PROGRAMA DE POSTGRADO EN ECONOMIA
ILADES / GEORGETOWN UNIVERSITY

Documento de Investigación #88
Octubre de 1995

International Price Signals in Agricultural Prices: Do Governments Care?

Jorge Quiroz
Raimundo Soto

Erasmo Escala 1835. Santiago de Chile. Tels: (562) 698-0046 / 696-8286.
Fax: (562) 696-4880
e-mail: economia@uahurtado.cl
www.ilades.cl/economia/publi.htm

INTERNATIONAL PRICE SIGNALS IN AGRICULTURAL MARKETS: DO GOVERNMENTS CARE?

Jorge A. Quiroz
GERENS

Raimundo Soto
ILADES/Georgetown University

October, 1995

Comments and suggestions to a previous draft by Paul McNelis, Tarhan Feyzioglu, XX and Marc Nerlove are gratefully acknowledged. This paper was written while the first author was under the visiting scholar program at The World Bank and the second worked at the Macroeconomics and Growth Division of the World Bank. The opinions expressed here are the sole responsibility of the authors and do not necessarily represent those of The World Bank.

Abstract

Widespread
international price signals. Based on information from 78 countries over the 1966-1991 period,

domestic markets. Unlike previous findings, we conclude that in 60 countries price transmission
takes place. A dynamic maximization model shows these results to be consistent with policy
makers
following international price signals. Results have implications for price instability in world

Keywords: Pricing Policies, Stochastic models, Agricultural markets

JEL classification: Q17, C15, F13

1. Introduction

Almost everywhere in the world, governments intervene intensively in domestic agricultural markets. Using a variety of trade distortions, pricing policies generate a gap between producer and international price levels, creating a "direct" price intervention.[1]This phenomenon, particular not only to recent history,[2] has been extensively documented for both developed and developing countries (Anderson and Hayami, 1986; Krueger, 1992; Schiff and Valdés, 1992).The widespread direct price intervention prevailing in many countries raises the issue of the degree of transmission of international price signals. If a negative supply shock occurs in the international market, prices will rise, reflecting the fact that the good has become relatively more scarce. Direct price intervention, however, may prevent this information from being transmitted to domestic agents making consumption and resource allocation decisions.[3]

The purpose of this paper is to analyze the extent to which new price information arising from international markets is transmitted to domestic agents in different countries. To this end we formulate and estimate, on a country-by-country basis, a dynamic econometric model for domestic agricultural prices conditional on international ones. The model is estimated for 78 countries, covering the last 30 years, and using a sample ranging from four to 15 products per country.The empirical findings are later contrasted to the simulated outcome of a dynamic maximization problem carried over by a hypothetical policy maker. This allows us to infer the policy objectives underlying the observed processes for domestic prices.

There are several reasons why studying the transmission of international price signals is important. First, there is the issue of world price instability. If a reduction in international prices

---

[1]This definition of direct price intervention follows that of Krueger, Schiff and Valdés (1992). It should be distinguished from indirect price intervention, which corresponds to relative price distortion induced by government intervention in non-agricultural markets and exchange rate misalignment.

[2] See Lindert (1989).

[3] It must be noted that the mere existence of direct price intervention does not necessarily imply that price information in international markets is not transmitted to domestic agents.For example, a constant tariff levied on an importable agricultural product, as long as it remains non-prohibitive, will fully transmit any increase (decrease) in international prices into an equivalent percentage change of domestic prices.Hence, in principle, the problem of international transmission of price signals is different from that of protection through direct price intervention.

is not fully transmitted to domestic prices, then reductions in world supply and increases in world demand - which would otherwise occur - will not take place, thereby making the price reduction more acute and prolonged. Hence, on a global scale, significant local-market isolation may induce augmented price fluctuations, a point emphasized by Johnson (1991) in a general context and by Josling (1980) in the specific case of the international wheat market. To some extent, this is a problem of "governments' coordination failure": widespread intervention in local agricultural markets implies increased international price instability, which in turn may induce governments to intervene with increasing trade distortions in order to isolate domestic prices from world market fluctuations.In a counterfactual scenario where no government market intervention prevails, one might conjecture that there would be less price instability.

From a policy perspective, another motivation for this paper relates to ongoing negotiations of trade liberalization agreements (such as NAFTA or MERCOSUR). To date, agriculture, along with a few other sectors, generally enjoys a special protection status. Negotiations usually pursue the objective of tariffication, that is, to convert all prevailing trade intervention instruments into trade taxes or subsidies; if successful, this would imply a complete elimination of quantitative restrictions on trade. With a constant import tariff (tax or subsidy in the case of exports), international price signals would be fully transmitted to domestic prices.This study on the current situation of international price transmission may be useful in unveiling how far are agricultural markets from the tariffication goal being pursued.

Another goal of this study is to examine the political economy underlying government intervention in agricultural markets. The way in which international price signals are allowed to affect (or not) domestic prices, might reveal some of the underlying objectives of government intervention in agriculture, in particular, the relative weight that producer-price stabilization carries in the objective function of policy makers.

The results provide a taxonomy of dynamic responses to price shocks, which we found to vary widely across different countries. While some countries tend to incorporate most international price movements in the very short run (Australia, Canada, New Zealand, and Uruguay among others), in a significant subset of countries the transmission to domestic markets of just half of any given international price change appear to take more than five years. Moreover, in an important number of countries -notably most members of the European Community (EC)- domestic producers appear virtually isolated from international price signals. These results contrast with the conclusion obtained by Mundlak and Larson (1992) from a similar data set, which suggested that most countries in the world followed international price signals in the long run. The major reason for this discrepancy is that in Mundlak and Larson, the relationship between domestic and international prices is analyzed in a static framework with disregard of dynamic considerations. Here, those dynamic considerations are found to be a core component of the transmission mechanism. From the political economy perspective, we found that a plausible interpretation of these results would be that policy makers put considerable weight on the objective of reducing between-years domestic price fluctuations.

The paper is organized as follows.The next section discusses the econometric approach utilized and presents the main empirical findings. Section 3 provides an interpretation of these results from the perspective of a dynamic optimization model in which a policy maker cares about both domestic price fluctuations as well as their relationship to international prices. Finally, Section 4 summarizes the main conclusions and suggests some policy implications.

2. An Econometric Model for International Price Transmission

The Data

Domestic "farmgate" producer prices in local currency were obtained from the FAO database and later converted into US$ using the average nominal exchange rate reported in the IMF database. We selected all agricultural commodities for which more than 20 observations were available; in all countries the data covers the period 1966-1991, although for some goods the information contained missing observations. Only goods which could be roughly identified as "commodities" were selected, thus including all cereals and most staples, but excluding horticulture products and fruits. This selection procedure generated a sample of 78 countries (23 developed and 55 developing countries) with a range of 4 to 15 products per country. Noticeably absent countries are Argentina and Brazil, for which the data base did not contain meaningful information on domestic producer prices, due to their high inflation rates during the eighties.

In the case of international prices at least two proxies can be used. One possibility, followed by Mundlak and Larson (1992), is to define the international price of each good as the world average of unit export values, i.e., the total value of exports of each good divided by its corresponding volume of exports. The alternative, used in this paper, consists in selecting the spot price of the most important free market in the world for each commodity. A free spot-market price reflects more appropriately the opportunity cost any given country has for its agricultural production. Average unit export-values may differ strongly from free-market prices --and from the real opportunity cost of goods-- as many contractual agreements, like those prevailing between the U.S. and Europe with some LDCs, strongly influence average export prices of some of the latter countries.[4] Appendix 1 describes the sources for international prices for each agricultural commodity included in the sample and the coverage of domestic prices.

A Static Approach

We first follow the simplest approach to analyzing the data, which is to run a static panel regression of domestic prices as a function of international ones. Let p_{it} denote the log of the domestic price of commodity **i** in year **t** and p^*_{it} its international counterpart, both expressed in US dollars. For each country the following econometric model is estimated:

$$p_{it} = \alpha_i + \phi p^*_{it} + \epsilon_{it} \qquad \textbf{(1)}$$

[4] The choice between average export prices as opposed to prices in specific spot markets makes an important difference in highly intervened commodities, where trade policies in countries like the U.S. and the EC distort average export prices of many LDCs, so that they no longer correspond to the real (marginal) opportunity cost. Krueger and Duncan (1993) present a detailed discussion of huge price differentials in the case of sugar as arising from the presence of country specific quotas and administrative restrictions. At its peak in 1985, the price of sugar in the U.S. was five times higher than world market prices.

which corresponds to a fixed-effects panel-data regression. Under the assumption that the error ϵ_{it} is an innovation at time **t**, the estimate of Φ provides the elasticity of domestic to international prices; for a full transmission of international price signals a value of Φ close to one should be expected. Note that we allow the intercept (α) to be commodity-specific. The results of this first exercise are recorded in Table 1.

