian data. An alternative test based on recursive eigenvalues is used to examine the stability of the long-term export functions; see section VIII below and also Annex F.

Graph 14a
Non-energy imports: Actual and simulated
(equation (I), logarithmic scale)

Graph 14b
Non-energy imports: Actual and simulated
(equation (II), logarithmic scale)

Graph 14c
Imports of manufactures: Actual and simulated
(equation (III), logarithmic scale)

significantly below 5%. This finding is consistent with the overprediction of these equations shown in Graphs 13a and 13b. Moreover, the data suggest that there was an upward shift in the relative price coefficient in the post 1992:Q1 period consistent with the prediction of the hysteresis model[57]. However, there is no evidence of a structural break in the equation for imports of manufactures in the post-1992:Q1 period. Is the apparent structural break affecting the forecasting performance of these equations? The answer is no. As can be seen in Table 7, the Chow forecasting test rejects instability in all equations. The evidence of these tests indicates that although there maybe instability problems in the post-1991 period for the aggregate, imports of non-energy goods, equation, this instability does not characterize the largest component of this aggregate, imports of manufactured goods.

A final, and stringent, stability test is the out-of-sample forecasting performance of the equations. Substantial over- or under-prediction out of sample constitutes evidence that a structural break has occurred; moreover, such evidence would be suggestive of hysteresis in import flows. Choosing 1992:Q2 as the breaking point, equations (I)-(III) were estimated over the sample 1970:Q1-1992:Q1 and the coefficients of the estimated dynamic equation, together with the estimated long-term cointegrating vector, were then used to produce dynamic simulations over the period 1970:Q1-1996:Q4.

The results of these simulations are presented in Graphs 14a-14c for each equation, respectively. It is clear that the equations track the data very closely over the estimation period but equations (I) and (II) miss the turning points during the 1993 recession and tend to over-predict for some quarters before converging again on the actual values from 1995 onwards.

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[57]   The value of the coefficient of the dummied relative price term (standard errors in parentheses) was 1.7 (0.47) in equation (I), 2.94 (1.70) in equation (II), and 4.68 (3.47) in equation (III); this shift was unambiguously significant only in equation (I), but not in equations (II) and (III).

However, equation (III), imports of manufactures, reproduces the data very well and shows no evidence of structural break. Since imports of manufactures represent the dominant component of non-energy imports, the apparent modest instability of the non-energy imports equations may be a reflection of instability in the remaining components of the import aggregate. Finally, the evidence from the out-of-sample simulations is consistent with test results for structural stability reported previously. It is possible to conclude, therefore, that there has been no hysteresis phenomena in French imports in recent years, and the simulations suggest that the flattening out of import growth in the 1990s is likely the result of the slowdown of economic growth.

VII. Determinants of exports: Specification and estimation results

The estimated equation for the long-run demand for exports, in a specification similar to that used for imports, has the following parsimonious form:

$$X_t = \alpha + \beta * Y_W + \gamma * REER_t \tag{16}$$

where all data are in logarithms, and X = volume of export goods, Y_W = world demand[58] and $REER$ = real effective exchange rate (or a relative price term). Equation (16) is assumed to represent the cointegrating, long-run relationship between exports, world demand and price competitiveness.

The data and mnemonics used in the regression are as follows (all data are in logarithms): $EXPNONE$ = non-energy exports; $EXPMQ$ = exports of manufactures; WDD = world domestic demand; $EXULC23$ = real exchange rate in terms of relative unit labour costs in total economy against 23 industrial countries; $EXGDP23$ = real exchange rate in terms of relative GDP deflators against 23 industrial countries; details of data sources are presented in Annex A.

Three different equations have been estimated on quarterly data over the period 1970:Q1-1996:Q3: equation (IV), where the dependent variable is non-energy exports and the real effective exchange rate is defined in terms of unit labour costs as a proxy for relative export prices[59]; equation (V), where the dependent variable is again non-energy exports but the real effective exchange rate is defined by relative GDP deflators against the IC-23; and equation (VI) where the dependent variable is exports of manufactures and the real effective exchange rate is defined by relative unit wage costs in manufacturing as a measure of competitiveness[60].

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[58]   The world domestic demand (WDD) has been computed as weighted average of the domestic demand in the following industrial countries: USA, the United Kingdom, Austria, Belgium (GDP), Denmark, Canada, Japan, Germany, Italy, Netherlands , Switzerland, Norway, Finland, Australia, New Zealand, Portugal and Spain. Because of lack of data, the industrialising countries have not been included in the average. Data have been taken from the WEFA-DRI database, and the main source is the OECD Quarterly National Accounts statistics. The weights have been calculated from IMF - Direction of Trade Statistics and are based on 1970-1995 average of each country's shares in France's total value of exports of goods.

[59]   A more appropriate index of relative prices or competitiveness, such as a real effective exchange rate based on export prices of manufactured goods or of a wider aggregate, is not available. The one calculated by the International Monetary Fund covers a more limited period than considered in our analysis. Also, Marquez (1992) shows that, from US trade functions, it is difficult to distinguish empirically between different cost and price indicators of international competitiveness.

[60]   In the Annex E are reported additional cointegration estimates for export equations (IV), (V) and (VI) based on alternative definitions of competitiveness.

The use of the latter seems more appropriate than that based on GDP deflators. Real effective exchange rate based on GDP deflators has been shown to be influenced by two factors: first, relative unit labour costs in the tradable sector and, secondly, the ratio of unit labour costs in tradeables and non-tradeables in each country[61]. Both definitions of competitiveness are not accurately representative of the true competitive

Table 8 Exports, world demand and relative prices Johansen likelihood ratio tests for cointegration, trace test	No cointegration	At most one vector	At most two vectors
Sample 1970:Q1-1996:Q3			
Eigenvalue	*0.225*	*0.029*	*0.009*
Non-energy exports, equation (IV)	31.06**	4.08	0.95
Eigenvalue	*0.224*	*0.034*	*0.012*
Non-energy exports, equation (V)	31.65**	4.98	1.31
Eigenvalue	*0.226*	*0.021*	*0.007*
Exports of manufactures, equation (VI)	29.89**	3.01	0.75

The null hypotheses are that there is no cointegration, that there is at most one cointegration vector, and that there is at most two cointegration vectors; the 5% critical values of the likelihood ratio statistic at the three largest eigenvalues reported for each equation in the Table are 24.3, 12.5 and 3.8; the 1% critical values are, correspondingly, 27.9, 16.3 and 6.5; * (**) rejects the null at the 5% (1%) level of significance; equations (IV), (V), and (VI) refer to those in Table 9.

position of export goods because they do not capture pricing-to-market behaviour. Not any change in costs will be automatically and completely passed through to prices of exported goods, but also the evolution of unit labour costs and export prices may diverge because of different product composition of exports and domestic production. Therefore relative unit labour cost and relative unit wage cost should be considered as ex-ante or underlying, potential, competitive position.

The lag length of the *VECM* has been set at two, which has proved to be enough to ensure Gaussian error terms. In order to decide the most appropriate way of modelling deterministic components in the *VECM* we have followed the procedure suggested in Johansen (1992b) and, as in the case of imports, based on Pantula principle (starting from the less restrictive model), testing the joint hypothesis of both the rank order and the deterministic components[62]. As result of the procedure, no deterministic variables has entered the cointegration space and the *VECM*.

The results of tests on the number of the cointegration vectors are given in Table 8. The reduced rank test statistics based on trace point to only one cointegrating relationship on the French export data. This result greatly simplifies the interpretation of the first vector β_1 as a long-run equilibrium relationship among export volume, world domestic demand and price competitiveness. The coefficient estimates of the cointegrating relation are reported in the upper panel of Table 8. It is worth noting that the estimated long-run world demand and relative price elasticities are consistent with those often found in other studies.[63]

The feedback coefficients in the first column of normalised α' (reported in Table 9) represent the weights with which excess demand for exports enters the three equations of our system and can be interpreted as the average speed of adjustment towards the estimated equilibrium relationship. The first weight, (α_1) is referred to the export demand equation and it implies that, in the case of model IV, a 1% deviation from equilibrium will be corrected by an adjustment of 0.05. % each quarter. The interesting result, also noted in the case of the import


[61] See Marston (1987).

[62] See Harris (1995).

[63] For similar results based on cointegration and error-correction methodology see Capet and Gudin de Vallerin (1993).

Table 9
Demand for exports
Restricted cointegration/error correction model for exports, world demand, and real effective exchange rate, 1970:Q1-1996:Q3
(equations (IV)-(VI), standard errors in parentheses)

Non-energy exports			Exports of manufactures	
	(IV)	(V)		(VI)
Cointegrating vector				
WDD	1.649 (0.154)	1.719 (0.139)	WDD	1.581 (0.168)
EXULC23	-0.621 (0.166)		EXULC23	-0.552 (0.181)
EXPGDP23		-0.697 (0.149)		
$\tilde{a}_1$	-0.051 (0.012)	-0.059 (0.015)	$\tilde{a}_1$	-0.045 (0.012)
$\tilde{a}_2$	-0.018 (0.004)	-0.022 (0.005)	$\tilde{a}_2$	-0.018 (0.004)
$\tilde{a}_3$	-0.005 (0.011)	-0.006 (0.013)	$\tilde{a}_3$	-0.000 (0.010)
Short-term adjustment model				
θ	-0.045 (0.013)	-0.050 (0.016)	θ	-0.027 (0.012)
d(EXPNONE(-1))	-0.276 (0.094)	-0.239 (0.093)	d(EXPMQ(-1))	-0.269 (0.095)
d(EXPNONE(-2))	-0.169 (0.091)	-0.162 (0.089)		
d(WDD)	0.788 0.282)	0.828 (0.282)	d(WDD)	1.017 (0.298)
d(WDD(-1))	0.902 (0.303)	0.908 (0.304)	d(WDD(-1))	0.746 (0.298)
d(EXULC23)	-0.139 (0.104)		d(EXULC23(-1))	-0.276 (0.111)
d(EXULC23(-1))	-0.169 (0.109)		d(EXULC23(-2))	-0.148 (0.012)
d(EXULC23(-2))	-0.146 (0.107)			
d(EXPGDP23(-1)		-0.260 (0.102)		
R²	0.268	0.260	R²	0.260
SER	0.020	0.020	SER	0.021
DW	1.92	2.02	DW	2.00
Jarque-Bera χ^2 (2)	4.704	2.963	Jarque-Bera χ^2 (2)	1.666
F(ARCH(4))	0.415 (0.80)	0.295 (0.88)	F(ARCH(4))	0.700 (0.59)
F(RESET)	1.200 (0.32)	1.518 (0.19)	F(RESET)	1.189 (0.32)
F(HET1)	0.972 (0.49)	1.043 (0.42)	F(HET1)	1.170 (0.32)
F(HET2)	0.799 (0.78)	0.832 (0.70)	F(HET2)	0.643 (0.90)
F(AR(4))	0.415 (0.80)	0.191 (0.94)	F(AR(4))	1.455 (0.22)

