Journal of Applied Economics, Vol. V, No. 1 (May 2002), 19-57

DETERMINANTS OF ARGENTINA'S EXTERNAL TRADE

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Following the liberalization reforms of the late 80s and early 90s, several emerging market economies have experienced large and persistent trade deficits. This paper focuses on the Argentine experience, examining the extent to which trade imbalances in the 1990s resulted from income and relative price movements, as well as from shifts in foreign trade elasticities associated with structural changes. New estimates of export and import equations are presented using a broader set of variables than previous studies and distinguishing between intra and extra MERCOSUR trade. We find that considerable export sensitivity to world commodity prices, domestic absorption, and economic activity in Brazil, combined with a high income elasticity of imports, are key determinants of Argentina's trade balance.

JEL classification codes: F11, F14, F31
Key words: Argentina, foreign trade elasticities, international competitiveness, MERCOSUR

I. Introduction

Following the financial and trade liberalization reforms of the late 1980s

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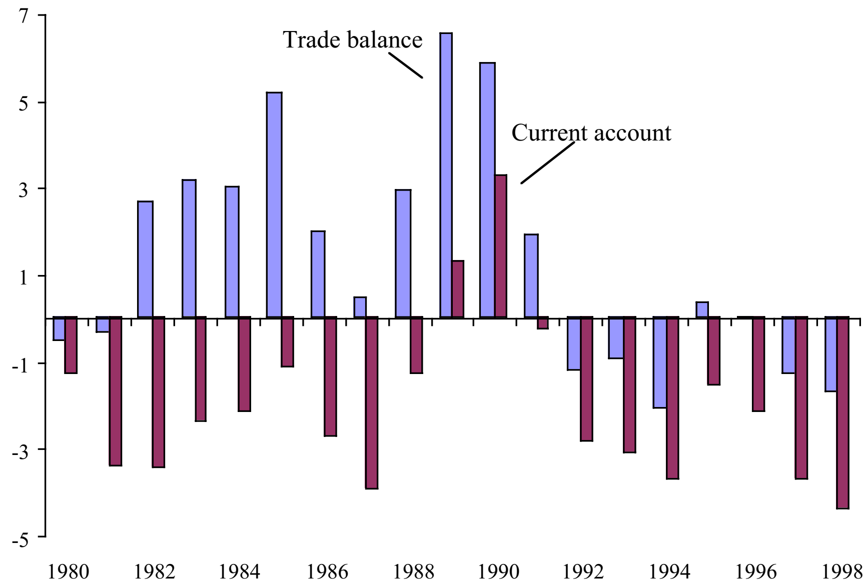
and early 1990s, a number of emerging market economies have experienced sizeable current account deficits. In countries such as Argentina and Mexico, where rapid disinflation was accomplished through the combination of an exchange rate peg and trade liberalization, soaring current account deficits appeared 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. At an initial stage of the reforms, expectations were created that such large 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. Even though it may be optimal for capital-scarce economies to run large external trade deficits for prolonged periods, it is nevertheless striking that trade and current account imbalances in several Western Hemisphere emerging economies continued to rise for years after economic and trade liberalization. In the particular case of Argentina, the current account deficit exceeded its previous (1994) cyclical 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- to long-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

Figure 1. Argentina: Trade Balance and Current Account



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.[1] 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 analysis, structural