At first glance, the results in Table 1 seem to suggest that a substantial degree of transmission of price signals takes place in most countries. For an overwhelming majority of them the models display high R^2 and Φ estimates significant and close to one. Furthermore, and contrary to the evidence that EC countries tend to intervene more in agricultural markets than other developed countries such as Australia or New Zealand (Anderson and Hayami, 1986), the evidence of the Φ estimates and the goodness of fit does not allow to conclude that EC countries perform any different than those countries in terms of transmitting international price signals. The results in Table 1 would seem to imply that government intervention in agricultural markets carries no consequence for the international price transmission process. Indeed, this is the type of evidence that led Mundlak and Larson (1992) to conclude that domestic agricultural prices tend to move closely in line with international ones.

However, the results do present a troublesome feature because the errors of the estimated equations cannot be truly described as "innovations." It can be seen that most Durbin-Watson (DW) statistics are extremely low, in most cases lower than the R^2, suggesting the possibility that these estimates may correspond to a 'spurious relationship', in the terminology of Granger and Newbold (1974). There is not a single case in all 78 panel estimations in which the DW statistic would be acceptable at standard confidence levels.[5]

Table 1
Static Panel Data Models

$$P_{it} = \alpha_i + \Phi P^*_{it} + \epsilon_{it}$$

Country	Φ	$t(\Phi)$	R^2	DW

[5] Mundlak and Larson (1992) do not report the DW statistics of their estimations which, unlike our estimations, use average unit export-values for proxies of international prices. Nevertheless, re-estimating the models in table 1 with this alternative proxy produces exactly the same symptoms: high R^2 and Φ's close to one but also low values for the DW statistics. Results are available upon request.

Angola	0.91	10.4	0.80	0.32
Australia	0.75	28.0	0.99	1.02
Austria	0.95	13.7	0.96	0.64
Bangladesh	0.53	9.0	0.96	1.44
Belgium	0.84	14.0	0.95	0.61
Burkina Faso	0.72	11.1	0.92	0.55
Burundi	0.96	19.9	0.95	0.62
C.African Rep.	0.78	9.2	0.86	0.50
Cameroon	0.85	13.5	0.82	0.42
Canada	0.77	17.5	0.86	0.86
Chile	0.75	8.0	0.80	0.87
Colombia	0.57	13.8	0.93	0.44
Costa Rica	0.67	19.4	0.96	0.64
Cote d'Ivoire	0.94	15.9	0.83	0.58
Cyprus	0.85	16.2	0.93	0.47
Denmark	1.08	13.5	0.96	0.84
Dominican Rep.	0.67	11.3	0.92	0.69
Ecuador	0.61	12.2	0.92	0.56
Egypt	0.66	11.4	0.90	0.59
Ethiopia	0.93	15.9	0.90	0.51
Finland	0.90	9.1	0.91	0.35
France	0.85	18.0	0.96	0.85
Germany	0.77	14.1	0.97	0.93
Ghana	1.08	8.8	0.74	0.42
Great Britain	0.77	13.6	0.94	0.63
Greece	0.89	15.3	0.91	0.60
Guatemala	0.60	15.9	0.96	0.58
Guinea	1.08	14.3	0.89	0.56
Honduras	0.60	14.8	0.95	0.56
India	0.40	12.2	0.95	0.45
Indonesia	1.13	17.9	0.86	1.06
Iran	1.32	13.5	0.81	0.43
Iraq	1.12	9.0	0.73	0.37
Ireland	0.84	12.8	0.96	1.00
Italy	0.77	14.9	0.92	0.58
Jamaica	0.81	10.4	0.89	0.41
Japan	1.29	14.9	0.88	0.54
Jordan	1.04	13.8	0.88	0.50
Kenya	0.68	18.9	0.97	0.52

Korea	1.65	19.1	0.84	0.58
Malawi	0.42	10.1	0.94	0.51
Malaysia	0.71	18.1	0.95	0.52
Morocco	0.77	11.0	0.90	0.58
Mauritius	0.88	13.8	0.95	0.49
Mexico	0.66	13.4	0.90	0.80
Netherlands	0.74	12.7	0.95	0.88
New Zealand	0.89	21.0	0.98	0.75
Niger	1.22	12.2	0.78	0.45
Nigeria	1.10	17.7	0.83	0.57
Norway	0.92	11.9	0.94	0.68
Pakistan	0.22	6.2	0.96	0.47
Panama	0.76	13.3	0.96	0.41
Paraguay	0.93	17.7	0.91	0.75
Peru	0.63	5.9	0.60	0.75
Philippines	0.57	16.0	0.95	0.88
Portugal	0.44	9.1	0.94	0.38
El Salvador	0.78	18.0	0.96	0.67
Senegal	0.84	13.2	0.91	0.59
Singapore	0.97	9.2	0.86	0.45
Somalia	0.51	4.5	0.65	0.73
South Africa	0.78	21.0	0.97	0.79
Spain	0.82	16.5	0.81	0.54
Sri Lanka	0.38	8.0	0.92	0.59
Sudan	1.13	11.3	0.82	0.84
Sweden	0.66	13.0	0.95	0.65
Switzerland	1.27	14.1	0.91	0.60
Syria	0.89	9.7	0.79	0.40
Tanzania	0.56	8.7	0.84	0.43
Thailand	0.80	18.8	0.95	0.58
Trinidad-Tobago	0.86	12.3	0.94	0.36
Tunisia	0.71	14.2	0.96	0.74
Turkey	0.40	6.7	0.87	0.37
Uruguay	1.00	17.3	0.91	0.73
USA	0.81	23.9	0.97	0.91
Venezuela	0.73	12.6	0.81	0.56
Yugoslavia	0.81	9.6	0.87	0.66
Zaire	0.42	2.2	0.25	0.71
Zimbabwe	0.70	15.2	0.94	0.61

There are three ways in which low DW statistics signal econometric problems. First and the most obvious case, if low DWs are actually being caused by autocorrelated errors, i.e, equation (1) is correctly specified but ϵ_{it} is correlated over time, t-tests tend to be overestimated, giving a false sense of accuracy to the estimates of Φ. Second, it is possible that both domestic and international prices are integrated processes, in which case a static regression, at best, would have to be interpreted as a cointegration relationship (Engle and Granger, 1987). A cointegration equation, however, is meaningful only to the extent that the errors (ϵ_{it}) are stationary, but the low DW statistics suggest that in many situations this is not the case. Furthermore, even if regressions were to be interpreted as cointegration relationships, the Φs would provide only the long run elasticity of domestic prices with respect to international prices, but no information whatsoever regarding the short run. In this context, a low DW could be a symptom of a very slow convergence to international prices. Third, it may be the case that the relationship between domestic and international prices is dynamic, in which case the DW would be simply signalling a dynamic misspecification of the model in equation (1).

Therefore, there is no way to deliver a sensible conclusion from the evidence of Table 1 without first dealing with the widespread and systematic problem of serial correlation. In turn, this sugests using a dynamic model to adequately capture the above mentioned potential dynamics.

A Dynamic Approach

In order to deal with the drawbacks of the static model, we postulate an error correction model in the form:

$$)p_{it} = *_{i} + ")p^{*}_{it} - \$(p_{it-1} - (p^{*}_{it-1}) + ,_{it} \qquad \textbf{(2)}$$

To a great extent, this equation captures the three cases outlined above when discussing the causes of low DW statistics in the static regression. Unlike the model in equation (1), its counterpart in equation (2) is a rather general formulation: it encompasses all one-lag dynamic linear models, as shown by Hendry and Richard (1986) and, in particular, it encompasses the possibility that the static model is correctly specified but that errors are serially correlated. On the other hand, if the cointegration interpretation is adopted and if domestic and international prices are in fact cointegrated, then an error correction model like equation (2) would be an appropriate dynamic representation (Engle and Granger, 1987) .

