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Revista de Economía y Estadística, Cuarta Época, Vol. 44, No. 1 (2006), pp. 13-38.

<http://revistas.unc.edu.ar/index.php/REyE/article/view/3823>



La Revista de Economía y Estadística, se edita desde el año 1939. Es una publicación semestral del Instituto de Economía y Finanzas (IEF), Facultad de Ciencias Económicas, Universidad Nacional de Córdoba, Av. Valparaíso s/n, Ciudad Universitaria. X5000HRV, Córdoba, Argentina.

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Cómo citar este documento:

Herrou-Aragón, A. (2006). How Far Have Commercial Policy Reforms of the 90s in Argentina Gone?. *Revista de Economía y Estadística*, Cuarta Época, Vol. 44, No. 1, pp. 13-38.

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How Far Have Commercial Policy Reforms of the 90s in Argentina Gone?

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Resumen:

El objetivo de este trabajo es evaluar el grado en que los cambios en la política comercial, en Argentina durante los años noventa, han contribuido a expandir el volumen de comercio del país, en comparación con los años ochenta. Las estimaciones indican que las políticas comerciales de fines de los 90, resultaron en una reducción en los impuestos y aranceles al comercio exterior del 80 por ciento, en la segunda mitad de los ochenta a un 20 por ciento. Esta reducción en la imposición al volumen de comercio, se estima que representa, el 68 por ciento del aumento total en el volumen de importaciones del período 1995-1999 con respecto al de la segunda mitad de los ochenta. Por otra parte, aumentos en la demanda agregada, en el producto real y términos de intercambios más favorables representan el 32 por ciento restante del aumento en las importaciones.

Palabra clave: Política Comercial.

Clasificación JEL F1

Abstract:

The purpose of this paper is to assess the extent to which changes in commercial policies in Argentina during the 90s have contributed to expand the country's volume of trade compared to those of the 80s. The estimates indicate that the commercial policies of the late 90s resulted in a reduction in the taxation of international trade from 80 percent in the second half of the 80s to 20 percent. This reduction in the taxation of the volume of trade

is estimated to account for 68 percent of the total increase in the volume of imports of the period 1995-99 over that of the second half of the 80s. On the other hand, increases in aggregate demand and in real output, and more favorable external terms of trade account for the remaining 32 percent of the increase in imports.

Keywords: International trade - Trade Policies

JEL Classification: F1.

Trade liberalization policies were introduced in Argentina during the 90s along with macroeconomic reforms, privatization of public sector enterprises, and de-regulation of the economy. Compared to the average of the 1985-89 period, the quantity of imports and exports in Argentina increased on average during the years 1995-1999 by 450 and 90 percent, respectively. Part of this increase in imports could be attributed to increases in real output and aggregate expenditure. During the second half of the 90s, real output and aggregate expenditure increased by about 30 and 40 percent, respectively, compared to their averages of the 1985-89 period. The question is thus the extent to which trade liberalization policies have contributed to the increase in the volume of trade.

This paper has the objective of estimating the magnitude of the tax on trade resulting from the commercial policies implemented during the 90s and the impact of this tax over the volume of trade of Argentina. In Section I, there is a description of the commercial policies in Argentina followed in the last three decades and equivalent uniform tax rates on trade are calculated for different time periods. In Section II, an import function is estimated in order to evaluate the impact of changes in this tax during the 90s on the volume of trade. In Section III, the increase in the quantity of imports during the 90s compared to that of the period 1985-89 is decomposed into several sources, namely, changes in commercial policies, in external terms of trade, and in real output and expenditure. The concluding remarks are in Section IV.

* This is a new version of the paper presented at the XXXV Meeting of the Argentinean Economic Association. The author thanks J. Berlinski (UTdT) and A. Navarro (National Academy of Economics) and my colleagues of the Department of Economics of the UESiglo21 for their helpful comments. The remaining errors are of the sole responsibility of the author.

I. COMMERCIAL POLICIES AND RELATIVE PRICES

As is well documented by Diaz-Alejandro (1970) and by J. Berlinski (2001), protection to import-competing activities arose in Argentina as a response to the crisis of the 30s. Trade restrictions intensified during the 40s and part of the 50s as the government pursued a policy of inward-looking industrialization. High tariffs, import licensing and prohibitions were extensively used along with subsidization of inputs. In the mid-60s, attempts were made to reduce the anti-export bias of commercial policies by establishing a drawback regime for exports of manufactured goods and replacing quantitative restrictions with import tariffs.

In 1967, the government reduced significantly maximum import tariffs from about 120 to 60 percent in order to offset, at least in part, the effects of a 40 percent currency devaluation on the domestic price level. At the same time, taxes on traditional exports were imposed in order to raise fiscal revenue as part of a price stabilization plan. These export taxes were gradually eliminated during the year to compensate producers of exportable goods for the rising inflation. In addition, most of import prohibitions were removed.

In 1971, multiple exchange rates discriminating against exports were re-introduced. In 1973, quantitative restrictions were re-introduced along with foreign exchange controls and import deposit requirements. During the Martínez de Hoz administration (1976-1981), these restrictions were gradually lifted along with reductions in import tariffs, exchange rates were unified, and export taxes eliminated. In 1978-1980, import tariffs were further reduced. In 1981-1982, quantitative restrictions were re-introduced in response to a balance of payments crisis, multiple exchange rates were resumed, and the program of trade liberalization abandoned.

The Alfonsín administration (1983-89) tried to tighten its control over import demand by increasing quantitative restrictions. Imports that were prohibited included most goods that were locally produced. Imports subject to prior approval required the consultation with domestic producers. A 15 percent import surcharge was introduced in 1985.

By 1987, it was becoming apparent that import-substitution policies had failed to foster economic growth in Argentina. Between 1987 and 1988, import licensing restrictions were relaxed along with reductions in import and export tariffs. These trade liberalization policies were intensified during the Menem administration (1989-1999). In early 1991,

import licensing was eliminated, the coverage of remaining quantitative restrictions significantly reduced, and import tariff rates reduced. In 1995, import tariffs were further reduced as the common external tariff of a regional preferential trade agreement (MERCOSUR) was adopted.

In an attempt to summarize the impact of the whole set of instruments of commercial policies on resources allocation, Díaz-Alejandro (1970) estimated the so-called “equivalent uniform import tariff” as the ratio of domestic prices of non-rural to exportable agricultural goods (adjusted by changes in external terms of trade). His estimates are in line with the view that the commercial policies between mid-40s and mid-50s introduced severe distortions in the economy compared to the policies of the late 20s. Larry Sjaastad (1981) estimated the equivalent import tariff as the ratio of domestic price of imported to exportable agricultural goods (corrected by changes in external terms of trade).

