Asymmetric Price Adjustments and Behavior Under Risk: Evidence from Peruvian Agricultural Markets (*)

Jorge M. Agüero
jmaguero@wisc.edu
University of Wisconsin-Madison
427 Lorch Street
Madison, WI 53706

April 26, 2004

Abstract

Most studies measuring asymmetric adjustments in vertical price transmissions fail to provide empirical support to explain such behavior. The literature invokes theoretical models, which derive asymmetric behavior based on variables that are difficult to measure such as oligopolies' coordination policies, market imperfections or menu costs. Therefore, with no empirical support explaining the asymmetries, these studies leave no room for policy implementation. In this paper I relate asymmetric price responses to a theory of behavior under risk driven by the perishable rate of the goods. Retailers of a perishable good facing an increment in the wholesale price may decide not to increase their prices for fear of being left with a spoiled product. Using three agricultural products with different perishable rates I reject the null hypothesis of symmetric adjustments in the most perishable product but fail to reject for the less perishable goods. The nonlinear responses are consistent with the prediction of the model. The test for asymmetries uses a threshold cointegration technique where the threshold level and the cointegration vector are estimated from the data instead of being imposed by the econometrician.

Selected Paper prepared for presentation at the American Agricultural Economics Association Annual Meeting, Denver, Colorado, July 1-4, 2004

Copyright 2004 by Jorge Agüero. All rights reserved. Readers may make verbatim copies of this document for non-commercial purposes by any means, provided that this copyright notice appears on all such copies.

---

[*] I would like to thank Bruce Hansen and Michael Carter for their comments and Javier Escobal from the Group of Analysis for Development (GRADE) in Peru for allowing me to use the data for this paper. All errors are mine.

1. Introduction

Do prices rise faster than they fall? In the last twenty years several studies (as well as the general public[1] and governmental offices[2]) have tried to answer this question. Focusing on the vertical transmission of prices from producers to retailers or from wholesalers to retailers, these studies cover a vast number of industries and countries. For example, Borenstein, et al (1997) study how gasoline prices in U.S. cities respond to international crude oil prices. Eckert (2002) conducts a similar study for Canada. Palaskas (1995) focused on the price transmission of agricultural products within the European Union. The recent article by Peltzman (2000) takes these studies to the limit. He tries to answer the above question analyzing 242 products in the U.S. He claims that the answer is yes: output prices tend to respond faster to input increases than to decreases. Therefore, he argues, the adjustments are asymmetric.

However, it is not clear in the literature what causes these asymmetries. Most of the papers present a variety of possible arguments but no evidence to support it. The arguments include search cost with locally imperfect market (Blinder et.al. 1998), trigger price models of oligopolistic coordination (Borenstein et.al. 1997) as well as menu costs (e.g. Azzam 1999; Blinder 1982)[3]. One of the reasons for the lack of empirical support for these theories is the difficulty to measure variables such as the coordination mechanism of oligopolies (assuming we can identify the oligopoly firms) or the extent of

[1] On January 9, 1999 the following headline appeared in The New York Times "The Great Pork Gap: Hog Prices Have Plummeted. Why Haven't Store Prices?" cited by Peltzman (2000).

[2] The United States Agricultural Marketing Act of 1946 directs the U.S. Department of Agriculture to measure, analyze, and disseminate farm-to-retail price spread data according to the 7 U. S. Code 1622(b)[2]. This Act motivated a stream of papers analyzing price spread (see Hahn, 1990).

[3] For an extensive discussion of other models generating asymmetric price transmissions see Cramon-Taubadel (1998)

the menu costs. In fact, when *proxies* for those variables were included in a regression to explain the degree of asymmetric responses of retail prices, the results show no statistical significance (Peltzman 2000). The major problem is that when evidence of asymmetric behavior is found but it is not possible to isolate or identify the causes of such behavior, we leave no room for policies to correct these asymmetries.

Ward (1982) presented an alternative explanation for asymmetric price adjustments based on behavior under risk. He argues that retailers of perishable goods may decide not to increase their prices for fear of being left with a spoiled product. For instance, when wholesale prices increase we should expect a less than proportional increment in retail prices, or a lagged increment, or maybe no changes at all. When, on the other hand, wholesale price, decrease, we should expect a faster or bigger reduction, creating an asymmetric adjustment. The perishability of a product is much easier to measure than the variables mentioned above. It is clear that with no refrigeration, a tomato will spoil faster than rice. This variable (the perishable rate of the product) was not included in the regression estimated by Peltzman. However, in the U.S. most of the retailers have access to refrigeration systems, reducing the spoil rate of products which could probably reduce the chances of having a significant impact on Peltzman's regression.

This is not the case in poor countries. In Peru, for example, most of the food for consumption is purchased fresh at street markets where there is almost no refrigeration. It is possible then to test for asymmetric price adjustments of agricultural products and relate the potential variation in asymmetric adjustments with the different perishable rates of these products. This is the objective of this paper.