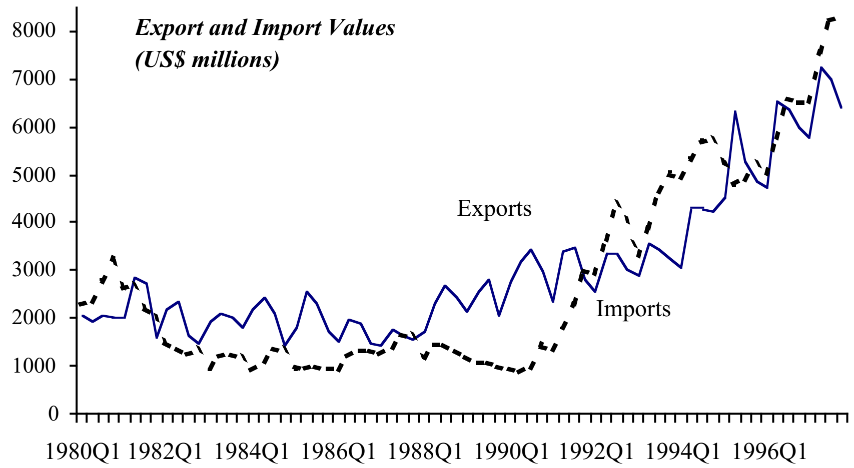
[1] See Goldstein and Khan (1985) for a comprehensive survey of earlier studies within this tradition.

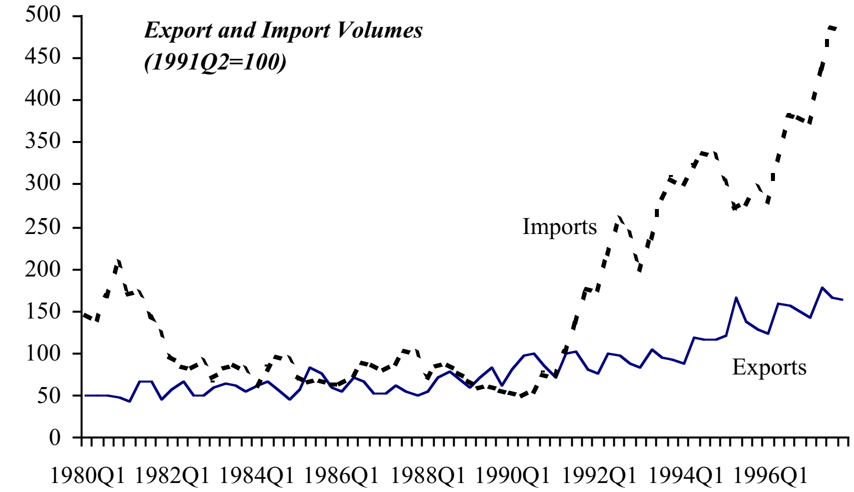
stability tests, and vector autoregressions (VARs) have provided a new foundation for modeling the complex interactions between cyclical and permanent components in time series and test for the existence of structural breaks in the underlying macroeconomic relationships. Some of these new econometric 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,[2] 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 70s and 80s, (averaging 5˘ percent a year between 1991 and 1998), inflation was weeded out, and both exports and imports trended up relative to the 1980s (Figure 2). 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.

[2] 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





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.

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, export volume growth not only accelerated markedly in the 90s relative to the 80s, but it also became less unstable, i.e., its standard deviation declined. Second, compared with other countries in the Western Hemisphere and Asia, the average growth rate of Argentina's exports in the 90s 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.

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[3] consist of crude oil, soya, wheat, vegetable oils, leather and meat-products which do not involve significant industrial

[3] 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	4.4	17.8	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

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

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° 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.

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 latter case, Argentina's policies and macroeconomic performance may play a key role, and as 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 the differences between 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 from the series on Argentina's total exports (primary commodity exports to Brazil, however, are left as part of this first group) 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. 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,[4] 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,[5] 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:

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

where x^s stands for export volume, px for the relative price of exports, k for the aggregate net capital stock, c for real domestic absorption, σ_{RER} for real exchange rate volatility, and ε for the residual. All variables are in logs except for the exchange rate volatility measure.

[4] Cf. Goldstein and Khan (1985). Recent empirical work within this tradition include Arize (1990), Reinhart (1995), and Giorgianni and Milesi-Ferretti (1997).

[5] 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).