Sjaastad finds that, in the 70s, the equivalent uniform import tariff was about 100 percent compared to that of less restrictive commercial policies of 1935-1939. J. Berlinski (2001) extended Díaz-Alejandro’s estimates to cover most recent commercial policy developments. One potential problem with this definition of relative prices of exportable goods is that the denominator may contain non traded goods that are dependent not only on commercial policies and external terms of trade but also on changes in domestic expenditure and output. Changes in domestic prices of domestic manufacturing goods may be isolated to some extent from changes in their world prices by quantitative restrictions or “water” in tariff rates¹. As a result, changes in prices of domestically-produced goods could reflect not only the changes in commercial policies and external terms of trade but also changes in aggregate expenditure and in real output.

In table 1, Sjaastad’s method of measuring the equivalent uniform import tariff is used to estimate the degree of trade restrictiveness of the commercial policies described above. As in Sjaastad’s paper, the wholesale price index of imported is used as a proxy for the domestic price of import-competing goods. A weighted average of wholesale prices of agricultural and food manufacturing commodities is used as a measure of the internal price of exportable commodities². This measure of relative

¹ In Sjaastad, L. and C. Rodríguez (1979), and in L. Sjaastad (1981), these goods are treated as non-traded goods and thus affected by commercial policies through substitution effects in production and consumption.

² See Annex A for a description of the data.

prices of importable compared to exportable goods (adjusted by changes in external terms of trade) is thus a proxy of commercial policies as it does not include the impact of aggregate expenditure and real output on prices of non-traded goods. The uniform equivalent import tariff rates are calculated using the average relative prices during 1935-39 as a benchmark because of data availability. Although this period was not characterized by free trade, commercial policies during this period were by far less restrictive than the ones implemented later on.

Table 1
Relative Prices of Importable Goods, Terms of Trade and
Equivalent Import Tariff
 (1993=1.00 for price indexes)

Period	Prices Relative (1)	External Terms of Trade (2)	Ratio (1)/(2)	Uniform Equivalent Tax (in %)
1935-39	0.50	0.60	0.85	--
1968-70	0.73	0.67	1.10	30.0
1971-79	1.01	0.69	1.47	73.1
1973-76	1.16	0.64	1.82	115.1
1979-80	0.75	0.76	0.99	17.3
1981-89	1.42	0.93	1.54	81.9
1981-84	1.17	0.76	1.54	82.1
1985-89	1.63	1.06	1.54	81.8
1985-87	1.59	1.04	1.59	81.4
1991-99	1.02	0.99	1.03	21.3
1991-94	1.07	1.02	1.05	23.4
1995-99	0.98	0.97	1.01	19.4

The estimates of Table 1 mirror very closely the aforementioned commercial policy developments. In particular, the aforementioned reversal of the 1968-70 trade policies during 1973-1976 is associated with a substantial increase in taxation of trade from 30 to 115 percent (an increase of 85 percentage points in the average equivalent import tariff). The most remarkable episodes of trade liberalization took place, according to the estimates, during 1979-80 and during the 90s when the overall taxation of trade was 17 and 21 percent, respectively. The measure of trade restrictiveness presented in this paper takes well into account the reversal of the trade liberalization policies of the 1979-80 period in subsequent years. As a result of the re-introduction of quantitative restrictions, overall taxation rate on trade is estimated to increase from an average 17 percent during the 1979-1980 liberalization period to an average of about 80 percent during 1981-89.

II. ESTIMATION OF THE IMPORT DEMAND FUNCTION

In this section, the parameters of an import demand function are estimated in order to quantitatively assess the magnitude of impact of changes in import taxation on the volume of imports (and of trade). The theoretical framework includes three goods, namely, exportable, importable and non-traded goods. The demand for imports depends upon the prices of importable goods compared to non-traded goods (P_m/P_h), the prices of importable compared to exportable goods (P_m/P_x), the level of real income (Y), and real aggregate expenditure (Y^e):

$$\ln M = a_0 + a_1 \ln \left(\frac{P_m}{P_h} \right) + a_2 \ln \left(\frac{P_m}{P_x} \right) + a_3 \ln Y + a_4 \ln Y^e \quad (1)$$

where a_1 and a_2 are <0 ,

$a_3 \leq 0$ or >0 , and

$a_4 > 0$

The a_3 coefficient can be positive or negative depending on the bias of the effect of changes in real output on imports. This bias is a function of the sources of economic growth under constant relative prices of goods and technology, and of the factor intensity in the production of goods. If an increase in real output is associated with an increase in the domestic output of import-competing activities at constant relative prices and

expenditure, then, the coefficient is going to be negative. If, on the other hand, there is a reduction in real output of import-competing activities, then the coefficient is going to be positive. The a_4 coefficient is positive as an increase in real aggregate expenditure is expected to increase the demand for imported goods.

The $\ln(P_m/P_h)$ variable is endogenously determined by the condition of clearance in the market for non-traded goods:

$$Y_h^d = D\left(\frac{P_m}{P_h}, \frac{P_m}{P_x}, Y, Y^e\right) \quad (2)$$

$$Y_h^s = S\left(\frac{P_m}{P_h}, \frac{P_m}{P_x}\right) \quad (3)$$

The equilibrium condition in the market for non-traded goods implies that:

$$\frac{P_m}{P_h} = g\left(\frac{P_m}{P_x}, Y, Y^e\right) \quad (4)$$

The g function can be specified as follows:

$$\ln\left(\frac{P_m}{P_h}\right) = b_0 + b_1 \ln\left(\frac{P_m}{P_x}\right) + b_2 \ln Y + b_3 \ln Y^e \quad (4')$$

$$0 < b_1 < 1$$

where $0 \leq b_2$ or $b_2 \geq 0$, and

$$b_3 < 0$$

Replacing (4') in (1), we get the reduced form of the import function to which an error term (ε) has been added:

$$\ln M = A_0 + A_1 \ln\left(\frac{P_m}{P_x}\right) + A_2 \ln Y + A_3 \ln Y^e + \varepsilon \quad (5)$$

where $A_0 = a_0 + a_1 b_0$

$$A_1 = a_2 + a_1 b_1$$

$$A_2 = a_3 + a_1 b_2$$

$$A_4 = a_4 + a_1 b_3$$

The coefficients of equation (5) take thus into account not only the elasticities of the demand for imports with respect to prices of importable goods compared to exportable goods, to real output and to aggregate expenditure of equation 1, but also the impact of these variables on the relative price of importable goods compared to non-traded goods.