Interpretation of the parameters in equation (2) is straightforward. For a long-run relationship between domestic and international prices to exist, both β and γ should be significant and positive. If β and γ are strictly positive, then, in an hypothetical steady state:

$$p_{it} = (p^{*}_{it} \tag{3}$$

and therefore γ corresponds to the long run elasticity of domestic prices with respect to international prices. Parameter β, on the other hand, measures the speed of convergence toward the long-run relationship. For example, a value of β of 0.25 implies that every year 25% of the gap in equation (3) inherited from the previous period is corrected through current domestic price changes. Finally, parameter α measures the contemporaneous impact of international price changes on domestic prices.

From an economic point of view, model (2) is more appealing than model (1) because it provides not only an estimate of the long run elasticity, but also a portrait of the short run dynamics and the adjustment path. For example, obtaining a positive α but β or γ equal to zero would imply that, although changes in domestic prices are associated with changes in international prices in the short run, price levels could diverge significantly over longer time periods. On the contrary, if γ is close to one but α and β are low, we obtain a case in which, despite having a long run relationship between domestic and international prices, convergence would be very slow. Finally, the case reflecting fast and substantial transmission of international prices would require α, β, and γ close to one.

Table 2.
Dynamic Panel Data Models
$P_{it} = \quad_i + \alpha \quad P^*_{it} + ß_1(P_{it-1} - \quad P^*_{it-1}) + ß_2(P_{it-2} - \quad P^*_{it-2})$

Country	α	t(α)	1	t(1)	2	t(2)		t( )	R²	SEE	DW
Angola	0.04	0.45	-0.14	-5.54			1.04	3.90	0.11	0.27	1.88
Australia	0.50	11.4	-0.51	-10.1			0.83	18.8	0.49	0.14	2.00
Austria	0.07	0.96	-0.04	-0.66	-0.08	-1.27	0.53	1.36	0.11	0.12	1.89
Bangladesh	0.003	0.004	-0.64	-12.0			0.08	1.27	0.43	0.24	1.82
Belgium	0.06	1.03	0.12	1.51	-0.25	-3.63	0.46	2.26	0.17	0.11	1.83
Burkina Faso	0.13	1.95	-0.14	-3.20			0.90	3.53	0.08	0.18	1.91
Burundi	0.15	3.61	-0.21	-7.88			0.87	8.61	0.23	0.12	1.92
C. Afric. Rep.	-0.02	-0.23	0.02	0.32	-0.14	-2.62	0.83	2.85	0.09	0.22	1.90
Cameroon	0.11	2.49	-0.07	-3.49			0.93	2.97	0.06	0.17	1.77
Canada	0.85	11.3	-0.41	-5.01			0.74	8.47	0.61	0.12	1.99
Chile	0.32	2.49	-0.41	-6.27			0.96	5.83	0.25	0.31	1.85
Colombia	0.15	4.25	0.29	5.66	-0.34	-6.45	-0.14	-1.44	0.21	0.15	2.12
Costa Rica	0.35	7.65	-0.25	-6.07			0.62	6.54	0.27	0.17	1.94
Cote d'Ivoire	0.15	2.53	-0.14	-5.01			1.16	5.53	0.09	0.22	1.78
Cyprus	0.06	0.68	-0.06	-1.25			-0.92	-0.60	0.07	0.19	1.74
Denmark	-0.01	-0.15	0.20	2.00	-0.31	-3.21	0.07	0.29	0.23	0.24	2.04
Dom. Rep.	0.41	4.54	-0.26	-5.17			0.66	3.87	0.19	0.31	2.02
Ecuador	0.28	4.45	-0.19	-5.30			0.74	4.23	0.13	0.26	1.99
Egypt	0.10	1.35	0.14	2.15	-0.40	-6.00	0.46	3.32	0.20	0.23	2.12
Ethiopia	0.15	1.57	-0.27	-4.97			0.63	3.35	0.12	0.31	1.96
Finland	0.14	2.87	-0.03	-1.59	-0.01	-0.66	2.86	2.37	0.17	0.24	1.37
France	0.26	4.21	-0.18	-3.60			0.58	3.10	0.17	0.19	1.79
Germany	-0.16	-2.10	-0.14	-2.47			-0.02	-0.04	0.18	0.14	1.82
Ghana	0.11	0.86	0.07	1.19	-0.29	-5.14	1.02	3.41	0.17	0.49	1.91
Great Britain	0.29	5.6	0.17	2.21	-0.41	-5.62	0.72	6.46	0.42	0.19	2.04
Greece	0.13	1.19	-0.17	-2.99			-0.05	-0.12	0.10	0.28	2.00
Guatemala	0.15	3.54	-0.10	-2.58			0.07	0.20	0.11	0.16	1.76
Guinea	0.26	2.97	-0.13	-2.94			0.25	0.55	0.11	0.30	1.79
Honduras	0.17	3.30	-0.09	-1.60	-0.21	-3.71	0.59	6.37	0.19	0.18	1.93
India	0.05	1.33	-0.16	-4.98			-0.01	-0.06	0.11	0.15	1.91
Indonesia	0.21	1.71	-0.32	-4.42			0.59	2.64	0.09	0.47	2.07
Iran	0.12	2.67	-0.02	-0.75	-0.03	-1.20	2.01	3.12	0.05	0.15	1.67
Iraq	0.04	0.61	-0.02	-1.32			4.60	1.54	0.04	0.24	2.29
Ireland	0.30	2.06	-0.02	-0.19	-0.39	-3.11	0.44	2.35	0.24	0.22	2.07
Italy	0.17	2.98	-0.16	-4.36			0.79	4.69	0.12	0.14	1.95
Jamaica	0.13	1.52	-0.07	-1.20	-0.10	-1.35	0.66	2.55	0.12	0.26	1.68
Japan	0.14	2.90	-0.05	-2.65			0.41	0.69	0.10	0.14	2.13
Jordan	0.04	0.52	-0.14	-3.77			0.74	2.78	0.11	0.16	1.86

Kenya	0.27	6.62	-0.21	-6.09			0.72	7.05	0.23	0.15	1.84
Korea	0.27	5.34	-0.09	-4.59			1.72	6.49	0.19	0.12	1.99
Malawi	0.08	1.72	-0.19	-5.17			0.54	4.12	0.11	0.16	1.96
Malaysia	0.18	3.55	-0.17	-4.24			0.36	2.01	0.12	0.20	1.79
Morocco	0.27	3.39	-0.19	-4.65			0.45	1.93	0.16	0.25	1.97
Mauritius	-0.02	-0.23	0.02	0.14	-0.19	-1.60	-0.07	-0.15	0.13	0.28	2.02
Mexico	0.35	4.55	-0.34	-7.90			0.74	6.37	0.20	0.31	2.06
Netherlands	0.19	2.91	0.06	0.76	-0.27	-3.39	0.73	4.47	0.19	0.11	1.93
New Zealand	0.16	9.51	-0.29	-4.00			0.85	7.35	0.46	0.17	1.76
Niger	0.14	2.46	-0.04	-1.71			1.94	2.20	0.05	0.23	2.17
Nigeria	0.24	4.08	-0.20	-7.25			1.07	6.95	0.19	0.14	1.95
Norway	0.07	1.49	-0.05	-0.19	-0.04	-0.95	0.92	3.42	0.09	0.19	1.72
Pakistan	0.03	0.78	-0.21	-5.43			0.28	2.61	0.11	0.14	1.87
Panama	0.16	4.97	-0.14	-6.67			1.14	8.73	0.30	0.10	1.95
Paraguay	0.35	4.87	-0.35	-8.75			1.04	9.41	0.25	0.28	2.22
Peru	0.14	1.07	-0.08	-1.90			-0.11	-0.10	0.02	0.54	1.97
Philippines	0.24	4.67	-0.35	-8.59			0.51	6.86	0.24	0.19	1.81
Portugal	0.08	1.97	0.01	0.28	-0.15	-3.19	0.01	0.06	0.15	0.15	1.74
El Salvador	0.09	2.69	-0.30	-5.91			0.68	5.97	0.15	0.23	1.87
Senegal	0.05	1.02	-0.09	-2.93			0.83	3.04	0.05	0.16	1.98
Singapore	0.11	0.43	0.20	1.86	-0.33	-3.31	0.43	1.90	0.19	0.16	1.96
Somalia	0.19	-0.30	0.18	2.05	-0.31	-3.38	-0.27	-0.49	0.09	0.45	1.97
South Africa	0.35	6.64	-0.33	-7.84			0.81	9.84	0.26	0.18	2.13
Spain	0.25	5.25	-0.10	-3.12			1.13	4.24	0.12	0.19	1.76
Sri Lanka	0.37	5.77	-0.25	-6.30			0.42	3.21	0.23	0.24	2.13
Sudan	0.54	2.65	-0.43	-5.97			0.87	3.80	0.18	0.61	1.87
Sweden	0.19	1.64	-0.04	-0.59	-0.11	-1.53	0.59	2.92	0.07	0.21	1.86
Switzerland	0.08	1.22	0.22	2.21	-0.28	-3.00	0.32	1.25	0.17	0.12	1.86
Syria	-0.05	-0.63	-0.15	-5.42			1.11	4.19	0.13	0.26	2.19
Tanzania	-0.02	0.39	0.41	9.11	-0.48	-10.2	-0.47	-4.03	0.29	0.21	2.11
Thailand	-0.02	8.10	-0.19	-5.39			0.88	6.56	0.26	0.17	2.05
Trinidad-Tobago	-0.05	5.15	-0.17	-6.03			1.25	7.59	0.32	0.15	1.80
Tunisia	-0.38	-0.41	0.09	1.06	-0.20	-2.49	-0.09	-0.36	0.19	0.10	1.94
Turkey	0.25	2.73	0.11	1.75	-0.25	-3.74	-0.23	-0.88	0.19	0.19	2.13
Uruguay	0.54	7.36	-0.33	7.54			1.16	9.38	0.33	0.24	2.04
USA	0.45	9.06	-0.32	-6.94			0.73	8.52	0.30	0.20	2.13
Venezuela	0.11	1.41	-0.17	-4.35			0.49	2.23	0.08	0.26	1.91
Yugoslavia	0.01	0.13	-0.13	-2.15	-0.10	-1.48	0.42	1.19	0.08	0.30	1.91
Zaire	0.03	0.04	-0.16	-3.94			-0.83	-0.91	0.05	1.05	2.17
Zimbabwe	0.08	1.38	-0.18	-4.56			0.02	0.08	0.14	0.20	1.74