This paper has three contributions to the literature testing for asymmetric price adjustments. First, I use data from Lima, the capital of Peru, for three agricultural products with different perishable rates: tomatoes, potatoes and rice. I focused on the wholesale to retail price adjustments and explore whether the existence of asymmetric adjustments are related to the perishability of these products. My findings provide empirical support to Ward's hypothesis. I found evidence of asymmetric price adjustments in tomatoes but not in rice or potato.

Second, the high perishable rate of tomatoes requires the use of high frequency data. I use a new released dataset, collected by the Peruvian Ministry of Agriculture, which includes daily prices from January 1, 1995 to July 18, 2001. The use of daily data allows me to capture a more accurate relation between wholesale and retail prices.

Third, this article uses a recent methodology developed to test for threshold cointegration where the threshold is estimated from the data instead of being imposed by the econometrician. Recent studies have shown that traditional approaches to test for asymmetric adjustments are inadequate because they do not take into account the cointegration or long run relation between wholesale and retail prices. Abdulai (2002) proposed to use the method described by Enders and Granger (1998) and he applies it to test for asymmetries in the Swiss pork market. Threshold cointegration methods allow us to test whether "increases in producer prices that lead to declines in marketing margins are passed more quickly to retail prices than decreases in producer prices that result in increases in the marketing margins." (Abdulai 2002, p. 679.) However, Enders and Granger (1998) assume that the cointegration vector and the threshold level are known by the econometrician. Instead, I used the estimation proposed by Hansen and Seo (2002) that allows for an estimation of both parameters from the data.

The rest of the paper is organized as follows: the next section describes the econometric methods comparing Enders and Granger (1998) and Hansen and Seo (2002). Section 3 describes the data used in this study. Section 4 shows that when the cointegration vector and the threshold parameter are estimated from the data I found evidence rejecting the hypothesis of symmetric adjustments in high perishable products (tomatoes) but the test fails to the reject it in long lasting products (such as potatoes and rice). The conclusions and possible extensions of the paper are discussed in section 5.

2. Measuring asymmetric price transmissions

Several studies attempting to measure asymmetric price transmission focused on the estimation of the following reduced form equation:

$$\begin{aligned}&\Delta x_{1t} = \alpha_0 + \sum_{i=1}^{I} \alpha_{1i} d_t \Delta x_{2t-i} + \sum_{i=1}^{I} \alpha_{2i} (1 - d_t) \Delta x_{2t-i} + e_{1t} \\ &\text{with} \quad d_t = 1(z_t \leq \gamma)\end{aligned} \qquad (1)$$

where Δx_{1t} and Δx_{2t} represent changes in retail and wholesale prices, respectively, and d_t is an indicator function that is equal to one when a variable z_t does not exceed the threshold parameter γ. Usually authors set $z_t = \Delta x_{2t-1}$ and $\gamma = 0$. In this case they want to test if the response of retail prices differs depending on whether wholesale prices increase or decrease. The null hypothesis of symmetric price responses is:

$$\sum_{i=1}^{I} \alpha_{1i} = \sum_{i=1}^{I} \alpha_{2i} \qquad (2)$$

Equation (1) and this test are used, for example, in Peltzman (2000).

This equation assumes that changes in retail prices are not affected by their past values. It also assumes that the *causality* (Granger's sense) goes from wholesale to retail prices only. It is possible, however, to test for both restrictions using a *vector autoregressive model* (VAR) instead. However, as some authors have noticed[4], when $x_t = (x_{1t}\ x_{2t})$ is a I(1) time series with a cointegration vector β, then the use of equation (1) to test for asymmetric price responses is not adequate. The existence of a cointegration process imposes a restriction on a multivariate version of equation (1) as follows:

$$\begin{aligned} &\Delta x_t = \pi_0 + \pi_1 w_{t-1} + \sum_{i=1}^{I} \pi_{2i} \Delta x_{t-i} + \varepsilon_t \\ &\text{with} \quad w_{t-1} = \beta' x_{t-1} \end{aligned} \qquad (3)$$

This equation is called a *vector error-correction model* (VECM) where w_t is defined as "the deviations from the long run equilibrium", with the vector-parameter π_1 measuring the adjustments to the long run equilibrium. As stated in equation (3) these adjustments are linear. Balke and Fomby (1997) present a model that allows for a non-linear adjustment to the long run equilibrium by introducing the concept of *threshold cointegration*. The presence of threshold cointegration will alter equation (3) in order to allow for two regimes:

$$\begin{aligned} \Delta x_t &= \pi_{01} d_t + \pi_{11} w_{t-1} d_t + \sum_{i=1}^{I} \pi_{21i} d_t \Delta x_{t-i} \\ &\quad + \pi_{02}(1-d_t) + \pi_{12} w_{t-1}(1-d_t) + \sum_{i=1}^{I} \pi_{22i}(1-d_t)\Delta x_{t-i} + \varepsilon_t \qquad (4) \\ \text{with} &\quad d_t = 1(z_t \le \gamma) \end{aligned}$$

As before, d_t is an indicator function that defines the different regimes. The goal of this paper is to test for asymmetric price transmission using the threshold cointegration

[4] See von Cramon-Taubadel (1998).

techniques. At least two different approaches have been proposed to estimate equation (4) and they differ on their assumption about the knowledge the econometrician has on the cointegration vector β and the regime threshold parameter γ. The methods I will present below allow z_t to be equal to w_{t-1} or Δx_{2t-1}. I will consider only the case of $z_t=w_{t-1}$ for ease of exposition[5].