To quantify the impact of trade policies on the volume of trade, equation (5) should be estimated along with an export supply function. However, the export supply and import demand equations are not independent. The reason for this is that the balance of trade at world prices is equal to the excess of aggregate demand (Y^e) over aggregate supply (Y) when the market for non-traded goods is in equilibrium. If $Y^e - Y$ is determined by monetary and fiscal policies, and by foreign capital inflows, the effects of, for instance, an increase in import tariffs is going to have symmetric effects on imports and exports. In other words, commercial policies that reduce the volume of imports also reduce the volume of exports and one thus needs to estimate either the import demand function or the export supply equation as they are not independent.

The import function is estimated with quarterly data covering the period 1970:1-1999:4. If the variables of equation (5) are non stationary, then each variable would show no tendency to return to their long run equilibrium levels. The existence of a long-run relationship as specified in equation (5) is then tested by checking the behavior of its residuals. If the hypothesis that these residuals follow a unit root process cannot be rejected, then there would not be any long-run equilibrium relationship between our variables as any departure from equilibrium would persist forever. If, on the other hand, the residuals were integrated of order zero, then a linear combination of the variables is also stationary and a long-run relationship among the variables can be estimated.

We test the order of integration of the variables using the Phillips-Perron (PP) unit-root test. The results of the test of the seasonally adjusted variables are presented in Table 2. Based on the results of the PP, the hypothesis that the variables are non-stationary cannot be rejected.

Table 2
Phillips-Perron Test for Unit Roots-1970:1-1999:4

First difference of the Variables	PP Statistic	Marginal probability of the test
$\ln M^{(1)}$	-0.67	0.85
$\ln (P_m/P_x)^{(2)}$	-0.05	0.70
$\ln Y^{(1)}$	-0.63	0.86
$\ln Y^e^{(1)}$	-0.88	0.79

Notes: (1) includes a constant.
(2) includes no deterministic terms

The results of the estimation of the equation in the levels of the variables are presented below (the t- statistics are the numbers in parenthesis):

$$\ln M = -55.08 - 0.37 \ln(P_m/P_x) - 1.93 \ln Y + 5.10 \ln Y^e \quad (6)$$

(-22.22) (-6.80) (-2.71) (8.52)

$$R^2 = .96 \quad D-W = 0.73 \quad Q(1) = 47.2 \quad Q(4) = 100.5 \quad \text{Breusch-Godfrey LM test}(1) = 47.2 \quad \text{Breusch-Godfrey LM test}(4) = 52.7$$

The least squares estimates show significant autocorrelation of residuals as indicated by the Ljung-Box Q-statistic with one and four lags of the residuals. The Breusch-Godfrey LM tests with one and four lags of the residuals also indicate that the hypothesis of autocorrelation of residuals cannot be rejected. As is well known, autocorrelation of residuals yields biased estimators of the standard errors of the coefficients and the tests of hypotheses based on these estimators are not reliable. Furthermore, the PP test statistic is calculated in -5.37 and the hypothesis of a unitary root is strongly rejected; thus, the variables are cointegrated and, consequently, a long-run relationship can be estimated with the data.

The presence of autocorrelation of residuals can be an indication that there is a short run adjustment of the variables to reach their long run equilibrium values. The short run adjustment process takes the form of an equilibrium correction representation of an autoregressive-distributed lag model. This representation is a more general specification of the serial

autocorrelation of residuals than the Cochrane-Orcutt method. Let this representation take the following form:

$$Y_t = \alpha_0 + \sum_{j=0}^n \alpha_{1j} X_{t-j} + \sum_{j=1}^n \alpha_{2j} Y_{t-j} \quad (7)$$

By estimating (7) by least squares, the long-run coefficient of vector X can be calculated from the short-run coefficients as follows:

$$\beta_1 = \frac{\sum_{j=0}^n \alpha_{1j}}{1 - \sum_{j=1}^n \alpha_{2j}} \quad (8)$$

Estimating equation (7) usually involves a high degree of multicollinearity among the variables so that their coefficients cannot be estimated with precision. That is, if X_t and X_{t-1} are correlated, then ΔX_t and X_t can be nearly orthogonal. Thus, equation (7) can be reparameterized to get the following specification³:

$$\Delta Y_t = \phi_0 + \sum_{j=0}^{n-1} \phi_{1j} \Delta X_{t-j} + \sum_{j=1}^{n-1} \phi_{2j} \Delta Y_{t-j} + \gamma (Y_{t-1} - \beta_1 X_{t-1}) \quad (9)$$

where γ can be interpreted as the average speed of adjustment of Y towards its stationary state, and β_1 is the long-run coefficient of the variable X .

In our case, the demand for imports is specified as follows:

$$\begin{aligned} \Delta \ln M_t = & \phi_0 + \sum_{j=1}^{n-1} \phi_{1j} \Delta \ln \left(\frac{P_m}{P_x} \right)_{t-j} + \sum_{j=1}^{n-1} \phi_{2j} \Delta \ln Y_{t-j} + \sum \phi_{3j} \Delta \ln Y_{t-j}^e + \\ & + \gamma \left(\ln M_{t-1} - \beta_1 \ln \left(\frac{P_m}{P_x} \right)_{t-1} - \beta_2 \ln Y_{t-1} - \beta_3 \ln Y_{t-1}^e \right) \end{aligned} \quad (10)$$

Equation (10) is estimated by ordinary least squares instead of constrained least squares because it is a reparameterization of equation (7) and, thus, no constraints on the parameters need to be tested. The results of the estimation of equation (10) with data covering the period 1970:2-1999:4 and the number of lags equal to one in the levels of the variables are presented in Table 3 below.

³ See Annex B for the derivation of equation (9).