Equation (2) was estimated for each country allowing the δ_is to differ among commodities, which amounts to estimating a non-linear fixed-effects panel regression for each country.[6] Following Phillips and Loretan (1991), whenever the DW statistic fell below the critical level 1.75, instead of using a lagged dependent variable term, an additional error-correction term was included.[7] In this case, equation (2) was modified to yield:

$$\Delta p_{it} = \delta_i + \alpha \Delta p^*_{it} - \beta_1 (p_{it-1} - \gamma p^*_{it-1}) - \beta_2 (p_{it-2} - \gamma p^*_{it-2})$$

The results of these alternative dynamic estimations appear in Table 2.

The Main Empirical Findings

The results indicate that the error correction model performs better than the static regression reported in Table 1. In the great majority of cases, the serial autocorrelation disappeared with this alternative specification. In a few cases, two lagged error correction terms had to be included, in particular for several EC countries.

The first question that can be elucidated from Table 2 is the existence of a long-run relationship between domestic and international prices, i.e, those cases where both the estimated β and γ are significant and have the correct signs.[8] Although a significant portion of the sample satisfied this criterion, we found that 30 out of the 78 countries did not. With the exception of Singapore and Morocco, which have t-tests in the neighborhood of 2, in the rest of the cases coefficients were unmistakenly non significant, so that we can unambiguously conclude that in 28 countries of the sample, domestic producer prices are virtually isolated from international price

[6] See Arellano and Bond (1991) for a primer on dynamic panel data estimation and Das (1993) for the fixed-effects estimator in non-linear models

[7] Bhargava, Franzini and Narendranathan (1982) provide critical values for a generalization of the DW statistic for panel data.

[8] In the cases where two error correction terms were included, we tested whether the ß with the correct sign had a t-test above 2.

signals in the long run. Countries belonging to this group are listed in the upper panel of Table 3. Notably, it includes most of the EC countries although is not restricted to them; other high-income economies like Japan are included, as well as low-income countries like Guinea and Zimbabwe. On the other hand, domestic prices in some EC countries like Spain, Great Britain, and France, do exhibit a long run relationship to international prices. Price isolation seems to be the rule, however, in northern continental European economies, but again, is not restricted to them: Greece stands out as an interesting exception.

Of the 30 countries which do not exhibit a long run relationship with international prices, nine did show a statistically significant relationship in the short run -α is significant- although the magnitude of the short run transmission mechanism was rather small. On the contrary, within the group of countries which did exhibit a long run relationship, the great majority also displayed a statistically significant contemporaneous transmission of international price signals, i.e, whenever international price signals matter in the long run, some contemporaneous transmission of prices also takes place.

A second issue concerns the speed of the transmission process. Of course, this question is relevant only for those countries displaying a long run relationship with international price signals. The speed of adjustment is jointly governed by the three parameters of interest, α, β, and γ, from which we compute the number of years it would take for a country to transmit a given fraction of any change in international prices. Based on the point estimates for these parameters, the lower panel of Table 3 classifies the countries according to the number of years they would require to transmit 50% of a shock. Of the 48 countries for which a long run relationship between domestic and international prices exists, 9 of them would never adjust even 50% of international price changes because the estimated γ is less than 0.5. Furthermore, 22 countries would require 5 years or more to transmit half the size of a shock. This leaves a total of 31 countries for which the adjustment can be labeled as very low. On the other extreme, we find that only 4 of the 48 countries -namely Australia, Canada, New Zealand, and Uruguay- would in principle transmit 50% or more of a price change within one year, and that 5 additional countries would do it in 2 years.

Table 3
Classification of Countries According to Responsiveness to International Price Signals

Category	Countries
No long run relationship	Austria, Bangladesh, Switzerland, Colombia, Cyprus, Germany, Denmark, Finland, Guinea, Greece, Guatemala, India, Iraq, Jamaica, Japan, Mauritius, Morocco, Netherlands, Norway, Peru, Iran, Portugal, Singapore, Somalia, Sweden, Tunisia, Turkey, Yugoslavia, Zaire, Zimbabwe [30]
Adjustment less than 50% of price changes in long run.	Belgium, Egypt, Ireland, Malaysia, Sri Lanka, Pakistan, Tanzania, Venezuela [8]
Takes more than 7 years to transmit 50% of a shock.	Cameroon, Central African Republic, France, Jordan, Malawi, Philippines, Senegal [7]
Takes between 5 and 7 years to transmit 50% of a shock.	Angola, Burkina-Faso, Ecuador, Ethiopia, Ghana, Great Britain, Honduras, Indonesia, Italy, Kenya, Niger, El Salvador, Spain, Syria [14]
Takes between 3 and 4 years to transmit 50% of a shock.	Burundi, Cote d'Ivoire, Costa Rica, Dominican Republic, Korea, Mexico, Nigeria, Panama, Thailand, Trinidad [10]
Takes 2 years to transmit 50% of a shock.	Chile, Paraguay, Sudan, U.S.A., South Africa [5]
50% or more of any given shock is transmitted within a year.	Australia, Canada, New Zealand, Uruguay [4]

Notes: The first panel of the table corresponds to the countries for which either the long run parameter γ or the adjustment parameter β in equation (2) had a t test below 2, according to the results in Table 2. Either of these two conditions implies the absence of a long run relationship between domestic and international prices. Marginal cases according to this criterion were Morocco and Singapore. Based on the point estimates of α, β, and γ, the lower panel of Table 3 computes the number of years that it would take for a given country to transmit 50% of a price shock to domestic prices. The sub-panel (2.1) includes countries with γ less than 0.5 which, correspondingly, will always transmit less than 50% of any price shock.

The picture emerging from these estimations is clear. In an overwhelming majority of cases transmission of international price signals in agriculture is either non-existent or low, by any reasonable standard; 30 of the 78 countries appear virtually isolated from international price signals in the long run, and 30 of the remaining 48 either take 5 years or more to transmit half of any price change or never adjust more than 50% altogether. EC countries stand out as either isolated in the long run from international price signals or belonging to the "low-speed" of convergence group.

It is interesting to observe the key role played by the dynamic formulation in this analysis. If we had ignored the serial correlation problem, which was pervasive to the estimation of model (1), we would have concluded that substantial transmission of international price movements took place in agricultural markets. In fact, when looking at the evidence presented in Table 1, no major difference arises between countries such as Austria and Uruguay; yet, the dynamic estimations of Table 2 point to a very different conclusion. While Uruguay incorporates most of foreign price shocks in the very short run, agricultural prices in Austria appear virtually isolated from international prices, both in the short and long run.

