Essays on Applied Econometrics: Agriculture and Development

by

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Abstract

This dissertation explores three topics on agriculture and development using different applied econometric techniques. In the first chapter, Non-linear regression models were used to estimate the effect of own and other taxa previous population levels, nitrogen application, and crop rotation on population dynamics of plant parasitic and non-parasitic nematodes using data from the Cullars rotation. Because field experimental data was used, a spatial component was included as populations in one plot were proved to be related to the population level of their neighbors. Own previous levels were found to be very important for all groups of nematodes and all of the groups had an interaction effect with at least one other group. Lesion and cotton root-knot nematodes were found to be competitive while Mononchidae, Dorylaimidae, microbivorous and lance nematodes were non-competitive. All of the populations showed high seasonality patterns having lower populations during winter, to then remain steady until September-October when there is a significant increase in the population of cotton root-knot, Dorylaimidae, microbivorous, and lesion nematodes. Nitrogen had a positive effect on Mononchidae, microbivorous, spiral, and cotton root-knot nematodes. The use of clover after cotton in the rotation crop program proved to be significantly better in reducing plant parasitic nematodes compared to other treatments.

The second chapter analyzed the market structure of Peruvian agricultural exports as Peru has become the largest fresh asparagus exporter, third in processed artichoke, and third in paprika. This may have generated market power but the exertion of it and towards whom has not

been studied yet. Pricing-To-Market (PTM) models tested for price discrimination in the Peruvian export market for these three goods. The results strongly suggested that Peruvian exporters were engaging in price discriminating behavior. Country markups were common and Peruvian exporters were stabilizing English pound prices in asparagus and Euro prices in paprika. Lastly, even though PTM found that fresh asparagus exporters were amplifying the effects of the exchange rate after the preferential free trade agreement between U.S. and Peru was established, the crucial assumption of constant marginal cost in the framework may be too restrictive to affirm that an incomplete exchange rate pass-through is happening.

The third chapter addressed a development topic related to the labor market in the U.S. In the U.S., Hispanic people carry a greater stigma regarding their immigration status compared to other ethnicities. This study focused on citizenship status' effect on wages using quantile wage regressions accounting for sample selection in the states of Alabama and Georgia. Wages for two groups were analyzed: Hispanics and non-Hispanics (all other ethnicities). Results suggested that using an average estimation of the effect of citizenship on wages for non-Hispanics does not reflect what is really happening in the different quantiles. For the Hispanic group, all the quantiles showed a significant positive coefficient (of having citizenship) and a higher magnitude in the lower one. However, this higher magnitude was not economically significant compared to the other quantiles. It is concluded that Hispanic are negatively affected by not having the proper immigration status regardless of the quantile they are in, thus an average effect regression provides a good fit for testing the impact of citizenship status on Hispanic people.

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List of Abbreviations

ANOVA	Analysis of Variance
EDM	Equilibrium Displacement Model
GMM	Generalized Method of Moments
HHI	Herfindahl-Hirschman index
IPUMS	Integrated Public Use Microdata Series
MC	Marginal Cost
MINAG	Peruvian Department of Agriculture (acronym in Spanish)
ML	Maximum Likelihood
OLS	Ordinary Least Squares
PTM	Pricing-to-Market
PTPA	U.S-Peru Trade Promotion Agreement
RE	Random Effects

Chapter 1:

Population Dynamics and Interactions between Plant Parasitic and Non-parasitic Nematodes: An Empirical Analysis

Abstract

Non-linear regression models were used to estimate the effect of own and other taxa previous population levels, nitrogen application, and crop rotation on population dynamics of Mononchidae, Dorylaimidae, microbivorous (Rhabditidae), lance (*Hoplolaimus galeatus*), spiral (*Helicotylencus dihystera*), stubby root (*Paratrichodorus minor*), lesion (*Pratylenchus zeae*), and cotton root-knot (*Meloigogyne incognita*) nematodes using data from the Cullars rotation, which is the oldest soil fertility experiment in the Southern United States. Because field experimental data was used, a spatial component was included as populations in one plot were proved to be related to the population level of their neighbors. Own previous levels were found to be very important for all eight groups of nematodes (all groups' current population relied heavily on its own previous population value) and all of the groups had an interaction effect with at least one other group. Lesion and cotton root-knot nematodes were found to be competitive while Mononchidae, Dorylaimidae, microbivorous and lance nematodes were non-competitive. All of the populations showed high seasonality patterns having lower populations during winter, to then remain steady until September-October when there is a significant increase in the population of cotton root-knot, Dorylaimidae, microbivorous, and lesion nematodes. Nitrogen had a positive

effect on Mononchidae, microbivorous, spiral, and cotton root-knot nematodes. The use of clover after cotton in the rotation crop program proved to be significantly better in reducing plant parasitic nematodes compared to other treatments.

1.1 Introduction

Agriculture is shifting from a classical/rigid farming system to the use of integrated crop production systems in order to become more efficient managing pests, financial risks, and environmental concerns. Integrated Pest Management is an important tool within these programs as it has been observed to impact positively agricultural yields and/or profits (Chavez et al., 2013; Fernandez-Cornejo, 1998; White and Wetzstein, 1995). In this context, knowledge of population dynamics of soil-borne organisms is critical to developing flexible, information-based farming systems that manage these organisms for maximum expected present value of profit over time.

Many published studies present statistical analyses of population dynamics of nematodes. However, the majority of these tend to use very simple procedures such as comparing the average of populations at different periods of time, or they base inferences on simple correlations. Multiple regression analysis is a statistically more powerful approach as it allows determination of partial effects of many variables that are believed to affect a dependent variable, thus better estimating impacts of treatments and decision variables. Examples of simple (one explanatory variable) linear regression models for the analysis of population dynamics of nematodes are (Jeger et al., 1993; McSorley and Gallaher, 1993), while examples of application of multiple regression models are (Bell and Watson, 2001; McGraw and Koppenhöfer, 2009; Taylor and Rodríguez-Kábana, 1999). An example of a dynamic decision model based on population dynamics estimated with a multiple regression model was published by Taylor and Rodríguez-Kábana (1999). Their decision model was based on dynamic population models for the root-knot nematode (*Meloidogyne arenaria*), Southern blight ("white mold") fungus (*Sclerotium rolfsii*), and microbivorous (Rhabditidae) nematodes estimated on the basis of fall observations in field experiments continued over several years.

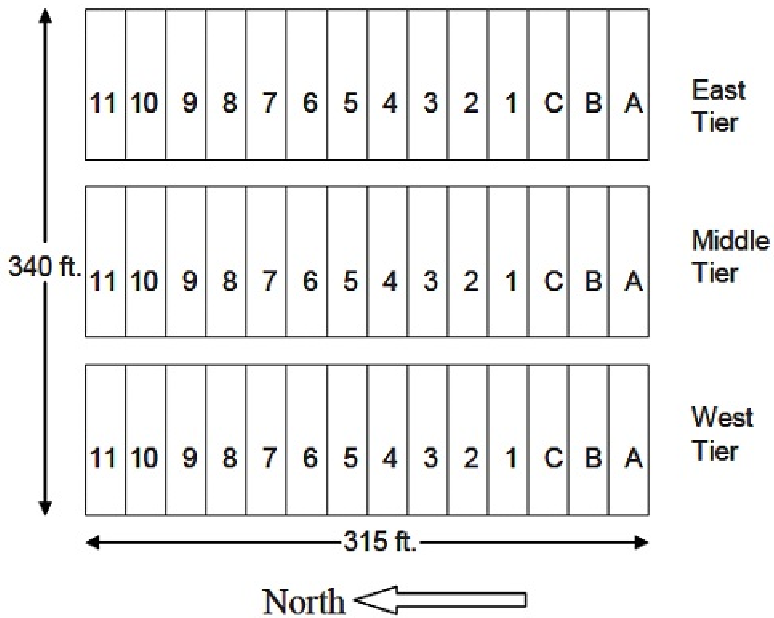
This article adds to the literature as it extends the population modeling approach introduced by Taylor and Rodríguez-Kábana (1999) to plant parasitic nematode species: *Helicotylenchus dihystera* (spiral), *Hoplolaimus galeatus* (lance), *Paratrichodorus minor* (stubby root), *Pratylenchus zeae* (lesion), and *M. incognita* (cotton root-knot) as well as to non-parasitic taxa within the Mononchidae, Dorylaimidae, and microbivorous (Rhabditidae) nematodes. In addition, the proposed statistical model is improved by including a spatial autocorrelation component in order to account for location effects between experimental samples. Dynamic models are estimated with field observations made approximately monthly from January of 1993 through April of 1996. Use of monthly observations, as opposed to annual observations, permits us to more closely examine the contemporaneous and dynamic interactions of the various genera of nematodes, and to analyze the seasonality of population growth and decline. .

The main objective of this study is to estimate the effect of the relevant factors affecting population dynamics of each of the aforementioned taxa. The hypothesis of this study is that nematode population dynamics are greatly affected by: crop rotation, seasonality, and its own and other nematode taxa previous population number.

1.2 Materials and Methods

1.2.1 Experimental Design

Nematode counts were taken from plots in the “Cullars Rotation”, which is the oldest soil fertility experiment in the Southern United States, and the second oldest continuous cotton experiment in the World (Auburn University, 2004). The experiment is located on Marvyn loamy sand (fine-loamy, siliceous, thermic Typic Kanhapludults) and, for our data set, used conventional tillage with moldboard plowing, disking, and regular cultivation. Soil samples for nematode analyses were obtained from four (A, B, 1, and 3) of the fourteen soil fertility regimes replicated three times in an ordered block design (see figure 1). The 3-year rotation sequence for all plots on the blocks was: (1) cotton followed by crimson clover, (2) corn grain followed by rye harvested the following season, and (3) soybean double cropped. Ten soil core-subsamples were taken from the rhizosphere of plants along the middle of each subplot at approx. 0.5 m spacing and to a depth of 25 cm with a 2.5-cm inner diameter Oakfield® soil probe (Model LS, Ben Meadows Co., Atlanta, GA). The cores were mixed thoroughly, and 100 cm3 of the mix was removed for nematode analysis using the "salad bowl" incubation technique (Rodriguez-Kabana and Pope, 1981). The soil fertility treatments for which nematode samples were taken all had lime, phosphorous and potassium added, but differed in nitrogen fertilization (N or no-N) and legumes (with or without) (see figure 1.1). Five subplots in each 6.1x30.2 meter plot were sampled. A total of 1680 observations (60 subplots during 28 periods) were used.