Table 3
Estimates of the Import Demand Function

Variables	1970 :2-1999 :4	1970 :2-1984 :4	1985 :2-1999 :4
C	-16.82 (-4.62)	-15.61 (-3.01)	-2.85 (-0.49)
$\ln(P_m/P_x)$	-0.18 (-1.82)	-0.24 (-1.71)	-0.40 (-2.70)
$\ln Y$	-2.54 (-2.75)	-1.50 (-1.20)	-4.61 (-3.14)
$\ln Y^e$	3.79 (5.28)	3.47 (3.76)	4.23 (3.63)
Υ	-0.27 (-4.50)	-0.35 (-3.24)	-0.23 (-3.03)
$\ln(P_m/P_x)_{t-1}$	-0.48 (-3.68)	-0.31 (-1.99)	-1.89 (-2.80)
$\ln Y_{t-1}$	-0.56 (-0.31)	-2.32 (-1.20)	-3.30 (-1.13)
$\ln Y_{t-1}^e$	4.12 (2.75)	4.94 (3.76)	4.64 (1.91)
R^2	0.51	0.54	0.59
Q(4)	12.32 (0.015)	3.39 (0.49)	3.22 (0.52)
Breusch-Godfrey LM Test (4 lags)	15.66 (0.004)	4.75 (0.31)	4.12 (0.39)
Number of observations	119	59	59

Notes: The numbers in parenthesis below the estimates of the coefficients are the t-statistics. The numbers in parenthesis in the rows of the Q- and of the Breusch-Godfrey statistics are the marginal probabilities of the values of the tests statistics.

The estimates of equation (10) for the whole period indicate that the hypothesis of uncorrelated residuals can be rejected according to the Q and the Breusch-Godfrey tests. As is well known, serial autocorrelation of residuals produce inconsistent estimators of the parameters of the distributed autoregressive representation. Residual autocorrelation can be the result of either the lack of stability of parameters over time, of

misspecification of the model, or of the lag length. A Wald test of the stability of parameters over time is performed by estimating equation (10) for two non-overlapping sub samples, namely, one covering the period 1970:2-1984:4 and the other for the period 1985:2-1999:4⁴ with the number of lags equal to one in the level of the variables⁵.

For the two sub samples, the PP tests indicate that the hypothesis that the variables are not stationary cannot be rejected. The results of the tests are presented below (see table 4). Furthermore, the hypothesis that the residuals of the least squares estimates of equation (5) follow a unit root autoregressive process in both sub samples can be rejected at 1 percent.

Table 4
Phillips-Perron Unit-Root Tests

First difference of the Variables	PP Statistic		Marginal probabilities	
	1970:2 - 1984:4	1985:2 - 1999:4	1970:2 - 1984:4	1985:2 - 1999:4
$\ln M^1$	0.14	-0.60	0.72	0.86
$\ln (P_m/P_x)$	-1.45	-1.62	0.55	0.47
$\ln Y$	-1.62	-0.53	0.47	0.88
$\ln Y^e$	-1.74	-0.73	0.41	0.83

Notes: (1) no deterministic terms for data of the 1970:84 sub sample.

According to the values of the Ljung-Box's Q and of the Breusch-Godfrey's statistics, the hypothesis of white noise residuals cannot be rejected for the two sampling periods. The values of the White test for heteroskedastic disturbances are 26.6 for the sample covering the period 1970:2-1984:4 and 27.7 for 1985:2-1999:4 with marginal significance levels of 27 and 9 percent, respectively, and the null hypothesis of homoskedasticity cannot thus be rejected. However, the autocorrelations

⁴ An alternative method is to introduce dummy variables to capture changes in the coefficients and in the constant term. As the model estimated under the hypothesis of no changes in the parameters show autocorrelation of residuals, the estimates of the coefficients of the restricted model are inconsistent and this invalidates the test of stability. In the case of the Wald test, there is no need to estimate the model for the whole sample.

⁵ The lag length is determined on the basis of the Q- statistic to test if the estimated residuals follow a white noise stochastic process.

of the squared residuals of the second sub sample suggest the existence of autoregressive conditional heteroskedasticity as the values of the Q(5) and Q(6) are 12.2 and 12.8, respectively, and they are statistically significant at 5 percent significance level (see below). Furthermore, if the White test statistic is calculated without cross terms, the hypothesis of homoskedastic disturbances is rejected by the data with a p-value of about 3 percent.

Autocorrelation of squared residuals:

	Lags							
	1	2	3	4	5	6	7	8
Q-stat	1.162	1.189	1.580	2.749	12.121	12.829	12.855	13.091
(p-values)	(0.281)	(0.552)	(0.664)	(0.601)	(0.033)	(0.046)	(0.076)	(0.109)

The import function is re-estimated with data of the sub sample 1985:2-1999:4 by specifying the behavior of the conditional variance as an ARCH process of order one⁶. The results are as follows:

$$\Delta \ln M_t = -2.793 - 0.189 \left[\ln M_{t-1} - (-1.991) \ln \left(\frac{Pm}{Px} \right)_{t-1} - (-3.235) \ln Y_{t-1} - 4.725 \ln Y_{t-1}^e \right]$$

(-0.837) (-4.378) (-3.394) (-2.090) (3.502)

$$-0.153 \left(\Delta \ln \left(\frac{Pm}{Px} \right)_t \right) - 3.706 (\Delta \ln Y)_t + 5.010 (\Delta \ln Y^e)_t$$

(-1.631) (-5.304) (7.336)

$$\sigma_t^2 = 0.0002 + 1.6055 e_{t-1}^2$$

(0.3480) (3.0880)

The covariance matrix of the coefficients is going to be consistently estimated if the residuals are normally distributed. To test this hypothesis, the Jarque-Bera (*J-B*) test statistic is calculated for the two sub samples. The *J-B* test statistic has an asymptotically chi-squared distribution and its good

⁶ Alternative specifications of ARCH models were estimated and the ARCH(1) is chosen on the basis of the values of the Schwarz and the Akaike criteria, and also on the basis of the value of the likelihood ratio test statistic. The three criteria reject higher dimensional ARCH models.

performance is highly dependent upon the use of empirical significance points as a result of slow convergence in the distribution to the chi-squared value. C. Urzúa (1996) suggests substituting the asymptotic means and variances of the standardized third and fourth moments by their exact means and variances to substantially improve the asymptotic convergence of the test. The adjusted *J-B* (*AJB*) test statistic is then calculated as follows:

$$AJB = \frac{(\hat{b}_1^{1/2})^2}{\text{var}(\hat{b}_1^{1/2})} + \frac{[\hat{b}_2 - E(\hat{b}_2)]^2}{\text{var}(\hat{b}_2)}$$

where $\hat{b}_1^{1/2}$ and \hat{b}_2 are the estimated standardized third and fourth moments of the distribution of residuals, $\text{var}(\hat{b}_1^{1/2}) = \frac{6(n-2)}{(n+1)(n+3)}$, $\text{var}(\hat{b}_2) = \frac{24n(n-2)(n-3)}{(n+1)^2(n+3)(n+5)}$, $E(\hat{b}_1^{1/2}) = 0$, and $E(\hat{b}_2) = \frac{3(n-1)}{(n+1)}$.

