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Western Hemisphere Department

Determinants of Argentina's External Trade

Prepared by Luis Catão and Elisabetta Falcetti[1]

Authorized for distribution by Thomas Reichmann

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Abstract

This paper presents new estimates of export and import equations for Argentina, using a broader set of variables than previous studies and distinguishing between intra- and extra-MERCOSUR trade. It measures the importance of relative price versus income effects in accounting for the higher trade deficit during the 1990s, and examines whether foreign trade elasticities have increased as a result of structural changes in the economy. It finds that the high income elasticity of imports and the responsiveness of exports to changes in world commodity prices, domestic absorption, and economic activity in Brazil have been key determinants of Argentina's trade balance.

JEL Classification Numbers: F11, F14, F31.

Keywords: Argentina, Foreign Trade Elasticities, International Competitiveness, MERCOSUR.

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I. Introduction

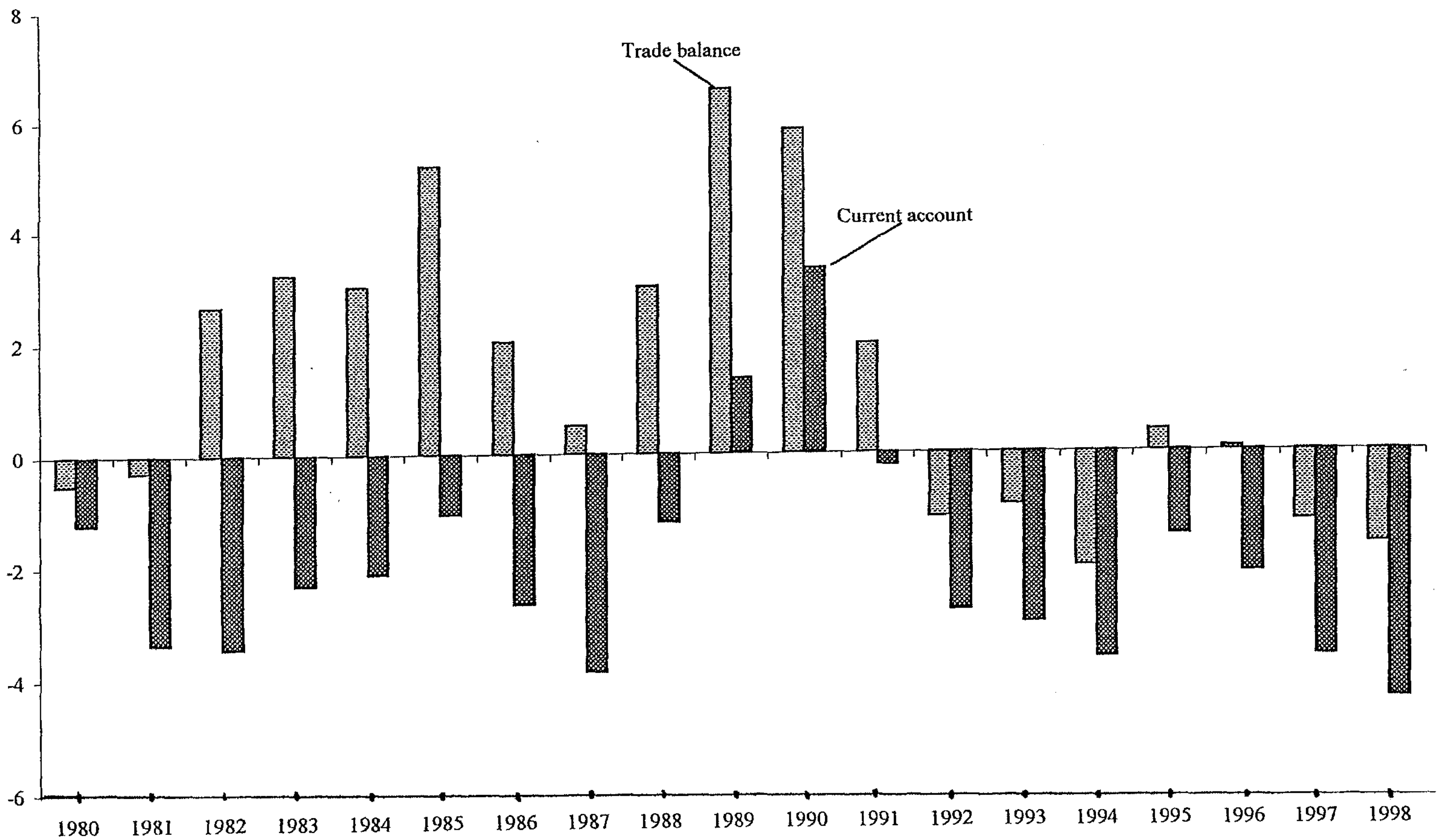
Following the financial and trade liberalization reforms of the late 1980s/early 1990s, a number of emerging market economies have experienced sizable current account deficits. In countries such as Argentina and Mexico, where rapid disinflation was accomplished through the combination of an exchange rate peg and trade liberalization, soaring current account deficits appeared to conform to the typical "stabilization cycle" (Kiguel and Leviatan, 1992; Végh, 1992): the marked real exchange rate appreciation and recovery of private consumption that followed macroeconomic stabilization spurred imports while inhibiting export growth, leading to sizeable trade deficits. At an initial stage of the reforms, expectations were created that such trade imbalances would be temporary: with inflation eradicated, consumption growth leveling-off, and domestic productivity enhanced through privatization and deregulation, trade deficits would tend to be gradually reversed, respecting the intertemporal budget constraint of a balanced current account in the long-run.

Thus far, these expectations are yet to be fulfilled. External deficits did narrow in all the emerging economies of the western hemisphere following the Mexican crisis of late 1994 but, as economic growth rebounded in 1996 and 1997, trade and current account imbalances reached new heights. In Argentina, the current account deficit exceeded its previous (1994) peak in the course of 1997 and rose to 5 percent of GDP in 1998, largely driven by a soaring trade deficit (Figure 1). With the dwindling of privatization revenues, the financing of such high deficits has become increasingly dependent on capital markets' assessments of the sustainability of the country's external position. A key question in this connection is whether and how, under current policies, the large trade imbalances of recent years will be eliminated in the medium-term .

The answer boils down to the likely response of the trade balance to changes in foreign and domestic demand, and in relative prices. If these income and price elasticities are relatively stable and can be accurately estimated on the basis of historical information, solid inferences can be made about the future evolution of the balance of trade. As with other areas of macroeconomics, however, the estimation of foreign trade elasticities has traditionally been plagued by econometric problems pertaining to dynamic specification, parameter stability, and the links between short-run adjustment and long-run equilibrium. The traditional approach to measuring trade elasticities has been to estimate least square regressions in levels assuming some sort of partial adjustment toward equilibrium.[2] As both experience and subsequent research have shown, this traditional approach imposes a very restrictive structure to the data, often producing biased estimates and misleading testing statistics. Advances in time series econometrics over the past decade have given rise to a more rigorous approach to dynamic specification of macroeconomic time series, enabling us to handle these problems more accurately. New developments associated with cointegration

[2]See Goldstein and Khan (1985) for a comprehensive survey of earlier studies within this tradition.

Figure 1. Argentina: Trade Balance and Current Account
(In percent of GDP)



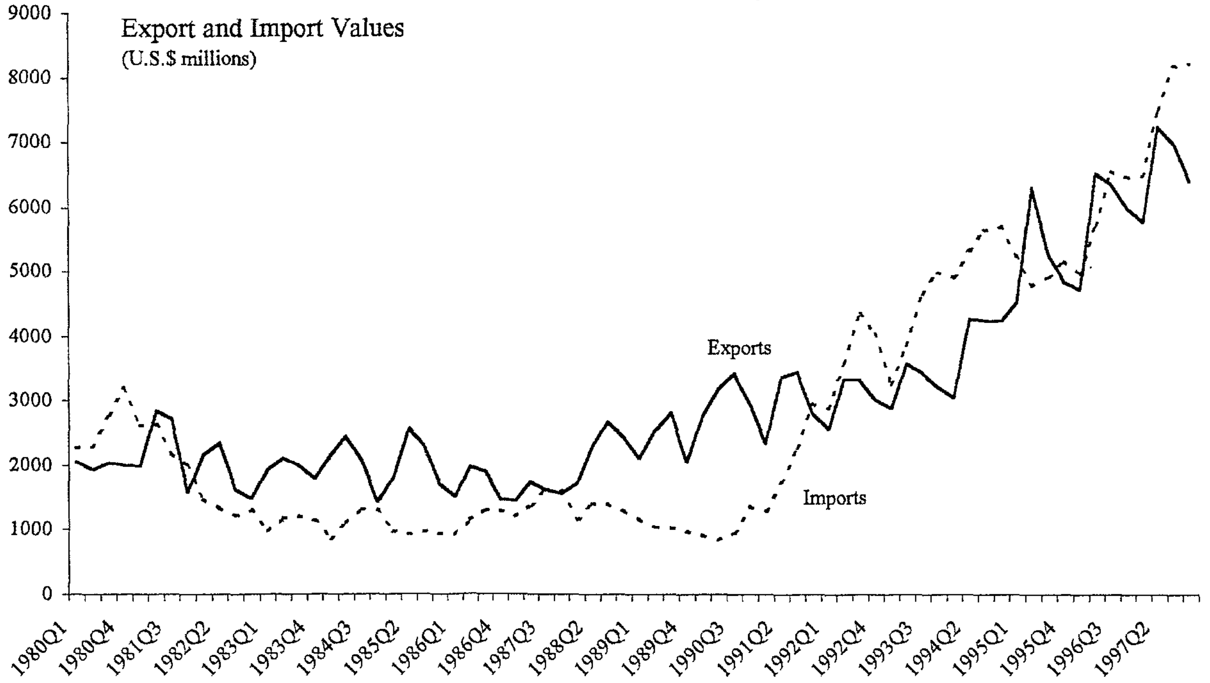
analysis, structural stability tests, and vector autoregressions (VARs) have provided a new foundation for modeling the complex interactions between cyclical and permanent components in time series and test for the existence of structural breaks in the underlying macroeconomic relationships. Some of these new econometric procedures have been fruitfully pursued in a number of recent studies on foreign trade (e.g. Clarida, 1994; Reinhart, 1995; Giorgianni and Milesi-Ferretti, 1997; Senhadji, 1998).