Plot	Description	Treatments N	P	K	S	Other	Soil pH
A	No N/+legume	0	+	+	+		6.1
B	No N/no legume	0	+	+	+	No legume	6.0
C	***No soil amendments***	***0***	***0***	***0***	***0***		***5.2***
1	No winter legumes/ + N	+	+	+	+	No legume	6.2
2	***No P***	+	0	+	+		6.2
3	Standard fertilization	+	+	+	+		6.1
4	4/3 K	+	+	+	+	Extra K	6.2
5	Rock phosphate	+	+	+	+	Rock P	6.0
6	***No K***	+	+	0	+		6.3
7	2/3 K	+	+	+	+		6.2
8	***No lime***	+	+	+	+	No lime	4.7
9	No S	+	+	+	0		6.2
10	Standard fertilization + micronutrients	+	+	+	+	+Zn,Cu, Mn, Fe, B & Mo	6.3
11	1/3 K	+	+	+	+		6.1

Figure 1.1 The Cullars Rotation.
Standard lime and fertilizer treatments:
- Limed to pH 5.8 to 6.5
- 101 kg. P_2O_5 per hectare per three-year rotation
- 269 kg. K_2O per hectare per three-year rotation
- 101 kg. N per hectare on cotton
- 134 kg. N per hectare on corn
- 67 kg. N per hectare topdress on small grain
- 45 kg. sulfate-S per hectare applied as gypsum to cotton and small grain

Source: Auburn University, 2004.

1.2.2 Population Model Specification

The functional form of the population model used for the empirical analyses is an augmented version of the model introduced and developed by Taylor and Rodríguez-Kábana (1999). The mathematical form of this population model is:

$$(\mathrm{P}_{it} + 1) = A_{im} \prod_{j=1}^{n} \left[\left(\mathrm{P}_{j,t-1} + 1 \right)^{b_{ij}} \right] e^{(\sum_k g_{ik} X_{ik} + u_{it})} \quad (1.1)$$

where P_{it} is population level of the i^{th} organism at time t; $\boldsymbol{\Pi}$ is the mathematical operator designating the product of the expression in brackets taken over all organisms, (i = 1, 2, ... n); X_{ik} is a set of k explanatory variables for the i^{th} organism, including nitrogen application; A_{im} is a set of monthly multiplicative intercepts to be estimated; b_{ij} and g_{ik} are other parameters to be estimated; and u_{it} is a random error term.

Since observations throughout the year were used to estimate the population models and since seasonality of nematodes is expected, the multiplicative intercept was allowed to vary depending on the month. As discussed in Taylor and Rodríguez-Kábana (1999), the multiplicative random error of this functional form does not allow for negative population levels and the multiplicative error specification allows higher populations of nematodes to randomly vary more compared to lower populations. Binary variables, X_{ik}, were included to represent the crop rotation sequence, nitrogen application, and whether the subplot was planted to a legume. The “+1” was added as it is possible to have organisms in period t even with zero organism in period t-1 and also because it allows equation (1.1) to be transformed into linear in parameters making the results easier to interpret. This transformation yields:

$$p_{it} = a_i + \sum_{j=1}^{n} b_{ij} p_{j,t-1} + \sum_k g_{ik} X_{ik} + u_{it} \quad (1.2)$$

where $p_{it} = \log(P_{it} + 1)$, $p_{j,t-1} = \log(P_{j,t-1} + 1)$, and $a_i = \log(A_{im})$.

An improvement to this model was made by adding a spatial lag of the dependent variable to the right-hand side of equation (1.2). The spatial lag enters the model as part of the error term. Spatial data contains information on the location of each one of the observations and how they interact with one another. There is spatial autocorrelation in a system if observations that are closer to each other in space have related values. This does not mean that nematodes "move" from one plot to another, it just means that plots that are contiguous share characteristics (such as soil, water, latitude/longitude, etc.) that make them similar. Tobler (1970) expressed this concept as "Everything is related to everything else, but near things are more related than distant things". All of the population models expressed in equation (1.2) tested positive for spatial correlation. The spatially lagged dependent model, also known as spatial autoregressive (SAR) model, is appropriate when it is believed that the values of the dependent variable in one unit i are directly influenced by the values of the dependent variable found in i's neighbors, and that the effect is not just some type of clustering (Ward and Gleditsch, 2008). A binary distance matrix (neighbors=1, not neighbors=0) was used for to statistically implement the neighbor effect into the regression model. Accounting for the spatial relationship, equation (1.2) becomes:

$$p_{it} = a_i + \sum_{j=1}^{n} b_{ij} p_{j,t-1} + \sum_k g_{ik} X_{ik} + v_{it} + \theta w_i p_{it} \qquad (1.3)$$

Where the previous random error term in equation (1.2) has been disaggregated into a stochastic component, v_{it}, and a spatial component, $\theta w_i p_{it}$, in which θ is the parameter to be estimated and w_i is the distance/connectivity matrix.

Interactions for a set of organisms can best be mathematically examined using matrix algebra, which for a set of equations based on equation (1.3), can be written as,

$$P_{(t)} = A + BP_{(t-1)} \qquad (1.4)$$

where $P_{(t)} = [p_{1t}\ p_{2t}\ p_{3t} \ldots p_{nt}]'$; $A = [a_1\ a_2\ a_3 \ldots a_4]'$;

and where
$$B = \begin{pmatrix} b_{11} & b_{12} & b_{13} \ldots & b_{1n} \\ b_{21} & b_{22} & b_{23} \ldots & b_{2n} \\ . & & & . \\ . & . & & . \\ . & & . & . \\ . & & & . \\ b_{n1} & b_{n2} & b_{n3} \ldots & b_{nn} \end{pmatrix}$$

Another advantage of this system of equations is that the "community matrix" often referred to in theoretical biological literature (Taylor and Rodríguez-Kábana, 1999) is simply the matrix $(B - I)$, where the i^{th} diagonal element of this matrix is (b_{ii} - 1), and the off-diagonal element is b_{ij}, with $(j \neq i)$[1]. Note that population levels in equation (1.4) are in logarithmic form. In equilibrium, $P_{it} = P_{i,t-1}$, which implies that $\log(P_{it}) = \log(P_{i,t-1})$, which in turn implies that $p_{it} = p_{i,t-1}$. That is, rather than solving for equilibrium population levels, we can equivalently solve for the equilibrium log-population values. This means that equilibrium population values can be obtained by solving a set of simultaneous linear equations rather than a set of non-linear simultaneous equations based on equation (1.1) above.

In matrix form, equilibrium means that $P_{(t)} = P_{(t-1)} = P^e$, where P^e is the vector of logarithms of equilibrium population levels. Substituting P^e for $P_{(t)}$ and $P_{(t-1)}$ in equation (1.4) gives,

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[1] A negative value for diagonal elements of the community matrix may seem wrong at first blush, but can be more easily seen in the case of a single population. The classical definition of the community matrix is associated with the "change" in population over time which is often theoretically put in terms of continuous time dP/dt, but can also be stated in terms of discrete time intervals as [P(t) – P(t-1)]. Consider a linear model for a single population: $P_{(t)} = a + bP_{(t-1)}$. $0 < b < 1$. Subtracting $P_{(t-1)}$ from both sides gives $\Delta P = [P_{(t)} - P_{(t-1)}] = a + bP_{(t-1)} - P_{(t-1)} = a + (b - 1)P_{(t-1)}$. Note that the coefficient on own lagged P is (b - 1), which should be negative under stability.

$$P^e = A + BP^e \quad (1.5)$$

Solution of (5) is given by,

$$P^e = (I - B)^{-1}A \quad (1.6)$$

Where I is the identity matrix.

Signs of the off-diagonal elements of the community matrix (that is, signs of the pair b_{ij} and b_{ji}) indicate the qualitative nature of how organisms in the community interact. Community interaction nomenclature given by Logofet (1993) and by Williamson (1972) is shown in Table 1.1. In this case community interactions refer to the direction of net effects, evaluated at equilibrium population levels, and not to the biological, chemical, or physical mechanism for such interaction.

Table 1.1 Types of Species Interactions :The Community Matrix

Sign of b_{ij}	**Sign of b_{ji}**	**Type of Interaction**
+	+	Mutualism or symbiosis
+	-	Prey-Predator, or Resource-Consumer, or Host-Parasite
+	**0**	Commensalism
-	-	Competition
-	**0**	Amensalism
0	**0**	Neutralism

Population models for each taxon of nematodes were estimated using panel (pooled cross-sectional and time-series) data. In statistical jargon, the data set is described as a strong balanced panel. This means that all subplots appear in all periods. The panel random effects (RE) model was chosen because the different plots are considered a random sample from a population; whereas in a fixed effects (FE) model, specific plots would have been chosen to see their impact on the dependent variable. Another reason to choose RE is that it allows inclusion of time invariant variables (in our case, nitrogen) whereas in FE models these variables are absorbed by the intercept. Based on econometric theory, the commonly used estimation approach for spatial regressions, Maximum Likelihood (ML), was discarded as the error terms of the regressions were not normally distributed and no manipulation of the data corrected the issue. The Generalized Method of Moments (GMM) estimation was used instead of ML as it represents a more appropriate choice when there is uncertainty about the form of the error distribution.