The first approach was introduced by Enders and Granger (1998) and it is used by Abdulai (2002) to test for asymmetric price transmission[6]. This methodology assumes that the cointegration vector β and the threshold parameter γ are both known by the econometrician. They also assume that $\pi_{21i}=\pi_{22i}$ for all i=1,…,I. This means that only the adjustments to the equilibrium change with the regimes while the autoregressive parameters of equation (4) remain constant.

To test for threshold models Enders and Granger (1998) propose the following steps given that β is known and setting $\gamma=0$:

Step 1: Regress w_t against a constant and/or a trend if *necessary*. The authors suggest the inclusion of a trend when "the data may have a clear trend" (p. 307). Let μ_t be the residuals of this regression.

Step 2: Define $d_t=1(w_{t-1}\leq 0)$ and run $\mu_t = d_t\rho_1\mu_{t-1} + (1-d_t)\rho_2\mu_{t-1} + \sum_{i=1}^{J}\Delta\mu_{t-i} + \xi_t$

Step 3: Using an F-test, evaluate the null hypothesis that μ_t is a random walk process by testing $\rho_1=\rho_2=0$. The authors presented critical values to evaluate this hypothesis.

[5] Enders and Granger (1998) call the process a momentum threshold autoregressive (M-TAR) when $z_t=\Delta x_{2t-1}$. Hansen and Seo (2002) say that this case is possible to be developed with their methodology but they focus only on $z_t=w_{t-1}$.

[6] Strictly speaking, Enders and Granger do not use the term *threshold cointegration*. They talk about threshold autoregressive (TAR) models.

Step 4: If the null hypothesis is rejected, test the restriction that the adjustments are symmetries (the null hypothesis $\rho_1=\rho_2$) using the usual F statistic. Enders and Granger suggest performing some diagnostic tests to assure that the estimate of ξ_t is correctly characterized as a white noise process. If not, they suggested repeating Step 2 using J+1 lags instead of J.

Step 5: If the null hypothesis of symmetry is rejected then estimate a VECM as defined in equation (4) imposing $\pi_{21i}=\pi_{22i}$ for all i=1,…,I as mentioned above.

Step 6: Evaluate the null hypothesis of symmetric price responses by testing $\pi_{11}=\pi_{12}$ using the likelihood ratio test. When this equality is true, adjustments to the long run equilibrium are the same whether the deviations from the long run equilibrium were positive or negative. If the null hypothesis is rejected there is evidence of asymmetric price responses.

The other approach to test for threshold cointegration is developed by Hansen and Seo (2002). Their methodology assumes β and γ are unknown parameters. Both parameters are estimated from the data. They do not impose the restriction that $\pi_{21i}=\pi_{22i}$ for all i=1,…,I. Therefore their proposed VECM is exactly the one described in equation (4) with $z_t=w_{t-1}$. They note that when β and γ are known the matrix-parameter π in (4) as well as the matrix $\Sigma=E(\varepsilon_t\varepsilon_t')$ can be estimated applying OLS to each price equation in (4). So they propose the following algorithm (Hansen and Seo 2002, p. 299):

Step I: Form a grid on $[\gamma_L,\gamma_U]$ and $[\beta_L,\beta_U]$ based on linear estimates of β[7] and the restriction that $\theta_0\leq Prob(wt-1\leq\gamma)\leq 1-\theta_0$, where $\theta_0>0$ is an arbitrary trimming parameter.

[7] They proposed the used of any consistent estimate of β, for example using Johansen MLE.

Step II: For each value of (β,γ) on those grids compute $\hat{\pi}(\beta,\gamma)$ and $\hat{\Sigma}(\beta,\gamma)$. Note that both sets of estimated parameters are functions of (β,γ).

Step III: Find $(\hat{\beta},\hat{\gamma})$ as the values of (β,γ) that minimizes the value of $\log|\hat{\Sigma}(\beta,\gamma)|$ which is the likelihood function after Step II.

Step IV: Set $\hat{\Sigma}=\hat{\Sigma}(\hat{\beta},\hat{\gamma}), \hat{\pi}=\hat{\pi}(\hat{\beta},\hat{\gamma})$ as the final estimated parameters including $(\hat{\beta},\hat{\gamma})$.

The advantage of this methodology is that (β,γ) are estimated from the data, not just assumed by the econometrician. The problem, however, is that Hansen and Seo (2002) do not provide a proof of consistency nor a distribution theory for the maximum likelihood estimation they proposed (p. 294). However they do provide the asymptotic null distribution of the Sup-LM test to evaluate the hypothesis of threshold cointegration and they also present two methods to calculate the p-values for the test (p. 301-305). If the null hypothesis of no threshold effect is rejected then:

Step V: Test for the null hypothesis of symmetric price adjustments ($\pi_{11}=\pi_{12}$). If rejected there will be evidence of asymmetric price transmission between retail and wholesale prices.