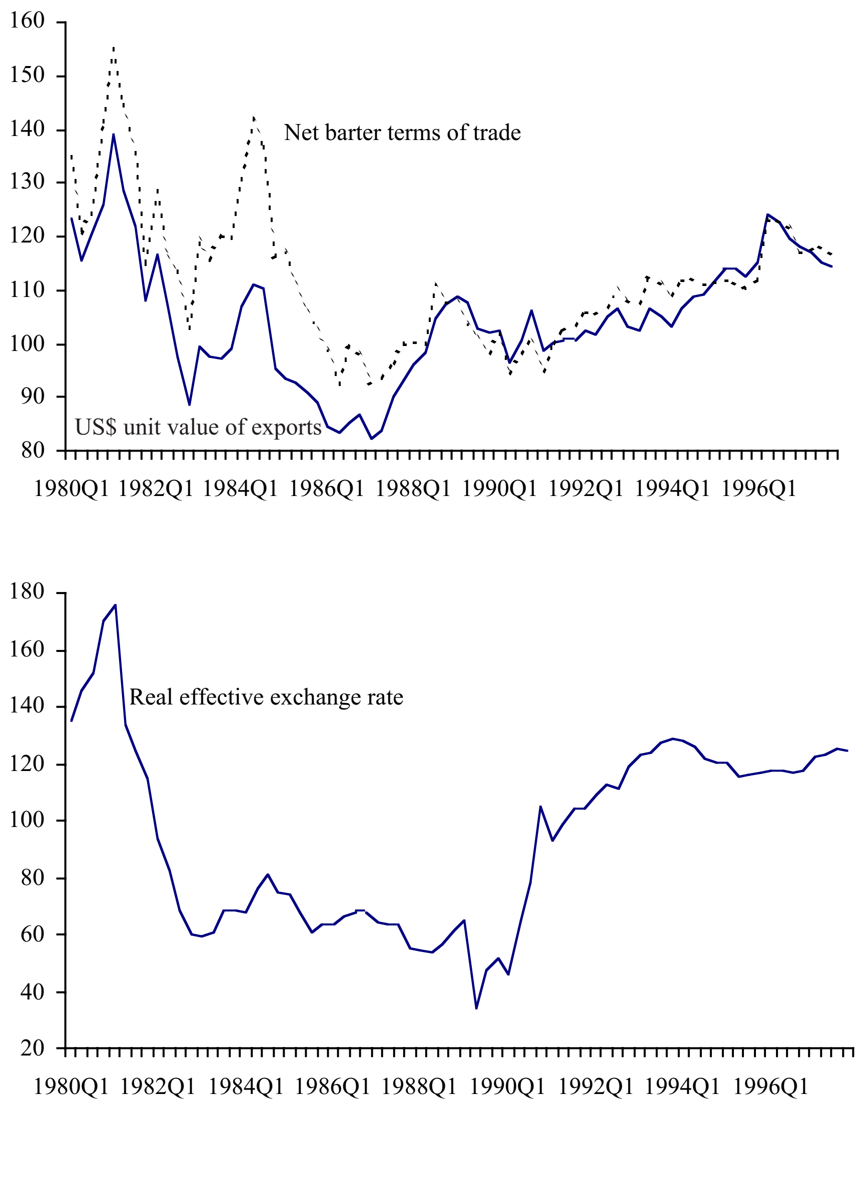
While there is a wide consensus on the basic functional form of aggregate export functions, a number of measurement issues remain controversial. These include: the choice of the relevant price indicator (e.g. whether to use the real exchange rate, relative unit labor costs, or simply the US dollar unit 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 distinct external competitiveness indicators (cf. Lipschitz and McDonald, 1991), most studies have followed an empiricist approach and simply pick the indicator which yields the best fit. We adopt the same strategy here, experimenting with a few relative price indicators which appear to share a common trend with that of exports, including the real effective exchange rate (defined as the ratio of the domestic consumer price index to the average foreign consumer price indices weighted by each foreign country's share in Argentina's trade and expressed in the same currency), the ratio of export price to domestic unit labor costs, the net barter terms of trade (the ratio of unit value of exports to the unit value of imports), the ratio of export price to domestic consumer price index, and the unit value of exports expressed in US dollars. Figure 3 depicts some of these price indicators.

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.