All models tested negative for spatial correlation. Population models were initially estimated using the full set of candidate explanatory variables, including the lagged population level of all other taxa. As expected, some of the estimated coefficients were not statistically significant. The final models, which are given in Table 1.2, were estimated in a backwards step-wise procedure, eliminating the least significant variables first and making sure that the variables taken out do not affect the dependent variable. This is known as restricted model specification. The statistical analysis was done using STATA 11.0.

1.3 Results

1.3.1 Own and Inter-taxa Effects

A dynamic relationship between same group successive population estimates is highly expected. The sign of the coefficient on lagged population level is expected to be positive; and the higher the coefficient the stronger is the effect of carryover of previous population on current population. From Table 1.2a, it can be seen that the lagged own-population (previous period own population) effect had a value between zero and one, thus satisfying the partial theoretical requirement for stability, and was statistically significant for all taxa except the microbivorous nematodes. In the model with the full complement of variables (not shown), the lagged microbivorous nematode population was significant at a 82% confidence level (Table 1.2) and had the correct sign and magnitude. Thus, we can conclude that own previous population values are highly important for all eight taxa population dynamics.

Estimated coefficients shown in Table 1.2a for the effects of population levels of other taxa should be interpreted as partial equilibrium interactions. As can be seen, many of the lagged population coefficients of other taxa are statistically significant, which may suggest an interaction of the various nematode taxa. Although examination of the signs of the partial equilibrium effects in Table 1.2a is informative, taxa interactions in equilibrium is more informative and is the traditional way in which population ecologists examine species interactions. Thus, we now turn to equilibrium considerations.

Table 1.3 shows equilibrium population levels for March and for October. These months were selected because the first one represents a good approximation of what happens year round, meanwhile the latter represents a very high seasonality period, as will discussed later. The computed equilibrium population levels are each well within the range of observed variation for that taxa in the experimental data set.

Table 1.2a Parameter Estimates for the Nematode Dynamic Population Models in the Cullars Rotation (Alabama, USA) - Taxa [a]

Variable	**Dependent Variable** (**Nematode taxon**)							
	Mononchidae	**Dorylaimidae**	**Microbivorous**	**Lance**	**Spiral**	**Stubby Root**	**Lesion**	**Root-Knot**
Lagged Mononchidae	.1970*** (.0236)	.1088*** (.0247)		.0378** (.0188)	-.0765*** (.0272)		.0789** (.0313)	
Lagged Dorylaimidae		.0591** (.0264)			-.0730** (.0282)			
Lagged Microbivorous		.1227** (.0578)		.0780* (.0443)		.1610*** (.0536)		
Lagged Lance				.3692*** (.0234)			.0621* (.0363)	
Lagged Spiral	-.0628*** (.0205)	-.0408* (.0211)	-.0186* (.0101)		.3284*** (.0243)		.1210*** (.0269)	
Lagged Stubby Root					.0851*** (.0290)	.1266*** (.0250)		-.0810** (.0391)
Lagged Lesion	-.0332* (.0189)						.2669*** (.0250)	-.1501*** (.0293)
Lagged Root-Knot					.0688*** (.0177)	-.0267* (.0153)	-.0649*** (.0205)	.3091*** (.0241)

Table 1.2b Parameter Estimates for the Nematode Dynamic Population Models in the Cullars Rotation (Alabama, USA) - Seasonality [a, b]

Variable	Dependent Variable (Nematode taxon)							
	Mononchidae	**Dorylaimidae**	**Microbivorous**	**Lance**	**Spiral**	**Stubby Root**	**Lesion**	**Root-Knot**
January	1.8344*** (.1961)	.8151*** (.1960)	.1027 (.0962)	.1741 (.1417)	-.1659 (.2326)	-.2648 (.2085)	.9878*** (.2783)	.5249 (.3358)
February	1.2866*** (.2070)	-.0716 (.2073)	-.4945*** (.1017)	.1083 (.1504)	.1094 (.2456)	-.4680** (.2195)	-.1711 (.2930)	.4922 (.3531)
March	2.2120*** (.1929)	.9591*** (.1929)	-.0169 (.0947)	.2200 (.1396)	-.1615 (.2289)	-.1317 (.2050)	.1783 (.2736)	.5071 (.3300)
April	.6207*** (.1926)	.5204*** (.1929)	.0568 (.0945)	-.0033 (.1398)	-.3572 (.2284)	-.0252 (.2044)	-.8297*** (.2728)	-.7083** (.3288)
May	-.0192 (.2321)	.0167 (.2323)	-.0872 (.1139)	-.0856 (.1681)	-.7068*** (.2752)	-.7543*** (.2467)	-1.1601*** (.3291)	-1.2403*** (.3971)
June	-.5671*** (.1838)	.3462* (.1838)	-.0730 (.0902)	.2515* (.1330)	.1057 (.2179)	-.2848 (.1953)	-.9048*** (.2606)	.3830 (.3145)
July	-.2179 (.1918)	.2126 (.1920)	.0452 (.0942)	.1655 (.1392)	.2052 (.2275)	.2296 (.2035)	-.8539** (.2716)	.3355 (.3275)
August	.3317 (.2343)	.4127* (.2345)	-.0919 (.1150)	.3172* (.1699)	-.4785* (.2779)	-.1753 (.2487)	-.7655** (.3319)	-.0532 (.4001)
September	-.6045*** (.2013)	.7316*** (.2014)	.1007 (.0988)	.1703 (.1457)	.2319 (.2387)	-.0550 (.2140)	.0900 (.2855)	1.2844*** (.3445)
October	-.2401 (.2061)	1.1278*** (.2062)	.4349*** (.1012)	.3903*** (.1493)	.4331* (.2444)	.0514 (.2188)	1.0256*** (.2920)	1.0081*** (.3521)
November	.6680*** (.2306)	.8284*** (.2304)	.1816 (.1132)	.2529 (.1665)	-.5038* (.2736)	.3774 (.2455)	.6323* (.3275)	.4215 (.3955)
December(base)	.1364 (.3459)	1.3257*** (.3556)	5.2134*** (.1699)	-.4464* (.2690)	.5995 (.4087)	.2520 (.3478)	1.7227*** (.4677)	2.1340*** (.5496)
Days since previous sample		-.0105*** (.0016)	-.0017** (.0008)	.0024** (.0011)		-.0078*** (.0017)		

Table 1.2c Parameter Estimates for the Nematode Dynamic Population Models in the Cullars Rotation (Alabama, USA) – Nitrogen and Crop Rotation [a]

Variable	**Dependent Variable** (**Nematode taxon**)							
	Mononchidae	**Dorylaimidae**	**Microbivorous**	**Lance**	**Spiral**	**Stubby Root**	**Lesion**	**Root-Knot**
Nitrogen	.1866** (.0791)			-.3901*** (.0620)	.7537*** (.1257)			
Legume			.2543*** (.0278)	.2681*** (.0477)	.4077*** (.0743)		-.2098*** (.0645)	.1857*** (.0667)
Soybeans						-.2828*** (.0875)	.3681*** (.1210)	-1.1675*** (.1397)
Corn					-.5661*** (.1143)		.2259* (.1260)	-.5350*** (.1449)
Clover	.3419** (.1560)				-.5608*** (.1895)	-.3519** (.1618)	-.9310*** (.2192)	-.9030*** (.2593)
Rye	.4291*** (.1522)		.1514** (.0760)		-.7046*** (.1874)		.9651*** (.2136)	
# of Plots	60	60	60	60	60	60	60	60
# of Observations	1620	1620	1620	1620	1620	1620	1620	1620
Adjusted R-squared	.6598	.2830	.3184	.4145	.5419	.3924	.5036	.4764

[a] The standard errors are shown in parentheses below each parameter estimate. Significance probabilities: *** p<.01, ** p<.05, * p<.10
[b] Note that including a full complement of monthly dummy (binary) variables would result in a singular model. The model used December as a baseline (regular intercept) so the coefficients on all other months are expressed in comparison to it. The complete baseline is: December, no nitrogen, no legume, and cotton. The full complement of monthly dummy variables was included in the final model, even if coefficients on some monthly variables were insignificant. One reasons for this is convenience, and the other reason is that inspection of the pattern of monthly effects is informative even with the insignificant coefficients included. Deletion of insignificant monthly dummy variables would not appreciably affect the magnitude or significance of the lagged population variables in the model.

Table 1.3 Equilibrium Population Levels Calculated Using the Nematode Dynamic Population Models in the Cullars Rotation (Alabama, USA)[a]

Taxon	**Equilibrium Population in:**	
	March	October
Mononchidae	20.9	.98
Dorylaimidae	28.5	34.47
Microbivorous	206.4	330.3
Lance	.81	1.06
Spiral	5.42	13.07
Stubby Root	3.01	3.71
Lesion	19.29	60.95
Root-Knot	33.45	68.72

[a]Equilibrium values (nematode count per sample) are computed for the plots using nitrogen without a legume cover crop, and using the intercept for indicated month.

Table 1.4 shows the community matrix associated with estimated coefficients given in Table 1.2. As expected, the diagonal elements of the community matrix are all negative since they are computed as one minus the coefficient on lagged population. The lagged coefficient should be positive and less than one for stable populations, thus making the diagonal elements negative. (see footnote 1 for further explanation). With regard to off-diagonal elements, two patterns stand out. First, note the shaded block in the lower right-hand corner of the community matrix, Table 1.4. All of the off-diagonal elements in this block are negative, indicating that the cotton root-knot and lesion nematodes are competitive. This block could be expanded to include stubby root nematodes, however the relationship between these and lesion nematodes was significant only at 76% (stubby root – lesion) and 68% (lesion – stubby root) level of confidence. The negative relationship is not surprising, as each of these nematodes is known to feed on plant roots and to need roots to survive. The other interesting block of interactions is the block highlighted in the upper left-hand corner of community matrix, Table 1.4. In this shaded block, the off-diagonal elements of the community matrix are zero or positive, indicating that the Mononchidae, Dorylaimidae, microbivorous, and lance nematodes have either a symbiotic, commensalism, or neutralism relationship. In other words, the estimated population models shown in Table 1.2 show that these four groups of nematodes are not competitive. As the population models are expressed in a double-log specification (in a strict sense it is a log-linear because of the binary variables), the coefficients are expressing own and inter-taxa elasticities. For example, having the Mononchidae nematodes as the dependent variable it can be said that for a 1% increase in its own population in time t-1, the population in time t will increase by .20%; or that for a 1% increase in the population of Spiral nematodes in t-1, the population of Mononchidae nematodes will be reduced by .06%.