$\tilde{a}_i$ refers to the value of the adjustment coefficient on the export function entering the first, the second (WDD) and third (relative price) equation, respectively, in the VAR; θ is the error correction term; d(Z) is the first-difference operator for Z; SER is the standard error of the regression; DW is the Durbin-Watson statistic; the Jarque-Bera statistic for the normality of the residuals, distributed as χ^2 with two degrees of freedom, has a 95% critical value of 5.99; F(ARCH(4)) of lag order 4 is the F test for autoregressive conditional heteroscedasticity (probability in parentheses); F(RESET) is Ramsey's F statistic for specification error run on five fitted terms; F(HET1) is White's F test for heteroscedasticity, no cross terms; F(HET2) is White's F test for heteroscedasticity, with cross terms; F(AR(4)) is the LM test for serial correlation of order 4; sample 1970:1-1996:4.
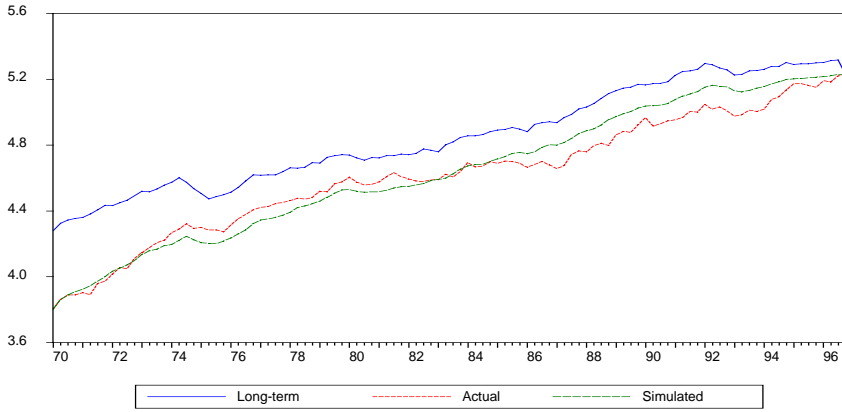
Graph 15a
Non-energy exports: Long-term, actual, and simulated
(equation (IV), logarithmic scale)
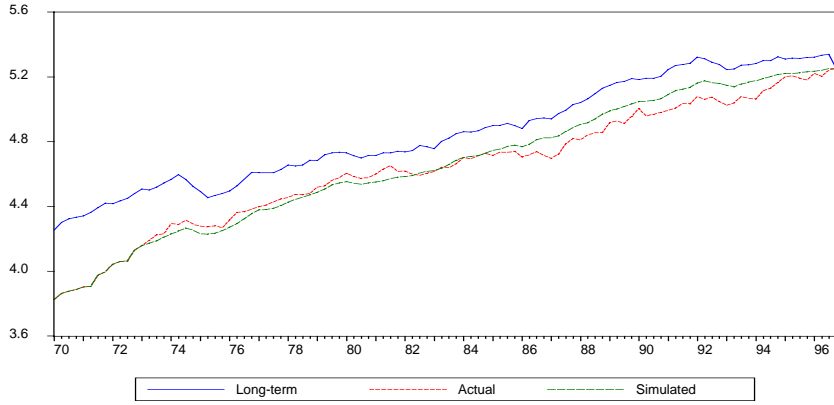
Graph 15b
Non-energy exports: Long-term, actual, and simulated
(equation (V), logarithmic scale))
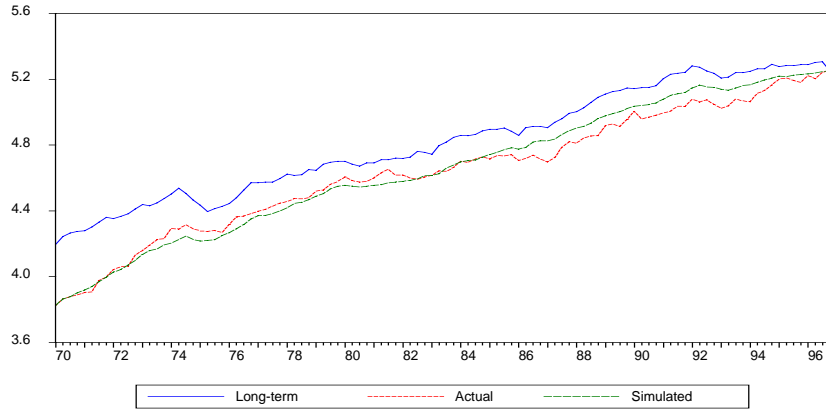
Graph 15c
Exports of manufactures: Long-term, actual, and simulated
(equation (VI), logarithmic scale)

equation, is that α_2 and α_3 are virtually negligible, indicating that a disequilibrium in export demand induces adjustment only in the first (export demand) equation. As in the import equation, this legitimises the use of a single equation estimation, supporting the assumption that world domestic demand and real effective exchange rates are weakly exogenous[64] for the parameters in the conditional export demand equation.

The estimated long-run export equations have the following form:

$$EXPNONE = 1.65*WDD - 0.62*EXULC23 \tag{17}$$

$$EXPNONE = 1.72*WDD - 0.70*EXPGDP23 \tag{18}$$

$$EXPMQ = 1.58*WDD - 0.55*EXULC23 \tag{19}$$

The results of the cointegrating equations imply that in the long-run the income elasticity of exports is quite high (one-percentage point growth in world demand increases manufactures exports by 1.58% and non-energy products by 1.65%), while the relative price elasticity is lower (respectively only -0.55 and -0.62%). The most significant implication of this result, in accordance with that found in the estimates of import functions, is that changes in France's trade balance will be dominated by real income (domestic demand) movements at home and abroad. The income responsiveness of exports is lower than the income elasticity of imports in all the three equations. This result, although problematic in terms of the external constraint on the rate of growth of economic activity in France, as will be discussed later (see section IX), is not surprising and is in accordance with the Houthakker-Magee (1969) asymmetry in the estimated elasticities found for many industrialised countries. Yet, the high value of estimated income elasticity may be due to the fact that the scale variable is likely to reflect also the impact of other non-price effects, such as product diversification, quality differences and above all supply-side effects. Furthermore, the high income elasticity may be also due to some upward bias reflecting the effect of tariff reductions and other trade liberalisation policies, which is difficult to estimate separately because of the lack of data.

In order to test the robustness of these findings with respect to the choice of the relative price term we have re-estimated the long-run relationship using different competitiveness measures. The results of the Johansen cointegration analysis, reported separately in Annex E, Table E1, turned out to be very similar to those reported in Table 9.

In the next step of the analysis, the estimated cointegrating vector was embedded in a dynamic error-correction model which was estimated by OLS. Its specification follows closely that of the imports equations (see equation (12) and the discussion in sections IV and V). Initially, four lags of the first-differenced variables were used, which provided a sufficient representation of the process generating the quarterly time series. Then, following a general-to-specific simplification procedure, we have eliminated negligible and insignificant effects through a sequential reduction. The final restricted conditional model resulting from the simplification procedure are reported in the lower panel of Table 9[65]. All regression coefficients show the expected signs and the diagnostic test statistics are quite satisfactory. The three models track the size and the direction of changes in the volume of non-energy and manufactured exports fairly well (see Graphs 15a-15c).

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[64] See comment in footnote 46 on the exogeneity status of variables.

[65] Following the approach suggested by Hendry we have taken into account statistics associated with the implied reductions, that is F statistic, Schwarz criterion and standard error of regression.

The parameter estimate on the error-correction term, that reflects the impact on Δx of having x_{t-1} out of the long-run equilibrium condition, has the right sign (which must be negative for the purpose of dynamic stability and cointegration) and appears significant at 5% level. Therefore, the deviation from the equilibrium due to random shocks represents a significant determinant of the export short-run dynamic behaviour. The estimates show a very slow adjustment to disequilibrium (3% each quarter for the export of manufactures and 4.5% for the non-energy exports equation when relative unit labour costs is used) which can arise from shocks in determinants of export demand, revealing the presence of large adjustment costs (such as transaction costs, search costs and costs linked to modifying capacity utilisation etc.) or other factors such as contract lags or order-delivery lags.

The estimates of the three equations satisfy all of the diagnostic statistics at their 95% critical value, so the models have approximately white noise, homoschedastic, normally distributed residuals, which can be considered innovations with respect to the current and lagged variables in the regressions. None of the regression coefficients displays any evidence of instability[66]. Therefore, the estimated models seem general enough to capture all the salient features of, and to provide an adequate representation of the process generating, the data; thus, it is possible to conclude that three satisfactory models for non-energy and manufacturing good exports have been achieved.

In view of the fact that the dynamic equations explain the rate of changes in the export variables, the adjusted R^2 value can be considered quite good (0.27 for equation IV and 0.26 for equations V and VI). The estimated standard errors (around 0.020) are also small. Taking into account the substantial fluctuations in the sample period, in particular those related to the two oil crises, the models explain the changes in export demand rather accurately (see also the plots of actual and fitted values in Graphs 15a-15c).

The evidence in Table 9 shows that both the short-run income and price elasticities are smaller than the long-run elasticities. The size of the coefficients shows large and immediate effect of a change in world demand on export demand for France's commodities (the impact elasticity, that is the response in the current quarter is around 0,80), but a smaller, and one- or even two-quarter lagged, impact of a change in competitiveness is suggested by the estimates. This is a common result in trade flows analysis where the data show that activity lags are much shorter than price or competitiveness lags.

The timing of adjustment and the size of the export price elasticity are both important for issues of exchange rate adjustments. The results of our estimate show that competitiveness has a predictable and systematic impact on export flows, but the elasticity is found to be low and below unity. A depreciation of 10 per cent, if completely passed-through into export prices, in our model will cause a reduction of almost 5 per cent in manufactures and non-energy export volumes but the reduction will take effect only after a quarter or two.

VIII. Hysteresis in exports and stability of the estimates

To investigate the issue of parameter constancy of our export models we turn now to results of recursive estimation. Structural instabilities may be interpreted as signalling hysteresis phenomena; in the event of competitiveness shocks, such as the real appreciation of the French franc through the first half 1990s, it is possible that this may have had adverse effects

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[66] Parameter instability statistics are based on the recent procedure suggested by Hansen (1992), which is aimed to test the stability of each parameter individually and the stability of all the parameters jointly; results of these tests are available upon request.

Table 10 Stability of export equations: Tests for structural break in 1992:Q2 (complete sample 1970:Q1-1996:Q4)				
	F	Prob	LR (χ²)	Prob
Non-energy exports: Equation (IV)				
Chow breakpoint test	1.248	0.246	-	-
Chow forecast test	-	-	23.039	0.189
Non-energy exports: Equation (V)				
Chow breakpoint test	1.227	0.264	-	-
Chow forecast test	-	-	21.417	0.208
Exports of manufactures: Equation (VI)				
Chow breakpoint test	1.135	0.348	-	-
Chow forecast test	-	-	220.817	0.288
Prob is the probability of drawing an F or a χ² statistic of the value shown in the Table from the respective distribution; the null hypothesis is that there in no structural break, and probability values of 5% or less can be taken to indicate rejection of the null.				

on export performance[67]. In particular, the appreciation of the real exchange rate for the French franc may have contributed to pricing French exporters out of the international markets where French products are imperfectly substitutable with other goods, and this is especially the case for exports of manufactures. On the other hand, it is possible that improved non-price competitiveness may have boosted exports above traditional trends and also above the level predicted by the economically relevant determinants.

As in the case of the import functions, we have first examined the stability of long-term relationship, but unlike the tests reported in section VI, we have chosen here to run a recursive *FIML* estimate of the *VECM* instead. The Graphs shown in Annex F plot the recursive eigenvalues associate with each eigenvector. These plots correspond to the standard graphs of recursive parameter estimates in the usual models, since non-constancy of the weighting elements α and/or in the elements of matrix β would be reflected in the non-constancy of the eigenvalues[68]. For the three long-term export models there is no evidence of parameter instability due to failure to account for structural breaks. Thus, the estimated cointegration vectors appear to be temporally stable.
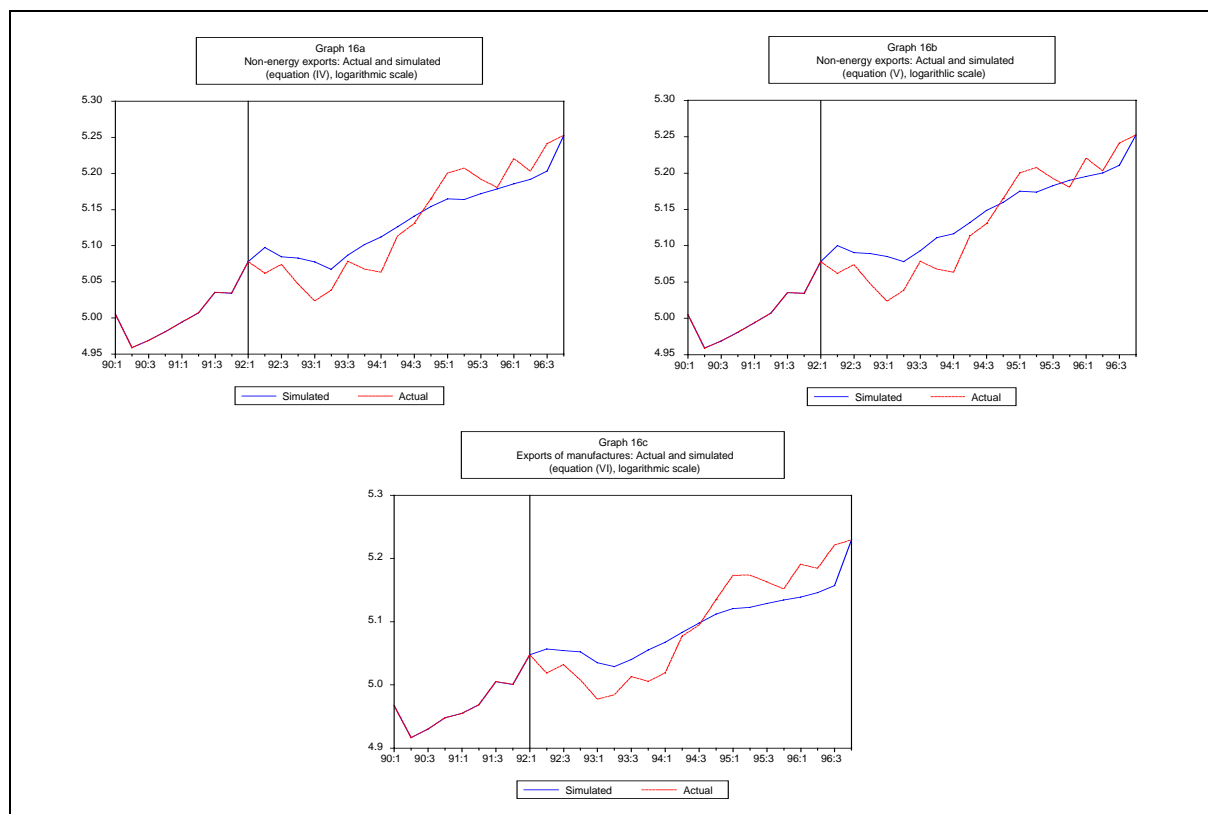
Regarding the stability of the dynamic error-correction model, the Graphs in Annex G show the result of the CUSUM and CUSUM of square tests; again, as in the case of imports, there is no evidence of structural instabilities in the dynamic equations.