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



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 seems 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 the real effective exchange rate (averaged over a one-year window) as a proxy for the effects of exchange rate uncertainty on exports.[6]

On the basis of the choice of indicators just discussed, we proceed with the estimation of the above export model in 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 g 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).

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 price variables considered appear to be stationary.

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[6] 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

Table 2. Unit Root Tests (Quarterly Data, 1980-97)

Variable	DF[1]	ADF[2]
Export volume (X)	-1.99	-1.20
Export price in dollars (PX^*)	-2.83	-4.13*
Net capital stock (K)	1.89	-0.41
Consumption (C)	-0.86	-2.81
Real effective exchange rate (REER)	-2.23	-3.35
Export price/domestic unit labor costs (PX/ULC)	-2.84	-3.47*
Export price/consumer price index (PX/CPI)	-1.29	-3.13
Import volume (M)	0.11	-2.91
Real GDP (Y)	-1.42	-2.53
Relative price of imports ($PM^*(1+T)/CPI$)	-1.36	-3.49*
Real interbank rate (RIR inter)[3]	-8.19	-3.63*
Real deposit rate (RIR dep)[3]	-8.15	-3.47*

Notes: [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. Sources: IMF, INDEC, and Ministry of Economy of Argentina.

Once we test for non-stationarity and find that some—if not all—the variables in (1) are I(1), we proceed to test them for cointegration. Here we employ two commonly used tests. One is the procedure due to Johansen (1988), which is based on the following vector autoregressive (VAR) system:

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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 REER over a one-year window appeared as a fair compromise.

$$\Delta X_t = \mu + \phi_t + \Pi\ X_{t-1} + \Gamma\ \Delta X_{t-1} + \textstyle\sum_2^n T_n \Delta X_{t-n} + w_t \qquad (2)$$

where X is a vector comprising all the I(1) variables entering equation (1), 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. Cointegration tests amount to testing whether the matrix Π is non-singular, 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[7] set of α cointegrating coefficients which renders the residual of a level equation such as (1) stationary.

An alternative test procedure has been advanced by Pesaran and Shin (1998) in the context of an auto-regressive distributed lag model (ARDL). This procedure allows for a mix of I(1) and I(0) variables in an equation such as (1), thus dispensing with the need of unit root pre-testing —an attractive feature given the well-known low power of unit root tests in small samples—. The Pesaran and Shin (1998) test consists of adding, in the first differenced version of equation (1), lags of first differences of the variables so as to orthogonalize the relationship between the explanatory variables and the residual term g. 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.[8]

The Johansen and Pesaran-Shin tests in Table 3 support the hypothesis there is one cointegrating vector tying the level of exports to the right-hand side variables in (1).

Once the existence of at least a cointegrating vector in (1) is established, consistent estimates of the long-run elasticities can be obtained within these respective testing frameworks. As discussed in Pesaran and Shin (1998), such

[7] The procedure allows for up to n-1 such vectors, where n is the number of I(1) variables.

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

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 Value	Trace Statistic	95% Critical Value
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 Test2

	F-Statistic	95% Critical Value
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 Value	Trace Statistic	95% Critical Value
r = 0	r = 1	52.62*	40.53	134.53*	102.56
r ≤ 1	r = 2	35.21*	34.40	81.90*	75.98
r ≤ 2	r = 3	20.01	28.27	46.69	53.48

Table 3. (Continued) Cointegration Tests

C. Import Demand Equation

Johansen's Maximum Likelihood Rank Tests[1]

Null	Alternative	λ-Max Statistic	95% Critical Value	Trace Statistic	95% Critical Value
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 Value
Model without a time trend	6.28*	4.05
Model including a time trend	7.30*	4.70

Notes: *Significant at 5 percent. [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.

estimates can be simply derived in an ARDL framework from the coefficients on the level variables, provided that orthogonalization between the residual term and the right-hand side variables is achieved by including a sufficient number of autoregressive terms, and residuals appear to be serially uncorrelated.[9] In the case of the Johansen method, the matrix of cointegration

[9] 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.

vectors can be obtained through a suitable factorization of the matrix Π whenever the latter is not full rank.[10]

Given the evidence supporting cointegration, ARDL estimates of the vector of long-run coefficients α are reported in part A of Table 4.[11] As discussed above, we experimented with several price variables. 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 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,[12] 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). Equations (A.1) and (A.2) should be clearly preferred.