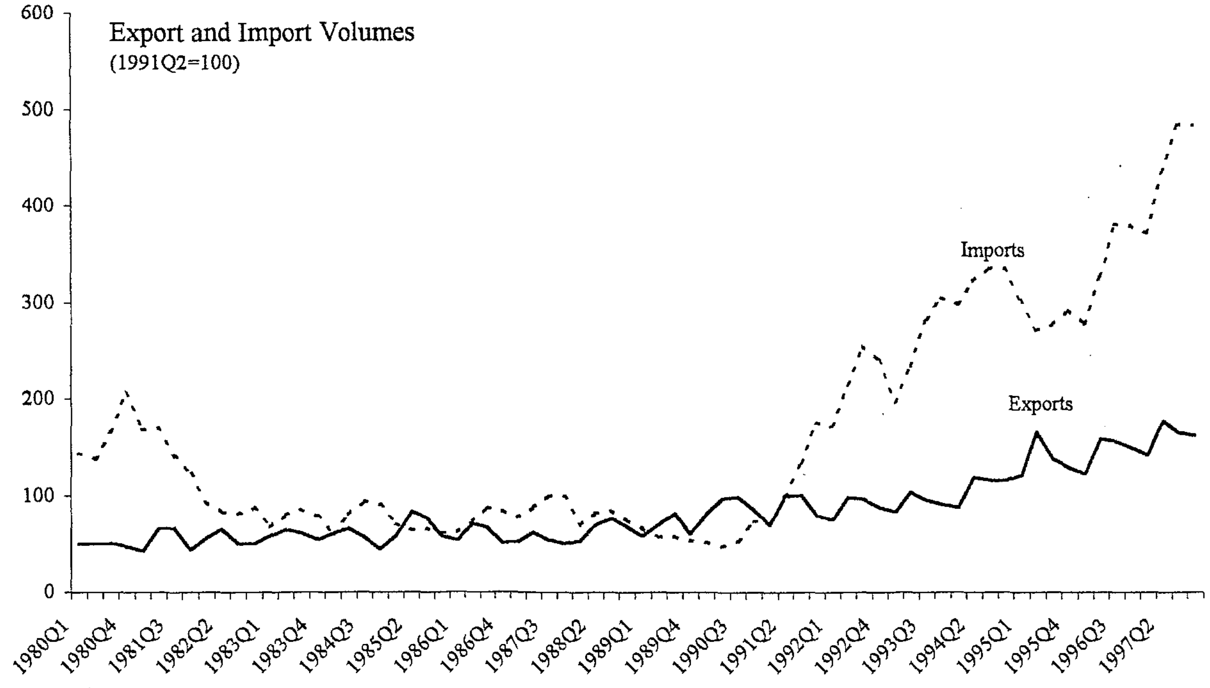
This paper uses a similar econometric framework to examine the determinants of Argentina's foreign trade. In contrast with previous studies,[3] we extend the analysis on three main fronts. First, in the specification of our export and import equations we consider a significantly broader set of explanatory variables, including the net domestic capital stock, real exchange rate volatility, unit labor costs, as well as a wide range of alternative indicators of relative prices. Second, given the growing importance of trade within the South American common market (MERCOSUR) in recent years and the fact that manufacturing exports to the region appear to be determined by a different set of factors, we separate out these exports from those to non-MERCOSUR countries in the estimation of our export equations. Third, we pay special attention to the issue of structural breaks over the sample period. This is important because, in many ways, the 1990s in Argentina stand in sharp contrast with the 1970s and 1980s. Following the liberalization of external transactions and the establishment of a currency board arrangement in early 1991, real GDP grew significantly faster than in the 1970s and 1980s, (averaging 5 ¼ percent a year between 1991 and 1998), inflation was weeded out, and both exports and imports trended up relative to the 1980s (Figure 2). However, as the growth of imports outpaced that of exports by a large margin, a sizeable external deficit emerged; and in contrast with the previous decade, when trade surpluses helped reduce current account deficits, trade deficits in the 1990s became a key component of such deficits. This *prima-facie* association between higher trade deficits, rapid economic growth, and a new foreign exchange regime underscores the need for a careful reappraisal of the relationship between these variables in recent years.

The remainder of this paper is divided into three sections. Section II lays out our econometric approach and applies it to a standard macromodel of supply and demand for exports. Short- and long-run elasticities of exports with respect to distinct supply and demand variables are estimated, and their stability over time is tested for. A similar approach is followed in the estimation of the import function in section III. Section IV summarizes the main findings and discusses some policy implications.

[3]There exist remarkably few systematic studies on the estimation of foreign trade equations for Argentina. Besides earlier work by Diaz-Alejandro (1970), recent studies by Ahumada (1994), Reinhart (1995), and Senhadji (1998) have estimated standard long-run demand functions for Argentina's exports using cointegration methods. The latter two authors, however, use panel data covering a large number of countries and devote little attention to the Argentine case.

Figure 2. Argentina: Foreign Trade





II. Supply and Demand for Argentina's Exports

A. General Considerations

Table 1 highlights three main "stylized facts" about the behavior of Argentina's export since the 1980s. First, volume growth not only accelerated markedly in the 1990s relative to the 1980s but also its standard deviation declined, i.e, export growth became less unstable. Second, compared with other countries in the Western Hemisphere and in Asia, the average growth rate of Argentina's exports in the 1990s cannot be considered outstanding but has been certainly respectable. Third, although growth instability has declined in recent years, Argentina's exports remain the most volatile among the largest economies in the Western Hemisphere countries, whereas among the selected Asian countries, only Indonesia has experienced higher export volatility than Argentina.

One apparent reason for this substantial export volatility is that, despite some diversification in recent years, Argentina's merchandise exports remain highly concentrated on a few raw materials and lightly processed primary products. In contrast with other emerging market economies which have become major exporters of manufacturing goods over the past two decades, Argentina's ten top export items[4] consist of crude oil, soya, wheat, vegetable oils, leather and meat—products which do not involve significant industrial processing and have been subject to large fluctuations in international prices.

Notwithstanding this dependence of overall exports earnings on few primary products, Argentina has experienced a rapid growth of its non-agricultural manufacturing exports to neighboring countries, *pari passu* with the lowering of tariffs and expansion of the South American customs union (MERCOSUR). Within this group, the growing share of manufacturing exports to Brazil stands out: while in 1990 manufacturing exports to Brazil accounted for 4 ½ percent of total Argentina's exports, in 1997 such a share rose to 16 percent. A conspicuous feature of Argentina's exports to MERCOSUR countries has been their prompt response to a set of government incentives and bilateral trade agreements which, *inter alia*, lowered tariff rates for key industries (notably, automobiles) and tied the export of these products to the partner country's imports of a similar good, with a view to keeping bilateral trade roughly in balance (Figueroa and Morales Rins, 1996; Kacef, 1998). An important implication of these arrangements is that, while the price of Argentina's exports of (raw or lightly manufactured) primary products are largely determined at the world market, the price and quantity of Argentina's industrial exports to MERCOSUR tends to be mainly determined by a different set of factors—namely, intra-bloc trade policies, geographical proximity, and income growth in the region.

[4]These account for nearly 40 percent of Argentina's exports. Other primary and agro-industrial products account for an additional 30 percent.

Table 1. Growth and Volatility of Export Volumes in Selected Countries

(Annual Percent Change)

	1980-90		1990-98	
	Mean	**Standard Dev.**	**Mean**	**Standard Dev.**
Western Hemisphere				
Argentina	6.8	14.0	9.3	9.0
Brazil	6.4	14.6	6.0	6.5
Canada	6.0	5.9	8.4	3.3
Chile	8.1	6.8	9.3	4.5
Mexico	8.7	8.6	15.4	8.5
USA	5.4	8.2	8.4	4.4
Asia				
Australia	6.0	4.7	7.8	5.3
Hong Kong, SAR	15.0	10.5	10.0	7.7
Indonesia	2.0	17.2	12.7	9.9
Korea	12.2	6.8	15.7	6.9
Malaysia	9.9	10.0	11.7	6.4
Thailand	13.9	10.6	10.7	6.7

Source: INDEC, Ministry of Economy of Argentina; and IMF.

In short, Argentina's exports comprise two different groups of products from the point-of-view of their economic determinants: on the one hand, exports of primary and lightly manufactured goods, for which Argentina is basically a price taker in international markets and subject to large fluctuations in the terms of trade of primary commodities on the other hand, a still relatively small but thriving group of manufacturing exports to MERCOSUR which are mostly influenced by trade policies, regional proximity, and regional macroeconomic developments. In this context, where Argentina's policies and macroeconomic performance play a key role and its exporters hold a substantial share of the foreign market, they no longer face an infinitely elastic demand schedule for their products; hence export prices become determined by the intersection of demand and supply variables. It is easy to see that failure to take this distinction into account may impart significant biases to estimates of Argentina's foreign trade elasticities.

Given these differences between the intra-MERCOSUR manufacturing trade and the remainder of Argentina's foreign trade, we use two distinct specifications for the export functions. In the first specification, we purge manufacturing exports to Brazil[5] from the series on Argentina's total exports and estimate a traditional export supply function for a small open economy, where commodity export prices are determined at the world market and thus exogenously given.[6] In the second specification, we introduce a standard aggregate demand function and solve the respective two-equation system for the equilibrium price and quantity of exports. We use this supply and demand system to examine the determinants of Argentina's manufacturing exports to Brazil.

B. Export Supply Excluding MERCOSUR Manufacturing Trade

Following a long-established tradition in the empirical literature,[7] we model supply of exportables as a positive function of the relative price of exports, as well as of a measure of domestic productive capacity such as the net capital stock, and a negative function of domestic absorption. Also in line with a burgeoning literature which emphasizes the potentially adverse effects of exchange rate uncertainty on trade,[8] we have included a measure of real effective exchange rate volatility as an additional explanatory variable, so that the export supply function can be written as:

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[5]Note, however, that primary commodity exports to Brazil are left as part of this first group.

[6]Although Argentina is a relatively large world producer of wheat and beef, for instance, in none of these markets is Argentina a price setter.

[7]See Goldstein and Khan (1985) for a survey. Recent empirical work within this tradition include Arize (1990), Reinhart (1995), and Giorgianni and Milesi-Ferretti (1997).

[8]See Côté (1994) for a survey. For a recent paper on the impact of real exchange rate volatility on exports in Brazil, see Gonzaga and Terra (1997).

$$x_t^s = \alpha_0 + \alpha_1 p_{x_t} + \alpha_2 k_t - \alpha_3 c_t - \alpha_4 \sigma_{RER_t} + \varepsilon_t \qquad (1)$$

where p_x stands for the relative price of exports, σ_{RER} for real exchange rate volatility, c for real domestic absorption, k for the aggregate net capital stock, and ε is the residual. All variables are in logs except for the exchange rate volatility measure, which is discussed in detail below.

While there is a wide consensus on the basic functional form of aggregate export functions, a number of measurement issues remain controversial. These include: the choice of the relevant price indicator (e.g. whether to use the real exchange rate, relative unit labor costs, or simply the US dollar unit value of exports); whether to use domestic consumption or simply GDP (gross or net of exports) as the scale variable for absorption effects; and which of the several possible measures of exchange rate volatility is the most appropriate.