The Jarque-Bera test statistic is calculated in 1.28 and this value amounts not to reject the null hypothesis of normality of residuals with a marginal significance of about 53 percent and, as indicated earlier, the covariance matrix of the coefficients is consistently estimated. All the coefficients of the import function estimated with the ARCH model are statistically different from zero at the usual significance levels. The autocorrelations of the squared residuals do not show any indication of remaining conditional heteroskedasticity of residuals as shown below.:

Autocorrelation of squared residuals:

	Lags							
	1	2	3	4	5	6	7	8
Q-stat	0.085	1.797	1.811	2.195	2.280	2.974	3.164	3.424
(p-values)	(0.771)	(0.407)	(0.612)	(0.700)	(0.809)	(0.812)	(0.869)	(0.905)

In order to test the null hypothesis of stability of parameters across the two sub samples, the least squares estimators of the coefficients and of their covariance matrix estimated with data of 1970:2-1984:4, and those of

the ARCH model are used to calculate the Wald test statistic⁷. The null hypothesis is strongly rejected by the data as the value of the test statistic, that is asymptotically distributed as $\chi^2(8)$, is calculated in 31.8 and this value has a marginal probability of 0.01 percent. In particular, the value of the Wald statistic to test the stability of the three coefficients of the error correction term is 9.70 and this amounts to reject the null hypothesis with a marginal significance of 2.1 percent.

According to the estimates of the long-run elasticity of the demand for imports with respect to the relative price of importable goods, this parameter has increased in absolute value from -0.30 during the period 1970:2-1984:4 to about -2.00 during the second sampling period. The calculated Wald test statistic of 7.68 to test the hypothesis of equality of these two coefficients indicates that the null hypothesis is strongly rejected at 0.5 percent significance level.

The long-run coefficients of the logarithm of real income ($\ln Y$) and of aggregate expenditure ($\ln Y^e$) estimated with data covering the period 1985:2-1999:4 are negative and positive, respectively. The negative coefficient of real output would indicate that, during this period, economic growth has been biased against the volume of international trade on the supply side.

The estimates of the coefficient of adjustment γ for the two sub samples are statistically different from zero and less than one. This is an additional confirmation that the variables are cointegrated as short-run disequilibria tend to be eliminated in the long-run. To test the hypothesis of stability of this parameter over time, the value Wald test statistic is calculated in 1.90 and this hypothesis cannot be rejected with a significance level of 17 percent. As the coefficient of adjustment is interpreted as the average speed of adjustment of disequilibrium towards the stationary volume of imports, the estimate of this parameter with data of the sub sample of 1985:2-1999:4 of about -0.20 means that the stationary state is achieved in about five quarters.

In the estimation of the demand for imports in its dynamic specification for the years covering the period 1985:2-1994:4, it has implicitly been assumed that the variables $\Delta \ln(P_m/P_x)$, $\Delta \ln Y$, and $\Delta \ln Y^e$ are weakly exogenous.

⁷ The Wald statistic is $(b_0 - b_1)'(V_0 + V_1)^{-1}(b_0 - b_1)$ and is asymptotically distributed as χ^2 under the assumption of normality and independence of the estimates of the coefficients with degrees of freedom equal to the number of parameters for which stability is tested. The vectors b_0 and b_1 are those containing the coefficients estimated with data of the two sub samples, and V_0 and V_1 are the estimates of the coefficient covariance matrixes.

As Harbo, Johansen, Nielsen, and Rahbek (1998) demonstrate, if these assumptions are violated, the estimators of the parameters of the cointegrating vector based on the partial model such as (10) are inefficient and the estimator of the speed of adjustment could be inconsistent. Harbo, et. al. (1998) recommend estimating the cointegrating vector in the partial model and then testing the null hypothesis of weak exogeneity by regressing $\Delta \ln(P_m/P_x)$, $\Delta \ln Y$, and $\Delta \ln Y^e$ on lagged changes of all the variables, a constant, and on the empirically derived cointegrating vector such that the errors appear iid Gaussian⁸.

The $\Delta \ln(P_m/P_x)$ variable is regressed against a constant, five lags of all the first differences in the variables, and the estimated cointegration vector in order to obtain iid Gaussian residuals. The calculated *t*-statistic under the null hypothesis of zero coefficient of the cointegrating vector is calculated in 0.51 with a marginal probability of 0.62 percent that amounts not to reject the null. The $\Delta(\ln Y)$ variable is regressed against a constant, three lags of the first differences of all the variables of the system, and the estimated cointegrating vector. The null hypotheses of normally distributed, non correlated, and homoskedastic residuals cannot be rejected under this lag specification. The calculated *t*-statistic calculated under the null hypothesis of weak exogeneity is 0.51 and this value cannot reject the null with a marginal significance of about 0.61 percent. Finally, the variable $\Delta(\ln Y^e)$ is regressed against a constant, three lags of the first differences of all the variables of the system, and the estimated cointegrating vector. The ARCH estimates indicate that the hypotheses of normally distributed and iid residuals cannot be rejected. The null hypothesis of weakly exogenous expenditure cannot be rejected as the calculated *t*-statistic has a marginal probability of 93 percent. Thus, the assumptions underlying the estimates of the equation (10) for the sample 1985:2-1999:4 cannot be rejected by the data.

Estimating the equation (10) by the Kalman filtering algorithm can provide additional insights into the pattern of changes of the coefficients over time. The model to be estimated is, in general, the following:

⁸ The alternative is, of course, to estimate the full system and testing the null hypotheses of weak exogeneity. The estimates of the full system with data of the sub sample 1985-1999 show that the estimated residuals are not normally distributed and heteroskedastic. These problems could be overcome by including dummy variables to deal with outliers but these variables affect the distribution of the trace statistics. This is the reason by which the estimation procedure indicated in the text has been chosen. The cost of this procedure is that the cointegrating rank of the system cannot be tested.

$$\Delta y_t = X_t \beta_t + u_t,$$

where y_t is $\ln M_t$, X_t includes $\ln M_{t-1}$, $\ln \left(\frac{P_m}{P_x} \right)_{t-1}$, $\ln Y_{t-1}$, $\ln Y_{t-1}^e$,

and the contemporaneous changes in $\ln \left(\frac{P_m}{P_x} \right)$, $\ln Y_t$, and $\ln Y_t^e$, u_t are $N(0, n_t)$ and the β_t coefficients are assumed to follow a random walk:

$$\beta_t = \beta_{t-1} + v_t,$$

where v_t is distributed as $N(0, \Omega_t)$, and $E(u_t v_t) = 0$.