1.3.2 Seasonality, Nitrogen, and Crop Rotation Effects

The estimated monthly intercept variables shown in Table 1.2b illustrates strong seasonality patterns for each of the eight groups of nematodes. These seasonal patterns are plotted in figure 1.2 for all taxa of nematodes. It can be seen that the populations tend to decrease during January-February but they go up quickly again during February-March to then go down again during late winter (excepting microbivorous nematodes who remains steady after March). Then they remain steady until September-October when there is a significant increase in the population of cotton root-knot, Dorylaimidae, microbivorous, and lesion nematodes.

Table 1.4 Signed Community Matrix for the Nematode Dynamic Population Models in the Cullars Rotation (Alabama, USA)

Nematode Taxon	**Nematode Taxon Affected**							
	Mononchidae	**Dorylaimidae**	**Microbivorous**	**Lance**	**Spiral**	**Stubby Root**	**Lesion**	**Root-Knot**
Mononchidae	-	+	0	+	-	0	+	0
Dorylaimidae	0	-	0	0	-	0	0	0
Microbivorous	0	+	-	+	0	+	0	0
Lance	0	0	0	-	0	0	+	0
Spiral	-	-	-	0	-	0	+	0
Stubby Root	0	0	0	0	+	-	0	-
Lesion	-	0	0	0	0	0	-	-
Root-Knot	0	0	0	0	+	-	-	-

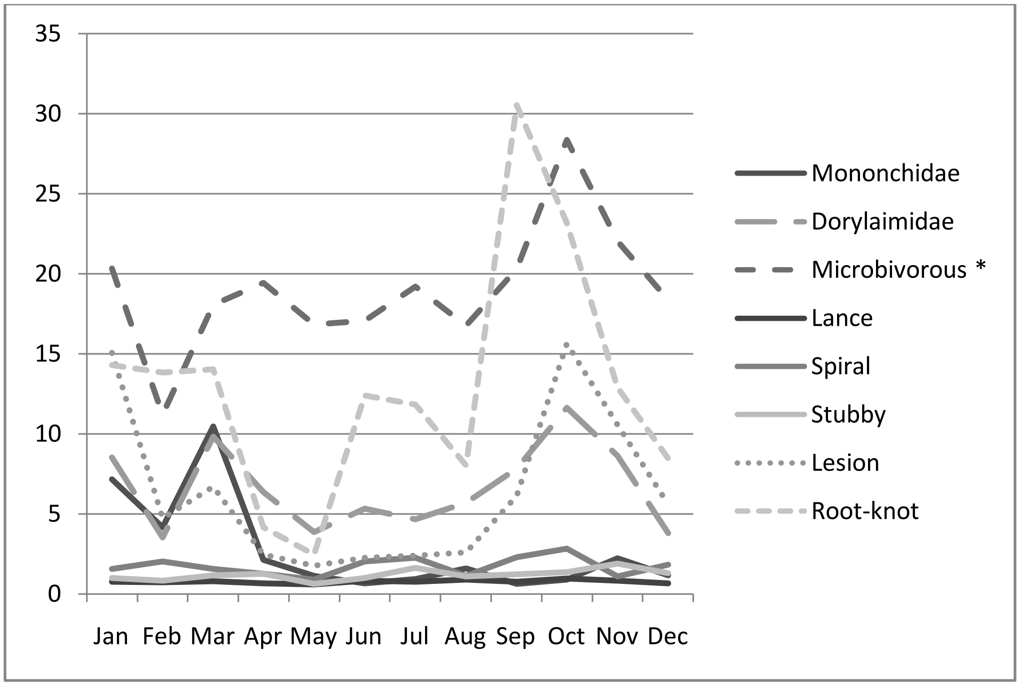

Figure 1.2 Monthly Intercepts in Simple Count for Each Nematode Taxon.
The monthly intercept values plotted here are in the multiplicative form given in equation (1), (Aim), and not in log form given in the estimated linearized model reported in Table 2. * denotes that the quantity was divided by 10 for formatting purposes.

Table 1.5 shows fertilization and crop rotation effects associated with the estimated coefficients given in Table 1.2. Nitrogen has a positive effect on Mononchidae and spiral nematode populations but affects negatively lance nematode population. Legume cultivation impacts positively the microbivorous, lance, spiral, and cotton root-knot nematodes but negatively lesion nematode population. Overall, a positive nitrogen effect can be attributed as the two negative coefficients may be related to the inherent nature of these two taxa being part of the plant feeder group (the increase on their competitors' populations can reflect a decrease in their populations). Regarding the interpretation of the nitrogen effect, it can be said that when nitrogen was applied to the plots, the population of Mononchidae nematodes increased, *ceteris paribus*, by 18.6%; and when legumes were present, the population of microbivorous nematodes increased, *ceteris paribus*, by 25.4%.

Table 1.5 Directional Effects of Fertilization and Crop Rotation on Current Nematode Population Levels in the Cullars Rotation (Alabama, USA)

Practice	**Population of:**							
	Mononchidae	**Dorylaimidae**	**Microbivorous**	**Lance**	**Spiral**	**Stubby Root**	**Lesion**	**Root-Knot**
Nitrogen	+	0	0	-	+	0	0	0
Legume	0	0	+	+	+	0	-	+
Soybeans **CSB**	0	0	0	0	0	-	+	-
Corn **CCN**	0	0	0	0	-	0	+	-
Clover **CCL**	+	0	0	0	-	-	-	-
Rye **CRY**	+	0	+	0	-	0	+	0

Table 1.5 also shows that, compared to the others, the best rotation for controlling plant feeder nematodes is clover following cotton. When clover is used in the rotation sequence, spiral, stubby root, lesion, and cotton root-knot nematode populations are reduced by 56.1%, 35.2%, 93.1%, and 90.3% respectively. The other rotation effects are as follows: Soybean is good for reducing stubby root and cotton root-knot nematodes, corn is good for reducing spiral and cotton root-knot nematodes, and rye following corn is good for spiral nematode population reduction. However, all of these three rotations increase lesion nematode populations.

1.4 Discussion

Previous own populations having an effect on current nematode populations is not surprising. Reports on plant-parasitic nematode populations such as *Meloigogyne* spp. (McSorley and Gallaher, 1993; Taylor and Rodríguez-Kábana, 1999) *Paratrichodorus minor* (McSorley and Gallaher, 1993), and *Pratylenchus* spp. (Bell and Watson, 2001), as well as non-parasitic nematode populations such as the Rhabditidae (McGraw and Koppenhöfer, 2009) and microbivorous (Taylor and Rodríguez-Kábana, 1999) agree with what is found here. Regarding the inter-taxa effects, this study found that for all taxa regression models at least one of the other taxa's coefficient was significant (which may imply causation). This has also been seen in other studies (Bell and Watson, 2001; Taylor and Rodríguez-Kábana, 1999) and makes sense as it is unlikely that forms of live with similar characteristics would have no influence on each other.

Also seen in the results presented here, plant-parasitic populations such as root-knot and stubby root nematodes were lower after the incorporation of soybeans in the rotation scheme. On cotton, plant parasitic nematodes populations, excepting lesion nematodes, were higher compared to soybeans, corn, clover and rye. The effects of crop rotation found in this study tend to agree with what was previously found in the literature. Rodríguez-Kábana (1982) and Rodríguez-Kábana and Truelove (1982) showed that enzymatic activities, such as xylanase and catalase, were higher after soybeans and lower after cotton. The authors established that higher enzymatic activity is negatively correlated with plant parasitic nematodes and positively correlated with the quantity of microorganisms living in the soil. This study also found that soybean, corn, and clover reduced the root-knot nematode population meanwhile rye did not have a significant effect on it. Also agreeing with our findings, Chen and Tsay (2006) showed that the incorporation of corn in the rotation in a strawberry field suppressed root-knot nematode populations and resulted in increased yields. It has also been discussed that the addition of velvet bean (*Mucuna deeringiana*), Bahia grass (*Paspalum notatum*) and partridge pea (*Chamaecrista* [*Cassia*] *fsciculata*) rotations helped reducing root-knot nematode populations in peanut fields (Rodríguez-Kábana et.al, 1992; Rodríguez-Kábana et.al, 1994; Rodríguez-Kábana et.al, 1995). All of these findings converge in that crop rotation sequences must be incorporated in the farmer's cultural practices in order to properly manage root-knot nematodes.

We have found that the majority of the plant parasitic nematode populations were higher between September and October. Several authors have agreed that root-knot nematode populations (*Meloidogyne spp.)* are very seasonal being higher around September and lower during late winter (McSorley and Dickson, 1990; Rodríguez-Kábana et.al, 1986; Rodríguez-Kábana and Robertson, 1987) thus agreeing with our study where a decreasing trend can be seen from January to May and the highest population count was during September to October. This also agrees with what Rodríguez-Kábana and Collins (1979) found on spiral nematode populations as they were low between December and June but high in October as summer crops developed. Our study also agrees with stubby root nematodes being low in number and having no visible seasonal peaks under cotton (baseline in this study).