With regard to the relative price variable, given the well-known pros and cons of each of the above mentioned indicators,[9] most studies have followed an empiricist approach and simply pick the indicator which yields the best fit. We adopt the same strategy here. As for the measurement of absorption effects, in countries which export substantial amounts of both consumer and producer goods, variables such as real GDP or absorption are obvious choices, as domestic demand for **both** consumer and investment good will tend to divert the production of these goods away from exports. However, since the vast bulk of Argentina's exports consist of either primary commodities or lightly manufactured goods that are widely consumed domestically, aggregate consumption seems a more suitable scale variable.

The literature on the effects of price uncertainty on external trade has used a number of different indicators for exchange rate volatility as a proxy for exporters' risk. The most commonly found measure is the unconditional standard deviation of the percentage change of the real exchange rate (e.g. Caballero and Corbo, 1989; Gonzaga and Terra, 1997; Dell'Ariccia, 1999). This measure implies that exchange rate uncertainty is zero when the exchange rate is either fixed or follows a deterministic trend, consistent with the assumption that policies based on a fully credible and constant rate of devaluation (or revaluation) would be fully anticipated by agents and would thus have no impact on export volume. The other property of this measure is the larger weight given to extreme observations. This is particularly suited to situations where the real exchange rate usually displays considerable instability, and where domestic firms are risk averse. The latter property seem appropriate to the Argentine context of the 1980s, whereas the former property squares well with developments since 1991. Moreover, given that other commonly used measures such as the difference between forward and spot exchange rates are unavailable, we shall use the unconditional standard deviation of quarterly percentage changes in

[9]For a discussion of the pros and cons of distinct external competitiveness indicators, see Lipschitz and McDonald (1991).

REER (averaged over a one-year window) as a proxy for the effects of exchange rate uncertainty on exports.[10]

On the basis of the choice of indicators just discussed, we proceed with the estimation of the above export model into two steps. The first step is to test for the existence of cointegration among the variables in equation (1), i.e., whether there appears to exist at least one equilibrium vector tying them up in the long-run. If such a stable long-run relationship exists, the residual term (ε) will be stationary or integrated of order zero [I(0)], even if some or all the variables involved are said to be "non-stationary" or first-order integrated [I(1)].[11] In this case, it is possible to obtain consistent and efficient estimates of the long-term elasticities of exports to the relevant price and quantity variables.

Once we test for non-stationarity and find that some—if not all—the variables in (1) are I(1), we proceed with testing them for cointegration. Here we employ two commonly used tests. One is the procedure due to Johansen (1988), which is based on a vector autoregressive (VAR) system comprising all the I(1) variables suspected of being cointegrated. While the variables enter the VAR in first-difference of the their log-levels, a matrix consisting of these variables in (log) levels is also added to the system; cointegration tests amount to testing whether such a matrix is not singular—i.e., whether its rank is different from zero. This is accomplished by comparing the largest eigenvalue of that matrix with tabulated critical values. If the former exceeds the latter, the test indicates the existence of at least one[12] set of α cointegrating coefficients which renders the residual of a level equation such as (1) stationary.

In addition, we also employ an alternative testing procedure advanced by Pesaran and Shin (1998) and Pesaran et al. (1996), which allows for a mix of I(1) and I(0) variables in the same cointegrating equation. In contrast with the Johansen multivariate VAR method, the Pesaran and Shin (1998) test consists of adding, in a single equation in first differences, lags of

[10]A question arises as to the appropriate choice of the frequency of observations (daily, monthly or quarterly) and temporal window period (one quarter, a year or several years). For instance, under certain circumstances it can be argued that quarterly export performance is significantly affected by weekly or daily changes in the exchange rate (Gonzaga and Terra, 1997). In our case, however, since we are mainly concerned with medium-term fluctuations in exports and base the remainder of the analysis on quarterly observations, the use of quarterly changes in the RER over a one-year window appeared as a fair compromise.

[11] I(1) variables can only be rendered stationary when first differenced in logs (or equivalently expressed in terms of percentage change). This is because the univariate autoregressive representation of such variable contains a unit root, which can only be eliminated by the first difference operator.

[12]The procedure allows for the existence of up to n such vectors, where n is the number of I(1) variables in the system

first differences of the variables so as to orthogonalize the relationship between the first differences of the explanatory variables and the residual term ε. Testing for cointegration then amounts to a F-test on the joint statistical significance of adding *level* regressors of the variables suspected to be cointegrated.[13]

Once the existence of at least one cointegrating vector in (1) is established, consistent estimates of the respective long-run elasticities can be obtained within the same testing framework. As discussed in Pesaran and Shin (1998) and Pesaran et al. (1996), such estimates can be obtained from an autoregressive distributed lag (ARDL) regression of the level of the dependent variable on the level of all other I(1) and I(0) variables. Once orthogonalization between the residual term and the right-hand side variables is achieved (by including a sufficient number of autoregressive terms),[14] and residuals appear to be serially uncorrelated, asymptotic Gaussian inference can be readily applied.

Having estimated a set of long-run elasticities, the second main step of our analysis consists of modeling the underlying short-run dynamics leading to the long-run, level equilibrium represented in (1). As shown by Engle and Granger (1987), if a cointegrating relationship among a set of variables exists, then there must also exist an "error correction" equation which relates the growth rate (or first difference of the logs) of these variables to their equilibrium relationships in levels. With the long-run cointegrating vector included as an additional explanatory variable in the first-difference representation of (1), OLS estimation of the latter allows us to recover the respective short-run elasticities which map how changes in the right-hand side variables impact on export growth. In the context of an error correction representation, one can then conduct a number of standard tests for alternative specification, structural change, and predictive power.

Tests for the non-stationarity of the variables entering equation (1) are provided in Table 2. They indicate that the hypothesis that the log level of exports, net capital stock, and aggregate consumption are all non-stationary cannot be rejected. In contrast, most relative price variables considered appear to be stationary. On the basis of this information, the

[13]Under the null hypothesis of no-cointegration, the distribution of such an F-statistic is non-standard; so the usual critical values employed in classical statistical inference do not apply. The relevant critical bounds have been tabulated by Pesaran et al. (1996) and are provided in Table 3.

[14]The optimal number of lags of the first-differenced variables can be determined by standard maximum log-likelihood based tests, such as the Akaike information criterion or the Schwarz Bayesian criterion.

Table 2. Unit Root Tests

(Quarterly Data, 1980-97)

Variable	DF[1]	ADF[2]
Export volume	-2.65	-1.56
Export price	-2.83	-4.13*
Net Capital Stock	1.89	-0.41
Consumption	-0.86	-2.81
Real Exchange Rate	-2.23	-3.35
Px/ULC	-2.84	-3.47*
Px/CPI	-1.29	-3.13
Import Volume	0.11	-2.91
Real GDP	-1.42	-2.53
Pm*(1+t)/CPI	-1.36	-3.49*
Real Interbank Rate[3]	-8.19	-3.63*
Real Deposit Rate[3]	-8.15	-3.47*

Source: IMF, INDEC, and Argentina's Ministry of Economy.

[1] Dickey-Fuller Statistic based on a log-level regression including an intercept but not a trend.

[2] Augmented Dickey-Fuller Test based on a log-level regression including a trend, with the order of the autoregressive term chosen by the Akaike information criterion.

[3] Deflated by one-period ahead inflation rate.

*Significant at 5 percent.

results of cointegration test are provided in Table 3. Both the Johansen and Pesaran-Shin tests clearly support the hypothesis of one existing cointegrating vector tying the level of exports to the right-hand side variables in (1). Given the evidence supporting cointegration, ARDL estimates of the vector of long-run coefficients α are reported in Table 4.

As discussed above, we experimented with a few relative price indicators which appear to share a common trend with that of exports , including the ratio of export price to domestic unit labor costs, the ratio of export price to domestic CPI, and the unit value of exports expressed in US dollars (Figure 3). Also, to allow for a possible structural break in the 1990s, we experimented with an intercept dummy as well as with a deterministic time trend kinked in 1991:q1. As shown in Table 4, only the equations using the US dollar unit value of exports as relative price indicator yielded an estimate of α_1 with the "right" sign; all other relative price indicators yielded α_1 coefficients with signs opposite that postulated by theory,[15] with very low (asymptotic) t-ratios, and produced estimates of long-term elasticities with respect to consumption and net capital stock which seem unrealistically high (equations A.3 and A.4 in Table 4). So, equations (A.1) and (A.2) should be clearly preferred.

Both equations point to a long-term price elasticity of exports around unity and estimated with considerable precision, as witnessed by the high t-ratios. Also close to unit are the estimated coefficients on domestic consumption and on real exchange rate volatility, both taking the expected negative signs. The estimated coefficient of about two for α_2, which was also estimated with considerable precision, indicates that Argentina's export volume has been particularly responsive to the rapid growth of the net aggregate capital stock in the 1990s (Figure 4). The estimated coefficient on the intercept dummy variable for the 1990s is small but yielded a high t-ratio, which bodes well with the hypothesis of a regime shift in the 1990s.[16]

[15]In principle, this could be due to a simultaneity bias stemming from an inverse causality running from exports to the exchange rate. In other words, higher exports could lead to higher consumer prices or labor costs expressed in US dollar terms and hence to a negative association between exports and the relative price variable px/pc. Using instrumental variables, Ahumada (1994) concludes that such a simultaneity bias does not appear to be significant, and thus cannot account for the lack of statistical significance of current levels of the real exchange rate in the export equation.

[16] The estimated coefficient on the kinked time trend for the 1990s yielded a rather low t-ratio while taking on the "wrong" sign, and was thus dropped from the reported regressions.

Table 3. Cointegration Tests

A. Export Supply Equation (Excluding Intra-MERCOSUR Manufacturing Trade)

Johansen's Maximum Likelihood Rank Tests[1]

Null	Alternative	λ-Max Statistic	95% Critical	Trace Statistic	95% Critical
r = 0	r = 1	36.65*	25.42	58.44*	42.34
r<=1	r=2	18.22	19.22	21.78	25.77
r<=2	r=3	3.56	12.39	3.57	12.39

Pesaran-Shin ADRL-based Test[2]

	F-Stastic	95% Critical
Model without a time trend:	21.28*	4.86
Model including a time trend:	20.92	5.87

B. Supply and Demand System for Manufacturing Exports to MERCOSUR

Johansen's Maximum Likelihood Rank Tests[2]

Null	Alternative	λ-Max Statistic	95% Critical	Trace Statistic	95% Critical
r = 0	r = 1	52.62*	40.53	134.53*	102.56
r<=1	r=2	35.21*	34.4	81.9*	75.98*
r<=2	r=3	20.01	28.27	46.69	53.48

C. Import Demand Equation

Johansen's Maximum Likelihood Rank Tests[1]

Null	Alternative	λ-Max Statistic	95% Critical	Trace Statistic	95% Critical
r = 0	r = 1	28.50*	19.22	34.28*	25.77
r<=1	r=2	5.78	12.39	5.78	12.39

Pesaran-Shin ADRL-based Test[3]

	F-Statistic	95% Critical
Model without a time trend:	6.28*	4.05
Model including a time trend:	7.30*	4.70

1/ Seasonal dummies and a restricted time trend included in the underlying fourth-order VAR.
2/ Seasonal dummies and a restricted intercept included in the underlying second-order VAR.
3/ Seasonal dummies included in the underlying ARDL regression.