How can the evidence presented in Table 1 be explained then? We contend here that the high elasticities and the high R^2s of the static regressions are mostly a spurious result that arises from the simple fact that both international and domestic prices behave closely to a random-walk with drift, being the drift in domestic prices higher than that of international prices. To test this assertion, we selected the most "controversial" cases from the group of countries for which we did not find evidence of long run relationship with international prices, i.e, those countries for which the static regression displayed an elasticity higher than 0.7 and a DW statistic higher than 0.6. For these cases we posited the simplifying assumption that both international and domestic prices could be described as random-walks with drift, allowing for some contemporaneous correlation between the annual growth rates.[9] That is, we postulate a simple data-generation-process (DGP) of the form:

$$)p_t^* = *_0 + ,_t^* \qquad \textbf{(5a)}$$

$$)p_t = (_0 + (_1)p_t^* + ,_t \qquad \textbf{(5b)}$$

[9] Assuming that prices follow random-walk processes is a convenient simplification for expository purposes. In the following section we return to this point to undertake a formal analysis.

where ϵ_t and ϵ_t^* are two independent uncorrelated random shocks. Table 4 presents the panel estimation of equations (5a and 5b) for these selected cases. The estimations display, in general, a low value of γ_1 - in the order of 0.1 - and values for δ_0 and γ_0 in the range of 0.03 and 0.05 respectively. The DW statistics suggest that this model can be reasonably taken as a parsimonious approximation to the data. We then used the estimated standard errors of the innovations plus the average point estimates for δ_0, γ_0, and γ_1 to run a Montecarlo simulation exercise. That is, using these estimated parameters of the DGP, we generated 1,000 artificial samples of size 30 each and on each of them we ran the following static regression:

$$p_t = c_0 + c_1 p_t^*$$

Across the 1,000 simulations, these static regressions displayed a mean value of c_1 of 0.6 with a t-test which on average was 5.7, a mean value of R^2 of 0.98, and a mean DW value of 0.6; that is, a configuration of results that resembles very closely the estimates presented in Table 1 for this group of countries.[10] Furthermore, we found that 45% of the cases displayed, simultaneously, a c_1 estimate higher than 0.6, with a "t" test higher than 5.0, an R^2 higher than 0.9, and a DW higher than 0.6. Hence, the "evidence" of Table 1, at least for all cases in which we concluded that no long run relationship with international prices applied, can be explained as a spurious result arising from regressing highly serially correlated prices with a stochastic trend.

It is interesting to note that in terms of the representation (5.a)-(5.b), even with γ_1 equal to zero, one would still be able to generate the bulk of results presented in Table 1; what matters for the spurious regression result are the stochastic trends δ_0 and γ_0, and in particular the fact that $\delta_0 > \gamma_0$. In fact, when modelling equation (1) to include a deterministic trend among the explanatory variables, we found that in all cases the trend was significant, and that point elasticities of ß were reduced by one-half to two-thirds of their original values, with most cases being below 50% (see Table 5). Nevertheless, the serial correlation did not disappear with the inclusion of the deterministic trend, further suggesting that the correct specification was of the form (5.b). Therefore, we unambiguously conclude that the evidence of the static regression is

[10] The possibility of obtaining a high R^2 in spurious regressions of this type is not a new result (see Granger and Newbold, 1974).

strongly misleading, and that in order to relate international agricultural prices with domestic ones, it is more appropriate to look at more general dynamic processes, as we do in this paper.

Table 4
Panel Data Estimation of First Differences Correlations Model

$$\Delta P^*_t = \delta_0 + \epsilon_t$$

$$\Delta P_t = \gamma_0 + \gamma_1 \Delta P^*_t + \epsilon_t$$

Country	δ_0	γ_0	γ_1	S.E. (5.a)	DW (5.a)	S.E. (5.b)	DW (5.b)
Austria	0.03 (1.56)	0.05 (4.04)	0.08 (1.33)	0.20	1.56	0.12	1.77
Switzerland	0.03 (1.56)	0.11 (1.81)	0.07 (5.23)	0.20	1.56	0.12	1.46
Germany	0.03 (1.78)	0.11 (1.50)	0.05 (3.31)	0.19	1.53	0.15	1.98
Denmark	0.03 (1.36)	0.09 (1.29)	0.06 (3.85)	0.21	1.52	0.14	1.54
Finland	0.03 (1.36)	0.08 (1.59)	0.05 (4.80)	0.21	1.54	0.10	1.12
Netherlands	0.03 (1.36)	0.16 (2.52)	0.04 (2.85)	0.20	1.54	0.13	1.69
Norway	0.03 (1.36)	0.04 (1.00)	0.06 (6.50)	0.21	1.54	0.08	1.66
Yugoslavia	0.03 (2.30)	0.07 (0.64)	0.01 (0.46)	0.19	1.70	0.29	1.89
Average	0.03	0.08	0.05	0.20		0.14	

Note: S.E. is the standard error of the estimates.

Table 5

Fixed Effects Panel Data Estimation of the Static Model
Including Deterministic Trend

$$p_{ti} = *_{i} + "t + \$p^{*}_{it}$$

Country	α	β	Corrected R^2	DW
Austria	0.05 (11.6)	0.29 (4.03)	0.98	0.51
Switzerland	0.06 (13.5)	0.36 (4.28)	0.97	0.41
Germany	0.04 (8.63)	0.60 (7.36)	0.98	0.67
Denmark	0.02 (6.94)	0.43 (6.37)	0.98	0.82
Finland	0.07 (22.7)	-0.04 (-0.77)	0.99	0.62
Netherlands	0.03 (8.97)	0.36 (5.87)	0.97	0.71
Norway	0.05 (23.5)	0.28 (7.01)	0.99	0.84
Yugoslavia	0.02 (2.7)	0.56 (4.49)	0.88	0.68

3. Modelling Price Intervention as a Result of Dynamic Optimization

In this section we present an interpretation of the previous econometric results in terms of plausible policy objectives. We formulate a stochastic dynamic-maximization problem which, if solved by an hypothetical policy maker, would give as a result the error correction mechanism found in the data.[11]

[11] Nickell (1985) works out several examples where an error correction model would appear as the decision rule outcome of a dynamic stochastic maximization process.

Within a general framework, we postulate that policy makers care about two main objectives. First, to maintain a given relationship between domestic and international prices, and second, to minimize year-to-year fluctuations in domestic producer prices. The goal of keeping domestic prices in line with international ones arises from the fact that systematic deviations from international price levels is not costless for the economy, a point emphasized by Mundlak and Larson (1992). For example, an exporter country cannot systematically give price incentives to producers which deviate from international price trends. However, there is a tension between this objective and the need to stabilize domestic price fluctuations, which calls for a certain degree of isolation from international price signals. In fact, one of the findings of Krueger, Schiff, and Valdés (1992) was that direct price intervention actually tended to generate less domestic price fluctuations than what would have occurred if no price intervention had applied.

More specifically, if p_t denotes the log of domestic prices and p_t^* the log of international prices, we conjecture that the policy maker is interested in minimizing the following quadratic intertemporal loss-function:

$$E_j \sum_{t=j}^{\infty} *^{t-j} \left((1-2)(p_t - p_t^* - s_t)^2 + 2(p_t - p_{t-1})^2 \right) \qquad \textbf{(6)}$$

by choosing domestic price levels contingent on a set of state variables including international prices. In this equation, δ is a discount factor; s_t is the possible time-evolving objective gap between domestic and international prices; E_j denotes the expectation operator conditional on information prevailing at time $t = j$; and θ is the relative weight of price stability in the objective function. So, the higher θ, the more the policy maker cares about domestic price stability.