The same authors (Rodríguez-Kábana and Collins, 1979) also stated that the number of spiral nematodes was higher when receiving complete fertilization and that the number of stubby root nematodes remained almost invariant regardless of fertilization regime, thus agreeing with what was found in this study as nitrogen had a positive effect on spiral but not significant effect on stubby root nematodes. Then, Rodríguez-Kábana (1982) and Rodríguez-Kábana and Truelove (1982) showed that legume cultivation also increases enzymatic activities thus reducing the amount of parasitic nematodes. This was not the case in this study as only lesion nematode populations were reduced. Legume cultivation did increase the population of microbivorous nematodes implying an increase in the quantity of microorganism in the soil as these nematodes are predominantly microorganism-feeders. Rodríguez-Kábana and Truelove (1982) stated that the addition of supplementary mineral nitrogen to a legume cultivation regime can reduce catalase activity. This study did not included an interaction between legume and chemical nitrogen application so the mixed results of an increase in spiral nematode populations and a reduction in lance nematode populations due to nitrogen application cannot be properly explained.

1.5 Conclusion

This article statistically estimated the effect of the several factors affecting the population dynamics of Mononchidae, Dorylaimidae, microbivorous, lance, spiral, stubby root, lesion, and cotton root-knot nematodes using non-linear regression models. A spatial component was included as populations in one plot were proved to be associated with the value of their neighbors. The results suggested that previous own population values are very important for all eight taxa and that all of taxa have an interaction effect with at least one other taxon. Lesion and cotton root-knot nematodes were found to be competitive; meanwhile Mononchidae, Dorylaimidae, microbivorous and lance nematodes were found to be non-competitive. Nitrogen (including legumes) had a positive effect on Mononchidae, microbivorous, spiral, and cotton root-knot nematode populations. The use of clover after cotton in the rotation crop program proved to be better in reducing plant parasitic nematodes compared to other treatments.

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Chapter 2:

Are Peruvian Non-traditional Agricultural Exports Price Discriminating?

Abstract

Over the past two decades, Peru has become the largest fresh asparagus exporter in the world and a major exporter of processed artichoke and paprika markets. Dominance in world trade of these three products may have given Peru market power but the exertion of it and towards whom has not been studied yet. This study uses Pricing-To-Market (PTM) models to test for market power as manifested by price discrimination in the Peruvian export market for these three goods. In the PTM model, market power is revealed by the adjustments that export prices make through a country destination effect and an exchange rate effect. Using panel data, 03/2005 to 08/2012, for the top importing countries, the results strongly suggest that Peruvian exporters are engaging in price discriminating behavior. Country markups were common and Peruvian exporters were stabilizing English pound prices in asparagus and Euro prices in paprika. Lastly, even though it was found that fresh asparagus exporters were amplifying the effects of the exchange rate after the preferential free trade agreement between U.S. and Peru was established, the assumption of constant marginal cost in this case may be too restrictive to definitively conclude that an incomplete exchange rate pass-through is happening.

2.1 Introduction

Trade theory suggests market concentration in international markets leads to imperfectly competitive prices (Goldberg and Knetter, 1997; Pall et.al, 2013). As trade usually happens between countries with different currencies, a realignment of importers' and exporters' currencies must occur. In this context, movements in exchange rates can have a big influence on an exporter's pricing decision. Exporters will try to maximize profits in local currency taking into account that the importing demand depends on the price in local currency of that specific importing country. This leads to any depreciation or appreciation of a currency being a crucial incentive for an exporter to price discriminate charging different prices to different countries.

The Pricing-To-Market (PTM) concept was developed by Krugman (1987) in the late 1980s. The PTM model recognizes that an exporter can adjust country markups to account for changes in their exchange rates, resulting in an incomplete pass-through. PTM occurs when export prices are maintained or increased as the foreign currency appreciates (Pick and Park, 1991). Under this premise, Knetter (1989) developed an empirical model that due to its simplicity and empirical applicability with publicly available data has become used worldwide. The model, however, relies heavily on the critical assumption of constant Marginal Cost (MC) to develop a testable hypothesis.

Peru has gone through a big economic growth period over the last two decades becoming one of the fastest growing economies in Latin America. An agricultural exporting "boom" has been happening over the last few years. Traditional agricultural exports have gradually lost importance and now have been replaced by non-traditional goods. Peruvian non-traditional agricultural export markets might not have a competitive structure as most of the goods are specialized and require specific comparative advantages. This study becomes important as it is the first to address the pricing behavior of Peruvian agricultural exports[2]. Now that Peru has become the largest exporter of fresh asparagus (USDA, 2007, USDA, 2005), the third in paprika (Lancaster, 2009, USAID, 2010) and processed artichoke (Boriss and Huntrods, 2013), the question about market power and pass-through effect gains importance. This study adds to the literature as: (1) it uses a PTM model to develop a testable hypothesis about the existence of market power and assess any differences that may have arisen due to the signing of the improved Free Trade Agreement (FTA) between Peru-U.S. in February 2009. (2) It discusses on the implications of the PTM model if the underlying assumption of MC is inappropriate.

The remainder of this article is structured as follows. The next section covers the literature review. The international fresh asparagus, paprika, and processed artichoke markets are then described. The following section describes the models, data used in the empirical estimations, and models insights. Results are discussed after that. The last section concludes.

2.2 Literature Review

Studies addressing price discrimination in an agricultural international trade context using PTM models have been widely used. One of the first studies of this type was the one made by Pick and Park(1991) where they examined the competitive structure of major U.S. agricultural exports. The authors concluded that cotton, corn, and soybeans exports did not show a non-competitive behavior. However, wheat exports showed a different result suggesting an incomplete pass-through to importers.

[2] There are no price discrimination studies by Peruvian agricultural exports yet according to searches in Econlit Academic Search Premier (including Spanish language publications). The reason is that the Peruvian government started publishing data on agricultural commodities after concerns about price manipulation a couple of years ago. Before this, data were scarce.

With this result of imperfect competition in the wheat market, Patterson and Abbott (1994) analyzed the export market structure and the pricing behavior of U.S. wheat exporting firms. They concluded that a positive markup is related to U.S. seller concentration which was measured by the Herfindahl-Hirschman index (HHI)[3]. Pick and Carter(1994) also estimated a PTM model for the same commodity using two exporters, U.S. and Canada. At the time, these two countries accounted for more than 50% of wheat exports. In addition to the regular PTM model, the authors tested for an effect of the interaction between these two exporter's currencies. It was found that price discrimination exists for both American and Canadian wheat exports. They also showed that the competitors' exchange rates influence exporter pricing decision.

Following this premise, two similar papers studying U.S. and Canadian agricultural exports were made by Carew (2000) and Carew and Florkowski (2003). The data were first analyzed for unit root (non-stationary) and then the PTM model was used. In several destination markets, Carew (2000) found evidence of pronounced price discrimination in only U.S. wheat exports and moderate price discrimination in Canadian wheat and pulse exports (the other commodity analyzed was tobacco). Carew and Florkowski (2003) found that U.S. exporters were more sensitive to exchange rates changes than Canadian exporters. This finding suggests that U.S. exporters tend to maintain stable prices in destination markets as a result of an appreciated U.S. currency. They also found that Canadian/U.S. exchange rate is not a significant variable influencing the pricing decisions of the exporters.

Miljkovic et.al (2003) explored the influence of exchange rates on U.S. meat export prices using a PTM model that included dummies for international trade agreements. In the case of beef, imperfect exchange rate pass-through occurred for exports to Japan, Canada, and

[3] The HHI index is a measure of the degree of competition in an industry that uses the market share of the firms and the number of firms present in that industry. It is defined as $H=\sum_{i=1}^{n} s_i^2$.

Mexico. GATT had positively influenced US beef export prices, while NAFTA demonstrated little impact. In the case of pork, incomplete pass-through occurred for Japan and GATT and NAFTA appear to have little influence. US poultry export prices declined in response to devaluations of Mexican and Hong Kong currencies. However, GATT appears to have positively influenced poultry export prices, while NAFTA appears to have had the opposite effect.

More recently, PTM was used in a study of Russian wheat exports because this country has achieved a relatively strong market position (11.2%) in the international wheat market during the last decade (Pall, et al., 2013).Pall et al (2013) suggested that Russia exercises PTM in some wheat-importing countries. However, the findings suggest that the structure of the Russian wheat exports could be more competitive than U.S. and Canadian wheat exports as they possess more of the wheat exports market share.

Most of these studies have used goods that are industrialized, so the requirement of constant MC in the market is plausible. However, some of these goods may not have a strictly elastic supply curve, meat for example, thus commenting on the implications of this assumption not holding would be valuable.

2.3 Peruvian Non-traditional Agricultural Exports

Peru is developing a strong worldwide position in producing and exporting certain fruits and vegetables. The factors that have contributed to this development are basically four: (1) a friendly investor environment and policy framework, (2) free trade agreements, (3) relatively cheap labor, and (4) geographic related, or comparative, advantages (USDA, 2010). Even though agricultural exports only account on average for 7.2 percent of the Peruvian merchandise trade (FAO, 2009), most of them are well positioned in the international trade market. Currently, non-traditional goods account almost for 90% of the Peruvian agricultural exports making them the foundation of the agricultural exports[4](Peru, 2013). The biggest markets for these goods are basically the United States and the European Union. This study focuses on three non-traditional goods that are produced year-round in Peru: asparagus, artichoke, and paprika.