Further insights on the presence of instability in the equations are provided by graphs in Annex H which plot the values of the recursively estimated coefficients of the dynamic model along with their (±2) standard errors, which provide an approximate 95% confidence interval. The intention here is to uncover evidence for instability with reference to the particular parameters of the equations. Again, all coefficients appear to remain inside the ex-ante standard errors for the entire period and vary little, confirming the structural stability of each of the three models.

As a further test for the presence of structural breaks, possibly arising from the real appreciation of the French franc since the beginning of 1992, we have used Chow test for predictive failure and Hendry's forecasting test; we implemented the test, as in the case of imports, by re-estimating the model up to 1992:Q1 and making one-step forecasts for the period ahead. Once again, there is no evidence of parameter instability (see the test statistics in Table 10). This is in contrast to the findings in the case of imports, where there appears to have been some instability problems, associated presumably with the competitiveness shock beginning in 1992:Q1 especially in the case of non-energy imports, equations (I) and (II) - see Table 7.

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[67] See Amano et. al. (1993) for a similar investigation of hysteresis in Canadian trade flows in response to the Canadian dollar's appreciation.

[68] See Harris (1995), p. 91.

Graph 16a
Non-energy exports: Actual and simulated
(equation (IV), logarithmic scale)

Graph 16b
Non-energy exports: Actual and simulated
(equation (V), logarithlic scale)

Graph 16c
Exports of manufactures: Actual and simulated
(equation (VI), logarithmic scale)

A final test of the stability of the estimated dynamic models was performed in terms of the out-of-sample predictive power of the equations. As in the case of imports, the equations were estimated up to 1992:Q1 and the coefficients so obtained were used to forecast out to 1996:Q3. The quarterly predictive performance is each equation is shown in Graphs 16a-16c. The equations forecast generally well, and clearly better than the companion import equations - compare with results of the present Graphs with those presented in Graphs 14a-14c. The three equations tend to over-predict exports in the 1993 recession but, overall, the results are very good since the deviations from actual are notably modest, especially considering the sharp movements in export volumes during this period. Thus, the evidence from the stability tests confirms that the estimated export functions are temporally stable.

IX. France's external constraint and economic growth

The estimated long-run price and income elasticities of France's trade flows allow a calculation of the severity of the external balance constraint on domestic policy choices and, more specifically, on the growth rate of domestic demand. Although France does not have a target for its external accounts, the underlying demand and competitiveness elasticities of trade flows impose constraints on demand growth given the state of the external accounts. Given these elasticities, demand or competitiveness shocks can explain movements in the external accounts such as the surplus of the 1990s. In particular, the external surplus will decline (increase), ceteris paribus, if domestic demand grows faster (slower) than foreign demand; alternatively, competitiveness elasticities, or the Marshall-Lerner condition, will determine, ceteris paribus, whether the nominal trade surplus will decline as a result of the exchange appreciation of the 1990s. Presently, the stylised model, first proposed by Johnson

(1958) and subsequently used widely in the literature[69] is the vehicle through which the analysis is prepared.

The standard foreign trade demand functions, on which we have carried out econometric estimates, can be expressed as follows:

$$X = Y_w^{\eta x,y} \ (P_i/eP^*)^{\eta x,p} \tag{20}$$

$$M = Y^{\eta m,y} \ (eP^*/P_i)^{\eta m,p} \tag{21}$$

where X = volume of exports, M = volume of imports, Y_w = world demand, Y = domestic demand, P_i = export or domestic prices, P^* = foreign prices in foreign currencies, e = nominal exchange rate, $\eta_{x,y}$ = (foreign) demand elasticity of exports, $\eta_{x,p}$ = relative-price elasticity of exports (absolute value), $\eta_{m,y}$ = income elasticity of imports, $\eta_{m,p}$ = relative-price elasticity of imports (absolute value) and $P_w = eP^*$ = foreign prices in domestic currency.

In terms of rate of change (X' and M'), equations (20) and (21) can be written as follows:

$$X' = \eta_{x,y} \ Y'_w - \eta_{x,p} \ P'_i \ + \eta_{x,p} \ P'_w \tag{22}$$

$$M' = \eta_{m,y} \ Y' - \eta_{m,p} \ P'_i \ + \eta_{m,p} \ P'_w \tag{23}$$

The trade balance in nominal terms is:

$$(Vx - Vm) = P_i * X - P_w * M \tag{24}$$

where Vx = value of exports and Vm = value of imports.

Differentiating equation (24) with respect to time, and noting this as $d = \delta f / \delta t$, we obtain:

$$d(Vx - Vm) = (XdP_i + P_idX) - (MdP_w + P_wdM) \tag{25}$$

where, using the definitions $X = Vx/P_i$ and $M = Vm/P_w$ and expressing the right hand in terms of rates of change, yields:

$$d(Vx - Vm) = Vx \ (P'_i + X') - Vm(P'_w + M') \tag{26}$$

Substituting from (23) and (24) we have:

$$\begin{aligned} d(Vx - Vm) &= Vx(P'_i + \eta_{x,y} \ Y'_w - \eta_{x,p} \ P'_i + \ \eta_{x,p} \ P'_w) \\ &\quad - Vm(P'_w + \eta_{m,y} \ Y' - \eta_{m,p} \ P'_i \ \eta_{m,p} P'_w) \\ &= Vx \ (P'_i \ (1 - \eta_{x,p}) + \eta_{x,p} \ P'w + \eta_{x,y} \ Y'w) \\ &\quad - Vm(\eta_{m,p} \ P'_i \ (1 - \eta_{m,p})P'w + \eta_{m,y}Y) \end{aligned} \tag{27}$$

[69] See Thirlwall (1978), Rollet (1983), Carone (1995) and Castaldo, Palmisani and Rossi (1986) from which material for this section has been drawn.

Imposing the invariance of trade balance condition, $d(Vx - Vm) = 0$, we can solve the (27) for the rate of growth of domestic demand which is consistent with such invariance :

$$Y' = (Vx/Vm) * (\eta_{x,y}/\eta_{m,y}) * Y'_w + P'_i * ((Vx/Vm) * (1-\eta_{x,p}) - \eta_{m,p}) * 1/\eta_{m,y}$$

$$- P'_w * (1-\eta_{m,p} - (Vx/Vm) * \eta_{x,p}) * 1/\eta_{m,y} \tag{28}$$

On the assumption that trade balance is initially in equilibrium, (28) can be rewritten as:

$$Y' = (\eta_{x,y}/\eta_{m,y}) * Y'_w + (P'_w - P'_i) * (\eta_{m,p} + \eta_{x,p} - 1) * 1/\eta_{m,y} \tag{29}$$

Equations (28) and (29) are the basic analytical formulations of the external constraint, and show the critical parameters which, given the evolution of world demand and of relative prices or competitiveness, determine the rate of growth of domestic activity compatible with the initial external imbalances (equation (28)), or with equilibrium in the external accounts (equation (29)). Correspondingly, for any rate of growth of domestic demand and evolution of world demand, these equations can be solved for the size of the change in competitiveness (that is in the nominal exchange rate or in relative costs or prices) which would be necessary to achieve a given trade balance position. Finally, the equations summarize succinctly the factors that impinge on external performance, and ultimately permit us to decompose the trade surplus according to the contributing factors.

On the assumption of unchanged competitiveness $(P'_i = P'_w)$, equation (28), in which the trade balance is in disequilibrium, becomes:

$$Y' = Y'_w (\eta_{x,y}/\eta_{m,y}) * (Vx/Vm) + P'_i * (Vx-Vm)/Vm * 1/\eta_{m,y} \tag{30}$$

while equation (30), where the trade balance is assumed to be in equilibrium $(Vx-Vm = 0)$, becomes:

$$Y' = (\eta_{x,y}/\eta_{m,y}) * Y'_w \tag{31}$$

or

$$Y'/Y'_w = \eta_{x,y}/\eta_{m,y} \tag{32}$$

It is clear that, ceteris paribus, the external constraint is less binding on domestic economic growth, the higher is the ratio of income elasticities of exports and imports, $(\eta_{x,y}/\eta_{m,y})$; moreover, in the case of equations (28) and (30) where $Vx \neq Vm$, the external constraint is less binding the higher is the export/import ratio, Vx/Vm.

Measured values of the critical parameters, based on estimates reported in Tables B1 and E1 of the Annexes (see the explanatory note in Table 11) are presented in Table 11. The ratio of export and import elasticities with respect to income is the dynamic Harrod foreign trade multiplier. In the case of France, and on the basis of our estimates, this ratio shows that the trade balance (both for manufactures and for non-energy products) would act as a constraint on income or domestic demand growth since had it been greater than world demand growth the external surplus (deficit) would be smaller (larger), as can be seen from $\eta_{x,y}/\eta_{m,y} < 1$. On the assumption of constant competitiveness and initially unbalanced trade, it is clear that, using the values of imports and exports in 1995, the maximum rate of growth of domestic demand in 1995 which would have permitted to sustain a given trade balance position was

around the 70-80% of those of the main trading partners (parameter (iii) in Table 11)[70]. In other words, the surplus of the 1990s is consistent with the slow growth in France relative to the rest of the world.

The critical factor $(\eta_{m,p} + \eta_{x,p} -1)$, which determines the direction of the impact of a change in relative prices on the trade balance, is the well-known Marshall-Lerner condition[71]. If the sum of price elasticities is greater than 1, then a depreciation (appreciation) of the real effective exchange rate will make possible a higher (lower) growth path for a given value of the trade balance. Computing these values with the long-term trade elasticities of the French trade flows, from the estimates cited in the note of Table 11, we see that sum of the price elasticities is somewhat smaller than unity. This implies that shocks in relative prices or competitiveness, such as an appreciation (completely passed-through to export prices), does not produce substantial changes in the nominal trade balance;
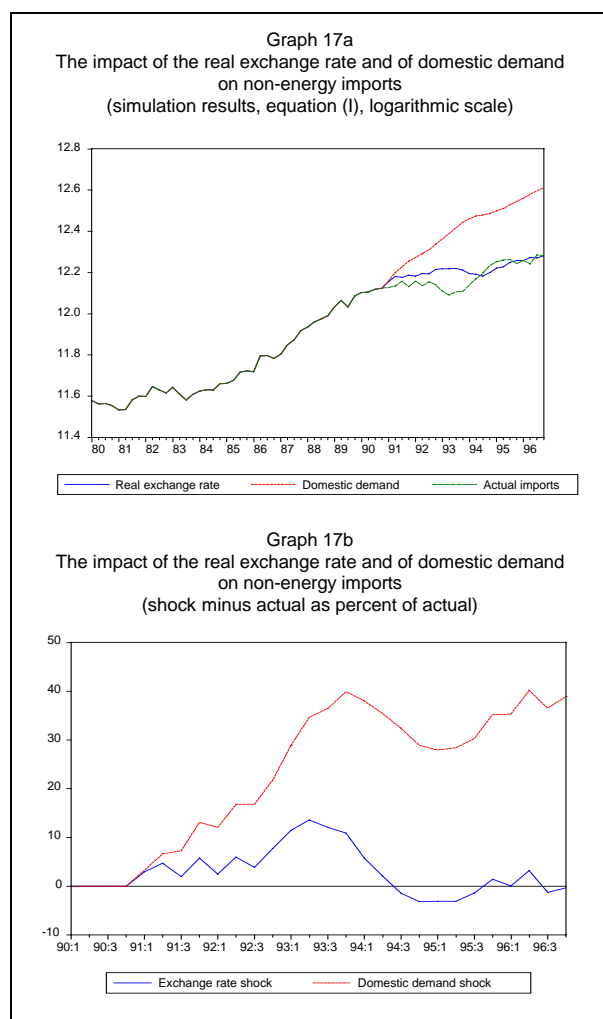
Table 11 — The external constraint and economic growth in France		
Critical parameters for the external constraint	Non-energy trade balance	Trade balance on manufactures
(I) (Vx/Vm)	1.08	1.05
(ii) $(\eta_{x,y}/\eta_{m,y})$	0.71	0.67
(iii) $(\eta_{x,y}/\eta_{m,y})* (Vx/Vm)$	0.77	0.70
(iv) $(\eta_{m,p} + \eta_{x,p} -1)$	-0.07	-0.06

Source: For uniformity, the calculations are based on estimates in equation (18) and the corresponding equation in Annex Table B1 in the case of non-energy trade, and the corresponding equations defined over EXUWC23 in Annex Tables B1 and E1 in the case of trade in manufactures.

that is, the real trade balance does not deteriorate sufficiently to offset the direct impact of the appreciation on the value of imports. Note, however, that this condition assumes that the trade balance is initially zero. The most general condition, which takes account of the starting position for the external balance, is given by $((Vx/Vm *\eta_{x,p}) + \eta_{m,p} -1)$. In the case of France in 1995, when the trade balance of both non-energy and of manufactured products were in surplus, the sum of the elasticities, with $\eta_{x,p}$ weighted by the positive ratio Vx/Vm and for a given $\eta_{m,p}$, would approach or increase above unity, ensuring that an appreciation would tend to cause a reduction (increase) in the surplus (deficit). Nevertheless, given the size of the export/import ratio in 1995, the effect of the appreciation (depreciation) on the trade balance would likely be quantitatively trivial. Finally, in view of this consideration, an unrealistically large nominal exchange appreciation would be necessary to reduce France's trade surplus.