If there is an estimate of β_{t-1} using all the information through $t-1$ ($\beta_{t-1/t-1}$) and its covariance matrix Σ_{t-1} , then the updated estimate is:

$$\beta_t = \beta_{t-1/t-1} + S_t X_t' (X_t S_t X_t' + n_t)^{-1} (y_t - X_t \beta_{t-1/t-1})$$

where

$$S_t = \Sigma_{t-1} + \Omega_t$$

$$\Sigma_t = S_t + S_t X_t' (X_t S_t X_t' + n_t)^{-1} X_t S_t$$

In order to update the estimates, the following information is needed:

- the initial vector of coefficients, $\beta_{0/0}$
- the initial covariance of the coefficients, Σ_0
- the variance of u_t , namely, n_t , and
- the variance of the change in the β coefficients, Ω_t .

In this paper, the initial estimates of the β coefficients, n and of Σ are given by the least squares estimates of (10) for the sub sample covering the period 1970:2-1982:4. The variance of u_t is assumed constant through the sample and equal to the estimated variance of innovations of equation (10) with data of the sub sample, and Ω is assumed to be proportional to Σ_0 , that is, $\lambda \Sigma_0$. The parameter λ is calculated by iterating different values over the interval (0,1) and choosing the value that maximizes the logarithm of the likelihood function that is equal to (disregarding constant terms)⁹:

⁹ For the derivation of the likelihood function, see Doan, T., R. Litterman, and C. A. Sims (1984) page 10.

$$\ln l = -0.5 \left[\sum_1^T \ln \sigma_t^2 + T \ln \left(T^{-1} \sum_1^T \ln \frac{u_{t|t-1}^2}{\sigma_t^2} \right) \right]$$

where $\sigma_t^2 = n_t + X_t (\Sigma_{t-1} + \Omega_t) X_t'$ is the variance of the one-step ahead forecast of Δy_t conditional upon previous information.

With the initial estimates given by the sub sample 1970:2-1982:4, the value of λ that maximizes the likelihood function is calculated in 0.35. The time path of the estimated long run coefficients of the demand for imports throughout the period 1983:1-1999:4 are reported in figures 1-4. The time paths of the recursively estimated coefficients tend to confirm the earlier finding of instability over time. In particular, the recursively estimated price elasticity of demand for imports increases over time in absolute value from the initial estimate of -0.25 to a range from -0.28 to -0.60. The output and expenditure elasticities are also subject to fluctuations over time. There are abrupt changes in the estimates of the β 's and in the γ coefficients that start taking place in mid-1991 that persist over the subsequent years. These changes in the parameters coincide with the announcement and the starting of the implementation of the so-called Convertibility Plan in the second quarter of 1991 that introduced deep economic reforms in the Argentinean economy that could have changed agents' expectations about the prospects of growth of the economy and induced a substantial contemporaneous increase in the volume of imports.



Fig. 1: Recursive estimates of γ . Initial estimates: 1970:2-1982:4. $\lambda = .35$, $n=0.00756$.

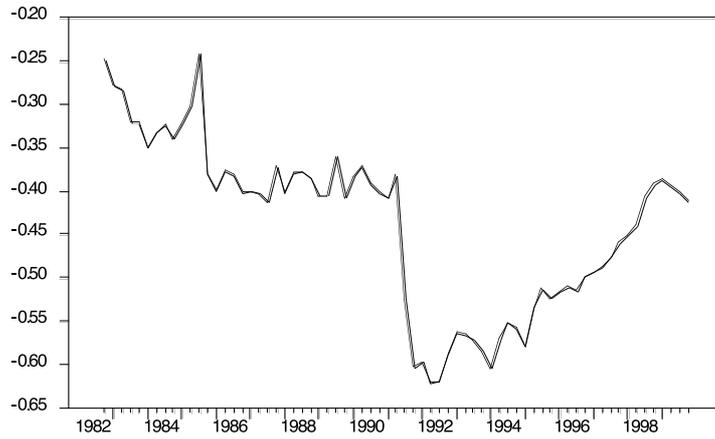


Fig. 2: Recursive estimates of β_1 . Initial estimates: 1970:2-1982:4. $\lambda = .35, n = 0.00756$.

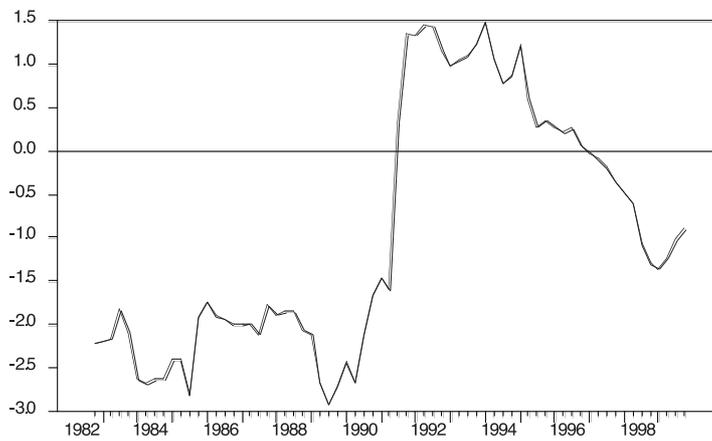


Fig. 3: Recursive estimates of β_2 . Initial estimates: 1970:2-1982:4. $\lambda = .35, n = 0.00756$.

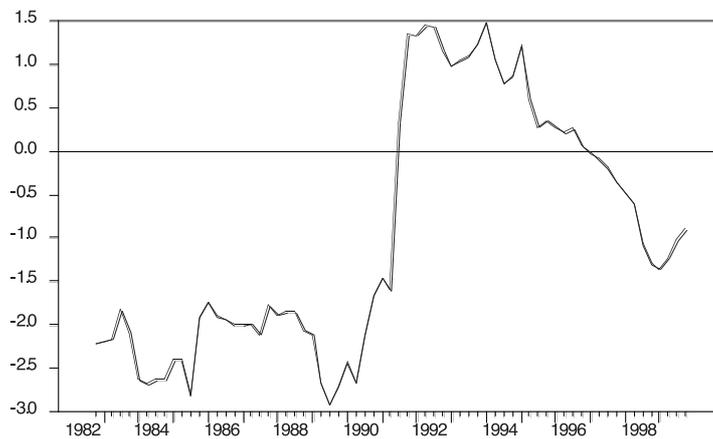


Fig. 4: Recursive estimates of β_3 . Initial estimates: 1970:2-1982:4. $\lambda = .35$, $n = 0.00756$.