In this paper I test for asymmetric price responses comparing both methodologies described above. Note that under these two methodologies the test is no longer whether retail prices react similarly to any positive or negative changes in wholesale prices as defined in equation (2). The test for asymmetric price transmission now is whether retail prices react different to positive or negative deviation from the long run equilibrium. This is the adequate measure when the series are integrated of degree one and have a cointegration vector. The data used to test for prices transmission is described in the next section

3. Data

The data used in this paper comes from the Peruvian Ministry of Agriculture (MINAG). Starting in January 1995 the MINAG put together a dataset that included retail and wholesale prices in 28 cities nationwide. In each city, an array of 52 agricultural products was included[8]. The original intention of the MINAG was to have a daily dataset for retail and wholesale prices for each of the 52 products in each of the 28 cities. This is a major goal for datasets collection in a poor country such as Peru.

Unfortunately, because the collection of information was administrated by local agencies of the MINAG, the quality of the data changes from city to city. In some of the cities visits to markets were not performed daily. It is because of this and by parsimonious reasons that I focus on only one city: Lima. Lima is the capital of Peru, harboring at least a third of the economic activity and is the first city in terms of population. The city of Lima also presents the lowest rate of missing data. In the years covered in this paper weekends alone (without including holidays) represent around 29% of the observations. The data available in Lima for the products included in this paper has less than 23% of the total number of observations lost because of missing data.

As discussed in the introduction, one goal of this paper is to evaluate if the existence of possible asymmetries are related to perishable rates. For these reason I used three products that have different perishable rates: tomatoes, potatoes and rice. It is

[8] The complete list of cities and products is available upon request. For access to the complete dataset please contact Javier Escobal at the Group of Analysis for Development (http://www.grade.org.pe).

important to note that most of the food traded for consumption in Peru is fresh with almost no refrigeration. Thus, while rice can be stored in order to speculate with the price, such an action will be riskier in the case of tomatoes when there is no refrigeration available.

The sample used in this paper contains retail and wholesale prices of tomato, potatoes and rice in the city of Lima. The data is daily from January 1, 1995 to July 18, 2001. Table 1 presents the basic statistics of the data. From this table and Figure 1 we can observe that tomatoes are the product with the highest volatility in both retail and wholesale prices, while the rice has the lowest. Table 2 presents two definitions of spread. First the *(absolute) spread* is equal to the difference between the retail and wholesale price. Second, the *relative spread* is equal to the absolute spread divided by the wholesale price multiplied by 100. The results of Table 2 show that the difference in volatility coincides with differences in spreads. Thus, rice has the lowest spread in absolute and relative terms while tomatoes present the highest. Potatoes prices are somewhere in the middle for both indicators. In the next section I use the data described here to test for asymmetric price transmission using the methods described in section 3.

Table 1
Basic Statistics: 1995-2001
(Prices per Kilogram in Nuevos Soles)

Product	Type	N	Missing values (%)	Mean	Std. Dev	Minimum	Maximum	Coefficient of variability
Tomatoes	Wholesale	1850	22.6	0.93	0.40	0.25	2.58	42.8
	Retail	1850	22.6	1.68	0.58	0.50	4.00	34.8
Potatoes	Wholesale	1882	21.3	0.59	0.24	0.21	1.82	41.4
	Retail	1882	21.3	0.92	0.29	0.46	2.33	31.1
Rice	Wholesale	1850	22.6	1.35	0.24	0.86	1.95	18.0
	Retail	1850	22.6	1.78	0.32	1.07	2.50	18.2

Note: Between 1995 and 2001 the average exchange rate was US $1=2.92 Nuevos Soles
Coefficient of variability = Std. Dev/Mean x 100

Table 2
Spreads by products: 1995-2001
(Prices per Kilogram in Nuevos Soles)

	Absolute Spread: Nuevos Soles [1/]				Relative
Product	Mean	Std. Dev	Minimum	Maximum	Spread (%) [2/]
Tomatoes	0.75	0.27	0.04	2.03	88.4
Potatoes	0.33	0.09	0.10	0.75	64.4
Rice	0.43	0.16	0.13	0.91	32.5

[1/] Retail minus wholesale price

[2/] Absolute spread divided by wholesale price x 100.