Equations (A.1) and (A.2) both 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

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[10] For a thorough discussion on the specification, statistical properties of VECMs and the factorization of the Π matrix, see Hamilton (1996). In this paper, all VECM estimates were obtained using the Pesaran and Pesaran (1997) Microfit program.

[11] Given the similar results yielded by both the Johansen VECM and the ARDL methods, only the ARDL estimates of the cointegrating vector underlying equation (1) are reported.

[12] 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 variables PX/CPI and PX/ULC. 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.

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	px^*	k	c	σ (Δrer)	Trend 90s	*px/cpi*	*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)		

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

B. Error Correction Representation (Dependent variable: first difference of the log of exports)

	Constant	$\Delta(px^*)_t$	$\Delta(px^*)_{t-1}$	k_t	c_t	$\sigma(\Delta rer)_t$	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)		

Notes: [1] Excluding manufacturing exports to Brazil during 1990-97. Seasonal dummies added to all regressions.

has been particularly responsive to the rapid growth of the net aggregate capital stock in the 1990s. 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.[13]

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 equilibrium represented by the level equation (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.

Estimates of the error correction representation of equation (1) are shown in part B of Table 4. The estimated coefficients 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 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. Equations (B.1) and (B.2) passed all the

[13] 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.

diagnostic statistics for residual autocorrelation, functional form, and heteroscedasticity.

In light of the marked structural change in the economy in the 1990s, structural stability tests were carried out. The estimates appear to be robust to underlying structural changes.[14] We also performed a number of variable addition tests to try to assess whether the model is robust to alternative specifications. These tests underscored the robustness of the original model.[15]

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:

$$x_t^d = \gamma_0 - \gamma_1 (1 + t^*)\, px_t^* / p_t^* + \gamma_3 y_t^* + u_t \tag{3}$$

$$x_t^s = \rho_0 + \rho_1 px_t^* - \rho_2 ulc_t + \rho_3 k_t - \rho_4 \sigma_{RER_t} + v_t$$

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[14] We based structural stability tests on the cumulative sum of the recursive residuals (CUSUM). The cumulative plot of recursive residuals falls well within the 5 percent significance band. 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 comment the results of the CUSUM test on (A.1).

[15] 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 proved to be statistically significant at 5 percent.

where px^* is the US dollar price of Argentina's exports net of the foreign tariff rate t^* relative to the foreign price index p^* (also expressed in dollars), and y^* is real GDP in the MERCOSUR partner countries, all in logs.[16] The supply equation differs from (1) in that *ulc* (the log of ULC deflated by nominal exchange rate) is included, domestic consumption is excluded,[17] and σ_{RER} now corresponds to the bilateral exchange rate between Argentina and the rest of MERCOSUR.

Strong simultaneity across equations, and the fact that most variables entering (3) appear to be non-stationary, call for an econometric specification that accommodates both features. This can be done re-writing (3) as a vector error correction model (VECM) as in (2). Maximum likelihood estimation of (3) yields a matrix of long-run coefficients Π bearing the cointegrating relations between the supply and demand variables (if any); estimates of Γ yield the short-run or impact elasticities. If (3) is a valid representation of the data generation process, one should find at least two cointegrating relationships—one for the supply and the other for the demand equation.

Results of cointegration tests for this supply and demand system are reported in part 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, the following identifying

[16] Since Brazil has accounted for 85 to 90 percent of Argentina's manufacturing exports to MERCOSUR, Brazil's real GDP and consumer price index (expressed in US dollars) 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).