* Significant at 5 percent.

Table 4. Export Supply Equation: Estimates of Long- and Short-run Coefficients[1]
(t-ratios in brackets; quarterly data for 1980-97)

A. ARDL Long-run Coefficients
Dependent Variable: log of exports

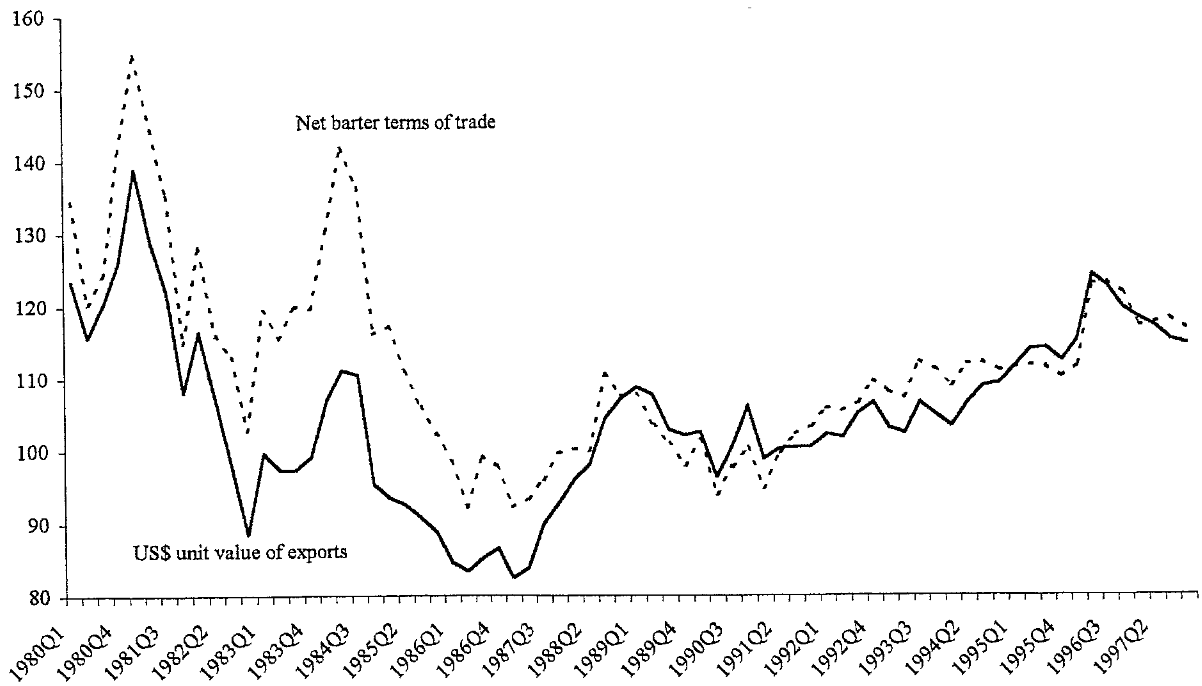
	Constant	**Log (p_x*)**	**Log (nk)**	**Log (c)**	**σ (Δrer)**	**Trend 90s**	**Log (px/cpi)**	**Log (px/ulc)**	**R^2**	**D.W.**
(A.1)	-11	1.18	2.28	-1.12	-1.02				0.95	2.06
	(-5.69)	(3.84)	(7.51)	(-2.32)	(-2.53)					
(A.2)	-6.4	1.1	2.07	-1.36	-0.78	0.17			0.96	1.77
	(-3.14)	(6.14)	(10.68)	(-4.40)	(-3.37)	(2.66)				
(A.3)	11.47		5.64	-6.81	-1.46		-0.36		0.96	1.88
	(0.80)		(2.05)	-1.53	(-1.19)		(-1.30)			
(A.4)	9.33		7.54	-8.83	-2.09			-0.34	0.96	1.92
	(0.58)		(1.63)	(-1.27)	(-1.11)			(-0.93)		

B. Error Correction Representation
Dependent Variable: first difference of the log of exports

	Constant	Δ Log $(p_x)_t$	Δ Log $(p_x)_{t-1}$	Log $(nk)_t$	Log $(c)_t$	σ $(\Delta rer)_t$	Δ Trend 90s	EC_{t-1}	R^2	D.W.
(B.1)	-4.53	0.68	-0.67	0.94	-1.28	-0.42	...	-0.41	0.83	2.06
	(-3.38)	(2.86)	(-3.05)	(3.83)	(-4.06)	(2.44)		(-3.70)		
(B.2)	-4.37	0.76	-0.63	1.41	-0.92	-0.53	0.11	-0.68	0.85	1.77
	(-2.80)	(3.30)	(-2.97)	(7.06)	(-4.56)	(3.45)	(2.52)	(-6.97)		

1/ Excluding manufacturing exports to Brazil during 1990-97. Seasonal dummies added to all regressions.

Figure 3. Argentina: Foreign Trade Prices and the Real Exchange Rate
(1991=100)



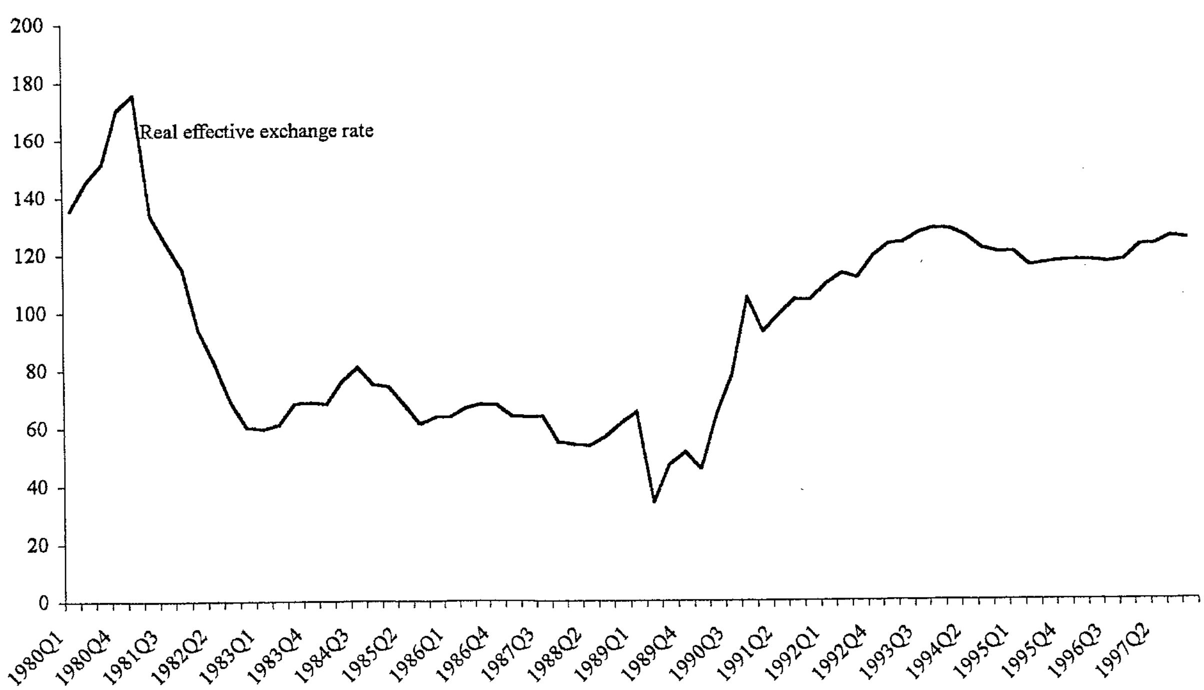
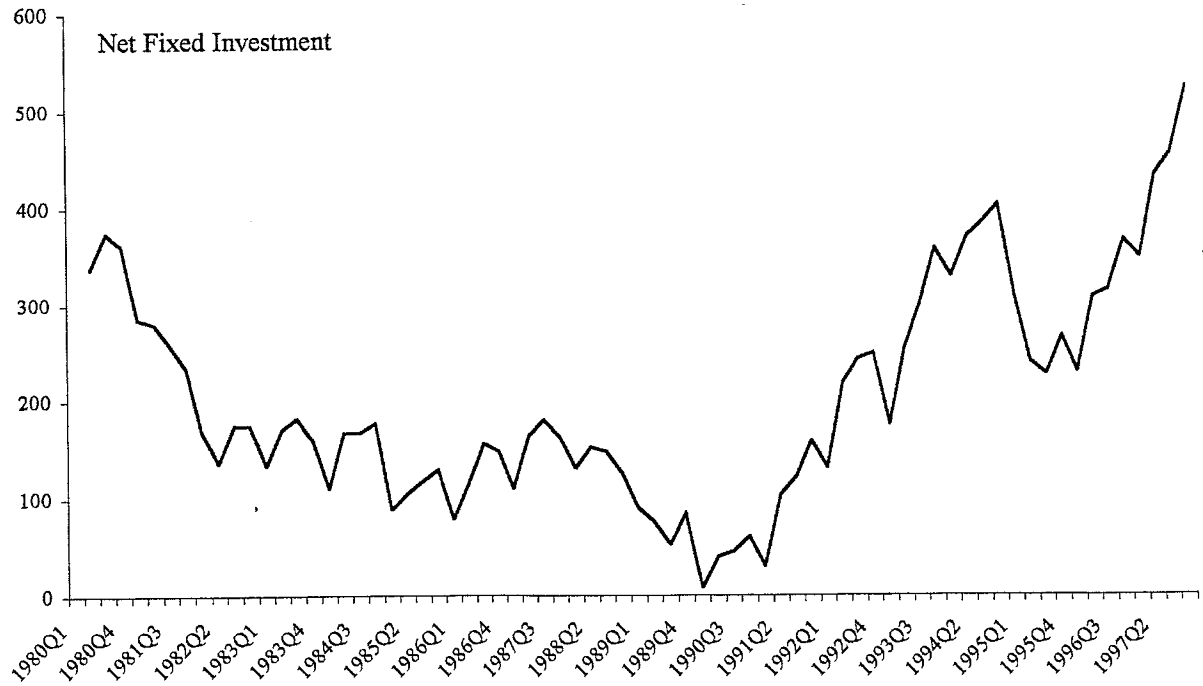
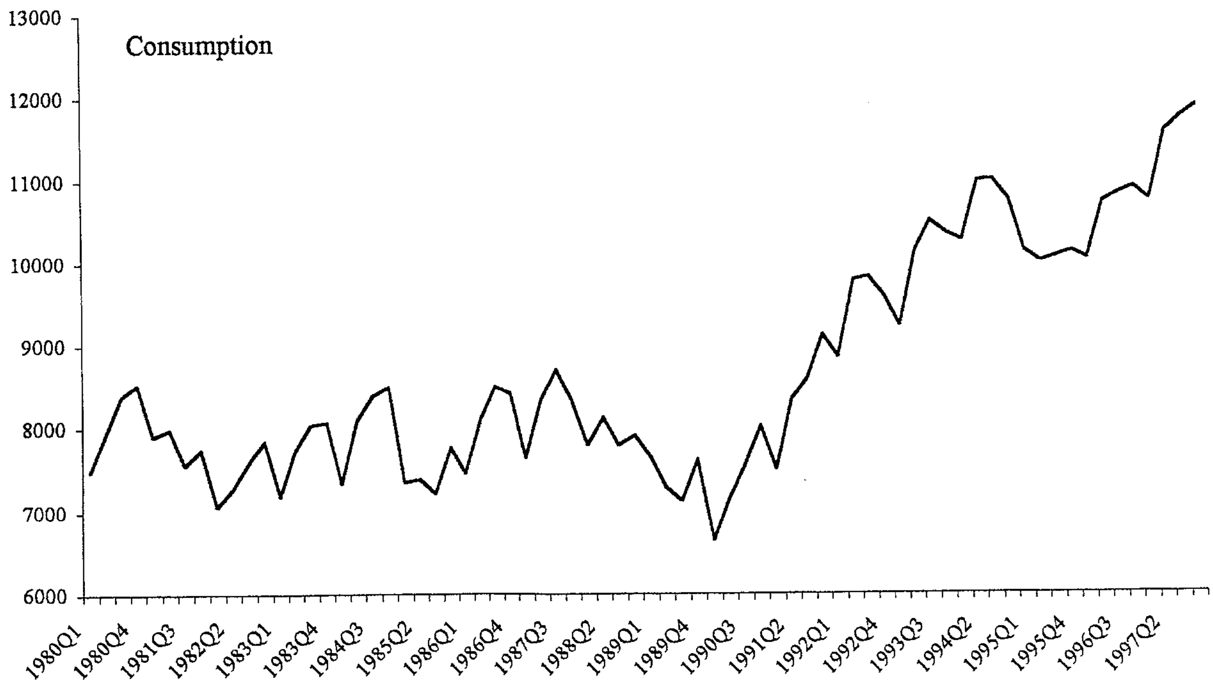