X. The trade surplus of the 1990s: Evidence from simulations

The factors contributing to the external surplus of the 1990s, as noted in the previous section, are domestic and external demand and movements in the real exchange rate; in this section the contribution of these factors is examined with the use of simulation analysis. Barring the construction of a full macroeconomic model of France, the simulations here represent ceteris paribus solutions to the import equation without regard for the consequent macroeconomic adjustments. The simulations permit us to see, under counterfactual conditions resembling the evolution of key domestic French and external competitiveness variables, what would have been France's external performance had historical trends of the 1980s prevailed in the 1990s.

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[70]   Similar results are reported in Castaldo et al. (1986).

[71]   The Marshall-Lerner (alternatively known as the Marshall-Lerner-Robinson) condition states that a nominal depreciation (appreciation) will improve (lead to a deterioration in) the external surplus if the sum of the price elasticities of domestic and foreign demands for imports exceeds unity.

Graph 17a
The impact of the real exchange rate and of domestic demand
on non-energy imports
(simulation results, equation (I), logarithmic scale)

Graph 17b
The impact of the real exchange rate and of domestic demand
on non-energy imports
(shock minus actual as percent of actual)

A representative import equation estimated previously is used to conduct the simulations. Since there is no evidence of instability in the export functions, it confirms that the source of the surplus cannot have been the superior export performance, because actual exports have evolved according to the prediction of the equations. It is, rather, that in the import side of the accounts that the source of the surplus can be found. Consequently, the simulations have been based on an estimated representative import equation.

Graph 12 and the discussion in sections V, VII and IX have suggested that the slow growth of domestic demand in France has likely been an important factor in the external surplus. One way to evaluate this is to assume that demand growth in France in the 1990s has followed the same trend as in the second half of the 1980s and, consistent with the evidence from the stability tests, that there is no break in the relationship between imports and domestic demand. Fitting a trend on domestic demand over the period 1985:Q1-1990:Q4 yields the following equation: DDO2 = 1077.899 + 31.35197*Time; this equation was used to forecast domestic demand over the period 1991:Q1-1996:Q4, and the latter variable was used to simulate the demand for imports over this period. Correspondingly, actual data for movements in the franc real exchange rate in the period 1990:Q1-1996:Q4 was used to construct the real exchange rate shock. Using as a representative demand for imports equation (I), demand for non-energy imports, the results shown in Graphs 17a and 17b are obtained. In Graph 17a the (logarithmic) level of non-energy imports in 1980 prices is presented, while Graph 17b shows the deviation of imports from their actual value as percent of actual according to each shock.

As can be seen in Graphs 17a and 17b, the appreciation of the franc real exchange rate has, ceteris paribus, contributed to raising the demand for non-energy imports by as much as 13% at its peak in 1993:Q2. However, during the overall period the impact of the appreciation has been modest; after 1993 import demand returns to its actual level and the impact of the real exchange rate shocks fades away, as can be seen from the convergence of actual and simulated imports under this shock.

Substantially more pronounced is the impact of the domestic demand shock. Had domestic demand followed the growth path of the second half of the 1980s, imports would have increased markedly in the 1990s, especially during the slow growth and the recession quarters of 1992 and 1993, as can be seen from Graph 17a. The recovery of domestic demand growth in 1994 and 1995, by bringing domestic demand closer to its trend value of the second half of the 1980s, contributed towards reducing the deviation between the simulated and actual

imports during this period. The results of this simulation show conclusively that a principal factor in the external surplus of the 1990s can be found in the import side of the accounts, where the recession and slow growth of domestic demand have depressed imports. As can be seen from Graph 17b, the volume of non-energy imports in this counterfactual would have been on average over 30% higher than actual in the post-1993 period thus, ceteris paribus, not only eliminating completely the external surplus but in fact causing the accounts to move into deficit.

XI. Concluding remarks

France's external performance since the beginning of the 1990s has been remarkable, with the merchandise trade and the current account balance rising to over 1.5% of GDP in 1996. These developments have taken place against a loss of price and cost competitiveness, and in the context of slow growth in France relative to its major partners. It has been argued that the external surpluses are indicative of changing fortunes, characterized by improving trade performance symptomatic of structural changes in import and export equations. Closer review of the data suggests that since 1990 the relatively larger rise in export than in import volumes dominates the contribution of the terms of trade in movements in both the non-energy trade balance and in the balance on manufactures.

In the paper, the behavioural content of the trade flow adjustments in response to competitiveness and output shocks was developed in a cointegration/error-correction framework. This framework permits the determination of the long-term equilibrium relationship between imports, and exports, and their economic determinants while, at the same time, it allows the characterization of adjustment towards equilibrium when deviations of predicted from actual trade flows emerge. This characterization is important since trade flows respond to shocks with variable lags and with varying weights. The results obtained in the sample of quarterly data starting in the beginning of the 1970s to the end of 1996 show that the coefficient adjustment in question is very large in the case of imports (where deviations from equilibrium are virtually eliminated with a year) but substantially slower in the case of exports. This implies that movements in the external accounts in response to shocks reflect predominantly adjustments in the import side of the accounts.

Both the cointegration relationship and the dynamic adjustment part of the estimated econometric models for imports, and for exports, of non-energy products and for manufactures appear to be temporally stable. However, there is some qualified evidence of instability in some specifications of the import equations. This instability, which appears to be temporary and does not affect the forecasting performance of the import equations, is likely related to a temporary break in the competitiveness elasticity in the beginning of 1992 when the large real appreciation of the French franc commenced. While instability is hinted at in the import equation for non-energy products, there is no evidence of instability in the case of the largest component of France's imports, imports of manufactures. In view of the broad stability of the import functions, and together with the absence of evidence of instability in the export functions, it is possible to reject the hypothesis that part of the surplus of the 1990s is in any way due to better performance in trade unrelated to historical relationships. The stability tests also provide some preliminary, but not conclusive, evidence against the hypothesis of hysteresis in French trade flows as a result of the competitiveness shocks of the 1990s.

The econometric results confirm that, consistent with earlier empirical work, the long-run demand elasticity of imports and of exports is substantially higher than the competitiveness

elasticity. By implication, the demand shocks associated with the recession and slow growth of the first half of the 1990s have played a dominant role in the growth of imports below the path observed during the high growth period of the second half of the 1980s. Simulations show that the dominant part of the surplus of the 1990s can be attributed immediately to the lower growth trajectory of domestic demand in France compared to its path in the second half of the 1980s which has led to a decline in the volume of imports below their trajectory over the second half of the 1980s. It is clear that, should recovery of growth at historical rates be achieved in coming quarters, the external accounts will undoubtedly see the disappearance of the surplus.

The importance of economic growth in determining movements in France's external accounts in the 1990s can also be evaluated from the perspective of how binding the external constraint for economic growth is. In particular, given the estimated trade flow elasticities with respect to demand and to competitiveness and given world demand growth, a given trade balance can be solved either for the implicit domestic demand growth consistent with this configuration; or for the level of competitiveness consistent with a given trade balance. The empirical estimates suggest that, not surprisingly, the external surplus is consistent with the slowdown on economic growth in France relative to her trading partners in the 1990s. Indeed, the estimates suggest that the good trade performance is consistent with demand growth in France in the range of 70% to 80% of demand growth abroad.

Much concern has been raised by the ERM turbulence of 1992 and 1993 in its potential adverse implications of the depreciations of several currencies for France's price and cost competitiveness, but the results obtained here do not support these concerns. France's trade flows are elastic with respect to competitiveness shocks although these elasticities are dwarfed by the demand elasticities. The estimates also suggest that there is a significant quantitative difference between the competitiveness elasticity of exports compared to that of imports, with the former being two to three times as large as the latter. However, for the trade balance, it is the sum of the competitiveness elasticities that matter, and the estimates suggest that the Marshall-Lerner condition in the sample under review is just met. Consequently, the evidence does not lend to this issue, which so exercized the French authorities in the aftermath of the ERM exchange rate crises of the early 1990s; however, although these shocks do not impinge significantly on the trade balance, they will have significantly more impact on the export-oriented sector of the economy.

A fundamental change in EMU will be that nominal exchange rate adjustments will not be available to affect temporarily competitiveness. Consequently, sustaining good competitiveness trends would require that, among others, favourable cost trends, or policies of competitive disinflation, continue to prevail. There are undoubtedly limits to such policies and, in an environment of low inflation and disciplined cost developments, such limits may be rapidly reached. At the same time, it is likely that the synchronization of business cycles across the EU Member States will be reinforced through policy co-ordination, with the result that France's demand-determined external account movements and the frequency of external imbalances would tend to be reduced. In these circumstances, sustaining a good trade performance would require principally policies enhancing modernization and innovation of the enterprise sector as well as improvements in the quality of factor inputs in the competitive sector of the economy.

ANNEX A

Sources and Time Series Properties of the Data

The principal source of the data is the INSEE databank *Séries Trimestrielles de Comptabilité Nationale de l'INSEE* and all constant-price variables are 1980 prices. Real exchange rate variables, also indexed at 1980=100, have been prepared in DG-II, Directorate D. Other variables, denoted as calculated in Table A1, have been defined as combinations of the original variables.

As noted in the text, it is essential that in order to obtain meaningful statistical results the time series used in the empirical work are stationary. Results of stationarity tests are presented in Table A1. All data are in logarithms. The equation used in the testing procedure was defined over an intercept alone; identical results were also obtained from an equation including a time

Table A1 Sources and time-series properties (stationarity tests) of the data (Sample 1970:Q1-1996:Q4, except prices 1970:Q1-1996:Q3)						
		Level		First difference		
Variable	Source	DF	ADF(4)	DF	ADF(4)	"I(z)
Domestic demand	PDI_V008	-1.90	-2.48	-8.25	-4.67	I(1)
Non-energy imports	P60_V008-P60_U038	-1.74	-1.82	-8.68	-5.83	I(1)
Imports of manufactures	P60_V098	-1.66	-1.76	-9.68	-6.37	I(1)
Import price/GDP deflator	Calculated	-0.25	-0.09	-6.78	-5.65	I(1)
Import price/domestic demand deflator	Calculated	-0.27	-0.17	-6.59	-5.86	I(1)
Manufactures, import/domestic price	Calculated	-0.97	-0.86	-8.91	-5.51	I(1)
Non-energy exports	P50_V0M8-P50_U038	-1.98	-1.88	-10.40	-4.35	I(1)
Exports of manufactures	P50_V098	-1.96	-1.69	-10.40	-4.66	I(1)
World domestic demand	Calculated	0.02	-0.19	-8.58	-3.98	I(1)
Real exchange rate IC-23, GDP deflators	Commission services	-2.22	-2.46	-7.56	-4.69	I(1)
Real exchange rate IC-20, GDP deflators	Commission services	-2.17	-2.37	-7.61	-4.66	I(1)
Real exchange rate EU-14, GDP deflators	Commission services	-2.49	-2.41	-8.75	-5.04	I(1)
Real exchange rate IC-23, unit labour costs	Commission services	-2.03	-2.11	-7.51	-4.82	I(1)
Real exchange rate IC-20, unit labour costs	Commission services	-1.97	-2.03	-7.56	-4.77	I(1)
Real exchange rate EU-14, unit labour costs	Commission services	-2.02	-1.98	-8.49	-4.76	I(1)
Real exchange rate IC-23, unit wage costs	Commission services	-2.66	-3.37	-7.49	-4.73	I(0)
Real exchange rate IC-20, unit wage costs	Commission services	-2.61	-3.28	-7.68	-4.69	I(0)
Real exchange rate EU-14, unit wage costs	Commission services	-2.58	-2.74	-8.59	-4.74	I(1)

DF and ADF statistics are the Dickey-Fuller and the augmented Dickey-Fuller statistics, respectively; ADF (4) is the ADF statistic of order 4; the critical value of the DF and the ADF statistics, calculated by the MacKinnon method, at the 99% level of significance, is -3.50, and at the 95% level of significance it is -2.89; the test statistics are obtained from an equation defined over an intercept; "I(z) signifies that the order of integration is z; all constant-price variables are defined in 1980 prices.

The principal source of the data is the INSEE databank *Séries Trimestrielles de Comptabilité Nationale de l' INSEE*, and each variable corresponds to the INSEE databank mnemonic indicated in the table; calculated values are from the original data; the world demand series was prepared at ISCO; the source of the remaining data is the Commission services and all measures of the real exchange rate have been prepared in DG-II, Directorate D-4.

trend. For the majority of the data the null of nonstationarity cannot be rejected in level form at either the 99% of the 95% level of significance. Correspondingly, the null of stationarity is overwhelmingly accepted when the data is first differenced; thus such variables are integrated of order one *(I(1))*. One exception to this is the data for real exchange rate defined in terms of relative unit wage costs in manufacturing against the IC-23 and the IC-20. For these

variables, only the *DF* statistics reject the null of stationarity in level form and these series appear to be *I(1)*; however, the *ADF (4)* statistics suggest that these series are *I(0)*. Finally, virtually all other variables are *I(1)* either by the *DF* or by the *ADF (4)* statistic, as can be seen from the results reported in the third and fourth columns of the Table.