The stochastic nature of this maximization problem is due to the presence of both international price changes and domestic policy objective shocks. The former originate because p_t^* fluctuates over time in response to changing conditions in international markets, while the latter arise with the evolution of s_t, usually influenced by various pressures and circumstances within and outside the agricultural sector.[12] Therefore, for the model to deliver testable implications, specific stochastic laws of motion for these state variables must be formulated. In order to keep the model as simple as possible, we postulate that the policy target s_t and the log of international prices p_t* evolve according to first and second order autoregressive processes respectively, that is:

$$\begin{aligned} s_{t+1} &= \mathrm{D}_0 + \mathrm{D}_1 s_t + ,_{1t+1} \\ p^*_{t+1} &= 8_0 + 8_1 p^*_t + 8_2 p^*_{t-1} + ,_{2t+1} \end{aligned} \qquad \textbf{(7)}$$

where $\epsilon_{j,t+1}$ is normal i.i.d., with variance σ_j^2, j = 1,2. The vector of states for this problem is defined as:

$$X_t = (p_{t-1}, p^*_t, p^*_{t-1}, s_t, 1)'$$

and the control variable by:

$$u_t = p_t - p_{t-1}$$

Under certain mild conditions regarding the laws of motion in (7), there exists a stationary Markov control-policy that solves the problem (Bertsekas and Shreve, 1978). Furthermore, since in this case the objective function is quadratic and the laws of motion of the states are linear, the decision rule that maximizes the problem will take the form of a linear mapping h: X → u (see Appendix 2 for a detailed description of the solution algorithm). The solution to the model then will be represented by a 1x4 matrix F, so that:

$$u_t = -FX_t \qquad \textbf{(8)}$$

---

[12] There is a large body of literature devoted to analyzing the political forces that determine a given sector to be protected or not, i.e, whether price intervention will imply domestic prices above or below international price levels, eg., Krueger (1992), Bates (1983), and Becker (1983). In order to focus on the dynamic implications of price intervention, we take those conditions as given outside the model.

Treating s_t as unobservable, the decision rule can always be reparameterized as an error correction mechanism like model (2) in the previous section. The linearity of the decision rule implies:

$$
\begin{aligned}
)\, p_t &= -f_1 p_{t-1} - f_2 p_t^* - f_3 p_{t-1}^* - f_4 s_t - f_5 \\
&= -f_2 \,)\, p_t^* - f_1 \left(p_{t-1} + \frac{f_3 + f_2}{f_1} p_{t-1}^* \right) - f_4 s_t - f_5
\end{aligned}
$$

so that we can link the parameters in equation (2) to the solution of the stochastic model as:

$$
\begin{aligned}
" &= -f_2 \\
\$ &= f_1 \\
(&= -\frac{f_3 + f_2}{f_1}
\end{aligned}
\qquad \textbf{(9)}
$$

Several interesting cases can be distinguished. If international prices are actually described by a second order autoregressive process - i.e, both λ_1 and λ_2 are different from zero - but ρ_1 is zero, then the implication of the model described by equations (6) and (7) is exactly an error correction model and the errors of an econometric equation which treats s_t as unobservable, as we implicitly did, will be serially uncorrelated. On the other hand, if ρ_1 is different from zero, then the disturbances of an error correction equation will still be serially correlated. Finally, if international prices are better described by first order autoregressive process ($\lambda_2 = 0$) then the error correction mechanism will be still valid, but it could be further simplified to a partial adjustment equation, because f_3 would be zero in this case, implying the following restriction:

$$" = (\,\$$$

In these three cases, and without a-priori knowledge regarding the laws of motion for s_t and p_t^*, the estimation of an error correction equation, as performed in the previous section, appears broadly consistent with these different alternatives.[13]

[13] In these three cases, the error correction estimation can be shown to yield asymptotically consistent estimators. However, with a-priori knowledge concerning the laws of motion in equation (7), more

Numerical Solution and Simulations

In order to contrast the implications of this model with the econometric findings of the previous section, we solved the model for alternative parameter configurations embedded in vector Ω:

$$S = (*, 2, D_0, D_1, 8_1, 8_2)'$$

which amounts to finding the solution matrix F for different choices of the parameters characterizing the objective function in (6) and the stochastic laws of motion in (7). Finally, the policy rule F is reparameterized as an error correction equation using equation (8).

The numerical solution of the model requires us to specify parameter values for the different elements in Ω. For some of them we have reasonable a-priori ranges; for example δ may fluctuate between 0.85 and 0.95, and λ_1 and λ_2 can be set at levels consistent with the time series evidence on international commodity prices (see below). Parameters describing the process of $\{s_t\}$, namely ρ_0 and ρ_1, are unobservable and so is our parameter of interest θ. In order to assess both the plausibility of this model and reasonable ranges for θ, the model is solved for different parameter choices of observable and unobservable variables. The resulting values for α, β, and γ in the error correction representation of the policy rule are later contrasted with the point estimates obtained from the data.

Table 6 presents the estimation of AR(1) and AR(2) models describing international commodity price processes for the 16 goods in the sample. For two commodities, wheat and wool, there is definite evidence of second order autoregressive processes: the null hypothesis of $\lambda_2 = 0$ can be rejected at standard confidence levels and the residuals of simple first order processes appear serially correlated. On the opposite side, there are four commodities - coffee, cotton, tea, and tobacco - for which a simple first order representation is clearly superior: the second order coefficient is statistically insignificant and negligible in terms of magnitude. For most other cases, however, the evidence is mixed, in the sense that despite the fact that t tests are not very high by

---

efficient estimators could be obtained. In the absence of that knowledge, the estimation performed in the previous section seems broadly consistent with the dynamic model formulated here.

conventional standards, the magnitudes of the coefficients are important. With the exception of banana prices, in all cases where the second order coefficient is statistically significant or important in terms of magnitude, it had a negative sign. On the other hand, the first order coefficient tended to be high: 11 out of 16 coefficients were equal or higher than 0.7, confirming the presumption in Section 2 that most parsimonious representations of international commodity prices cluster around unit-root representations.[14] The average point estimate for λ_1 was 0.79 and -0.17 for λ_2. If we exclude the four commodities for which a second process is definitely rejected, the average point estimates change to 0.82 and -0.24. If the true parameter configuration were actually in the neighborhood of a unit-root process, then the sample distribution of λ_1 would be skewed toward values below one. This suggests solving the model with values of λ_2 in the order of -0.20 and λ_1 in the range of 0.75 and 1.2.

Table 7 presents the values of α, β, and γ implied by the stochastic model for alternative parameter choices in Ω. A base scenario, which collects the previous discussion on the appropriate size of parameters, was characterized by choosing $\lambda_1 = 0.75$, $\lambda_2 = -0.20$, a discount rate of 10% per annum (i.e., $\delta = 0.91$), and a mean-zero i.i.d. process for $\{s_t\}$. If in this base simulation the value of θ is fixed equal to 0.5, the simulated values of α, β, and γ --presented in the first row of the table-- would be consistent only with countries such as Canada or New Zealand, which exhibit a very high responsiveness to international price signals. In other words, had policy makers given equal weight to international price signals and year-to-year domestic price changes, then the parameters of the error correction models estimated above would have all looked like those found for Canada or New Zealand. The fact that most of them did not, suggests that values of θ above 0.5 may be underlying the price transmission estimates found in the data. The next three rows of Table 7 present the simulated parameter configurations which result from increasing the weight given to domestic pricing policies. It is evident that only when θ reaches a level of 0.9 or above, do simulated parameters begin to be consistent with the bulk of previous econometric estimations.

[14] It is clearly outside the scope of this paper to discuss the stochastic processes governing commodity prices in any depth, but only to obtain a reasonable parsimonious representation that can be useful for the purposes of delivering empirical implications out of the model. The simple processes adopted here are broadly consistent with the findings of Gersowitz and Paxon (1990) and Cuddington (1992), but ignore the complexities discussed by Deaton and Laroque (1992).

In other words, speeds of transmission corresponding to the first half panels in Table 3 would be consistent with domestic pricing policies assigning 90% or more to the objective of minimization of year-to-year price changes and only 10% to international price signals.