[17] 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.

restrictions are imposed: exclude y^* and p^* from the supply equation and *ulc* and *k* from the demand equation, and make the coefficient on px^* equal to the negative of the coefficient on $p^*/(1 + t^*)$ in the demand equation. Since only two restrictions are needed for exact identification, our system becomes overidentified. This allows us to test the extra restrictions and assess the robustness of the postulated theoretical model.

The estimated vector of long-run coefficients with all these restrictions imposed is reported in 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 respect to foreign income and relative prices 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 estimated very imprecisely, as indicated by its low t-ratio. In contrast, the estimated elasticity of export supply with respect to the capital stock is rather high and estimated with considerable precision.[18] Somewhat surprisingly, Lagrange multiplier tests (not reported) indicated that the real exchange rate instability did not add any significant explanatory power to the VECM and therefore was dropped from the regressions.[19] This contrasts with the results of section II.B which indicate that real exchange rate instability tended to undermine Argentina's commodity exports, particularly in the 1980s when Argentina's operated a flexible exchange rate regime and experienced extreme relative price volatility.[20]

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[18] This may reflect the inadequacy of our proxy for productive capacity in manufacturing, which here is taken to be the aggregate capital stock.

[19] It lacked statistical significance 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 theory.

[20] The statistical insignificance of the exchange rate volatility term in the equation on manufacturing exports to MERCOSUR, albeit surprising, echoes the findings of other

Table 5. Manufacturing Exports to Brazil: VECM Estimates of the Equation System (t-ratios in brackets, quarterly data for 1980-97)

A. Long-run Coefficients [1]

	px^*	ulc	k	y^*	p^*
Supply Equation	1.29 (1.75)	-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 [p val. = 0.23]*

B. Error Correction Representation[1]

	$\Delta (ulc)_t$	$\Delta(px^*)_{t-4}$	ECM^s_{t-1}			R^2	D.W.
Supply Equation	-0.82 (-2.04)	0.51 (1.76)	-0.33 (-3.77)			0.51	2.11

	$\Delta (x)_{t-1}$	$\Delta ((1+t^*) px^*/p^*)_{t-1}$	$\Delta (y^*)_{t-1}$	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

	$\Delta (ulc)_t$	$\Delta ((1+t^*) px^*/p^*)_{t-1}$	$\Delta (y^*)_{t-1}$	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

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

The associated error correction equations show that the above variables also play an important role in the short-run, with the notable exception of the 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 (Table 5, part B).[21]

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 the effect of variables other than income in fostering import demand, this suggests 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; Senhadji, 1998).

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 insignificant. 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).

[21] In addition to standard tests for residual autocorrelation, heteroscedasticity, and functional form, we submitted all error correction equations to CUSUM tests for parameter stability. The plot of the respective cumulative sum of residuals fell well within the 5 percent significance boundaries, suggesting the estimated parameters of the model are relatively stable over time.

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 Senhadji (1998) for the period 1960-93. A number of factors may explain such a mismatch. One is sampling. Since both authors worked with annual observations spanning over twenty 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.[22] Another possible explanation is the type of explanatory variables included in their model. Neither Reinhart (1995) nor Senhadji (1998) incorporate any tariff effect in their price variables. 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, the important structural changes in the Argentine economy in the 90s may have raised the income elasticity of imports, 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"

[22] 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.

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, 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

$$m_t = \beta_0 + \beta_1 y_t - \beta_2 pm_t^*(1+t)/cpi_t - \beta_3 RIR_t + \xi_t \qquad (4)$$

where m stands for total import volume, y for real gross domestic product, pm^* for US dollar unit value of imports, t for the import tariff rate,[23] cpi for the consumer price index deflated by nominal exchange rate, RIR for the real interest rate, and ξ for the error term. All variables, except for the real interest rate, are in logs.

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

[23] Here we measure average import tariff rate as the ratio of total tariff revenues by total imports. Although this can be an inaccurate 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).

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 vs. 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 have tried both indicators.