ANNEX B

Further Cointegration Results, Import Equations (I) - (III)

This annex presents additional FIML cointegration estimates for import equations (I)-(III) based on alternative definitions of competitiveness. The estimation results reported in sections V and VI were based on the price of imports relative to the relevant domestic demand deflator, a strategy intended to capture pricing-to-market behaviour and similar phenomena. According to this view, competitiveness approximated by conventional real exchange rate variables would not be accurately representative of the true competitive position of exports and importers. However, competitiveness defined by real exchange rates and by relative costs are important concept in themselves, widely used in empirical trade functions, and their role deserves investigation in quantitative import functions.

A review of the empirical importance of real exchange rate variables for French imports was undertaken over the sample 1970:Q1 - 1996:Q4, and the results are presented in Table B1. The equations are cointegration vectors for non-energy imports and imports of manufactures and domestic demand and the real exchange rate. In the first panel of the Table, referring to non-energy imports, the real exchange rate variable is defined by the double export-weighted value of an index in terms of the relative GDP deflators; the second panel, also referring to non-energy imports, shows the same cointegration results for the real exchange rate defined in terms of the relative unit labour costs in total economy; and the third panel presents cointegration results for imports of manufactures where the real exchange rate term is defined by the relative unit wage costs in manufacturing. Three sets of results are reported in each panel: first, results for the real exchange rate against the IC-23; secondly, results for the real exchange rate against the IC-20; and, third, results for the real exchange rate against the 14 EU Member States. For each equation the relevant cointegration vector is also written out. The mnemonics used in the Table are as follows (all variables in logarithms):

$IMPNONE$ = non-energy imports; $DDO2$ = domestic demand including stocks; $EXPGDPj$ = real exchange rate in terms of relative GDP deflators, j indexes the country group against which the real exchange rate is measured, = 23, 20, 14; $EXULCj$ = real exchange rate in terms of relative unit labour costs in total economy, j = 23, 20, 14; $IMPMQ$ = imports of manufactures; and $EXUWCj$ = real exchange rate in terms of relative unit wage costs in manufacturing, j = 23, 20, 14. All variables are in 1980 prices and the real exchange rate indices have a value of 100 in 1980. The statistical properties of the variables used are reported in the previous Annex, Table A1.

In the first panel, the result do not reject cointegration between non-energy imports, domestic demand and the real exchange rate at the 95% level of significance, and at the largest eigenvalue the vector is unique. The long-term import equation has an identical domestic demand and real exchange rate elasticity (2.4 and 0.2, respectively) irrespective of whether the real exchange rate is defined in terms of the IC-23 or the IC-20 industrial group. In the case of imports of manufactures, however, the vector is determined at the 99% level of significance and, while the domestic demand elasticity is not different from its value estimated in the previous two equations, the real exchange rate elasticity against the EU-14, at 0.5, is more than twice as high as that reported in the previous equations.

Table B1 Imports, domestic demand and the real exchange rate Johansen likelihood ratio tests for cointegration (sample 1970:Q1 - 1996:Q4)			
	No cointegration	At most one vector	At most two vectors
Eigenvalue	*0.158*	*0.074*	*0.038*
Non-energy imports, equation (I)	30.22*	12.13	4.05*
Cointegrating vector: IMPNONE = 22.12 + 2.42*DDO2 + 0.23*EXPGDP23			
Eigenvalue	0.157	0.074	0.037
Non-energy imports, equation (I)	29.91*	12.04	4.01*
Cointegrating vector: IMPNONE = 22.08 + 2.42*DDO2 + 0.21*EXPGDP20			
Eigenvalue	*0.205*	*0.084*	*4.529*
Non-energy imports, equation (I)	37.88**	13.78	4.53*
Cointegrating vector: IMPNONE = 23.84 + 2.46*DDO2 + 0.49*EXPGDP14			
Eigenvalue	*0.175*	*0.065*	*0.038*
Non-energy imports, equation (II)	31.60*	11.21	4.08*
Cointegrating vector: IMPNONE = 22.01 + 2.45*DDO2 + 0.27*EXULC23			
Eigenvalue	*0.173*	*0.062*	*0.038*
Non-energy imports, equation (II)	31.08*	10.88	4.14*
Cointegrating vector: IMPNONE = 21.95 + 2.45*DDO2 + 0.26*EXULC20			
Eigenvalue	*0.171*	*0.067*	*0.034*
Non-energy imports, equation (II)	30.89*	11.07	3.72
Cointegrating vector: IMPNONE = 21.99 + 2.44*DDO2 + 0.27*EXULC14			
Eigenvalue	*0.209*	*0.104*	*0.038*
Imports of manufactures, equation (III)	40.17**	15.56*	4.03*
Cointegrating vector: IMPMQ = 23.77 + 2.51*DDO2 + 0.29*EXUWC23			
Eigenvalue	*0.207*	*0.103*	*0.037*
Imports of manufactures, equation (III)	39.80**	15.43*	3.96*
Cointegrating vector: IMPMQ = 23.68 + 2.51*DDO2 + 0.27*EXUWC20			
Eigenvalue	*0.211*	*0.076*	*0.044*
Imports of manufactures, equation (III)	37.90**	12.96	4.70*
Cointegrating vector: IMPMQ = 23.43 + 2.51*DDO2 + 0.22*EXUWC14			
The null hypotheses are that there is no cointegration, that there is at most one cointegration vector, and that there is at most two cointegration vectors; the 5% critical values of the likelihood ratio statistic at the three largest eigenvalues reported for each equation in the Table are 29.68, 15.41 and 3.76; the 1% critical values are, correspondingly, 35.65, 20.04 and 6.65; * (**) rejects the null at the 5% (1%) level of significance; equations (I), (II), and (III) refer to those in Table 5 of the main text.			

These estimates appear to suggest that price competitiveness in manufacturing against the EU-14 group is quantitatively more important than compared to the other trading groups because intra-EU real exchange rate shocks influence imports of manufactures significantly more than other competitiveness shocks.

The results reported in the second panel suggest that while cointegration is generally supported by the data, this occurs at only the 95% level of significance; no vector is significant at a more stringent level. Thus, the definition of cost competitiveness by the relative unit labour costs in the total economy finds somewhat weaker support in the long-term import equations. Nevertheless, at this level of significance, the estimated domestic demand elasticity of imports is virtually identical in the three equations. Moreover, the elasticity of imports with respect to the real exchange rate is identical in the three equations.

The third panel of the Table presents results for imports of manufactures with the cost competitiveness term approximated by relative unit wage costs in manufacturing. In all equations the presence of cointegration is supported at the 99% level of significance, irrespective of whether the trading group is the IC-23, the IC-20, or the EU-14. A second cointegration vector is supported at the 95% level of significance and arises from the stationarity of the competitiveness term. The practical implication of including a stationary variable in a *VECM* is that the cointegration rank will increase and the additional cointegration vector in β contains only this variable. These alternative cointegration vectors had no economic meaning and, as a result, only one vector is reported. The results show that the domestic demand elasticity of imports is estimated to be 2.51 across the three equations for imports of manufactures. The real exchange rate elasticity, on the other hand, is lower in the case of imports of manufactures against the EU-14 competitiveness .

The results of this Annex provide support for using alternative measures of competitiveness even if they are not conceptually as defensible as the real exchange rate variables used in the main part of the paper. They also suggest that the assessment of competitiveness may need to be made on the basis of alternative measures since no one measure can be isolated as the best among those available.

ANNEX C

Stability of Dynamic Import Equations (I)-(III)

Graphs C1-C3 present CUSUM and CUSUM of Squares tests and corresponding 5% significance levels for import equations (I)-(III).


Graph C1: Non-energy imports, equation (I)

Graph C2: Non-energy imports, equation (II)

Graph C3: Imports of manufactures, equation (III)
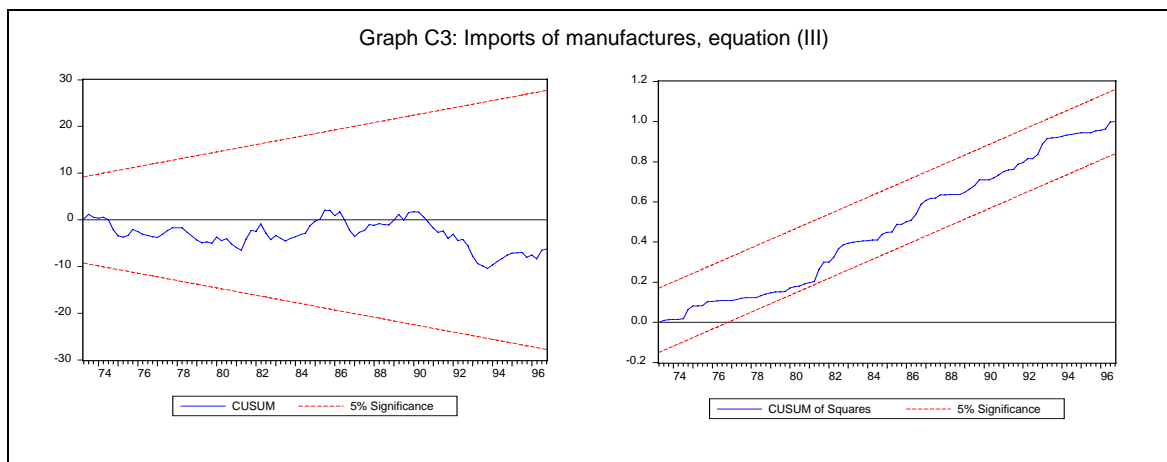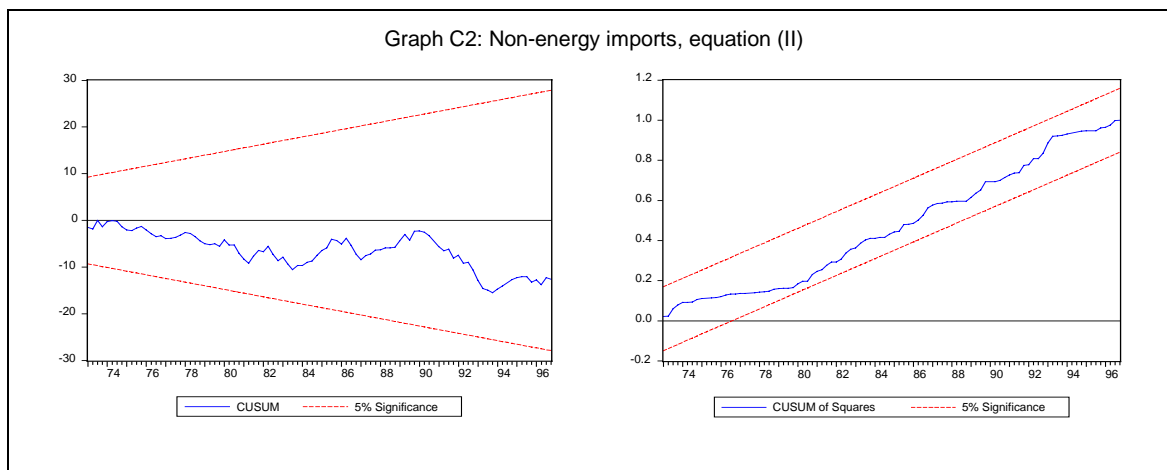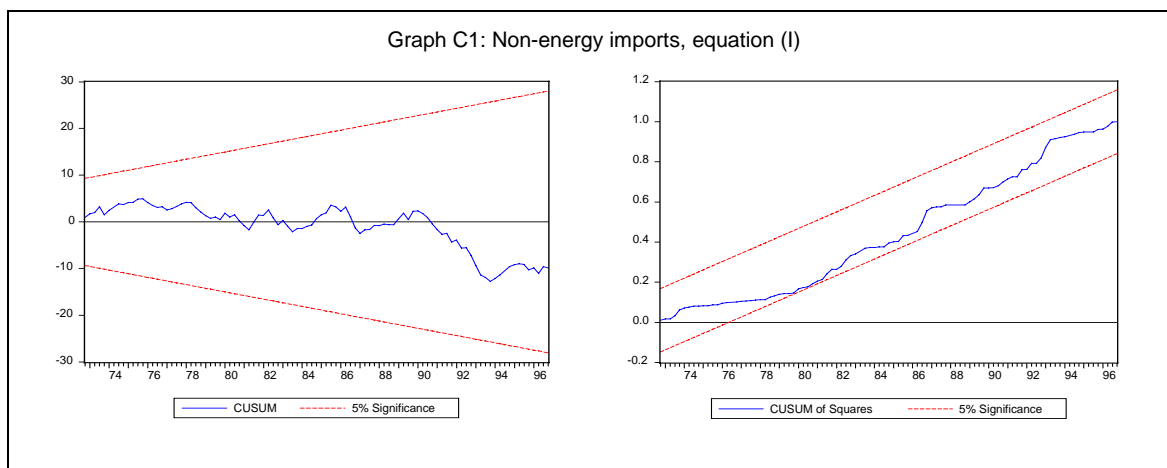
ANNEX D

Stability of Coefficients of Dynamic Import Equations (I)-(III)

Stability tests for the coefficients of the short-term adjustment model of equations (I)-(III) are presented, correspondingly, in Graphs D1-D3. The graphs show the recursive estimates for each coefficient for the period 1970:Q1-1996:Q4, except the constant term which shows no sign of instability and was omitted for reasons of economy, and two standard errors. The estimates confirm that the import equations are temporally stable.