Asparagus(*Asparagus officinalis*) is the leading, by value, non-traditional agricultural export in Peru (table 2.1).Asparagus is a high-value, labor-intensive perennial crop that can be eaten raw or used to prepare more complex dishes (Boriss and Brunke, 2006). Historically, the top producer of fresh asparagus has been China followed by Peru, United States, and Mexico. According to the Food and Agriculture Organization, these countries produced 587,500, 186,000, 102,780 and 67,247 tons respectively in the year 2004 (USDA, 2005). However, the ranking changes if exporting quantities and not total production are compared. In this case Peru becomes the largest exporting country with 73,038 tons accounting for almost 50% of total world exports (USDA, 2005). Also, according to the Global Trade Atlas and U.S. Census Bureau statistics, only Peru's export market share has grown each year over the last several years. On the other side, the top fresh asparagus importers in 2007 by quantity were in first place, by far, the United States, followed by the European Union (EU), and Japan (USDA, 2007). In 2005, the U.S. was the destination of 67% of total Peruvian fresh asparagus exports, which represented 54% of the total Unites States fresh asparagus supply (Díaz Rios, 2007). For the EU, the only countries that produce asparagus in significant quantities are Germany and Spain; however, this production is not even enough to satisfy their own demand (Díaz Rios, 2007, Garde Adrian, 2010). So it is safe to state that Peruvian imports are a significant part of EU fresh asparagus supply.

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[4] In general, traditional goods are those who have a long history of being exported by a specific country. In Peru, this category includes minerals, gas, fish meal, etc. However, according to the Peruvian department of agriculture, there are only three traditional agricultural (crops) exports: coffee, sugar cane, and cotton.

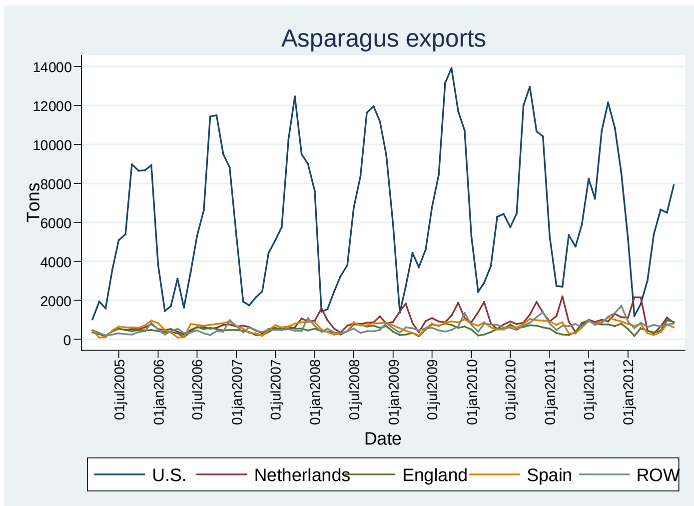
Table 2.1 Selected Peruvian fruit and vegetable exports in 2009

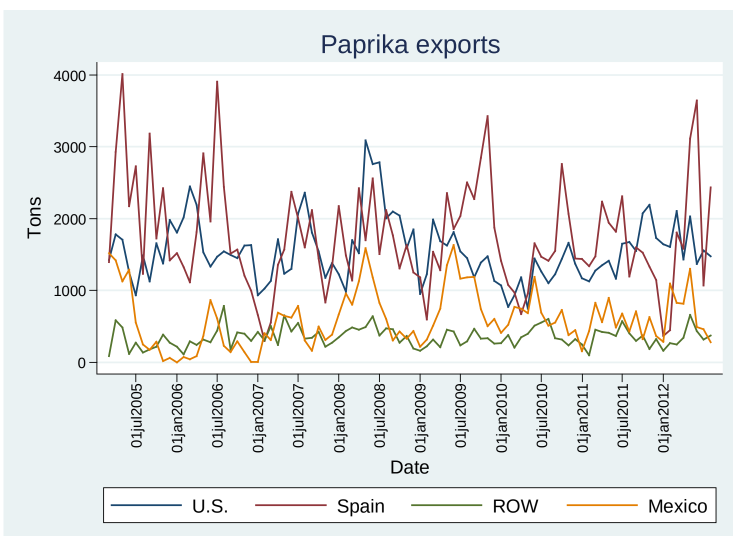
Product	Value *(U.S. $ million)*	Share shipped to selected destinations: United States	EU	Northeast Asia	Rest of world
		Percent			
Asparagus	373.9	50.2	44.1	1.3	4.4
Grapes	126.6	27.8	31.6	18.9	21.7
Paprika	101.3	37.8	39.1	0.1	23.0
Mango	81.2	29.9	62.1	1.4	6.6
Artichoke	72.3	59.7	36.8	0.0	3.5
Avocado	64.1	0.2	95.2	0.0	4.6
Bananas	51.6	24.3	63.4	11.7	0.5
Citrus	43.4	26.7	52.5	0.0	20.7
Onion	28.1	77.4	4.0	0.0	18.5

Source: USDA, 2010

Paprika is a powder that is made from grinding *Capsicum annuum* peppers. This product, which is used as a spice or as a food colorant, occupies the third place in importance in the Peruvian non-traditional agricultural exports basket (table 2.1). Peru has emerged as an important supplier for North America and Europe over the last years. For example, revenues for paprika from the U.S. increased from US$5.9 million in the year 2000 to US$95.3 million in the year 2005, representing an annual growth rate of about 85% (Lancaster, 2009). In 2005, Peru was the largest supplier of paprika to the EU, shipping 26,167 tons valued at $36 million (USAID, 2010). Peruvian exports represent a very important share in total U.S. paprika supply as domestic production of Paprika in the U.S. is very small (Kebede, 1990).

Artichoke (*Cynaracardunculus var. scolymus*) is a perennial plant that has an edible portion. It occupies the fifth place in the value of Peruvian non-traditional agricultural exports as in 2009 (table 2.1). As in 2011, Peru occupies the fourth place in artichoke production after Italy, Egypt, and Spain; However, Peru occupies the second place in processed artichoke production only beaten by Spain (USDA, 2010). Peruvian processed artichokes are mostly exported to the U.S. and the European Union; representing 60%, 21%, and 8% the U.S., Spanish, and French market (USDA, 2010). The Peruvian exports to the different destinations are shown in figure 2.1. Peruvian processed artichokes imports are an important part of these countries' total consumption, especially in the U.S where there is no artichoke-processing industry. Peruvian imports accounted for 45% of the U.S. processed artichoke consumption in 2009 (USDA, 2010). Spain and France are big artichoke producers but they are also big importers because of their big consumption and strict seasonality (USDA, 2010).





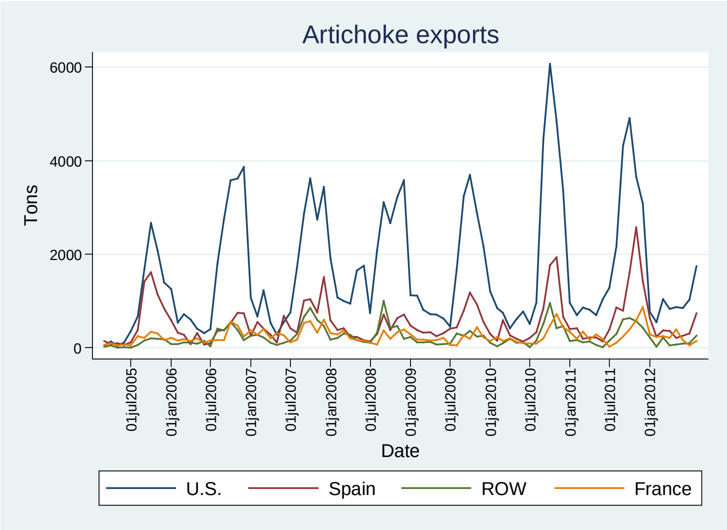


Figure2.1 Peruvian non-traditional agricultural exports by country. March 2005 – August 2012.

From an international trade perspective, the dominant countries here are Peru as a seller and the United States as a buyer. Peru accounted for 58% of total U.S. asparagus imports while Mexico accounted only for 38% in the year 2006 (USDA, 2007). Peru is also the largest supplier of paprika to the US accounting for 59% of the U.S. market share in 2009 (USAID, 2010). Lastly, The United States imported processed artichokes valued at $153.1 million in 2012 with Peru the leading supplier, followed by Spain (which was the leading supplier until 2007)(Boriss and Huntrods, 2013).

The big amount of agricultural exports from Peru to the United States is a result of the 1991 Andean Trade Preference Act (ATPA) which aimed to reduce Peruvian coca production as part of U.S. anti-drug efforts. Due to ATPA, agricultural exports are allowed to enter the United States duty-free (Boriss and Brunke, 2006, USDA, 2010). The ATPA was schedule to expire in 2006 and was annually renewed until a better agreement is signed. In February 2009, the U.S-Peru trade promotion agreement (PTPA) entered into force. The PTPA is a FTA that grants preferential access to the U.S. permanently(USDA, 2010). At the end of 2012, Peru signed a

FTA with the European Union (European-Comission, 2012). However, this time frame does not overlap with the available data and cannot be empirically analyzed at this time.