Figure 4. Argentina: Aggregate Investment and Consumption
(1986 constant prices)





Estimates of the error correction representation of equation (1) are shown in the bottom part of Table 4. The estimated coefficients again took on the signs predicted by theory and both the capital stock and domestic absorption appear to be main determinants of export performance even in the short-run. Current changes in export prices (Δp_x) also have a significant positive impact, but this is largely offset by the impact in the previous quarter changes—such an offset being commonly observed in cases where inventory adjustment plays a crucial role in the short-term response of export volume to price signals. Also worth noting in this connection is the relatively high coefficient on the error correction term (EC_{t-1}), which implies that 41 to 68 percent (depending on which specification one chooses) of deviations from long-term equilibrium are corrected for within the quarter.

The R-square statistic shows that equations (B.1) and (B.2) yield a good fit, explaining 83 to 85 percent of the percentage change in exports. Both regressions passed all the diagnostic statistics for residual autocorrelation, functional form, and heteroscedasticity. Moreover, in light of the marked structural change in the economy in the 1990s, structural stability tests were carried out. One such structural stability tests is based on the cumulative sum of the recursive residuals, the so-called CUSUM test.[17] As shown in Figure 5, the cumulative plot of recursive residuals falls well within the 5 percent significance band, thus indicating that the estimates appear to be robust to underlying structural changes. Finally, we performed a number variable addition tests to try and assess whether the model is robust to alternative specifications. Among potentially significant variables, we considered current and lagged values of relative unit labor costs, REER, world real GDP, with a view to capture the impact of greater openness and structural reforms on exports if not fully captured before. None of the F-statistic on this set of variables, however, proved to be statistically significant at 5 percent, thus underscoring the robustness of the original model[18]

C. Supply and Demand of Manufacturing Exports to MERCOSUR

As explained above, Argentina cannot be considered a small country in the MERCOSUR. So, in addition to the supply-side variables already discussed above, one needs to introduce a demand equation for the joint determination of price and volume of Argentina's manufacturing exports to MERCOSUR. As in most of the literature, our demand function is log-linear in foreign income and the relative price of exports, so that export price and quantity are jointly determined in the long-run by the following model:

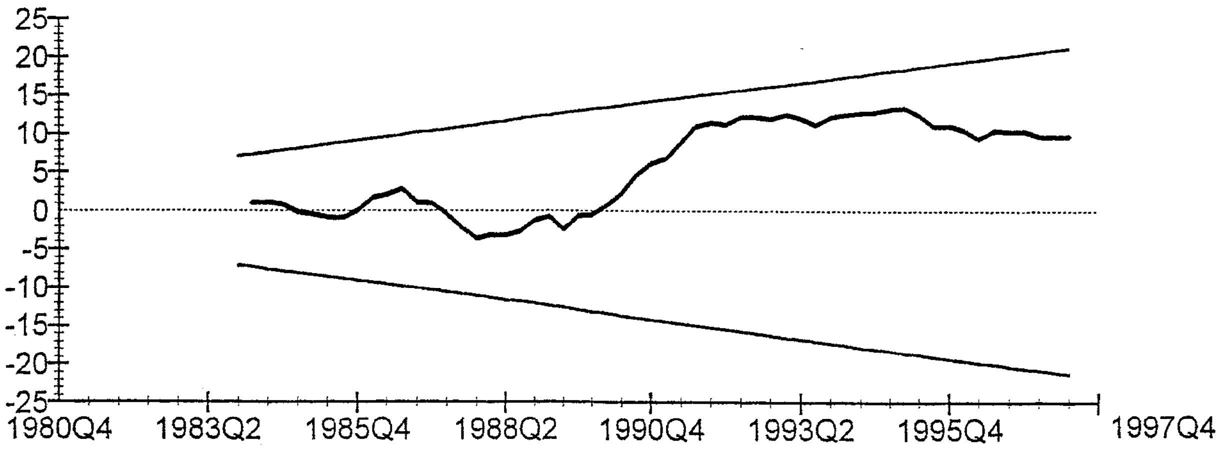
---

[17]Because regression (A.2) contains an intercept dummy defined as zero between 1980 and 1990, this test can only be computed for the post-1991:q1 period. Thus, we only report the results of the CUSUM test on (A.1).

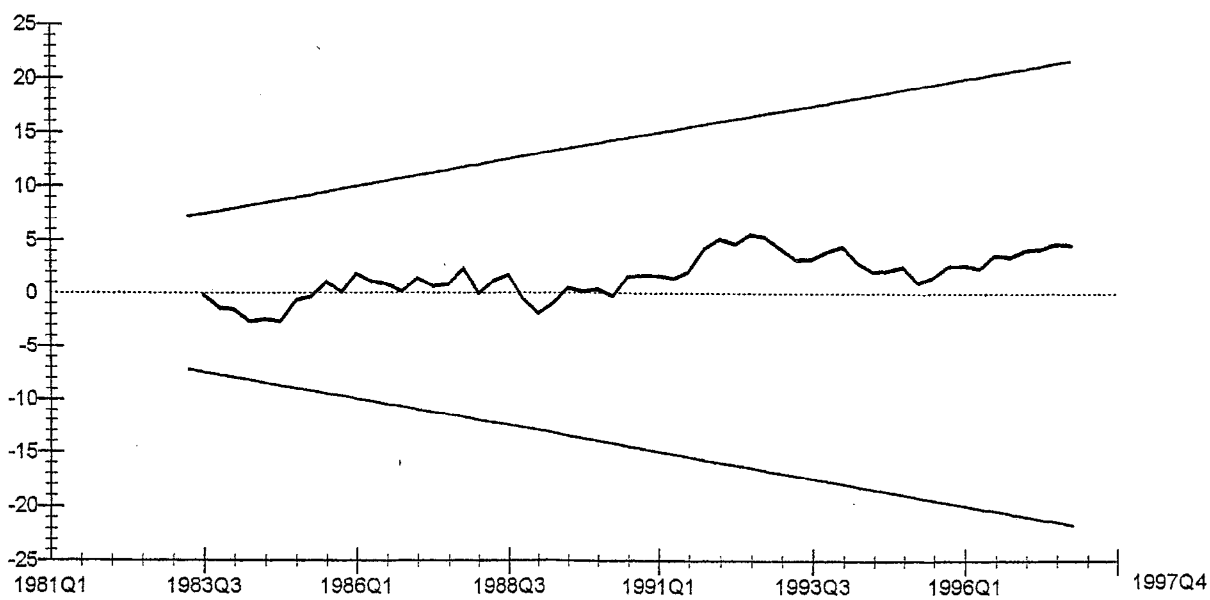
[18]Not reported but available from the authors upon request.

Figure 5. Plot of Cumulative Sum of Recursive Residuals

A. Export Supply Equation



B. Import Demand Equation



The straight lines represent critical bounds at 5% significance level

$$x_t^d = \gamma_0 - \gamma_1 (1+t^*)\, p_{x_t}/p^*_t + \gamma_3 y_t^* + u_t \qquad (2)$$
$$x_t^s = \rho_0 + \rho_1 p_{x_t} - \rho_2 ULC_t + \rho_3 k_t - \rho_4 \sigma_{RER_t} + v_t$$

where y* is real GDP in the MERCOSUR partner countries and p_x is the price of Argentina's exports net of the foreign tariff rate (t*) relative to the foreign price index, p*.[19] All the variables entering the supply equation are the same as in (1), with two only differences: one is the exclusion of domestic consumption as a potentially significant variable;[20] the other is that σ_{RER} now corresponds to the bilateral exchange rate between Argentina and the rest of MERCOSUR and UCL stands for unit labor costs in domestic manufacturing.[21]

Strong simultaneity across equations, and the fact that most variables entering (2) appear to be non-stationary, call for an econometric specification that accommodates both features of the data generation process. This can be done by re-writing (2) as a vector error correction model (VECM)

$$\Delta X_t = \mu + \phi_t + \Pi X_{t-1} + \Gamma \Delta X_{t-1} + \Sigma T_s \Delta X_{t-s} + w_t \qquad (3)$$

where X is a vector comprising all the I(1) variables entering (2), μ is a vector of constant terms, ϕ is a vector of exogenous I(0) variables, and w is a vector of serially independent but possibly contemporaneously related error terms.[22] Maximum likelihood estimation of (3) will yield a matrix of long-run coefficients Π bearing the cointegrating relations between these variables, if any; estimates of Γ will yield the respective short-run or "impact" elasticities. If equation (2) is a valid representation of the data generation process, one should expect to find at least two cointegrating relationships—one for the supply and the other for the demand equation.