Graph D1
Coefficient stability of equation (I)
(Recursive estimates and two standard errors)
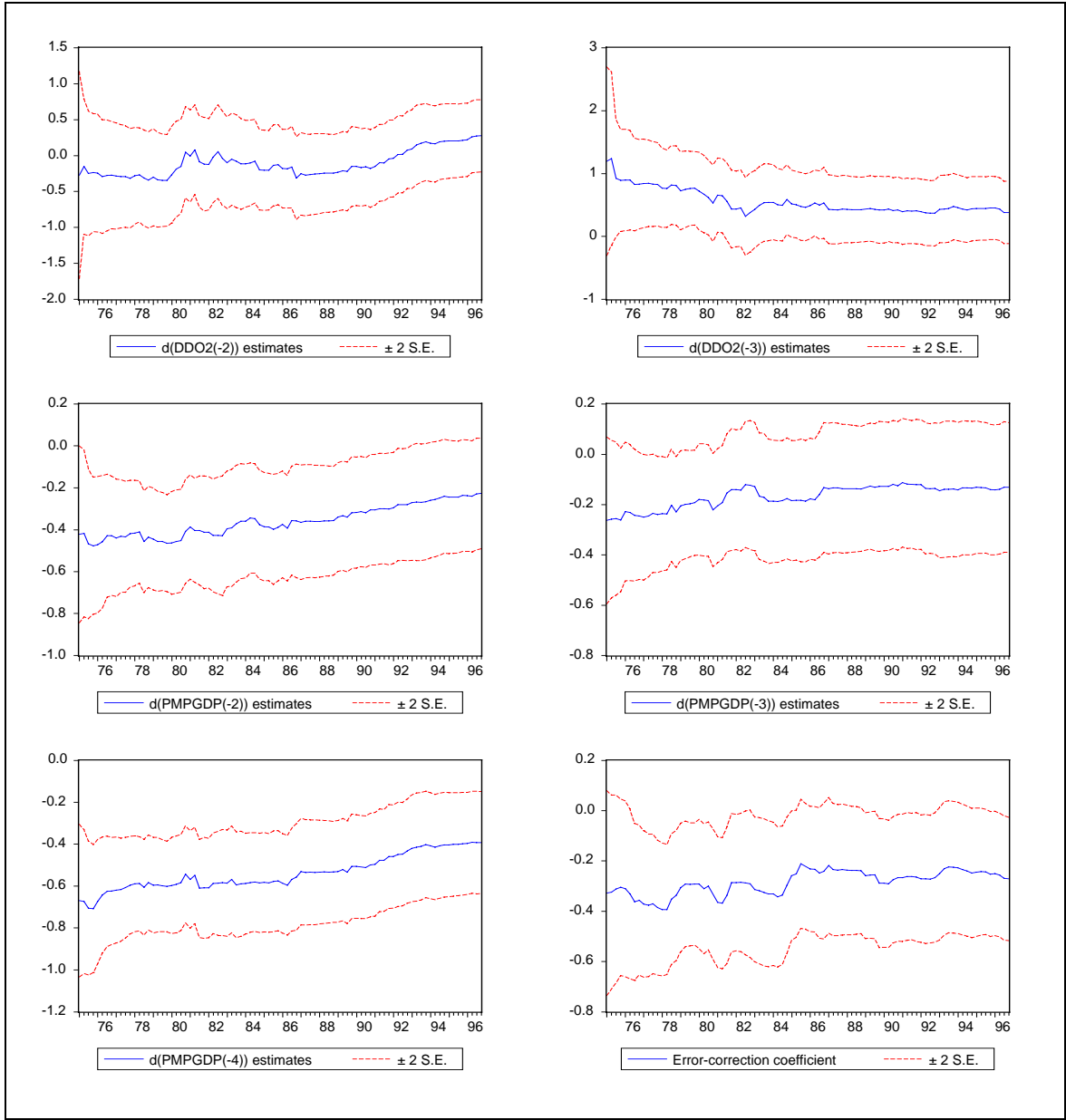
Graph D2
Coefficient stability of equation (II)
(Recursive estimates and two standard errors)
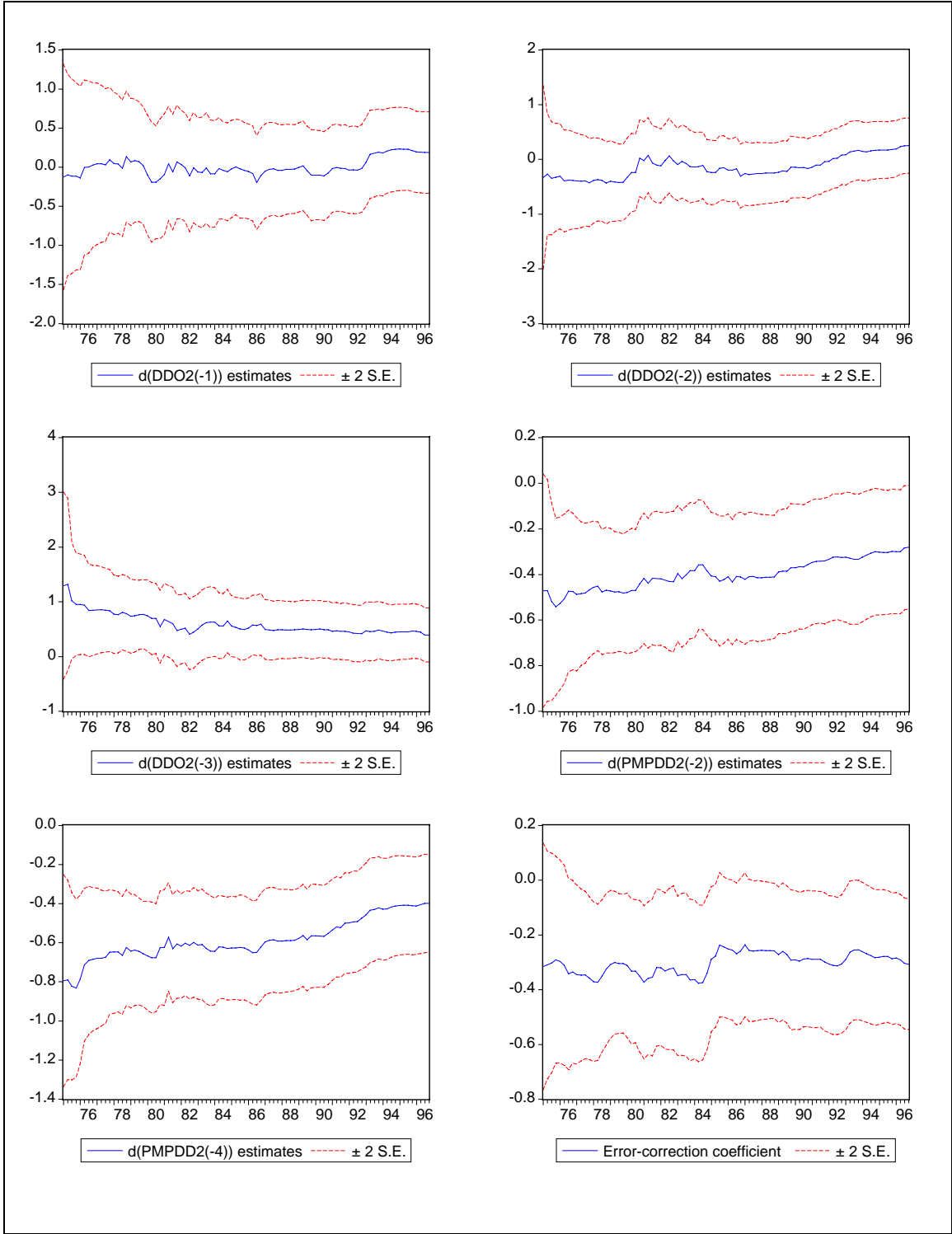
Graph D3
Coefficient stability of equation (III)
(Recursive estimates and two standard errors)
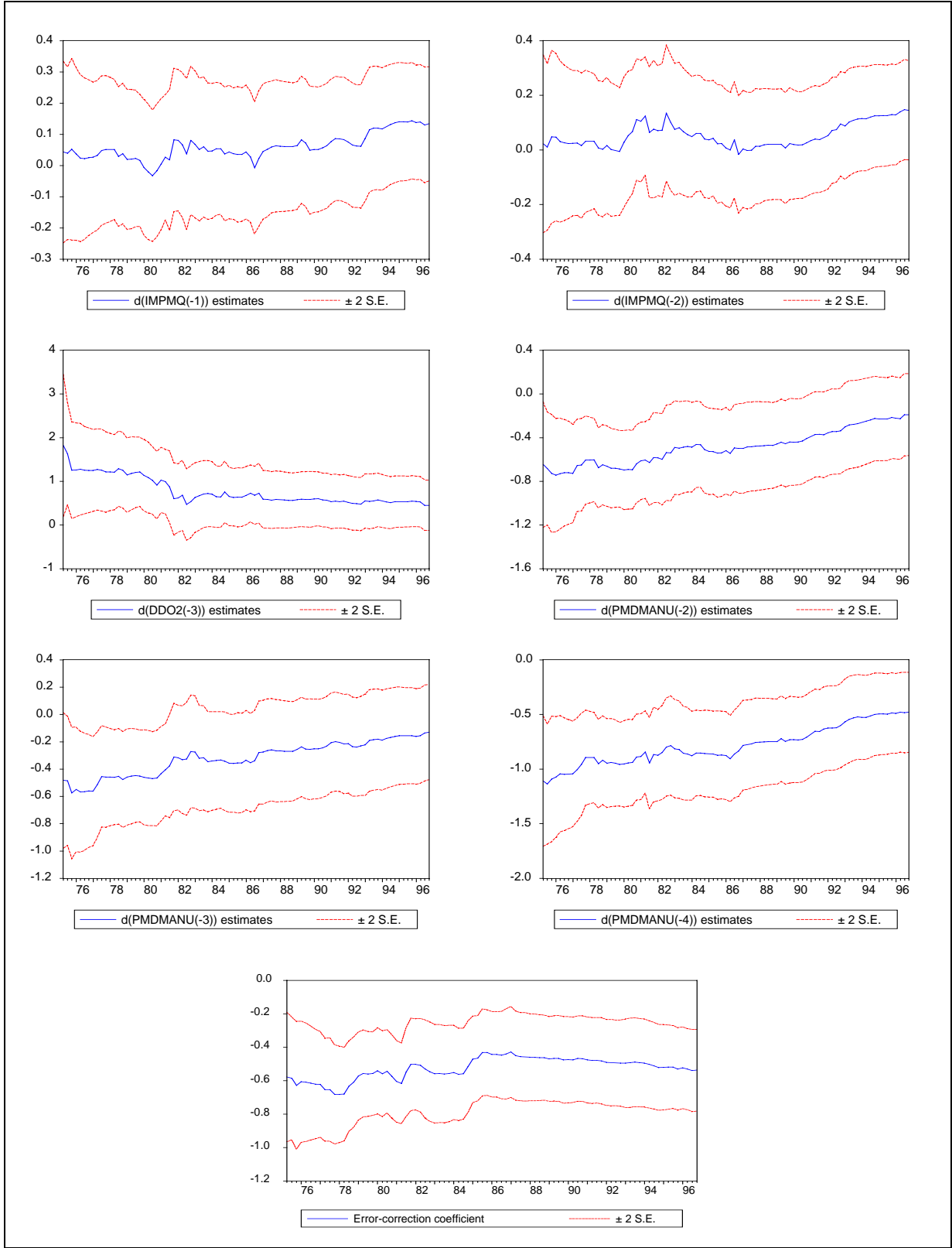
ANNEX E

Further Cointegration Results, Export Equations (IV)-(VI)

Additional results from the FIML estimation of cointegrating vectors for export equations (IV) -(VI) based on somewhat different definitions of competitiveness are reported in the present Annex. For purposes of comparison results of the estimates of the equations presented in Table 9 of the text are also reported in Table E1. For both the aggregates under investigation(non-energy exports and exports of manufactures) cointegration testing and estimation have been carried out using as index of competitiveness measures of the real exchange rate based on relative unit labour costs in France against the IC-23, the IC-20 and the EC-14; on relative GDP deflators against the same groups of countries; and on relative unit wage costs in the manufacturing sector also against the same groups of countries. More details on these variables are in Annex A, and further comments on the specification of the cointegration vectors can be found in Annex B where similar results for the import equations are reported.

The mnemonics used in Table E1 are as follows (all variables in logarithms):

$EXPNONE$ = non-energy exports; WDD = world domestic demand; $EXPGDPj$ = real exchange rate in terms of relative GDP deflators, j indexes the country group against which the real exchange rate is measured, = 23, 20, 14; $EXULCj$ = real exchange rate in terms of relative unit labour costs in total economy, j = 23, 20, 14; $EXPMQ$ = exports of manufactures; and $EXUWCj$ = real exchange rate in terms of relative unit wage costs in manufacturing, j = 23, 20, 14. All variables are in 1980 prices and the real exchange rate indices have a value of 100 in 1980. The statistical properties of the variables used are reported in the previous Annex, Table A1.