2.4 Model and Data Description

2.4.1 Model

As it was shown on Pick and Park (1991) model, the basic PTM model applied to agricultural commodities is as follows. Consider an agent who exports to N different foreign destinations with individual import demand in each destination, $i = 1, \dots, N$. Then the demand for the good is:

$$q_i = f_i\left(e_i P_i\right) \qquad (2.1)$$

where q_i, is quantity demanded by destination market i, P_i is theexport price in terms of the exporter's currency, and e_i is the exchange rate that converts the exporter's currency into the importers' currency in market i. Specifically, $\mathrm{e}_i = \frac{\mathrm{FCU}_i}{\mathrm{DCU}}$ where FCU_i is the foreign currency unit of the ith importer, and DCU is the domestic currency unit of the exporter. Thus, for example, if the United States sells soybeans to Japan, e = FCU/DCU = YEN/US$ and an increase in e implies domestic or U.S. currency strengthening. Letting $\widetilde{\mathrm{P}}_i = \mathrm{P}_i \cdot \mathrm{e}_i$ be the export price in the buyer's currency (exclusive of transportation and tariff wedges), an increase in e_i will cause $\widetilde{\mathrm{P}}_i$ to increase, i.e., domestic currency strengthening raises the cost of exports to foreign buyers when the product is priced in their local currency. A profit function is obtained having the cost structure of the exporter to be a function of the quantity exported[5]. The first-order conditions for profit maximization indicate that the firm will allocate output levels across destination markets to

[5] $\pi = \sum P_i\, q_i - C(\sum q_i)$

equate marginal revenue in each market with the common marginal cost (MC)[6]. Thus, export prices charged to each destination market are comprised of the product of the common marginal cost and a destination-specific markup denoted by:

$$P_i = MC_i \left(\frac{\eta_i}{\eta_i - 1}\right) \tag{2.2}$$

where η_i(>0) is the absolute value of price elasticity of demand faced by the exporter in the destination market *i*. Equation (2.2) shows that the export price in home currency is influenced by the perceived elasticity of demand in the different destinations. In the case where the market structure is perfectly competitive, the demand elasticities are infinite and independent of destination. In that case the maximizing agent would equate MC to world price. The marginal cost to market *i* is represented by equation (2.3).

$$MC_i = MC(q_i) \tag{2.3}$$

Following Knetter's (1989) implementation of the PTM model, the consistency of non-traditional goods was tested for price discrimination by the means of a panel fixed effects model, specified as:

$$\log P_{it} = \theta_t + \gamma_i + \beta_i log\ e_{it} + \mu_{it} \tag{2.4}$$

Where θ_t is the time intercept, γ_i is the country effect (fixed effect), and μ_{it} is the error term. The time intercept provides a measurement of the marginal cost as it measures the common price in each time period. The country effect will measure the markups to the different destinations including geography, trade policy, tastes and institutional features that vary across countries but are constant over time. Under perfect competition, export prices will be the same for all destinations thus $\gamma = 0$ as there is no country effect and $\beta = 0$ as changes in the bilateral

[6] In order to have a testable hypothesis, the assumption of MC being exogenous, constant, and equivalent in all destination markets is required.

exchange rates do not affect export prices. As with most agricultural goods, it is expected that lagged prices have a statistical relationship with current prices. This is consistent with market price dynamics, thus, a first lagged dependent variable will be added to the right hand side of equation (2.4).

Following Carew (2000), there are two market power possibilities by which exchange rates and country effects may influence exporters' price markups. The first possibility is to have price discrimination across the different destination markets. Under price discrimination with constant elasticity of demand in each foreign market, the price charged in each destination is a fixed markup over marginal cost. So, price changes in each destination market will be affected by the time effect and the country effect. Though exchange rates will not affect the exporters' optimal markup ($\beta = 0$), country effects that measure markups are likely to vary across destination markets ($\gamma \neq 0$). However, a significant country effect does not strictly imply imperfect competition as constant quality differences (if any) may be reflected in this parameter (Knetter, 1989). The second possibility is to have imperfect competition and varying elasticity of demand. In this case, the MC is not well measured by the time effect therefore the optimal markup will be influenced by exchange rates and prices will be different in each country ending up with $\beta \neq 0$ and $\gamma \neq 0$. These two possibilities were well explained by Knetter (1989) "At a given price in the exporter's currency, a depreciation of the importer's currency raises the local currency price paid by the importer. If demand has constant elasticity with respect to price, the optimal markup charged by the exporter will not change ($\beta = 0$) as exchange rate changes increase the price paid by the importer. If, however, demand elasticities change with changes in the local currency price, then export prices will depend on exchange rates". Table 2.2 illustrates the possible outcomes under the PTM model.

Table 2.2 Relationship between the estimated parameters and market scenarios

γ	β	Market scenarios
Not significant	Not significant	Perfect competition, imperfect competition with common markup
Significant	Not significant	Constant elasticity of demand which leads to constant markup, which can differ across countries.
Not significant ⁄ significant	Significant	Varying elasticity of demand which leads to varying markup, which can differ across countries.
	Positive	Amplification of exchange-rate effects
	Negative	Local currency price stability

Source: Pall, et al., 2013

2.4.2 Data

Data come from two sources. The export nominal price charged (in Peruvian currency or "Nuevos soles") and the quantity exported (in tons.) to several import markets for each of the three goods comes from the Peruvian Department of Agriculture (MINAG is the Spanish acronym). Monthly data from March 2005 to August 2012 (90 periods) is available on the top importers. For fresh asparagus the top importers are: the U.S., Holland, Spain, and England; for paprika the importers are: U.S., Mexico, and Spain; and for artichoke the countries are: U.S., France, and Spain. The price used in this study was Free on Board (FOB) as most transactions on these goods are under contract. Country specific nominal and real exchange rates (nominal deflated by the CPI) were collected from the U.S. Economic Research Service (ERS) website. Export nominal prices and real exchange rates were used for the empirical analysis. This study first estimates the PTM for the whole series and then it divides the data in two sections: before the PTPA (46 months) and after the PTPA (44 months).

2.4.3 Model Insights

Despite PTM still being considered as the "workhorse" of modeling exchange rate pass-through effects, this model has potential weaknesses besides the constant MC assumption. Some of these potential weaknesses are: (1) that its foundation lays on firm level economics and using aggregated level data might not be appropriate (Atkeson and Burstein, 2008), (2) that it is not useful when product differentiation is present (Lavoie and Liu, 2007, Mallick and Marques, 2012), and (3) that the results are not reliable if the series are not first tested/corrected for stationarity (Abbott, et al., 1993, Larue, 2004, Miljkovic, et al., 2003). The first point can be addressed by showing that within the Peruvian non-traditional export industry there is market concentration. Using data from the Peruvian Integrated System of Foreign Trade Information (Sistema Integrado de Información de Comercio Exterior (SIICEX)), Table 2.3 shows the market share of the firms involved in exporting these three goods as of 2012.

It can be seen that the paprika and artichoke exporting industries have a relatively high HHI indicating moderate to high concentration; meanwhile the HHI in the asparagus sector shows a fairly un-concentrated industry. Second, these products can be considered as essentially homogenous goods. In the case of artichoke and paprika, goods are processed to then become industrialized /standardized. In the case of fresh asparagus, producer associations (which are very common in Peru especially for exporting produce) propose a standard quality to be considered part of the group. So the idea of findings indicating product differentiation can be assumed away, thus facilitating the economic analysis proposed here. Third, the exchange rate data was examined for time series properties (stationarity) using the Augmented Dickey-Fuller (ADF) test for unit root. For this task, the proper number of lags was selected using the higher value

Table2.3 Market share and HHI indexes as of 2012

Asparagus	
Firm	participation
COMPLEJO AGROINDUSTRIAL BETA S.A.	18%
CAMPOSOL S.A.	6%
DANPER TRUJILLO S.A.C.	6%
SOCIEDAD AGRICOLA DROKASA S.A.	5%
SANTA SOFIA DEL SUR S.A.C.	5%
AGRICOLA LA VENTA S.A.	4%
GLOBAL FRESH S.A.C.	4%
AGRO PARACAS S.A.	4%
PEAK QUALITY DEL PERU S.A.	4%
Other firms (91)*	44%
HHI	**0.053**
Paprika	
Firm	participation
CORPORACION MISKI S.A.	43%
ECO - ACUICOLA SOCIEDAD ANONIMA	16%
EXPORTADORA NORPAL S.A.C	11%
AGROINVERSIONES MISTUL SAC	9%
CORPORACION CAPAS S.A.C.	6%
CORPORACION PERUNOR S.A.C.	6%
CONSORCIO LA CHACRA S.A.C.	5%
EXPORTADORA AJM SOCIEDAD ANONIMA	4%
HHI	**0.242**
Artichoke	
Firm	participation
SOCIEDAD AGRICOLA VIRU S.A.	35%
DANPER TRUJILLO S.A.C.	24%
DANPER AREQUIPA S.A.C.	10%
ALSUR PERU S.A.C.	8%
CAMPOSOL S.A.	8%
OPEN WORLD EXPORT SAC	4%
AGROINDUSTRIAS DEL MANTARO S.A.C.	4%
CYNARA PERU S.A.C.	3%
AGROINDUSTRIAS AIB S.A	2%
Other firms (13)*	4%
HHI	**0.2282**

Source: CIISEX, 2013

*In order to calculate the HHI it was assumed that the other firms had the same market share.

between the Schwarz's Bayesian information criterion (SBIC), the Akaike's information criterion (AIC), and the Hannan and Quinn information criterion (HQIC). The null hypothesis of nonstationarity was rejected only for the U.S. and the Mexican exchange rates. For the others, as non-stationary processes were found to be present, first differences were used instead (Carew, 2000, Falk and Falk, 2000). The statistical package used was STATA 11.0

The assumption of constant MC is highly plausible in this study as these three goods are high value crops. High value crops go on the best land and, until the quality of land becomes limiting, MC should be approximately constant. On the other side, expansion of major crops occurs on marginal land with lower yields, higher costs, etc.

It is importance to point out what are the implications of assuming that marginal cost is constant. An Equilibrium Displacement Model (EDM) for a net exporter will aid in explaining this[7]. Here it is shown that the model previously established (equation (2.4)) is valid only if the marginal cost curves are horizontal, i.e., only if $\frac{\partial MC_i}{\partial Q_i} = 0$, which implies export supply to market *i* is perfectly elastic. If this assumption does not hold, the partial derivative $\frac{\partial P_i}{\partial e_i}$ is always negative regardless of market structure.