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[19]Since Brazil has accounted for 85 to 90 percent of Argentina's manufacturing exports to MERCOSUR, Brazil's real GDP and consumer price index will be used as proxies for y* and p*, respectively. Quarterly figures for t* were obtained by interpolation from annual data on Brazil's average external tariff as provided in Garriga and Sanguinetti (1995).

[20]In contrast with the exports of traditional staples such as wheat and beef, domestic consumption tends to have a less significant bearing on manufacturing exports to MERCOSUR due to a variety of tax and tariff exemptions which lower the relative cost of exporting to neighboring countries relative to selling in the domestic market. Indeed, the inclusion of aggregate consumption yielded statistically insignificant coefficients in all the equations on Argentina's manufacturing exports to Brazil.

[21]As already noted, this will be proxied by Argentina's real exchange rate with Brazil.

[22]For a thorough discussion on the specification and statistical properties of VECMs, see Hamilton (1996).

Results of cointegration tests for this supply and demand system are reported in section B of Table 3. They indicate the existence of exactly two cointegrating vectors among these variables, consistent with the theoretical model. To map from the unrestricted estimates underlying this VAR to the supply and demand model postulated above, some identifying restrictions need to be imposed. As is clear from (2), these consist of excluding y* and p* from the supply equation and ULC, k, and c from the demand equation; moreover, the model states that the coefficient on p_x is equal to the negative of the coefficient on $p^*/(1+t^*)$ in the demand equation. Since only two restrictions are needed to exact identification, once all these restrictions are imposed, our system becomes overidentified. This allows us to test the extra restrictions and thus assess the robustness of the postulated theoretical model.

The estimated vector of long-run coefficients with all these restrictions imposed is reported in the bottom part of Table 5. The respective likelihood ratio (LR) test does not reject the additional restrictions at 5 or 10 percent levels. The signs are as predicted by theory, while the magnitude of the respective parameters point to an elasticity of export demand with respective to foreign income and relative price around 2 and 1.3, respectively. On the supply side, the elasticity of Argentina's exports to export price (US dollar denominated) is of similar magnitude but that on unit labor costs is much lower (less than 0.2) and imprecisely estimated, as indicated by its low t-ratio. In contrast, the estimated elasticity of export supply with respect to the capital stock is rather high.[23]

Somewhat surprisingly, Lagrange multiplier tests (not reported) indicated that the real exchange rate instability variable did not add any significant explanatory power to the VAR and therefore was dropped from the regressions.[24] This contrasts with the results of section II.B which indicate that real exchange rate instability tended to undermine Argentina's

[23]It is possible that this reflects the inadequacy of our proxy to productive capacity in manufacturing which here is taken to be the aggregate capital stock.

[24]It lacked statistical significant for both the level and first-difference equations. Indeed, inclusion of this variable in the error correction supply equation yielded a positive (though statistically insignificant at any conventional level) coefficient, contrary to the theory.

Table 5. Manufacturing Exports to Brazil: VECM Estimates of Equation System (2)

(t-ratios in brackets, quarterly data for 1980-97)

A. Long-run Coefficients [1]

	p_x*	ULC	k	Y*	p*
Supply Equation:	1.29 (1.78)	-0.17 (-0.38)	5.07 (7.27)	--	--
Demand Equation:	-1.24 (-5.39)	--	--	2.03 (1.56)	1.24 (5.39)

LR test of overidentifying restrictions: χ^2 *(2) : 4.28 [pval = 0.23]*

B. Error Correction Representation [1]

	Δ ULC_t	Δ p_x*_{t-4}	ECM^s_{t-1}			R^2	D.W.
Supply Equation:	-1.00 (-2.73)	0.51 (1.76)	-0.33 (-3.77)			0.51	2.11

	Δ x_{t-1}	Δ $(T*px/pc)_{t-1}$	Δ $Y*_{t-1}$	ECM^d_{t-1}		R^2	D.W.
Demand Equation:	-0.22 (-1.97)	-0.93 (-3.88)	4.49 (3.24)	-0.28 (-3.19)		0.58	2.13

	Δ ULC_t	Δ $(T*px/pc)_{t-1}$	Δ $Y*_{t-1}$	ECM^s_{t-1}	ECM^d_{t-1}	R^2	D.W.
Reduced Form Equation:	-0.99 (-2.98)	1.45 (1.94)	3.17 (2.70)	-0.3 (-3.76)	-0.19 (-2.40)	0.63	2.29

[1]Obtained on the basis of a second-order VAR (selected by the Schwarz bayesian criterion), with x and p_x* as endogenous variables. Seasonal dummies and restricted intercept included. All variables in logs.

commodity exports, particularly in the 1980s when Argentina's operated a flexible exchange rate regime and experienced extreme relative price volatility.[25]

The associated error correction equations show that all the above variables also play an important role in determining exports in the short-run, with the notable exception of capital stock. The estimated coefficients on the latter were statistically insignificant at 5 or 10 percent levels for the alternative specifications we tried, perhaps reflecting the fact that capital buildup effects take long to impact on manufacturing exports. In contrast, unit labor costs and economic activity in Brazil proved to be a much more important determinant of Argentina's manufacturing exports in the short-run, yielding high and statistically significant coefficients in the first-difference equations (Table 5, section B). In addition to standard tests for residual autocorrelation, heteroscedasticity, and functional form, we have submitted all error correction equations to CUSUM tests for parameter stability.[26] The plot of the respective cumulative sum of residuals fell well within the 5 percent significance boundaries, thus suggesting that the estimated parameters of the model appear to be relatively stable over time.

III. Import Demand

A. General Considerations

As pointed out in the introduction, the marked acceleration in import growth has been a main factor behind Argentina's current account deficits in the 1990s. While greater integration within the world economy has led imports to grow somewhat faster than real GDP in most countries, between 1991 and 1997 Argentina's import grew four and a half times as fast as real GDP. Barring on the effect of variables other than income in fostering import demand, this simple correlation suggest that the income elasticity of imports in Argentina has been significantly higher than those estimated for non-industrial countries (Houthakker and Magee, 1969; Reinhart, 1995; Giorgianni and Milesi-Ferretti, 1997; Sehadji, 1998).

The much faster growth of imports relative to real GDP in the 1990s fails to square with previous estimates of the income elasticity of imports in Argentina. Using annual data for the period 1970-91, Reinhart (1995) estimates a long-run income elasticity of imports in Argentina slightly above unity; a similar value was obtained by Sehadji (1998) for the period 1960-93. A

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[25]The statistical insignificance of the exchange rate volatility term in the equation on manufacturing exports to MERCOSUR, albeit surprising, echoes the findings of other empirical studies. Gonzaga and Terra (1997) find that the effect of real exchange volatility on Brazilian exports—a large share of which consist of manufacturing goods—is also statistically significant. Evidence for the European Union covering mostly manufacturing trade indicates that the effects of exchange rate volatility on trade are statistically significant but small (Dell'Ariccia, 1999).

[26]Not reported here but available upon request from the authors.

number of factors may explain such a mismatch. One is sampling. Since both authors worked with annual observations spanning over nearly thirty years, their estimates are likely to capture more closely the steady-state features of the relationship between imports and real GDP, which entail an income elasticity of imports close to unity.[27] Another possible explanation is the type of explanatory variables included in their model. Neither Reinhart (1995) nor Sehadji (1998) incorporate any tariff effect in their price variables, and since average tariff rates have varied widely during the period, the relative price indicator used in their regressions is bound to be a misleading indicator of relative import prices. Last but not least, there were important structural changes in the Argentine economy in the 1990s which may have raised the income elasticity of imports, thus biasing estimates based on data from previous decades.

Be that as it may, it is clear that a re-assessment of these previous estimates is needed. In this section we provide new estimates of income and price elasticities of imports on the basis of quarterly data that include the 1990-97 years and using a more comprehensive set of explanatory variables. In addition, we look at the stability of such estimates over time so as to examine two competing hypotheses, namely: whether there was indeed a "permanent" structural break in the underlying relationship between GDP growth and imports as a result of the regime change of the 1990s; or, alternatively, whether the much faster growth of imports relative to GDP in the 1990s reflects mainly temporary cyclical developments, rather than a permanent upward shift in income and price elasticities.

B. Estimation Issues

As in most of the relevant empirical literature, our starting point is a log-linear import function in real output, the relative price of imports (inclusive of import tariffs) and the real interest rate, so as to capture both inter- and intra-temporal substitution effects. Since these variables appear important to explain imports from MERCOSUR as well as from non-MERCOSUR countries, and there is no obvious reason to split the two groups of imports as we did for exports. The aggregate import function can thus be written as

$$mq_t = \beta_0 + \beta_1 GDP_t - \beta_2 (1+t)\, pm_t / CPI_t - \beta_3 R_t + \xi_t \qquad (4)$$

---

[27]Here it is important to note the difference between the theoretical concept of steady-state where an income elasticity of imports significantly above unity is ruled out by assumption (as it would entail explosive behavior of the share of imports in GDP), and the working definition of "long-run" underlying this paper. In the latter, long-run is simply defined as a time span covering nearly two decades (the 1980s and the 1990s). In the case of Argentina, this definition is not only more relevant for the purpose of current policy analysis, but also avoids the pitfalls of estimating steady-state relations on the basis of a data sample spanning over several decades which are subject to uneven data quality and the existence of major structural breaks.

where mq and pm stand for total import volume and US dollar unit value of imports, respectively, t for the import tariff rate,[28] CPI for the consumer price index, R is the real interest rate, and ξ is the error term. All variables are in logs, except for the real interest rate which is defined as explained below.

As in the case of the export equation, questions arise about the best empirical proxies for the relative price variables in (4). While the relative price of imports to CPI (inclusive of tariffs) is a relatively uncontroversial indicator for intra-temporal substitution effects, different measures of the real interest rate have been used in the literature. Particularly in the case of Argentina, there is no obvious proxy for agents' opportunity cost of consuming one unit of importables today v. consuming it tomorrow. While studies on other countries have used broad measures of short-term interest rates such as the 3-month T-bill rate (e.g. Ceglowsky, 1991), a similar series for Argentina is not available for the whole 1980-97 period. The only (unregulated) interest rate instruments for which data is readily available on a quarterly or monthly frequency are the interbank call rate and the time deposit interest rate. In the estimation of equation (4), we shall try both indicators and check the sensitivity of the results to the different choices.

Also controversial is the measurement of expected inflation, needed to convert nominal to real rates. Only previous or current period inflation are observable by agents, but in a context of high and volatile inflation of the period 1980-90 expected inflation is likely to have had a significant forward-looking component. Thus, backward-looking inflation measures would tend to underestimate expected inflation by a substantial margin in periods when inflation was rising (such as in the late 1980s), and overestimate expected inflation when the latter was rapidly declining (as during 1991-93). In light of these short comings the use of actual **future** inflation as a measure of inflationary expectations seem to be more appropriate. So, real interest rate measures based on a one-period ahead actual inflation should be preferred to those of constructed on the base of past or current inflation. Thus, in the estimation of (2) we give greater prominence to the use "forward-looking" real interest rate measures, though also check the sensitivity of the results to the use of "backward-looking" measures.[29]

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[28] Here we measure average import tariff rate as the ratio of total tariff revenues by total imports. Although this can be an inacurate proxy of the "true" protection costs (specially when certain import items are subject to quantitative restrictions as was the case in Argentina until the late 1980s), it has the advantage of being derived from observed data and appears to be the only measure of protection costs for which a consistent series is available on a quarterly basis over the entire 1980-98 period. For a discussion of different measures of tariff protection and evidence on the correlation between actual tariff revenues and official (or ex-ante) tariff rates, see Pritchett and Sethi (1994).