Note: Between 1995 and 2001 the average exchange rate was US $1=2.92 Nuevos Soles

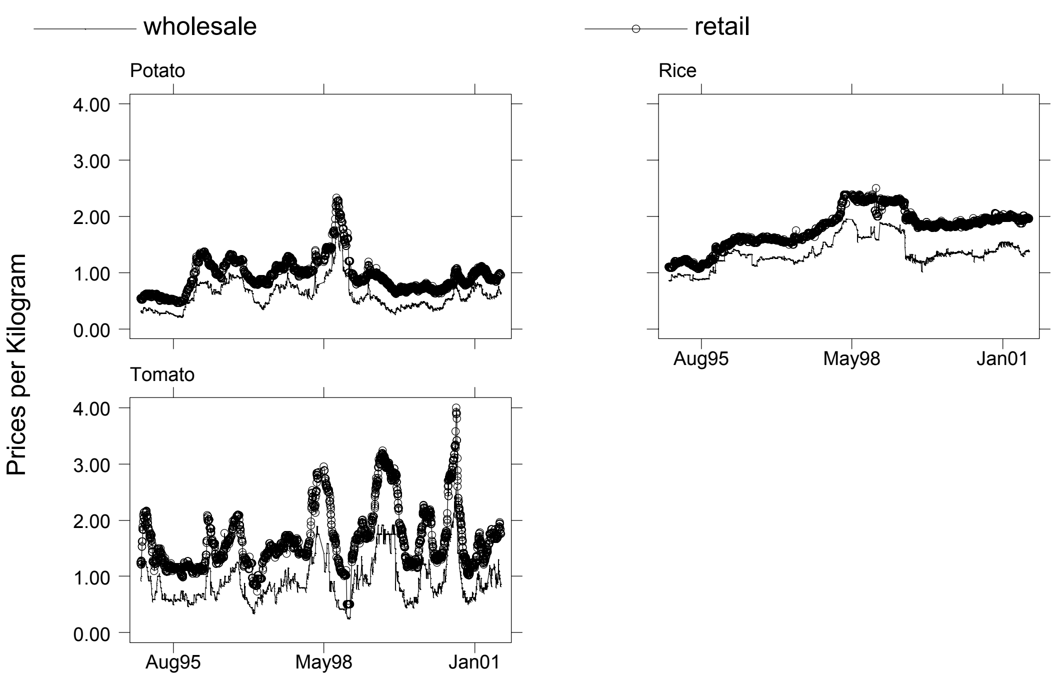

Figure 1. Evolution of Prices: 1995-2001

4. Results[9]

As described above, the major problem with the traditional methodology described in equations (1) and (2) to test for asymmetric price responses occurs when the retail and wholesale prices are integrated of order one. The first step, then, is to test for the presence of unit root process in each of the six variables under analysis.

To perform this test I used the Augmented Dickey-Fuller (ADF) statistic selecting the autoregressive lags by minimizing the Akaike and Schwartz information criteria (AIC and BIC, respectively). The test was performed in three different ways, but always including a constant. In the first case a trend was added, in the second case eleven dummy variables were included to account for seasonal effects; and in the third case a trend and the seasonal dummies were included. In all cases the test failed to reject the null hypothesis of a unit root. I also found that in all cases the ADF test rejected the null hypothesis that the first differences of the variables follow a unit root process, so all the variables are integrated of degree one only. I also test if the vector $\beta=[1,-1]$ is a cointegration vector using the Engle and Granger (1987) approach. For all the three products I reject the null hypothesis of no cointegration. It is important to keep in mind that throughout all the paper I define x_{1t} as the retail price and x_{2t} as the wholesale price, therefore, $\beta=[1,-1]$ is a cointegration vector meaning the (absolute) spread is stationary. The results of the unit root and cointegration tests are not reported here to save space but are available upon request.

---

[9] In developing the code to perform the test described in Hansen and Seo (2002) I benefit from the authors ' generosity to make their Gauss code publicly available. Their code can be downloaded from http://www.ssc.eisc.edu/~bhansen

An important element of this series is the seasonality they present. When each variable was regressed against a constant an eleven seasonal dummies I reject the null hypothesis that the dummy variables are jointly equal to zero[10]. Therefore, I test for asymmetric adjustment using the seasonally-adjusted data, that is, the residual of the above regression.

I present now the results of applying the method of Enders and Granger (1998). Table 3 follows the steps described in section 2. As previously discussed this procedure takes w_t -the deviation from the long run equilibrium- as given. I first estimated the cointegration vector β using Engle-Granger two-step estimator. I call this *Step 0*. Step 1 was not necessary in this case.

As shown in Step 3, for all products the null hypothesis of no cointegration ($\rho_1=\rho_2=0$) is rejected. Step 4.1 strongly rejects the null hypothesis of no threshold cointegration ($\rho_1=\rho_2$) in the case of potatoes and rice but fails to reject in the most perishable product in the sample: tomatoes[11]. These results contradict Ward's argument that the most perishable products will display more asymmetric adjustments. However, I obtained the opposite results –asymmetries in the most perishable goods- when the threshold parameter is not imposed but estimated using Hansen and Seo (2002).

[10] The results of the tests are reported in table A.1 in the appendix.
[11] As display in Step 4.2, diagnostic tests over the residuals show no evidence against a white noise process using the Ljung-Box statistics as recommended by the authors.