The results of the econometric analysis suggest that cointegration is generally supported by the data at 99% level of significance and in all cases the vector is unique. This contrasts with the results for the cointegration tests for the equations for non-energy imports, where the null of no cointegration was rejected at the 95% but not at the 99% level of significance for the reasons discussed in Annex B- see Table B1.

The results also show that there are only slight differences in the estimated elasticities for the equations based of alternative definitions of competitiveness. Income elasticities range from a high of 1.72 in the case of the equation defined over $EXPGDP23$ to a low of 1.60 in the case of exports of manufactures defined over $EXUWC14$. The price elasticities range from 0.55 in the case of exports of manufactures defined over $EXUWC14$ to 0.69 in the case of non-energy exports defined either over $EXPGDP23$ or over $EXPGDP20$.

The lowest competitiveness elasticities in the long-term export demand functions are in the case where the real exchange rate refers to the EC-14 group. As noted in Annex B, alternative measures of competitiveness give consistently good estimates of the elasticity of trade flows with respect to the real exchange rate and no a priori measure of competitiveness can be singled out as being the best. Although on theoretical grounds some measures can be rejected, the broadly similar estimates that the various measures examined here yield suggest that, depending on the use of the equations, little information is lost when alternative measures of the real exchange rate are used.

Table E1 Exports, world demand and the real exchange rate Johansen likelihood ratio tests (trace test) for cointegration (sample 1970:Q1 - 1996:Q3)	No cointegration	At most one vector	At most two vectors
Eigenvalue	*0.223*	*0.035*	*0.012*
Non-energy export, equation (IV)	31.64**	5.07	1.30
Cointegrating vector: EXPNONE = 1.72*WDD - 0.69*EXPGDP23			
Eigenvalue	*0.224*	*0.034*	*0.012*
Non-energy export, equation (IV)	31.65**	4.98	1.31
Cointegrating vector: EXPNONE = 1.71*WDD - 0.69*EXPGDP20			
Eigenvalue	*0.221*	*0.037*	*0.006*
Non-energy export, equation (IV)	30.84**	4.62	4.36
Cointegrating vector: EXPNONE = 1.64*WDD - 0.60*EXPGDP14			
Eigenvalue	*0.225*	*0.029*	*0.009*
Non-energy export, equation (V)	31.06**	4.08	0.95
Cointegrating vector: EXPNONE = 1.64*WDD - 0.62*EXULC23			
Eigenvalue	*0.227*	*0.027*	*0.009*
Non-energy export, equation (V)	31.07**	3.98	1.02
Cointegrating vector: EXPNONE = 1.64*WDD - 0.62*EXULC20			
Eigenvalue	*0.226*	*0.033*	*0.005*
Non-energy export, equation (V)	31.02**	4.13	0.58
Cointegrating vector: EXPNONE = 1.64*WDD - 0.62*EXULC14			
Eigenvalue	*0.229*	*0.024*	*0.008*
Exports of manufactures, equation (VI)	30.91**	3.48	0.82
Cointegrating vector: EXPMQ = 1.68*WDD - 0.65*EXUWC23			
Eigenvalue	*0.230*	*0.023*	*0.008*
Exports of manufactures, equation (VI)	30.93**	3.38	0.86
Cointegrating vector: EXPMQ = 1.68*WDD - 0.65*EXUWC20			
Eigenvalue	*0.227*	*0.030*	*0.006*
Exports of manufactures, equation (VI)	31.01**	3.92	0.65
Cointegrating vector: EXPMQ = 1.60*WDD - 0.55*EXUWC14			

The null hypotheses are that there is no cointegration, that there is at most one cointegration vector, and that there is at most two cointegration vectors; the 5% critical values of the likelihood ratio statistic (trace test) at the three largest eigenvalues reported for each equation in the Table are 24.31, 12.53 and 3.84; the 1% critical values are, correspondingly, 29.75, 16.31, and 6.51; *(**) rejects the null at the 5% (1%) level of significance; equation (IV), (V), and (VI) refer to those in Table 9 of the main text.
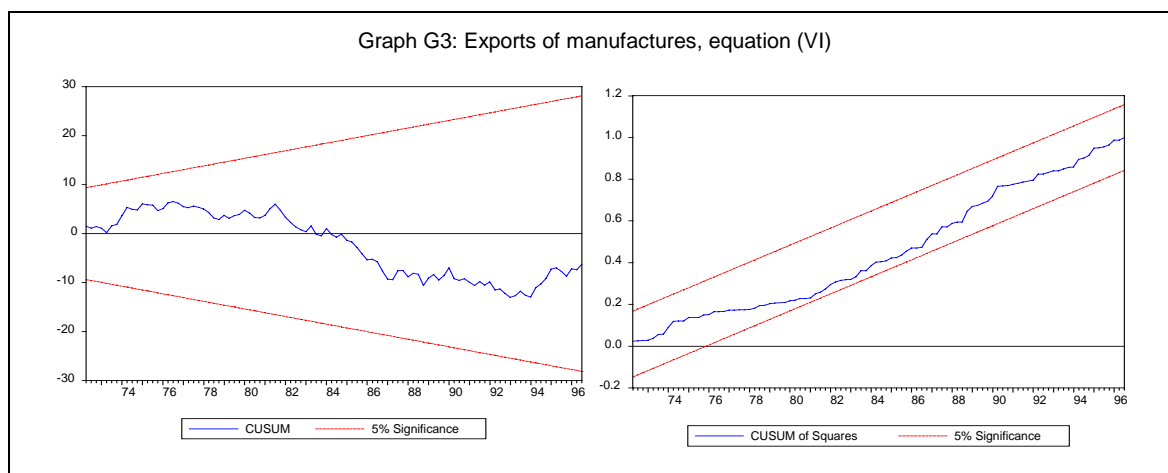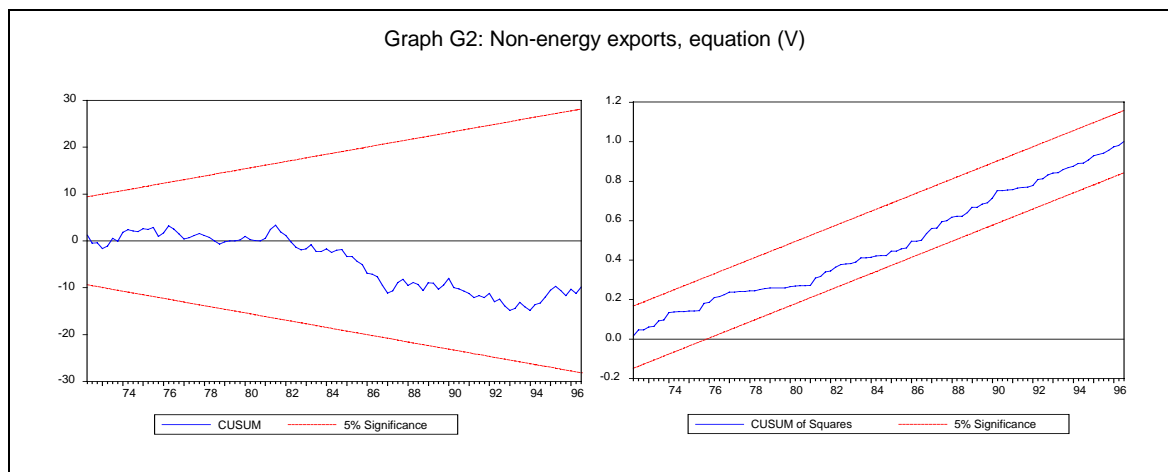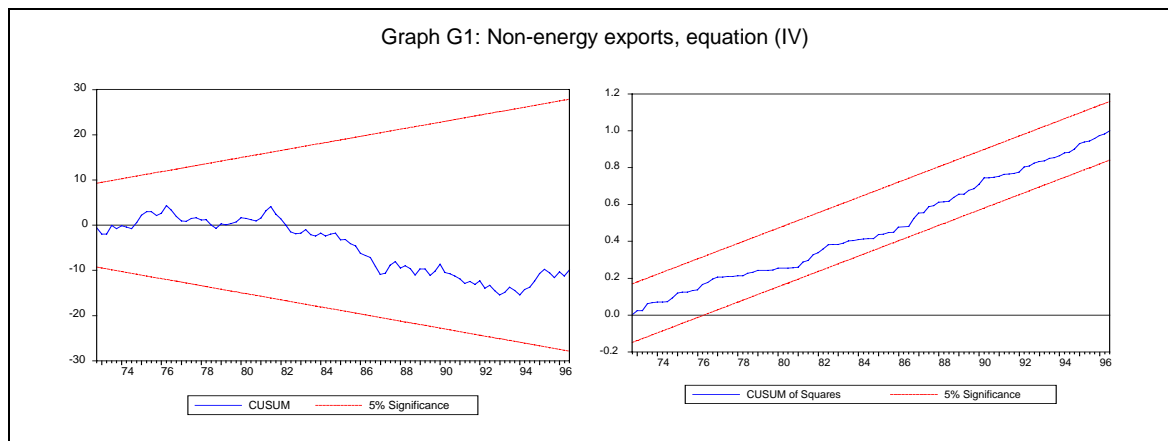
ANNEX F

Stability of Cointegration Vectors, Export Equations (IV)-(VI)
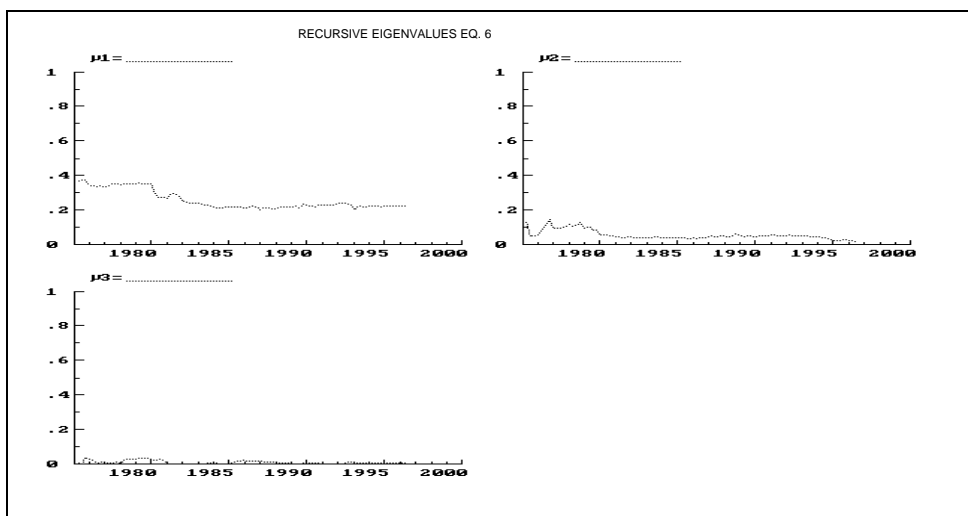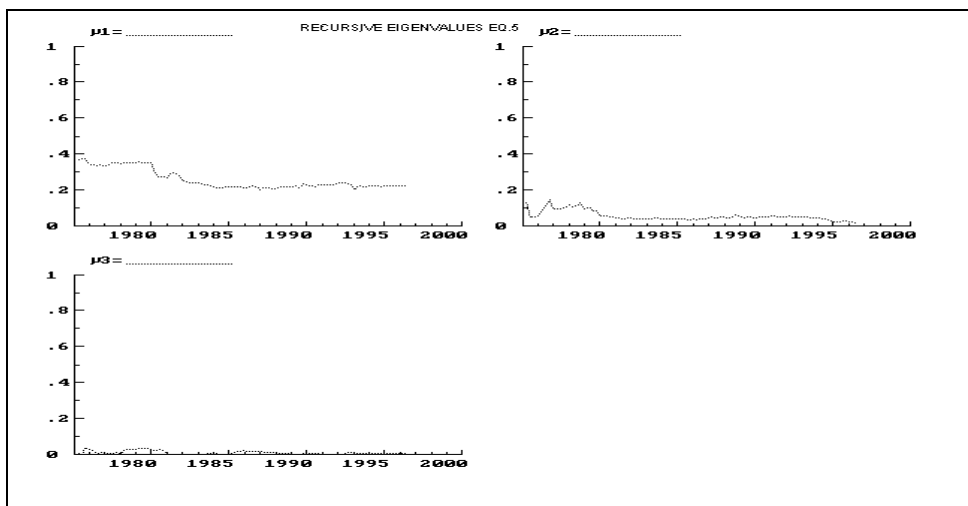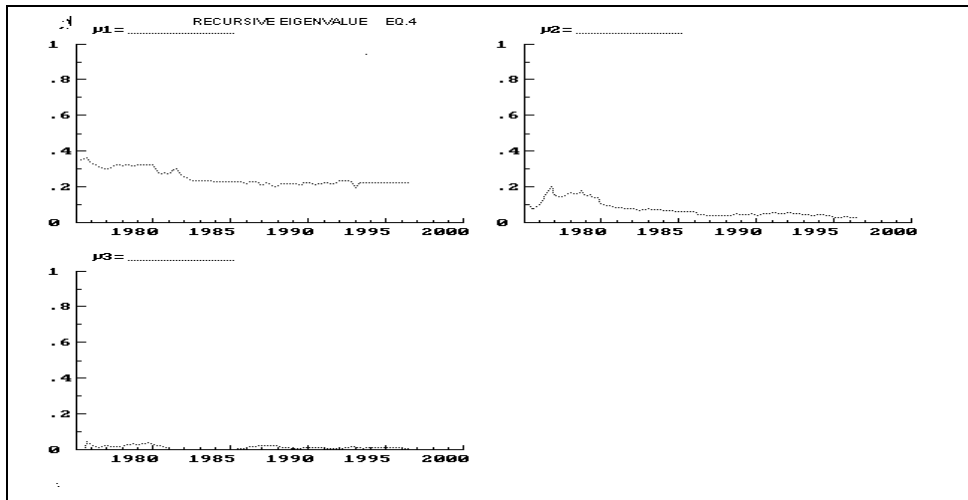
ANNEX G

Stability of Dynamic Export Equations (IV)-(VI)

Graphs G1-G3 present CUSUM and CUSUM of Squares tests and corresponding 5% significance levels for export equations (IV)-(VI).