[29] In the high inflation environment of the 1980s nominal interest rate were quoted on a monthly basis; so, the respective real rate was obtained by deflating the nominal figure by the one-month ahead actual inflation and the annualized. With the advent of macroeconomic stabilization in the 1990s, domestic lending institutions resumed quoting interest rates on an annual basis which

(continued...)

C. Results

Dickey-Fuller tests reported in Table 2 cannot reject the unit root hypothesis for both the log levels of import volume and real GDP, while rejecting the non-stationarity of the relative price of imports and the real interest rate. Thus, the variables entering equation (4) can be said to represent a long-term equilibrium relationship only if imports and real GDP are cointegrated with each other.

Second, we test for cointegration using the testing procedures advanced by Johansen (1988) and Pesaran and Shin (1998), already discussed in section 2. As shown in Table 3, the results of both tests indicate the existence of one cointegrating vector tying together all the variables in (4), consistent with the existence of a long-run import function as postulated above. Given cointegration, we then move to the estimation of the respective long- and short-run elasticities using the Pesaran et al. ARDL methodology, as we did for exports.

Table 6 reports the long-run coefficients of the estimation of (2) by the ARDL method, with the lag structure of the autoregressive selected by the Schwarz Bayesian Criterion. The coefficients on the real interest rates are very small irrespective of what real interest rate indicator one chooses, though they are estimated with some precision (t-statistic above 2). While this suggests that intertemporal substitution is of relatively minor importance in determining imports—as one would expect over a fairly long time window—the coefficients on income and relative import price are sizeable. Depending on the specification (using a forward or backward-looking real interest rate measure, or with or without a trend among the regressors), the coefficient on real GDP yields an elasticity in the 2.0-2.5 range; the relative price elasticity gravitates in the neighborhood of 0.7. Very high t-statistics are attached to both coefficients, indicating that they are estimated with considerable precision.

were then deflated by the 12-month ahead inflation. Quarterly real interest rates were computed as arithmetic averages of the monthly rates.

Table 6. Import Demand Equation: Estimates of Long- and Short-run Coefficients
(t-ratios in brackets; quarterly data for 1980-1997)

A. ARDL Long-run Coefficients
Dependent Variable: Log of Imports

	Constant	Log (GDP)	Log (pm(1+t)/cpi)	RIR inter1/	Trend 90s 1/	RIR dep 2/	RIR dep 3/	R^2	D.W.
(A.1)	-15.2	2.42	-0.8	-0.14 E(-6)	...	...	...	0.99	2.34
	(12.0)	(5.95)	(-7.56)	(-2.28)					
(A.2)	-13.3	2.17	-0.70	-0.15 E(-6)	0.18	...	...	0.99	2.37
	(-3.03)	(4.80)	(-5.90)	(-2.35)	(1.40)				
(A.3)	-15.2	2.42	-0.79	...	...	-0.16 E(-4)	...	0.99	2.32
	(-2.22)	(5.97)	(-7.55)			(-3.67)			
(A.4)	-18	2.72	-0.79	...	...	...	-0.85E(-2)	0.99	1.89
	(-9.34)	(-14.34)	(-15.79)				(-2.33)		

B. Error Correction Representation
Dependent Variable: First Difference of the Log of Imports

	Constant	Δ log (GDP) $_t$	Δ log (GDP) $_{t-1}$	Δlog (pm(1+t)/pc)	Δ RIR $_t$	Δ RIR $_{t-1}$	EC $_{t-1}$	R^2	D.W.
(B.1)	-4.78	1.92	0.75	-0.25	0.24 E(-7)	0.37 E(-7)	-0.31	0.83	2.34
	(2.77)	(13.37)	(4.58)	(-6.99)	(2.18)	(3.20)	(-6.69)		

Other Diagnostic Tests

Functional Form: $\chi^2 = 0.88$ [p=0.35]
Heteroscedasticity: $\chi^2 = 2.20$ [p = 0.14]
Variable Addition: I) Level Dummy for the 1990s ==> F (1, 62) = 0.06 [p=0.80]
II) 1990s Slope Dummy on Real GDP ==> F(1, 62) = 2.48 [p=0.12]
III) 1990s Slope Dummy on Relative Import Price ==> F(1, 62) = 1.25 [p=0.27]

1/ Time trend defined as 1 for 1991:q1-1997q4 and zero otherwise.
2/ Time deposit interest rate deflated by the 12-month ahead inflation rate and expressed as % p.a.
3/ Time deposit interest rate deflated by the 12-month current inflation rate and expressed as % p.a.

At the bottom part of Table 5 we report the "short-run" or error correction models associated with the long-run specification (A.1). All explanatory variables entering the model were statistically significant in these error correction specifications and yielded the signs predicted by theory, with the exception of changes in interest rate. Yet, the estimated coefficients on the latter were again rather small and so practically insignificant. The estimates point to "impact" income elasticity of imports around 2 and a cumulative short-run elasticity (over a two quarter period) of 2 ¾, while the relative price elasticity is ¼. This clearly suggests that income effects are particularly strong in the short-run, while relative price or substitution effects take time to unravel (J-curve effects) until their full effect, estimated in the level equation above, is felt. The coefficient on the error correction term of 0.3 indicates a moderate speed of adjustment toward equilibrium, similar to that observed for other countries (see, e.g., Giorgianni and Milesi-Ferreti, 1997). Diagnostic statistics for the estimated equation were very good overall, with an R-square of 0.83 and no evidence of residual autocorrelation.

We have undertaken a number of tests to gauge whether and, if so, the extent to which a structural change in the determinants of import demand took place during the 1990s. First, we carry out a CUSUM test for parameter stability in the ARDL model estimated above. As explained in section 2, according to this test the hypothesis of parameter stability should be rejected whenever the cumulative sum of the recursive residuals exceeds a critical boundary, usually set at a 5 percent significance level. The sum of residuals, together with the respective 5 percent critical bounds, are depicted in the bottom panel of Figure 5. As the cumulative plot of recursive residual falls well within the 5 percent bounds, this test does not reject the null hypothesis of parameter stability.

A second set of test comprises those for variable addition. If there was no structural break in the relationship between imports and its income and relative price determinants, then the inclusion of variables such as intercept or slope dummies (defined as zero during 1980-90 and 1 henceforth), or a time trend kinked in 1991-97, should not add any significant explanatory power to the model. The results of such F-tests for variable addition are reported in Table 5 None of the intercept or slope dummy variables added to the regression significantly improved its fit. Likewise, F-tests for the inclusion of a time trend for the 1990s do not support the hypothesis of a structural change.

Finally, we conducted a (stronger) test for structural change based on the level estimation of (4) by recursive OLS. The latter allows the estimated coefficients to change over time and so should be able to reveal whether they have witnessed a marked increase or decrease in the 1990s. The resulting elasticity is graphed in Figure 6.[30] As expected, the

[30] Lags of the variables in levels were included in the regressors so as to reproduce the ARDL representation underlying the long-run estimates of Table 5. The autoregressive structure, selected by the Schwartz Bayesian criterion, added one lag of the (log) level of the dependent variable, two lags of the (log) level of real GDP and two lags of the real interest rate.

Figure 6. Argentina: Recursive OLS Estimates of the Income Elasticity of Imports

income elasticity of imports appears to have a clear cyclical component, as witnessed by the upturn of 1991-92 which was then followed by a gradual decline through early 1995, and then by a slightly upturn since. However, the main point is that, by the end of the sample period, the estimated elasticity was roughly back to its mid-1980s level, indicating that in the income elasticity of imports in the longer-term has been roughly stable.

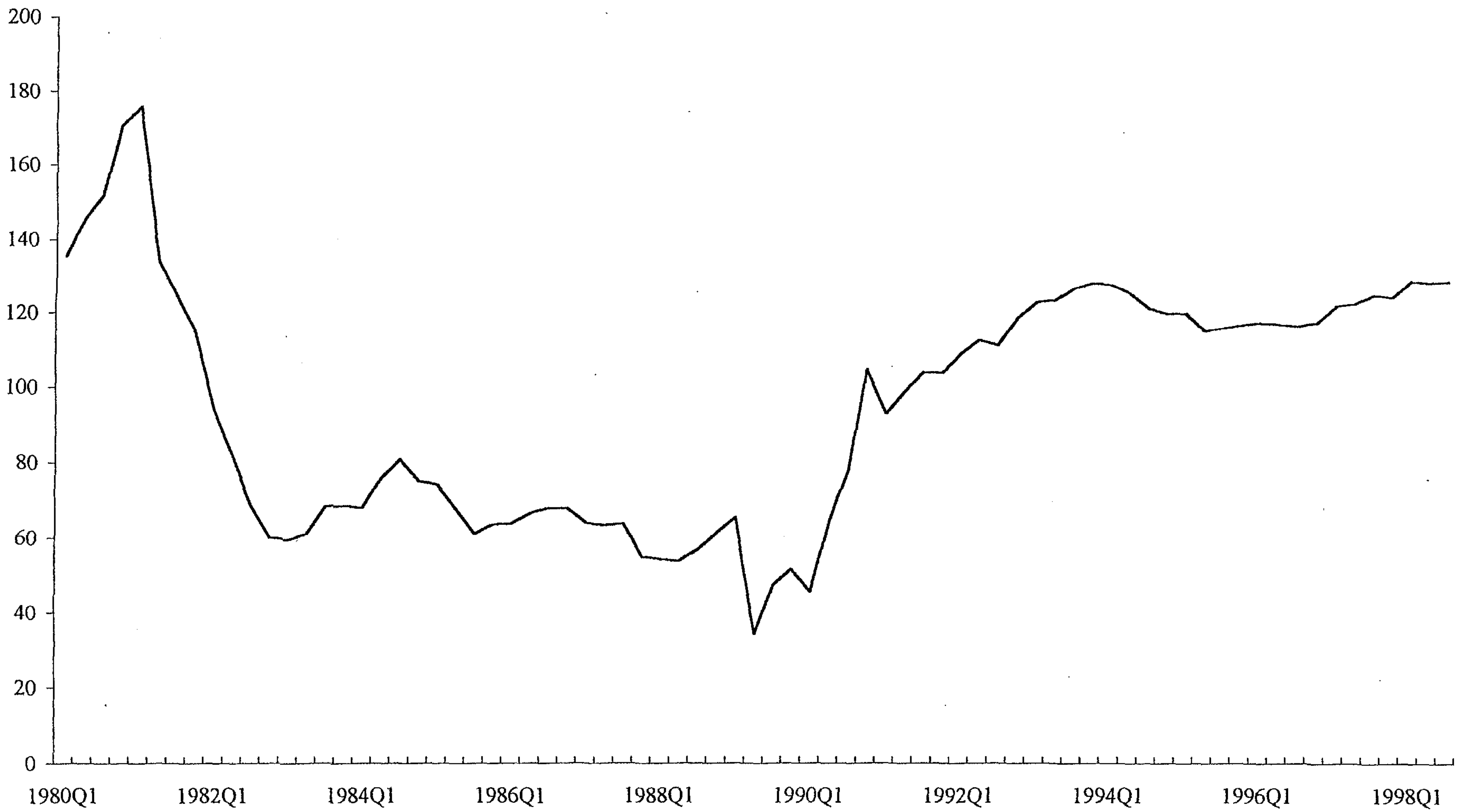
The conclusion which emerges from tests is that the estimated coefficients underlying equation (4) appear to be robust to structural change. Although there are signs of a pro-cyclical behavior of the income and price elasticity of imports, the underlying long-run determinants of Argentina's aggregate imports show no evidence of a breakdown in the 1990s, despite the far-reaching structural changes in the economy. In sum, the very rapid import growth of the 1990s should be seen as combination of a high income and price elasticities of imports, rapid real GDP growth, and relative price movements (via tariff reductions and real exchange rate appreciation) that favored importables relative to domestic goods.