Table 3
Tests for threshold cointegration: Enders-Granger

Step	Estimates	Products Tomatoes	Potatoes	Rice
0	Cointegration vector β	1.329	1.133	1.259
	(Standard errors)	(0.013)	(0.009)	(0.012)
2	ρ_1	-0.108	-0.060	-0.015
	(Standard errors)	(0.016)	(0.015)	(0.009)
	ρ_2	-0.083	-0.056	-0.033
	(Standard errors)	(0.013)	(0.013)	(0.009)
3	F-test (ρ_1=ρ_2=0)	40.235	59.975	38.534
4.1	F-test (ρ_1=ρ_2)	1.459	52.695	54.284
	(p-value)	(0.227)	(0.000)	(0.000)
4.2	Ljung-Box statistic (1)	0.016	0.026	0.003
	(p-value)	(0.901)	(0.871)	(0.960)
	Ljung-Box statistic (7)	5.957	0.264	9.901
	(p-value)	(0.545)	(1.000)	(0.194)
	Ljung-Box statistic (14)	10.203	5.875	17.760
	(p-value)	(0.747)	(0.970)	(0.218)

Note: Null hypothesis of the Ljung-Box (q): First q autocorrelations of the residuals are jointly equal to zero. The Critical values for F-test (ρ_1=ρ_2=0) are 3.04, 3.75 and 5.36 for 10%, 5% and 1%, respectively. See Enders and Granger (1998)

Recall that the first three steps of Hansen and Seo (2002) are related to setting the grid for β and γ and finding their values that maximize the likelihood function[12]. Those optimal values are reported in Table 4. In Step III, the point estimate of γ is different from zero. Even though Hansen and Seo did not provide a way to compute the standard deviations of this parameter the indicator function $d_t=1(w_{t-1}\leq\gamma)$ could differ substantially for various values of γ. Furthermore, in the case of tomato, when $\gamma=0$ the mean value for d_t, i.e. the proportion of observations in the first regime, is around 0.56, while when γ=-

[12] I used Johansen MLE as a consistent estimator for β and set the trimming parameter $\theta_0=0.05$.

0.308 as reported in Table 4, the mean is around 0.09. Hansen and Seo (2002) found similar skewed regimes when applied to the term structure of interest rates. This alternative estimation of γ obtained from the data is generating very different regimes than when assuming γ=0 a priori as in Enders and Granger (1998)[13].

Table 4
Tests for threshold cointegration: Hansen - Seo

		Products		
Step	Estimates	Tomatoes	Potatoes	Rice
III	Cointegration vector β	1.558	1.174	0.742
	Threshold parameter γ	-0.308	0.119	0.233
V	SupLM test	27.620	18.614	24.071
	Fixed Regressor	26.571	25.607	26.278
	(p-value)	(0.046)	(0.464)	(0.116)
	Residual Bootstrap	27.565	27.374	26.202
	(p-value)	(0.050)	(0.548)	(0.138)

Note: The grid search for β and γ was performed over 100 gridpoints for each parameter. The Bootstrap values were computed with 500 simulation replications. The null hypothesis for the SupLM test is no threshold effects.

Step IV corresponds to the estimation of the VECM parameters. I will focus first on the test for threshold cointegration. In Step V, I reject the null hypothesis of no threshold cointegration for tomatoes and fail to reject for potatoes and rice. These results support the hypothesis that the more perishable products will display more asymmetric adjustments.

Table 5 shows the VECM estimates for tomatoes in the presence of threshold cointegration[14]. The parameters do not change much between regimes, mostly because

[13] Figure A.1 in the Appendix shows that γ=0 is "far" from maximizing the likelihood function.
[14] The VECM estimation for the other products is available upon request.

they are statistically equal to zero, but Hansen and Seo (2002) recommend that since there is no formal distribution theory for these estimates and their standard errors "the results should be interpreted somewhat cautiously" (p. 311).

Table 5
Vector Error-Correction Model: Tomatoes

	Dependent Variables			
	Δ Retail price $_t$		Δ Wholesale price $_t$	
Variables	Regime 1	Regime 2	Regime 1	Regime 2
Intercept	-0.043	-0.001	-0.021	0.001
	(0.035)	(0.002)	(0.033)	(0.002)
W_{t-1}	-0.195	-0.081	-0.047	0.007
	(0.082)	(0.010)	(0.079)	(0.011)
Δ Retail price $_{t-1}$	-0.042	-0.125	-0.037	0.008
	(0.147)	(0.037)	(0.155)	(0.028)
Δ Retail price $_{t-2}$	0.180	-0.012	0.219	-0.041
	(0.143)	(0.025)	(0.242)	(0.025)
Δ Wholesale price $_{t-1}$	-0.049	-0.006	-0.343	-0.053
	(0.100)	(0.034)	(0.128)	(0.045)
Δ Wholesale price $_{t-2}$	-0.017	-0.022	-0.114	0.021
	(0.112)	(0.027)	(0.099)	(0.030)
Obs. in Regime 1 (%)	8.6			
Obs. in Regime 2 (%)	91.4			
No. Observations	1847			
Log Likelihood	-9934.2			

Hypothesis	Wald test	P-value
Equality of Autoregressive coefficients	7.03	0.534
Equality of Error Correction coefficients	1.95	0.378

Note: Eicker-White standard errors in parenthesis.
Regime 1: W_{t-1}<=-0.308, Regime 2: W_{t-1}>-0.308

Notice that the adjustments to the long run equilibrium are driven more through the retail prices than the wholesale prices. The error-correction has minimal effects on ΔWholesale$_t$ in both regimes and on ΔRetail$_t$ in the first regime; the parameter is on the borderline of statistical significance, but this is not the case in the second regime. The dynamics are minimal too. The authors found similar behavior in their applications. This would suggest that prices changes are close to a white noise process, so the prices in levels are close to random walks with no drift, at least for wholesale prices.