Table 6
Time-Series Models
for International Commodity Prices
$P^*_t = \lambda_1 P^*_{t-1} + \lambda_2 P^*_{t-2} + \epsilon_t$

Commodity	λ_1	$t(\lambda_1)$	λ_2	$t(\lambda_2)$	R^2	Q(7)
Bananas	0.58	2.81	0.16	0.82	0.58	2.10
Cocoa	0.96	5.01	-0.29	-1.43	0.53	4.06
Coffee	0.71	4.06			0.36	1.65
Copra	0.59	3.04	-0.26	-1.33	0.28	7.52
Cotton	0.59	3.61			0.32	4.05
Maize	0.93	4.60	-0.17	-0.85	0.63	4.21
Palm Oil	0.75	3.79	-0.14	-0.69	0.46	11.4
Rice	1.01	5.19	-0.32	-1.64	0.62	3.21
Rubber	0.59	3.00	-0.18	-0.92	0.29	5.18
Sorghum	0.99	4.93	-0.20	-0.97	0.69	2.78
Soybean	0.95	4.32	-0.24	-1.22	0.50	10.27
Tea	0.70	5.48			0.52	7.72
Tobacco	0.81	7.51			0.68	2.71
Sugar	0.92	5.09	-0.27	-1.48	0.58	5.02
Wheat	0.99	5.28	-0.39	-2.02	0.56	4.33
Wool	0.72	4.21	-0.53	-3.14	0.97	3.48

The lower panels of Table 7 examine the robustness of this conclusion to alternative parameter specifications. The central panel shows that increasing the size of λ_1 results in α, β, and γ being higher, which implies an increase in responsiveness to international prices, both in the short and long run. The intuition is clear: a higher λ_1 implies that current price shocks are expected to last longer so that, for a given θ, a higher degree of responsiveness to international prices is required. Nevertheless, under this alternative assumption, one would again need values of θ of 0.9 or higher to reproduce most of the previous econometric results. Observe that high values of γ, in the order of 0.8 or more, appear sporadically in the data (see Table 3) in conjunction with small short run responses (low α and β). That particular configuration can be replicated by the model under the assumption of a close-to-random-walk process in international prices ($\lambda_1 = 0.95$, $\lambda_2 = 0$) concurrent with a high weight to domestic price stability ($\theta = 0.95$). In that case, the result is $\alpha = 0.22$, $\beta = 0.25$, and $\gamma = 0.90$.[15] The random-walk certainly creates a tension in the presence of high θ here: it implies that policy makers effectively perceive prices as being very persistent over time, but they are reluctant to adjust quickly to international price signals due to domestic price stabilization objectives. The maximization solves this tension through a high long run response simultaneously with low short run responsiveness, a pattern observed in a few cases, notably in several African countries (see in Table 2 the evidence for Burundi, Cote d'Ivoire, Ghana, Burkina Faso, Nigeria, Panama, and Syria). In all cases, hence, to replicate the low short run response found in the data, values of θ in the order of 0.9 or above are needed.

Finally, the lower panel of Table 7 shows that simulated error correction coefficients do not depend in any significant way on the nuisance parameters ρ_0 and ρ_1 nor on the discount rate δ when they vary within plausible ranges. This invariance of the estimated parameters to the unobservable process $\{s_t\}$ suggests that the conclusion of θ being in the order of 0.9 or higher for most countries is robust.

---

[15] With $\lambda_1 = 0.97$ and $\theta = 0.99$ one obtains $\gamma = 0.91$, $\beta = 0.15$, and $\alpha = 0.13$. As the process converges to the unit root case, the long run elasticity converges to unity.

Table 7
Error Correction Parameters Implied by the Dynamic Stochastic Maximization Model

Parameter Configuration	α	β	γ
Base Scenario, $\theta = 0.5$	0.66	0.73	0.87
$\theta = 0.8$	0.37	0.48	0.70
$\theta = 0.9$	0.23	0.35	0.58
$\theta = 0.95$	0.14	0.25	0.48
Base Scenario, $\theta = 0.9$			
$\lambda_1 = 0.85$	0.25	0.35	0.64
$\lambda_1 = 0.95$	0.28	0.35	0.71
$\lambda_1 = 1.10$	0.34	0.35	0.86
$\lambda_1 = 1.15$	0.36	0.35	0.93
$\lambda_1 = 0.95$, $\lambda_2 = 0$, $\theta = 0.95$	0.22	0.25	0.90
Base Scenario, $\theta = 0.5$	0.66	0.73	0.87
$\delta = 0.95$	0.66	0.73	0.86
$\delta = 0.85$	0.66	0.72	0.87
$\rho_1 = 0.50$	0.66	0.73	0.87

Notes: The table shows the values of α, β, and γ of an error-correction representation like equation (2) that would come out of the dynamic maximization model in the text under alternative choices of the parameters. The base scenario parameters are $\delta = 0.9$, $\lambda_1 = 0.75$, $\lambda_2 = -0.20$, $\rho_0 = \rho_1 = 0$. In the first column, each row specifies the particular parameter value that was changed in the based scenario.

As a summary of this exercise, we conclude that the dynamic maximization model posited here is broadly consistent with the econometric findings of the previous section and strongly suggests that, when it comes to agriculture, most governments in the world put substantially higher weight in the objective of domestic price stability rather than international price signals. This result appears consistent with previous empirical findings pointing to the conclusion that a major outcome of agricultural pricing policies was to achieve greater domestic price stability (Schiff and Valdés, 1992; Hazell et al., 1990). However, it is important to observe that in the context of this model, the fact that pricing policies assign greater priority to minimizing year-to-year domestic price fluctuations does not necessarily means that ex-post producer prices will always fluctuate less than international ones. The extent to which this happens will depend on the variance of the exogenous policy objective, s_t, vis a vis the variance of international prices. Under the maintained assumption that s_t is i.i.d., the standard deviation of s_t can be identified by the standard error of the error correction equation in Table 2.[16] EC countries, for example, will have a low ex-post variance of domestic prices because they tend to be isolated from international price signals and also they exhibit an error correction equation with a low standard error. This will probably not be the case with countries similarly isolated from international price signals such as Peru and Somalia, for which the high standard errors of the equation in Table 2 suggest wide fluctuations in the exogenous policy objective s_t. For these countries, it is possible that ex-post prices would have been less volatile if agricultural pricing policies had closely followed international price signals without further government intervention.

4. Conclusions

This paper analyzes the extent to which international price movements of agricultural goods are transmitted to domestic prices. We study a data set including a sample of 78 countries, covering approximately the last 30 years and a range of products between four to 15 commodities per country. We concluded, unambiguously, that in an overwhelming majority of cases - 60 out of 78 - the transmission of international price signals was either non-existent or extremely low by any reasonable standard. From an economic perspective, we show that this behavior is consistent with policy makers assigning a 90% weight or more to the objective of minimizing year-to-year price changes, as opposed to keeping domestic incentives in line with international price signals.

Several implications follow from these results. A reduction in international prices, if not transmitted to domestic markets worldwide, will fail to generate substantial increases in world demand and reductions in world supply, thereby making the price reduction more acute and prolonged, a point earlier emphasized by Johnson (1991). The finding that most governments tend to follow a leaning-against-the-wind policy, i.e, that they are reluctant to transmit new information on international prices in any significant way, implies that a significant part of agricultural price fluctuations in international markets may be "man-made", that is, induced by government intervention rather than the outcome of truly exogenous shocks. Based on these results, an

---

[16] Indeed, under the maintained assumptions of the model, the error of that equation corresponds to s_t times the parameter f_4 of the optimal decision rule.

interesting future line of research would be to quantitatively asses the "excess" world price volatility generated by this pattern of government intervention.

A second implication deals with the ongoing negotiations in the context of economic integration. When successful, negotiations will probably imply sizable impact on the levels of protection to agricultural goods in most countries, as a result of increasing reliance on tariffication as opposed to quantitative and/or international price-contingent trade restrictions. A constant import tariff, no matter how large it is as long as it remains non-prohibitive, will generate a pattern of price transmission substantially different from what we found in most countries, implying full transmission of international price movements to domestic prices. As a consequence, our results imply that a successful Uruguay Round in agriculture will generate a dramatic regime change - in the structural sense - in the processes governing domestic agricultural prices in most countries in the world. This great departure from what has been the historical behavior may help us to understand the various reactions against this proposal by key players in international agricultural trade.

Finally, the fact that most cases appear consistent with policy makers assigning great importance to the reduction of between-years price fluctuations, as opposed to international price signals, suggests that in most countries, price risk management policies will have to be at the core of any trade liberalization program in agriculture.

Appendix 1

Data Sources and Structure

International prices for 15 agricultural commodities were obtained from the International Financial Statistics database of the IMF, for the 1966-1991 period. The goods and the markets from which price quotations were obtained are: Bananas (US ports), Cocoa (New York and London), Coffee (New York), Copra (European ports), Cotton (US, 10 markets), Maize (US Gulf ports); Palm Oil (Europe); Rice (Bangkok), Rubber (New York), Sorghum (US Gulf ports), Soybeans (Rotterdam); Sugar (New York), Tea (London), Tobacco (US), Wheat (US Gulf ports) and Wool (Australia).

The following table summarizes the number of goods per country used in this study. All data were obtained from the FAO Producer Prices database and corresponds to the "close-to-farmgate-price" price definition. Specific country data is available from the authors upon request.