IV. Conclusions

Like many other emerging market economies in the 1990s, Argentina has experienced a marked deterioration in its foreign trade balance as economic growth picked up. Shifts in the trade balance have displayed a pronounced cyclical behavior, moving sharply into deficit during cyclical upturns and improving only moderately during recessions. To identify the determinants of these fluctuations, this paper has estimated aggregate export and import functions for Argentina on the basis of cointegration analysis and error correction models, and using a considerably broader set of indicators than previous studies.

The widening of Argentina's trade balance during 1991-98 reflects the substantially lower growth of the *volume* of exports relative to that of imports. Although export growth did accelerate in the 1990s relative to the 1980s, it fell short of achieving the double-digit rates observed in other fast growing emerging market economies. We show that this can be traced to the conflicting balance of four distinct variables. First, the bulk of Argentina's exports still consists of raw materials or lightly manufactured primary products; accordingly, we found that world commodity prices has a nearly one-to-one effect on the volume of exports in the long-run. So, while the favorable terms-of-trade in the first half of the 1990s helped foster exports, the latter have been clearly hindered by the decline in world commodity prices since 1996. Second, we find that Argentina's exports are highly elastic to net aggregate investment but also to domestic consumption with an (negative) elasticity close to unity. Thus, while the rapid expansion of the net capital stock associated with the structural reforms have clearly contributed to enhance the country's export capacity through the 1990s, a growing share of the production of exportables has been absorbed domestically by buoyant consumer demand. Finally, manufacturing exports have been highly sensitive to economic activity in MERCOSUR as well as to the real exchange rate between Argentina and Brazil. As the Argentinian peso appreciated relative to the Brazilian Real (Figure 7) and

Figure 7. Argentina: Real Bilateral Exchange Rate with Brazil [1]
(1991=100)



[1] Based on consumer price indexes.

economic activity in Brazil slowed down in the second half of the decade, regional demand for Argentina's exports tapered off accordingly.

On the import side, domestic absorption has been the single most important factor behind shifts in the trade balance. We have estimated the income elasticity of imports to be around 2¼ in the longer run and close to 3 in the short-run. Moreover, although no obvious structural changes could be detected in the long-term relationship between imports and real GDP over the 1980-97 period, we find evidence of some pro-cyclical behavior of the import elasticity, rendering the trade balance particularly sensitive to the domestic business cycle. With regard to the elasticity of imports to relative prices, we find it to be nearly a third lower than the income elasticity in the longer-run, and even less significant in the short-run. Thus, while the substantial appreciation of the real exchange rate in the 1990s explains part of the rapid growth of imports during the period, the buoyancy of economic activity was clearly the main driving force.

These results have some relevant policy implications. First, the impact of high domestic demand on the trade balance cannot be easily offset by a depreciation of the real exchange rate. A simple calculation illustrates the point. If potential real GDP growth is 5 percent a year, the relative import price (or similar measure of the real exchange rate) depreciates by 2 percent a year,[31] and the external terms-of-trade are unchanged, the volume of exports would have to grow by some 10 percent a year to prevent a deterioration in the trade balance from a given initial position. Should export growth fail to grow at such a rapid pace, any potentially positive effect of a one-off devaluation—even if the latter could be *sustained* in practice—[32] would be gradually overturned by the cumulative negative effect of rapid income growth on the trade balance. Second, in light of these considerations and given the relatively high elasticity of the trade balance with respect to domestic consumption, a possibly less traumatic way to help restore trade balance equilibrium is through an increase in aggregate savings. Although the aggregate savings ratio in Argentina has risen since the onset of the convertibility regime in 1991, it remains well below the levels found in many other emerging markets.[33] Third, since a substantial part of Argentina's exports was shown to be responsive to unit labor costs, structural measures to keep labor costs below that of main trading partners (including through a reduction in Argentina's high labor taxes and job market rigidities) seem crucial to improve export performance and thus bring the trade balance into equilibrium without requiring a substantial slowdown in economic growth.

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[31]Consistent with the hypothesis of a fixed nominal exchange rate, zero inflation at home, and a trade basket weighted foreign inflation of 2 percent.

[32]Which is not trivial given Argentina's previous experiences with hyperinflation and high degree of real wage resistance.

[33]For a discussion of the determinants of aggregate savings in Argentina and of policy measures conducive to higher savings, see Edwards (1996) and López Murphy and Navajas (1998).

References

Ahumada, Hildegart, 1994, "Argentine Exports: Long-run Behavior and Dynamics," *Anales de la Associacion Argentina de Economia Politica*, XXIX Reunion Annual, Universidad Nacional de la Plata.

Arize, Augustine, 1990, "An Econometric Investigation of Export Behavior in Seven Asian Developing Countries," *Applied Economics*, Vol. 22, pp. 891-904.

Caballero, Ricardo and Victorio Corbo, 1989, "The Effect of Real Exchange Rate Uncertainty on Exports: Empirical Evidence," *World Bank Economic Review*, 3, pp. 263-78.

Ceglowsky, Janet, 1991, "Intertemporal Substitution in Import Demand," *Journal of International Money and Finance*, Vol. 10, pp. 118–130.

Clarida, Richard H, 1994, "Co-integration, Aggregate Consumption, and the Demand for Imports: A Structural Econometric Investigation," *American Economic Review*, Vol. 84, pp. 298–308.

Cóté, A, 1994, "Exchange Rate Volatility and Trade: A Survey," Working paper 1994–1995, Bank of Canada.

Dell'Ariccia, Giovanni, 1999, "Exchange Rate Fluctuations and Trade Flows: Evidence from the European Union," *Staff Papers*, forthcoming

Diaz Alejandro, Carlos, 1970, *Essays in the Economic History of the Argentine Republic*. Yale.

Figueroa, Edgardo and J. Morales Rins, 1996, "El Perfil Exportador de la Argentina," Fundacion Andina, Buenos Aires.

Edwards, Sebastian, 1996, "Why are Latin America's Savings Rates So Low? An International Comparative Analysis," *Journal of Development Economics*, Vol. 51, pp. 5–44.

Garriga, Marcel and Pablo Sanguinetti, 1995, "?Es el Mercosur un Bloque Natural? Efectos de la Politica Comercial y la Geografia Sobre el Intercambio Regional," *Estudios*, XVIII, No. 73, pp. 59–68.

Giorgianni, L. And Gian Maria Milesi-Ferretti, 1997, "Determinants of Korean Trade Flows and their Geographical Destination," International Monetary Fund, Working Paper WP/97/54.

Goldstein, Morris and Mohsin S. Khan, 1985, "Income and Price Effects in Foreign Trade," in Jones, R. And Kenen, P. (eds.) *Handbook of International Economics*, Vol. II, pp. 1041–105, Amsterdan.

Gonzaga, Gustavo M. and Maria Cristina Terra, 1997, "Equilibrium Real Exchange Rate, Volatility, and Stabilization," *Journal of Development Economics*, Vol. 54, pp.77–100.

Hamilton, James, 1996, *Time Series Analysis*. Princeton.

Houthakker, H.S. and Stephen P. Magee, 1969, "Income and Price Elasticities in World Trade," *Review of Economics and Statistics*, Vol. 51, pp. 111–25.

Johansen, Søren, 1988, "Statistical Analysis of Cointegration Vectors," *Journal of Economic Dynamics and Control*, Vol. 12, pp. 1551–80.

Kacef, O. L, 1998, "La Penetracion de las Exportaciones Argentinas en el Brasil," *Estudios de la Economia Real*, No. 4 (May), Ministerio de Economia y Obras y Servicios Publicos. Buenos Aires.

Kiguel, Miguel A. and Nissan Liviatan, 1992, "The Business Cycle Associated with Exchange Rate-Based Stabilization Policies," *Word Bank Economic Review,* no.6 (May), pp. 279–305.

Lipschitz, Leslie and Donogh McDonald, 1991, "Real Exchange Rates and Competitiveness: A Clarification of Concepts," International Monetary Fund Working Paper, WP/91/25, March 1991.

López Murphy, R. And Fernando Navajas, "Domestic Savings, Public Savings And Expenditures on Consumer Durable Goods in Argentina," *Journal of Development Economics*, Vol 57, pp. 97–116.

Pesaran, M. H. and Y. Shin, 1998, "An Autoregressive Distributed Lag Modelling Approach to Cointegration Analysis," in Strom, S., A. Holly, and P. Diamont (eds.), *Centennial Volume of Ragnar Frisch*, Cambridge.

Pesaran, M.H., Y. Shin, and R.J. Smith, 1996, "Testing for the Existence of a Long-run Relationship," DAE Working Paper No. 9622, Department of Applied Economics, University of Cambridge.

Pitchett, Lant and Geeta Sethi, 1994, "Tariff Rates, Tariff Revenues and Tariff Reforms: Some New Facts," *World Bank Economic Review*, 8(1), pp.1–16.

Reinhart, Carmen. M., 1995, "Devaluation, Relative Prices, and International Trade," *IMF Staff Papers*, Vol. 42, No. 2, pp. 290–312.

Senhadhi, Abdelhak, 1998, "Time-Series Estimation of Structural Import Demand Equations," *IMF Staff Papers*, Vol. 45, No.2, pp. 236–268.

Végh, Carlos A., 1992, "Stopping High Inflation: An Analytical Overview," *IMF Staff Papers*, Vol. 39, pp. 626–95.