Finally, following Ward (1982), this paper argues that the reasons to find asymmetric price responses in perishable goods is that agents in possession of these products may decide not to increase their prices for fear of being left with a spoiled product. I argue that if that is the case when wholesale prices increase we should expect a less than proportional increment in retail prices, or a lagged increment, or even no changes at all. When, on the other hand, wholesale prices decrease, we should expect a faster or bigger increment creating an asymmetric adjustment. Figure 2 shows these patterns.

Figure 2
Price Response to Error-Correction: Tomato

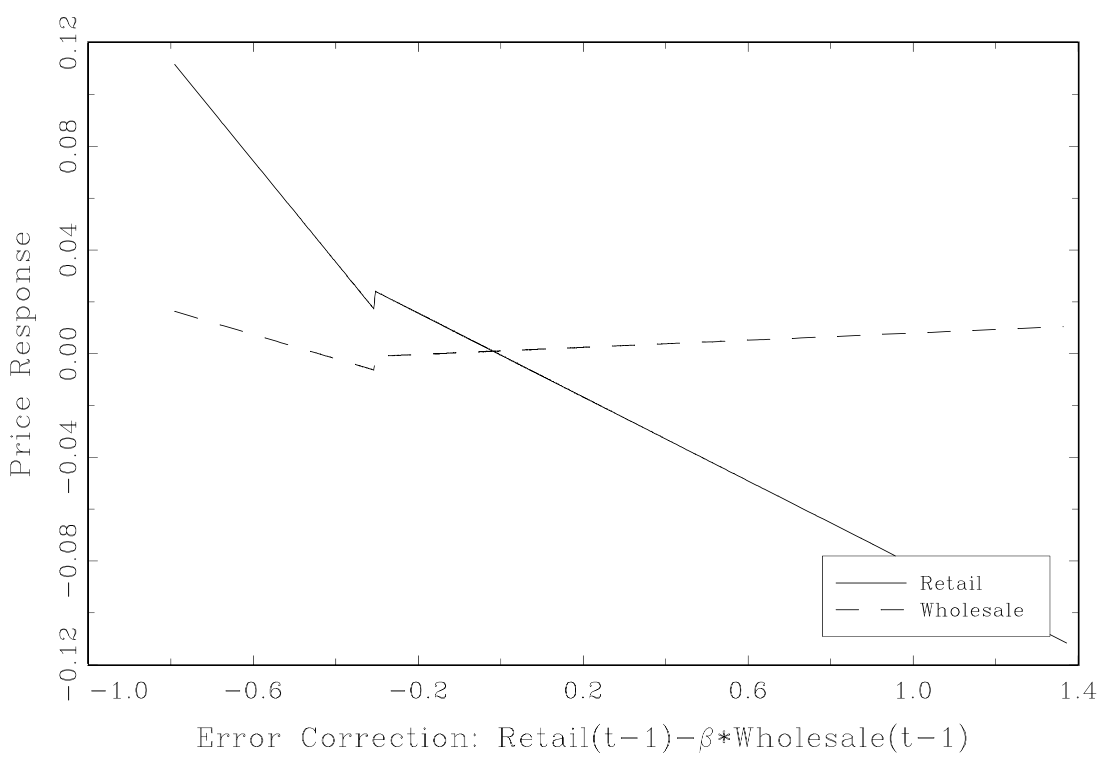

Figure 2 plots the predicted values of $\Delta Wholesale_t$ and $\Delta Retail_t$ as a function of w_{t-1} using the parameter reported in Table 5, holding all the autoregressive variables constant. First, as discussed above, in both regimes the parameters associated to the (lagged) error-correction are not statistically significant different from zero in the regressions for $\Delta Wholesale_t$. Thus, the dotted line should be seen as "flat". Second, Table 5 shows that in the first regime for $\Delta Retail_t$, the adjustments to the equilibrium are barely significant[15], so the retail price response when $w_{t-1} \leq -0.308$ is close to zero. This is not the case in the second regime where the t-test is -8.17. Notice also that the second regime ($w_{t-1} > -0.308$) retail prices decrease as shown in Figure 2. That means that if at t=1, *ceteris paribus*, the wholesale price of tomatoes is above the long-run equilibrium relation (i.e., where in regime 1) the retail price of tomatoes at t=2 will remain similar to its level at t=1. But when at t=1 wholesale is below the long-run equilibrium relation (regime 2) retail price will decrease significantly at t=2. These different changes describe an asymmetric adjustment. In the case of less perishable products (potatoes and rice) I found no evidence of asymmetric behavior.