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 shortcomings, the use of actual future inflation as a measure of inflationary expectations seems 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 (4) we give greater prominence to the use of "forward-looking" real interest rate measures, though also check the sensitivity of the results to the use of "backward-looking" measures.[24]

[24] 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 then annualized. With the advent of macroeconomic stabilization in the 1990s, domestic lending institutions resumed quoting interest rates on an annual basis which were then deflated by the 12-month ahead inflation. Quarterly real interest rates were computed as arithmetic averages of the monthly rates.

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.

Second, we test for cointegration using the testing procedures advanced by Johansen (1988) and Pesaran and Shin (1998), already discussed in Section II. As shown in part C of Table 3, 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 and Shin (1998) ARDL methodology.

Table 6, part A, reports the long-run coefficients of the estimation of (4) by the ARDL method, with the lag structure of the autoregressive selected by the Schwarz Bayesian Criterion.[25] 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 (absolute t-statistic above 2). 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, regardless of the specification (using a forward or backward-looking real interest rate measure, 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 0.7 to 0.8 range. The high t-statistics attached to both coefficients indicate they are estimated with considerable precision.

In part B of Table 6 we report the short-run or error correction models associated with the long-run specification (A.1). All explanatory variables

[25] Results using the Johansen VECM approach are similar to the ARDL model.

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, *m*)

	Constant	*y*	*pm** *(1+t)/cpi*	RIR inter[1]	Trend 90s[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)		

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

B. Error Correction Representation (Dependent Variable: First Difference of the Log of Imports, Δm)

	Constant	$\Delta (y)_t$	$\Delta (y)_{t-1}$	$\Delta(pm^*)(1+t)/cpi)_t$	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 90s==> F (1, 62) = 0.06 [p = 0.80]

II) 90s Slope dummy on real GDP ==> F(1, 62) = 2.48 [p = 0.12]

III) 90s Slope dummy on relative import price ==> F(1, 62) = 1.25 [p = 0.27]

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

entering the model were statistically significant and yielded the signs predicted by theory, with the exception of changes in interest rate. Yet, the estimated coefficients on the latter were rather small and practically insignificant. The estimates point to an impact income elasticity of imports around 2 and a cumulative short-run elasticity (over a two quarter period) of 2ʃ, while the relative price elasticity is ˘. 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 no evidence of residual autocorrelation.

We undertook 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 90s. First, we carried out a CUSUM test for parameter stability in the ARDL model estimated above which does not reject the null hypothesis of parameter stability.[26] A second set of tests 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 6. 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.

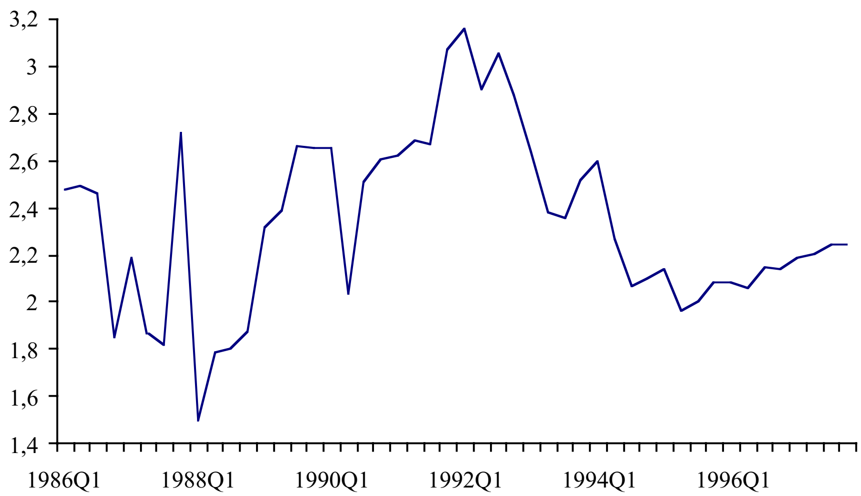
Finally, we conducted a stronger test for structural change based on the level estimation of (4) by recursive OLS. The latter allows the estimated coefficients to change over time and so should be able to reveal whether they

[26] The cumulative sum of residuals falls well within the 5 percent critical bounds.

have witnessed a marked increase or decrease in the 1990s. The resulting elasticity is graphed in Figure 4.[27] As expected, the 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 slight 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 these tests is that the estimated

Figure 4. Argentina: Recursive OLS Estimates of Income Elasticity of Imports



[27] Lags of the variables in levels were included among the regressors so as to reproduce the ARDL representation underlying the long-run estimates of Table 6. 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.