As this paper is aimed at assessing the contribution of trade liberalization policies to the increase in the volume of imports during the sub sample 1995-1999 compared to 1985-1989, the behavior of the long-run parameters over this period is further analyzed to inquiry over their stability over time. Recursive estimates of these parameters are obtained over the period 1991:1-1999:4 with data covering the time period 1985:2-1990:4 used to obtain the initial estimates. The value of λ that maximizes the value of the likelihood function is 0.005. The recursive estimates are reported in figures 5-8 and they show a similar pattern of behavior as before with abrupt changes in the coefficients after mid-1991. The results also indicate that the recursive estimates of the price elasticity increasing in absolute value from the initial estimate of -0.36 to a range from -2.12 to -1.76 over 1995-1999. The estimates obtained earlier with data covering the period 1985:2-1999:4 by least squares (-1.89) and by the ARCH method (-1.99) are within this range. The output and expenditure elasticities are subject to substantial changes over time. These changes might be the result of innovation outliers following the launching of the reform plan of 1991 that affect subsequent estimates of the parameters through the dynamics of the model and not necessarily the result of structural changes.

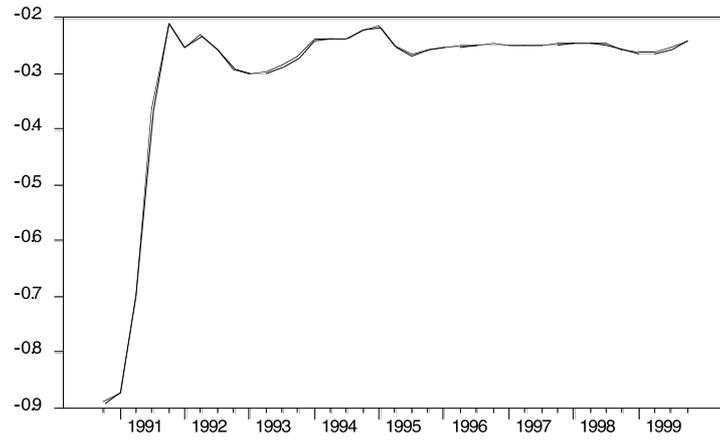


Fig. 5: Recursive estimates of γ . Initial estimates: 1985:2-1990:4. $\lambda = .005$, $n = 0.00256$.

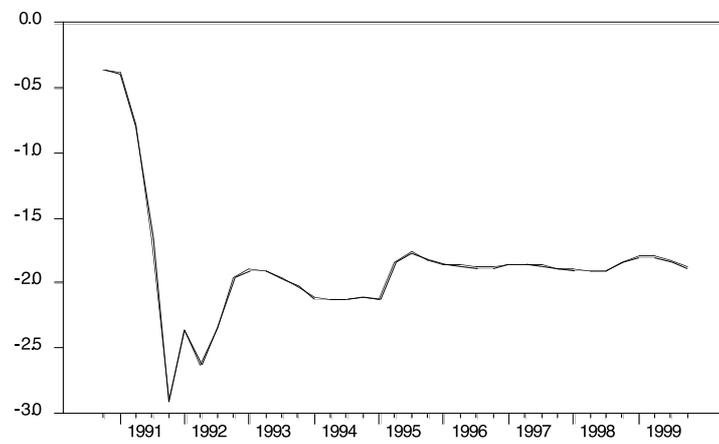


Fig. 6: Recursive estimates of β_1 . Initial estimates: 1985:2-1990:4. $\lambda = .005$, $n = 0.00256$.

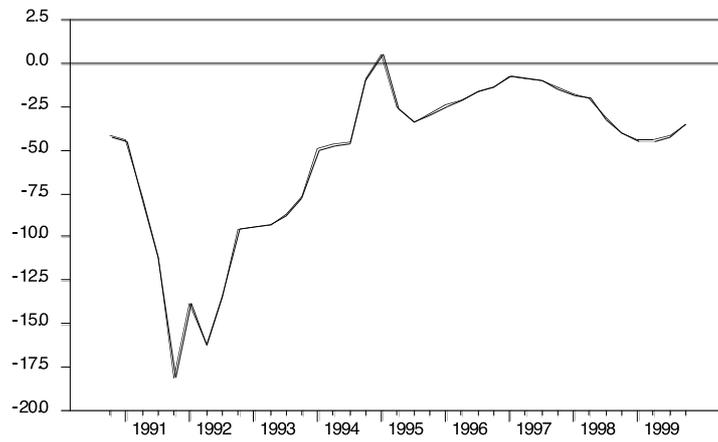


Fig. 7: Recursive estimates of β_2 . Initial estimates: 1985:2-1990:4. $\lambda = .005$, $n = 0.00256$.

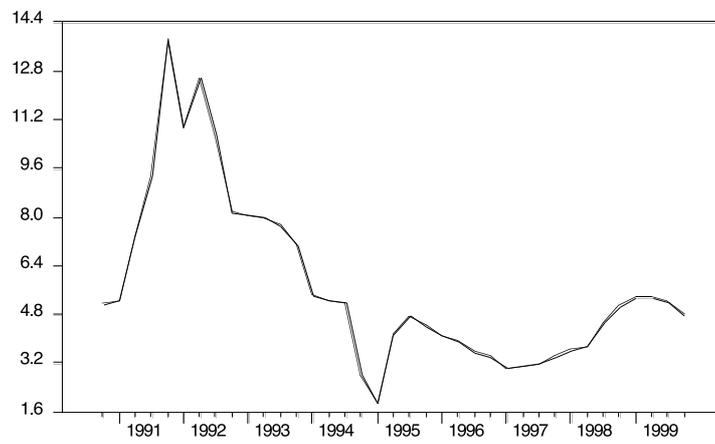


Fig. 8: Recursive estimates of β_3 . Initial estimates: 1985:2-1990:4. $\lambda = .005$, $n = 0.00256$.