5. Conclusions

This paper tests for asymmetric price adjustment in Peruvian agricultural markets. The major difference with respect to the previous literature is the intention of relating the existence of possible asymmetries to theories predicting such behavior. Most papers testing for nonlinear price adjustments fail to provide the corresponding empirical support to explain such behavior. The theories invoked by these papers rest their analysis on variables that are complicated to measure, such as oligopolies' coordination policies,

[15] The fact the sample is large here (1847 observation) requires the use stricter significance levels.

market imperfections or menu costs. Attempts to relate rough measures of these variables failed to show a significant correlation with asymmetric responses as reported in Peltzman (2000). This lack of identification about the causes of such asymmetries leaves policy makers with no instruments to correct them.

In this paper I relate asymmetric price responses to a theory of behavior under risk driven by the perishable rate of the goods. Ward (1982) argues that retailers of these products may decide not to increase their prices for fear of being left with a spoiled product. I found empirical evidence that can support this idea. Using three agricultural products that differ in their perishable rate I found evidence of asymmetric price adjustments to the long-run equilibrium in the case of tomatoes (the most perishable good in the sample) but not for the less perishable goods (potatoes and rice).

To test for these asymmetries I used threshold cointegration techniques, but where the cointegration vector and the threshold parameter are estimated from the data instead of being arbitrarily chosen by the econometrician.

The results presented in this paper can be extended to more products in order to have a more accurate identification of the relation between the risk induced by trading with perishable goods and the nonlinear price adjustments. Also, the technique used here needs a distribution theory for the parameters estimated as discussed in Hansen and Seo (2002). Finally, assuming the existence of only one threshold parameter and therefore only two regimes could be quite restrictive. It might be more adequate to estimate the VECM described in equation (5) in a semi-parametric way or in a nonlinear way as under the conditions described by De Long (2001). However, in both cases, the cointegration vector needs to be known ex-ante.

References

Abdulai, A. (2002) "Using threshold cointegration to estimate asymmetric price transmission in the Swiss pork market." *Applied Economics*, 34:679-687.

Azzam, A. (1999) "Asymmetry and rigidity in farm-retail price transmission." *American Journal of Agricultural Economics*, 81:525-533.

Balke, N. and Fomby, T. (1997) "Threshold cointegration." *International Economic Review*, 38: 627-645.

Blinder, A. (1982) "Inventories and sticky prices: more on the microfoundations of macroeconomics." *American Economic Review*, 72: 334-48.

Blinder, A.; E. Canetti; D. Lebow and J. Rudd (1998) *Asking About Prices: A New Approach to Understanding Price Stickiness*, Russel Sage Foundation, New York.

Borenstein, S. A., Cameron, A. C. and Gilbert, R. (1997) "Do gasoline prices respond asymmetrically to crude oil prices?" *Quarterly Journal of Economics*, 112:305-339.

De Long (2001) "Nonlinear estimation using estimated cointegration relations." *Journal of Econometrics*, 101: 109-122.

Eckert, A. (2002) "Retail price cycles and response asymmetry." *Canadian Journal of Economics*, 35(1):52-77.

Elitzak, H. (1999) "Food cost review, 1950-97." *Agricultural Economic Report No. 780* Food and Rural Economics Division, Economic Research Service, U.S. Department of Agriculture.

Enders, W. and Granger, C. W. J. (1998) "Unit-root tests and asymmetric adjustment with an example using the term structure of interest rates." *Journal of Business and Economic Statistics*, 16:304-11.

Hahn, W. F. (1990) "Price transmission asymmetry in pork and beef markets." *The Journal of Agricultural Economic Research*, 42:21-30.

Hansen, B. and B. Seo (2002) "Testing for two-regime threshold cointegration in vector error-correction models." *Journal of Econometrics*, 110: 293-318.

Palaskas, T. B. (1995) "Statistical analysis of price transmission in the European Union." *Journal of Agricultural Economics*, 46:60-69.

Peltzman, S. (2000) "Prices rise faster than they fall." Journal of Political Economy, 108(3): 466-502.

von Cramom-Taubadel, S. (1998) "Estimating asymmetric price transmission with the error-correction representation: An application to the German pork market." *European Review of Agricultural Economics*, 25:1-18.

Ward, R. (1982) "Asymmetry in retail, wholesale and shipping point prices for fresh fruits and vegetables." *American Journal of Agricultural Economics*, 62: 205-221.

Appendix

Table A.1
F-test. Null hypothesis: no seasonal effects

Product	Price	F-test	p-value
Tomato	Retail	27.65	0.000
	Wholesale	24.93	0.000
Potato	Retail	5.64	0.000
	Wholesale	7.14	0.000
Rice	Retail	2.11	0.017
	Wholesale	15.27	0.000

Note: Each variable was regressed against a constant and 11 seasonal dummies.

Figure A.1
Concentrated Negative Log Likelihood