Table A1
Country Data

Country	Number of Goods	Country	Number of Goods	Country	Number of Goods
Angola	10	Guatemala	11	Paraguay	12
Australia	10	Guinea	9	Peru	14
Austria	4	Honduras	11	Philippines	12
Bangladesh	10	India	14	Portugal	9
Belgium	6	Indonesia	12	El Salvador	10
Burkina Faso	6	Iran	10	Senegal	8
Burundi	10	Iraq	9	Singapore	6
C. Africa Rep.	10	Ireland	3	Somalia	8
Cameroon	12	Italy	8	S. Africa	11
Canada	4	Jamaica	7	Spain	10
Chile	7	Japan	8	Sri Lanka	11
Colombia	14	Jordan	6	Sudan	8
Costa Rica	11	Kenya	11	Sweden	6
Cote d'Ivoire	12	Korea	8	Switzerland	4
Cyprus	8	Malawi	10	Syria	9
Denmark	5	Malaysia	12	Tanzania	15
Dom. Rep.	4	Morocco	8	Thailand	11
Ecuador	14	Mauritius	6	Trin. Tob.	6
Egypt	9	Mexico	13	Tunisia	6
Ethiopia	9	Netherlands	4	Turkey	9
Finland	4	New Zealand	5	Uruguay	9
France	7	Niger	7	USA	12
Germany	5	Nigeria	12	Venezuela	11
Ghana	11	Norway	4	Yugoslavia	8
Great Britain	5	Pakistan	10	Zaire	14
Greece	8	Panama	7	Zimbabwe	11

Appendix 2

Using equation (8), observe that problem (6)-(7) can be written as:

$$MAX \quad E_0 \sum_{t=0}^{\infty} *^t (X_t^{\prime} Q X_t + u_t^{\prime} R u_t + 2X_t^{\prime} W u_t)$$

subject to:

$$X_{t+1} = AX_t + Bu_t + C\,,_{t+1}$$

where:

$$Q = \begin{pmatrix} 1 & -1 & 0 & -1 & 0 \\ -1 & 1 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 \\ -1 & 1 & 0 & 1 & 0 \\ 0 & 0 & 0 & 0 & 0 \end{pmatrix}$$

$$W = (1,-1,0,-1,0)^{\prime}$$

$$R = -1 - \frac{F}{2}, \qquad \frac{F}{2} = \frac{1}{2}\frac{2}{1-2}$$

$$A = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 \\ 0 & (_1 & (_2 & 0 & (_0 \\ 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & D_1 & D_0 \\ 0 & 0 & 0 & 0 & 1 \end{pmatrix}$$

$$B = (1,0,0,0,0)^{\prime}$$

and where ϵ_{t+1} has been defined as a 2x1 random i.i.d. vector of standard normal disturbances.

The solution strategy for problems of this kind has been widely discussed in the literature. Following McGrattan (1990), let:

$$C = \begin{vmatrix} 0 & 0 \\ F_1 & 0 \\ 0 & 0 \\ 0 & F_2 \\ 0 & 0 \end{vmatrix}$$

$$\begin{aligned} \tilde{A} &= \sqrt{*}(A - BR^{-1}W') \\ \tilde{B} &= \sqrt{*}B \\ \tilde{Q} &= Q - WR^{-1}W' \end{aligned}$$

the matrix F will be given by:

$$F = (R + \tilde{B}'P\tilde{B})^{-1}\tilde{B}'P\tilde{A} + R^{-1}W' \qquad (*)$$

where P is the steady-state solution to the matrix Ricatti difference equation:

$$= \tilde{Q} + \tilde{A}'P_{t+1}\tilde{A} - \tilde{A}'P_{t+1}\tilde{B}(R + \tilde{B}'P_{t+1}\tilde{B})^{-1}\tilde{B}'P_{t+} \qquad (**)$$

For different parameter choices in:

$$S = (*, 2, D_0, D_1, 8_1, 8_2)'$$

F can then be obtained by first directly iterating on (**) and then using (*).

REFERENCES

Anderson, K. and Hayami, Y. (1986). *The Political Economy of Agricultural Protection*. Allen and Unwin, Sydney, Australia.

Arellano, M. and Bond, S. (1991). 'Some Tests of Specification for Panel Data: Montecarlo evidence and an application to employment equations.' *Review of Economic Studies*, vol. 55, pp. 277-297.

Bates, R. H. (1983). *Essays on the Political Economy of Rural Africa*. University of California Press, Berkeley, California.

Becker, G. (1983). 'A Theory of Competition among Pressure Groups for Political Influence.' *Quarterly Journal of Economics*, vol. 98, pp. 371-400.

Bertsekas, D. P. and Shreve, S. (1978). *Stochastic Optimal Control: The Discrete Time Case*. Academic Press, New York.

Bhargava, A., Franzini, L. and Narendranathan, W. (1982). 'Serial Correlation and the Fixed Effects Model.' *Review of Economic Studies*, vol. 49, pp. 533-549.

Cuddington, J.T. (1992). 'Long-run Trends in 26 Primary Commodity Prices: A disaggregated look at the Prebisch-Singer hypothesis.' *Journal of Development Economics*, vol. 39, pp. 207-227.

Das, S. (1993). 'The Asymptotic Distribution of the Fixed Effects Estimator for Non-linear Regression.' *Journal of Statistical Planning and Inference*, vol. 37, pp. 269-277.

Deaton, A. and Laroque, G. (1992). 'On the Behavior of Commodity Prices.' *Review of Economic Studies*, vol. 59, pp. 1-23.

Engle, R. and Granger, C.W.J. (1987). 'Cointegration and Error Correction: Representation, estimation and testing.' *Econometrica*, vol. 55, pp. 251-276.

Gersowitz, M. and Paxon, C. (1990). 'The Economies of Africa and the prices of their exports.' *Princeton Studies in International Finance*. vol. 68, Princeton University, N.J.

Granger, C.W.J. and Newbold, P. (1974). 'Spurious regressions in econometrics.' *Journal of Econometrics*, vol. 2, No. 2, pp. 111-120.

Hazell, P.; Jaramillo, B.R. and Williamson, A. (1990). 'Relationship between World Price Instability and the Prices Farmers receive in Developing Countries.' *Journal of Agricultural Economics*, vol. 41, pp. 227-241.

Hendry, D. and Richard, J.F. (1983). 'Econometric Analysis of Economic Time-Series.' *International Statistical Review*. vol. 51, pp. 111-163.

Johnson, D.G. (1991). *World Agriculture in Disarray*. Second edition. St. Martin Press, New York.

Josling, T. E. (1980). 'Developed-country Agricultural Policies and Developing-country Food Supply: the case of wheat.' *Research Report*. International Food Policy Research Institute, Washington, DC.

Krueger, A. (1992). 'A Synthesis of the Political Economy in Developing Countries'. in A. Krueger, M. Schiff and A. Valdés (eds.), *The Political Economy of Agricultural Pricing Policy*, vol. 5, Johns Hopkins University Press, Baltimore, Md.

Krueger, A. and Duncan, R. (1993). 'The Political Economy of Controls: Complexity', *Working Papers Series # 4351*, NBER.

Krueger, A.; Schiff, M. and Valdés, A. eds. (1992), *The Political Economy of Agricultural Pricing Policy*. Johns Hopkins University Press, Baltimore, Md.

Lindert, P. (1989). 'Historical Patterns of Agricultural Policy.' in C.P. Temin (ed), *Agriculture and the State: Growth, Employment, and Poverty in Developing Countries*. Cornell University Press, Ithaca, NY.

Mc Grattan, E. (1990). 'Solving the Stochastic Growth Model by Linear Quadratic Approximation.' *Journal of Business and Economic Statistics*, vol. 8, pp. 21-36.

Mundlak, Y. and Larson, D. (1992). 'On the Transmission of World Agricultural Prices.' *World Bank Economic Review*, vol. 6, pp. 399-422.

Nickell, S. (1985). 'Error Correction, Partial Adjustment and all that: an expository note.' *Oxford Review of Economics*, vol. 47, pp. 119-129.

Phillips, P.C.B. and Loretan, M. (1991). 'Estimating Long run Economic Equilibria.' *Review of Economic Studies*, vol. 55, pp. 407-436.

Schiff, M. and Valdés, A. (1992). 'A Synthesis of the Economics in Developing Countries.' In Krueger, A.; Schiff, M. and Valdés, A. (eds.), *The Political Economy of Agricultural Pricing Policy*', vol. 4, Johns Hopkins University Press, Baltimore Md.