III. A QUANTITATIVE ASSESSMENT OF THE TRADE POLICIES OF THE 90s

In this section, the increase in the 1995-1999 average quantity of imports over the average of 1985-89 is decomposed according to its sources using the ARCH estimates of the long-run parameters of the import demand function for the period 1985:3-1999:4 along with the measures of the equivalent uniform import tax rates for 1985-89 and for 1995-99:

- (i) Aggregate expenditure and real output effects holding constant the average 1985-89 domestic relative prices of importable vis-à-vis exportable goods and the average external terms of trade for this sub sample;
- (ii) Changes in the external terms of trade holding constant aggregate expenditure and real output at their 1995:99 levels and the commercial policies of 1985:89;
- (iii) Changes in commercial policies holding constant the external terms of trade and aggregate expenditure and real output at their average values of 1995-99.

The results show that the trade liberalization policies of the 90s are the main factor behind the increase in the volume of imports. The estimated changes in the equivalent uniform import tax from about 80 percent in the second half of the 80s to about 20 percent in the second half of the 90s and the estimated price elasticity of the demand for imports of about -2.00 resulted in an increase in the volume of imports of about 130 percent if the average real expenditure, holding constant real output and the external terms of trade at their average values of the period 1995-99. This increase in the quantity of imports is estimated to account for about 68 percent of the actual increase in the total volume of imports¹⁰. On the other hand, changes in real expenditure and in real output account for about 23 percent of the increase in the volume of imports. The remaining 9 percent is estimated to be accounted for by changes in the external terms of trade.

¹⁰ The average quantity of imports increased from 29.2 in 1985-89 (1993=100) to 155.5 in 1995-99, representing an increase of about 430 percent.

IV. FINAL REMARKS

According to our estimates, in-depth trade liberalization policies were implemented during the 90s as indicated by the reduction in the estimated equivalent uniform tax on international trade of about 60 percentage points compared to the average tax of the period 1985-89. The equivalent uniform tax rate on the volume of international trade was reduced from about 80 percent in the second half of the 80s to about 20 percent in the 90s.

The estimates provided in this paper suggest that if the effects of trade liberalization policies of the 90s are isolated from changes in aggregate expenditure and in real output, and from changes in external terms of trade, then, these policies have led to a substantial increase in the quantity of imports. According to the estimates, the increase in the volume of imports in response to trade liberalization policies would represent about 68 percent of the total increase in the volume of imports over the period of more severe trade restrictions of the late 80s. The remaining 32 percent would be attributed mostly to increases in aggregate expenditure and real output (23 percent), and to more favorable external terms of trade (9 percent).

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ANNEX A: DATA DESCRIPTION

The volume of import (M) is an index of the quantity of imports (1993=100). For the period 1993:1-1999:4, the data is from the Institute of National Statistics and Census (INDEC). Data for the periods 1970:1-1999:2 was obtained from national accounts at constant prices. These data were linked using the rate of change of the quarterly value of imports at constant prices. (P_m / P_x)

Relative prices of imports are defined as ratio of the wholesale price index of imports to the price of exports from 1960:1 until 1999:4. P_x is a weighted average of the prices of agricultural and food manufacturing (excluding beverages). The weights of the components of P_x are those of the 1993-1999 wholesale price index, in which prices of agricultural and food manufacturing receive weights of 47 and 53 percent respectively. Alternatively, a moving average of the two prices was constructed using the shares of exports agricultural and manufactures of agricultural origin in the total of these exports. The results of the estimation are very robust to changes in the definition of relative prices. The price ratio is extrapolated backwardly to the year 1939 with the ratio of the wholesale price of imported commodities to that of agricultural commodities. For 1935-1939, the wholesale price of non-agricultural products is used as a proxy for imported commodities.

Real GDP (Y) is the quarterly real gross domestic product at 1993 prices. The same procedure to construct the data of imports is followed.

Real expenditure (Y^e) is the quarterly aggregate demand at constant 1996 prices.

External Terms of Trade data are from Mundlak, Y., D. Cavallo, and R. Domenech, "Estadísticas de la evolución económica de Argentina: 1913-1984", *Estudios*, IX, N° 39, Julio-Septiembre 1986. These data are extrapolated with estimates of the World Bank from 1984 until 1988 and with INDEC figures afterwards.

ANNEX B: DERIVATION OF EQUATION (7)

Consider the following autoregressive representation of a variable Y:

$$(A1) \quad Y_t = a0 + a1Y_{t-1} + a2Y_{t-2} + a3Y_{t-3} + b0X_t + \\ + b1X_{t-1} + b2X_{t-2} + b3X_{t-3} + \varepsilon_t$$

Subtracting Y_{t-1} from both sides of (A1), and adding and subtracting specific terms, the following equation is obtained:

$$(A2) \quad Y_t - Y_{t-1} = a0 + a1Y_{t-1} - Y_{t-1} + (a2Y_{t-1} - a2Y_{t-1}) + \\ + (a3Y_{t-1} - a3Y_{t-1}) + a2Y_{t-2} + \\ + (a3Y_{t-2} - a3Y_{t-2}) + a3Y_{t-3} + \\ + b0X_t + (b0X_{t-1} - b0X_{t-1}) + b1X_{t-1} + \\ + (b2X_{t-1} - b2X_{t-1}) + (b3X_{t-1} - b3X_{t-1}) + \\ + b2X_{t-2} + (b3X_{t-2} - b3X_{t-2}) + b3X_{t-3} + \varepsilon_t$$

Rearranging (A2) and collecting terms, the following equations are obtained:

$$(A3) \quad \Delta Y_t = a0 + \left(\sum_{j=1}^3 a_j - 1 \right) Y_{t-1} + \sum_{j=0}^3 b_j X_{t-1} - \\ - (a2 + a3) \Delta Y_{t-1} - a3 \Delta Y_{t-2} + \\ + b0 \Delta X_t - (b2 + b3) \Delta X_{t-1} - b3 \Delta X_{t-2} + \varepsilon_t$$

$$(A4) \quad \Delta Y_t = a0 + \left(\sum_{j=1}^3 a_j - 1 \right) \left(Y_{t-1} - \left(\frac{\sum_{j=0}^3 b_j}{1 - \sum_{j=1}^3 a_j} \right) X_{t-1} \right) - \\ - (a2 + a3) \Delta Y_{t-1} - a3 \Delta Y_{t-2} + \\ + b0 \Delta X_t - (b2 + b3) \Delta X_{t-1} - b3 \Delta X_{t-2} + \varepsilon_t$$

A generalization of (A4) to more variables is equation (10) in the text.