Graph G1: Non-energy exports, equation (IV)

Graph G2: Non-energy exports, equation (V)

Graph G3: Exports of manufactures, equation (VI)

ANNEX H

Stability of Coefficients of Dynamic Export Equations (IV)-(VI)

Stability tests for the coefficients of the short-term adjustment model of equations (IV)-(VI) are presented, correspondingly, in Graphs H1-H3. The graphs show the recursive estimates for each coefficient for the period 1970:Q1-1996:Q3 and two standard errors. The estimates confirm that the export equations are temporally stable.
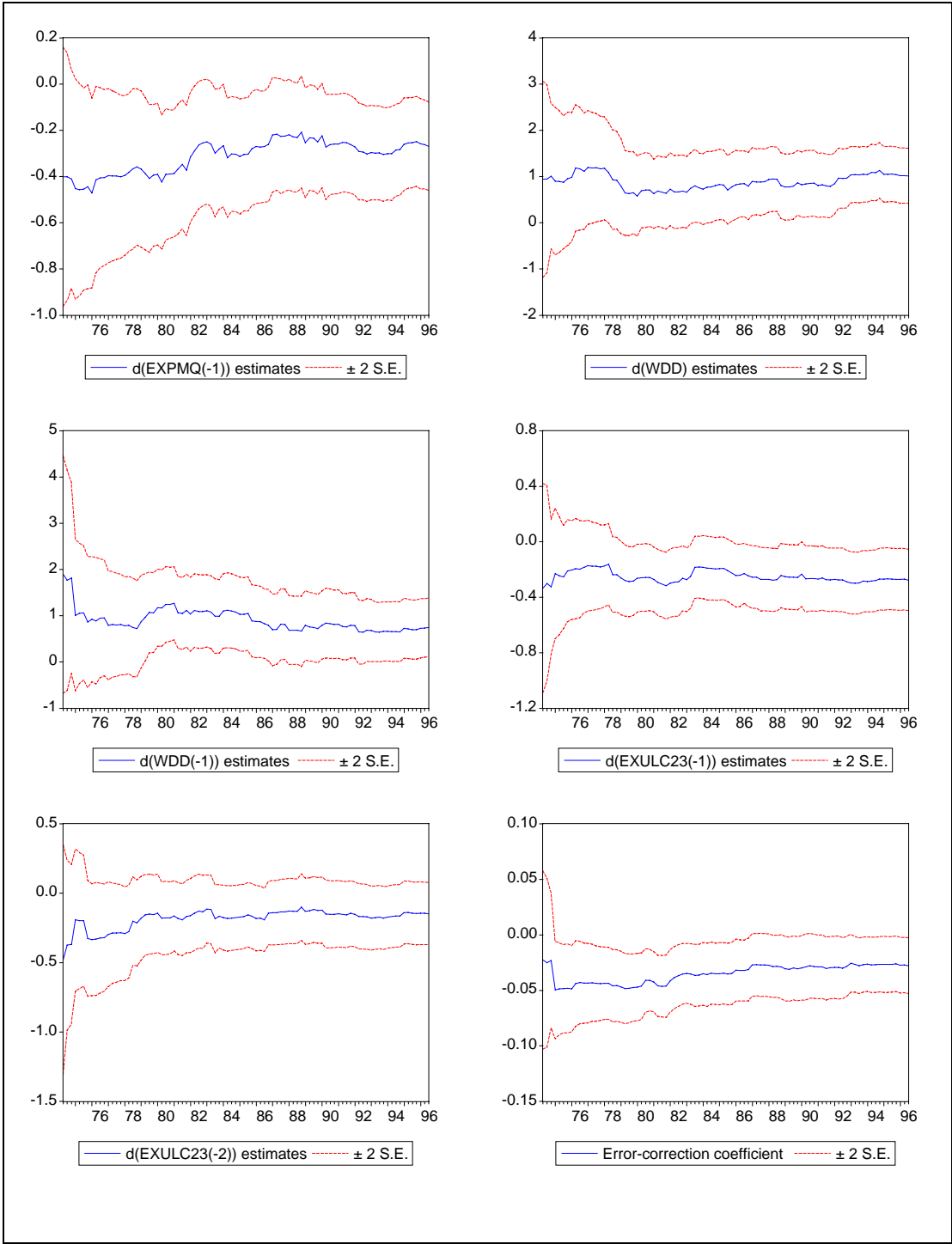
Graph H1
Coefficient stability of equation (IV)
(Recursive estimates and two standard errors)
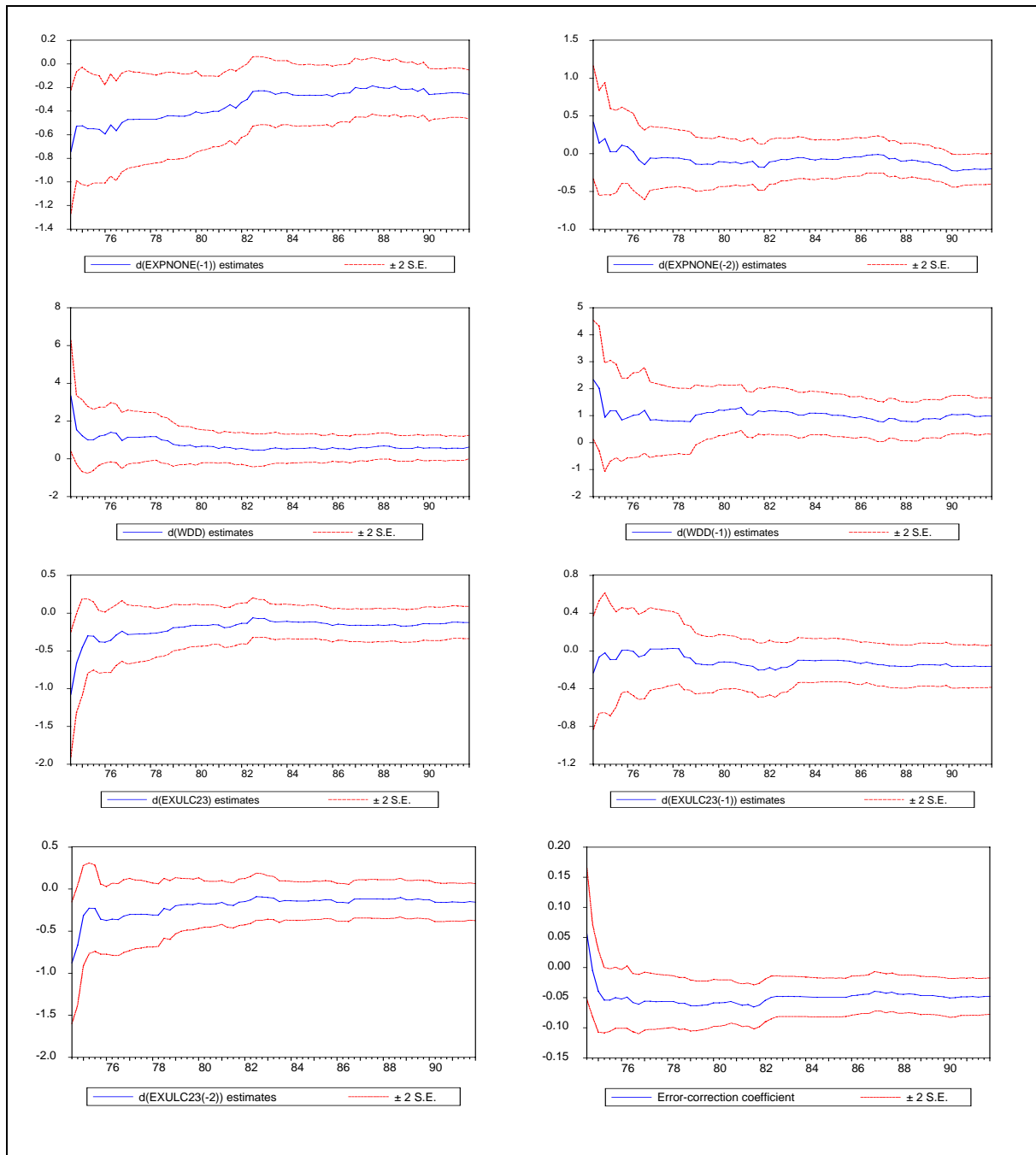
Graph H2
Coefficient stability of equation (V)
(Recursive estimates and two standard errors)
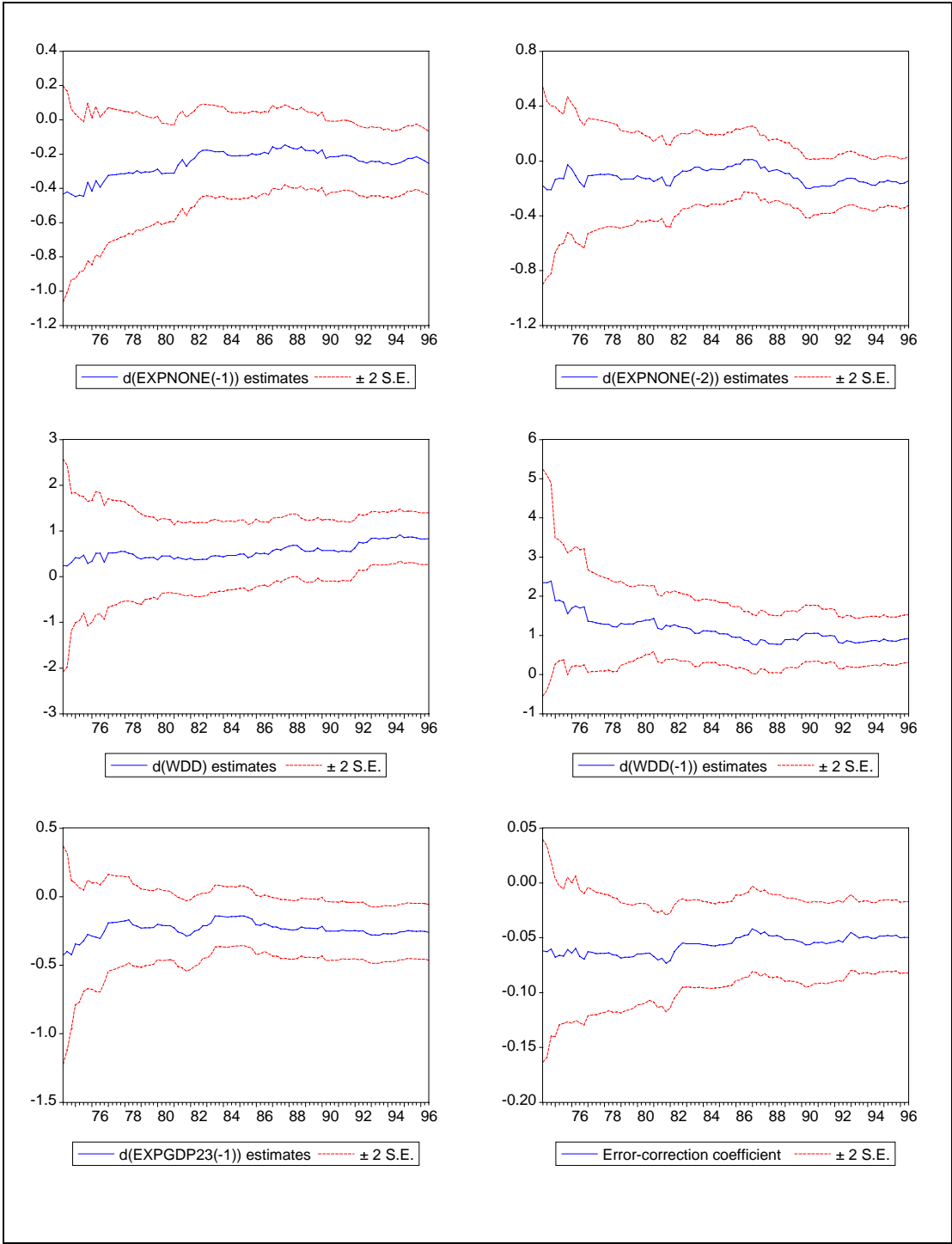
Graph H3
Coefficient stability of equation (VI)
(Recursive estimates and two standard errors)

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