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 1991-98 period 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

A key policy issue faced by several emerging market economies in recent years has been the marked widening of external trade and current account deficits following the economic and trade liberalization reforms of the late 1980s and early 1990s. The persistence of such large deficits raises concerns about the sustainability of high average rates of economic growth and increases the vulnerability of those economies to swings in the supply of external finance. These concerns are all the more serious in countries with hard currency pegs, insofar as the latter imposes clear limits on the speed and extent of relative price adjustments that may be needed to compensate for losses in external competitiveness stemming from adverse shocks. Against this background, this paper has tried to identify the extent to which Argentina's rising trade deficit in the 1990s resulted from income and relative price effects, and whether such effects might have been exacerbated by the country's adherence to a fixed exchange rate regime and structural changes in the economy. Aggregate export and import equations were estimated using a considerably broader set of macroeconomic and relative price indicators than previous studies, and distinguishing between intra- and extra-MERCOSUR trade.

Two main sets of factors lay behind the widening of Argentina's trade balance during 1991-98. One is the relatively low growth of the volume of exports, both relative to that of imports as well as relative to the export performance of other fast growing emerging market economies. This, in turn, can be traced to the interplay of four variables. First, the bulk of Argentina's exports still consists of raw materials or lightly manufactured primary products that bear a unit elasticity to world commodity prices. Thus, while the favorable terms-of-trade in the first half of the 1990s helped foster the volume of exports, the latter appear to have been hindered by the decline in non-oil commodity prices since 1996. Second, while Argentina's exports appear to be highly elastic to net aggregate investment, they also appear to be reasonably elastic to domestic consumption. This implies that while rapid investment growth following the structural reforms of the early 1990s has helped enhance the country's export capacity, a growing share of exportable goods has been absorbed domestically due to buoyant consumer demand. Finally, Argentina's manufacturing exports tend to be 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 and economic activity in Brazil slowed down in the second half of the decade, regional demand for Argentina's exports tapered off accordingly.

The other main cause of the deterioration of Argentina's trade balance is the high income elasticity of import demand. We estimate the latter to be 1.9 on impact and around 2˘ in the longer run.[28] Moreover, while the income elasticity of imports does not appear to have risen significantly in the 1990s as far-reaching structural changes took place in the economy, it has displayed a slightly pro-cyclical behavior. This renders the trade balance all the more

[28] As noted above, care must be taken not to interpret the 2˘ figure as a long-run equilibrium elasticity, as it would imply that (ceteris paribus) all GDP would end up being eventually imported. Rather, this figure should be interpreted as an estimated elasticity over a time window that is not so long to encompass several generations but which is long enough so that that elasticity is likely to be "persistent" enough not to be significantly affected by short-term developments. Other studies using a similar econometric methodology and data series that span over a twenty to thirty year horizon, also find "long-run" foreign trade elasticities significantly higher than one in several emerging market economies (see Reinhart, 1995; Senhadji, 1998; Senhadji and Montenegro, 1999).

sensitive to the domestic business cycle, exacerbating the deficit problem during cyclical upswings. Regarding the price elasticity of imports, we find it to be much lower, being nearly a third of the income elasticity in the longer-run and only 0.25 on impact. Thus, while the substantial appreciation of the real exchange rate surely explains part of the rapid growth of imports during the 1991-98 period, its impact was not as large as one might think; the buoyancy of economic activity appears to have been 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,[29] 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[30] — 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 more sustainable 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.[31] Third, since a substantial part of Argentina's exports

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

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

[31] 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).

was shown to be responsive to unit labor costs, structural measures to keep labor costs below that of main trading partners (including 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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