2.4.3.1 Comparative Statics

The relationship between P_i and e_i in imperfectly competitive equilibrium can be determined by solving the structural model (equations (2.1) – (2.3)) for the reduced-form elasticity $\frac{P_i^*}{e_i^*}$, where the asterisk (*) indicates proportionate change (X* = dX/X). The first step is to express equations (2.1) – (2.3) in proportionate change form as follows:

$$Q_i^* = -\eta_i(P_i^* + e_i^*) \qquad (2.5)$$

$$P_i^* = MC_i^* - \mu_i\eta_i^* \qquad (2.6)$$

[7] Special thanks are given to Dr. Henry W. Kinnucan, professor in agricultural economics at Auburn University, for providing helpful comments and suggestions on the developing of the EDM.

$$MC_i^* = \frac{1}{\varepsilon_i} Q_i^* \qquad (2.7)$$

where $\mu_i = \frac{P_i - MC_i}{MC_i} \geq 0$ is the monopoly markup over marginal cost, and $\varepsilon_i = \left(\frac{\partial Q_i}{\partial MC_i}\frac{MC_i}{Q_i}\right) > 0$ is the export supply elasticity with respect to marginal cost.

Equations (2.5) – (2.7) contain three endogenous variables (Q_i^*, P_i^*, and MC_i^*) and two exogenous variables (e_i^* andη_i^*). Although the demand elasticity is apt to change with the price level, which implies η_i^* is endogenous, it is sufficient for the purposes at hand to treat η_i^* as exogenous.

Solving equations (2.5) – (2.7) simultaneously for P_i^* in terms of the exogenous variables yields:

$$P_i^* = -\left(\frac{\eta_i}{\varepsilon_i + \eta_i}\right) e_i^* - \left(\frac{\mu_i \varepsilon_i}{\varepsilon_i + \eta_i}\right) \eta_i^*. \qquad (2.8)$$

An isolated increase in exchange rate reduces export price, as does an isolated increase in the demand elasticity *provided the exporter exercises market power*, i.e., provided μ_i is strictly greater than zero. Recalling that $\tilde{P}_i = P_i \cdot e_i$ is the export price in the *buyer's* currency, the reduced-form elasticities for export price and exchange rate pass-through are:

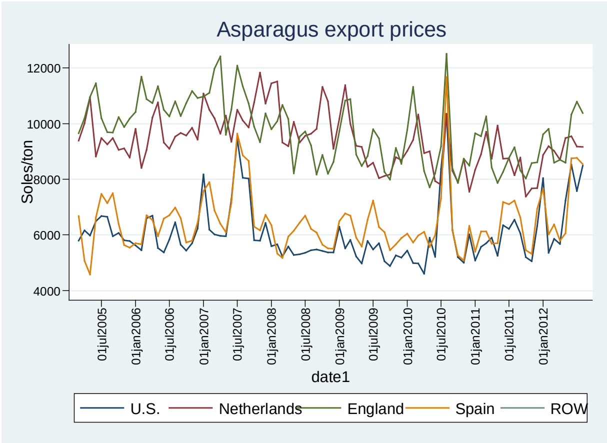
$$\frac{P_i^*}{e_i^*} = \frac{-\eta_i}{\varepsilon_i + \eta_i} \leq 0 \qquad (2.9)$$

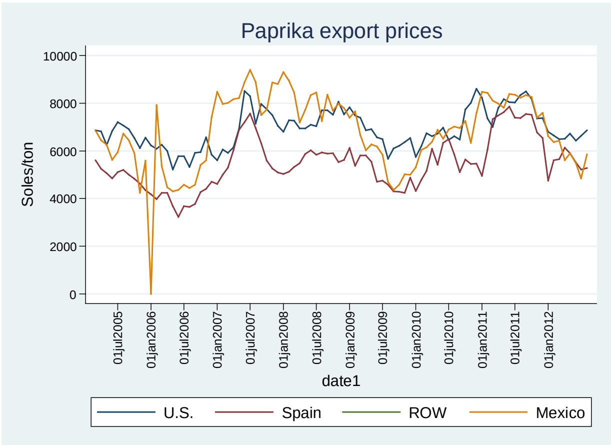
$$\frac{\tilde{P}_i^*}{e_i^*} = \frac{\varepsilon_i}{\varepsilon_i + \eta_i} \geq 0. \qquad (2.10)$$

Clearly, the only instance in which exchange rate pass-though is complete $\left(\frac{\tilde{P}_i^*}{e_i^*} = 1\right)$ is when export supply is perfectly elastic ($\varepsilon_i = \infty$), in which case $\frac{P_i^*}{e_i^*} = 0$. Thus, a finding from equation (2.4) that $\beta_i = \frac{P_i^*}{e_i^*} < 0$ could mean PTM, or it could mean simply that the export supply curve is upward sloping.

2.5 Results and Discussion

Figure 2.2 shows the differences in export prices per ton (in Peruvian soles) by country and by date. Inspection of the data suggests presence of price discrimination at least in asparagus as there are definitively different prices for two groups: a lower price for U.S. and Spain and a higher price for Holland and England. Also, the U.S. having the lowest average price in asparagus and artichoke was not unexpected as this country imports a very big amount (some buyer power may being exerted) and has a FTA with Peru.





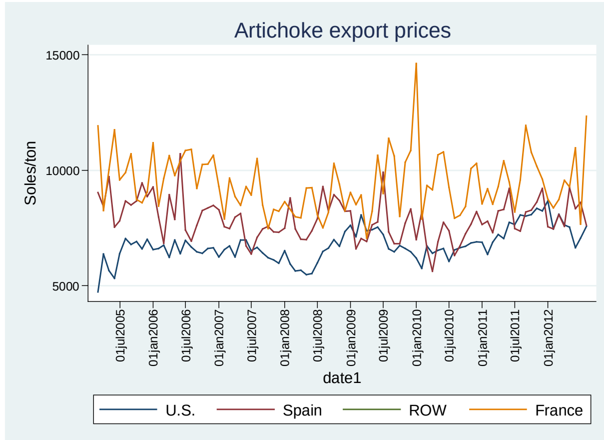


Figure 2.2 Peruvian non-traditional agricultural export prices by country. March 2005 – August 2012

Using the Durbin's alternative test, it was shown that there was no serial autocorrelation of the errors. Regarding the lagged dependent variable, its coefficient was positive and very significant. The PTM results for the whole series are shown in table 2.4. Through model diagnostics, it was shown that 85%, 61%, and 69% of the variance in the regression is due to differences across panels for asparagus, paprika and artichoke respectively (making country dummies very important in the model). In order to avoid singularity, the model has T-1 time effects and N-1 country dummy effects. The baseline country used was the U.S. The first lagged value of price had the correct sign (positive) and was significant in all models.

For asparagus, it can be seen that there is a country effect for all importers (compared to the U.S.). There is also an exchange rate effect for England. The country effects suggest that Spain, Netherlands, and England importers were paying relatively higher prices compared to what U.S. importers were paying. The negative significant exchange rate elasticity for England (-0.65) suggests that Peruvian exporters price-to-market this country to stabilize English pound prices in order to maintain their market share. This is easier to understand acknowledging that any depreciation in the foreign currency will increase the price in local currency so adjusting is related to local currency price stability. The result of the U.S. having relatively low prices in asparagus compared to all other importers is not surprising as it has by far the biggest share of Peruvian exports (see figure 1) thus possibly exerting buyer power. In the paprika market, it can be seen that there is a country and exchange rate effect only for Spain. Spain has been paying a relatively lower price compared to the U.S. and again Peruvian exporters were stabilizing Euro prices to maintain market share (elasticity of -0.60). Finally in the artichoke market, there are only country effects. French and Spanish importers were paying relatively higher prices compared to the U.S. As in the asparagus case, this can be a result of some buyer power being exerted as the quantity bought by this country is significant (see figure 2.1).

Table 2.5 shows the model divided by before and after the PTPA. Before the PTPA, there was a country effect for the Netherlands and England in asparagus, Spain in Paprika, and France and Spain in the artichoke market. Netherlands and England (asparagus) and France and Spain (paprika) were getting higher prices compared to the U.S. However, Spain was getting lower prices in the paprika market (compared to the U.S.). The negative exchange rate elasticity for England (-0.86) and the U.S. (-0.34) suggest that Peruvian exporters were stabilizing the English pound in asparagus and the U.S. dollar in artichokes. Now onto the after the PTPA results, there are country effects for all importers in the three markets. Compared to the U.S., all other countries paid higher prices in the asparagus and artichoke markets but lower prices in the paprika market. Peruvian paprika exporters were stabilizing Mexican peso prices as the exchange rate elasticity was -0.42. The most interesting finding of this study is that after the PTPA was enforced, the exchange rate elasticity for the U.S. became positive and significant (1.16) meaning that Peruvian asparagus exporters were pricing-to-market the U.S. by amplify the effect of the exchange rates.

Table 2.4 Effect of country dummies and exchange rate on Peruvian non-traditional export prices.

	Asparagus					
	Country effect			Exchange rate effect		
Destination	Coefficient		t-value	coefficient		t-value
Spain	0.032576	**	2.35	0.352768		0.63
Netherlands	0.216608	***	7.75	0.117772		0.32
England	0.234416	***	7.92	-0.65399	*	-1.74
U.S.				0.009026		0.06
	Adj. R-squared=0.8790					
	# of observations=90					
	Paprika					
	Country effect			Exchange rate effect		
Destination	coefficient		t-value	coefficient		t-value
Mexico	-0.01402		-1.2	-0.04655		-0.51
Spain	-0.07913	***	-4.51	-0.60422	*	-1.7
U.S.				0.082362		1.01
	Adj. R-squared=0.8680					
	# of observations=90					
	Artichoke					
	Country effect			Exchange rate effect		
Destination	coefficient		t-value	coefficient		t-value
France	0.233801	***	8.47	0.391535		0.68
Spain	0.104868	***	5.56	0.491015		1.13
U.S.				0.110754		1.36
	Adj. R-squared=0.6471					
	# of observations=90					

Level of significance: *** p<0.01, ** p<0.05, * p<0.1

Table2.5 Effect of country dummies and exchange rate on Peruvian non-traditional export prices: Pre and post Peru-U.S PTPA

	Pre PTPA						Post PTPA					
Asparagus												
	Country effect			Exchange rate effect			Country effect			Exchange rate effect		
Destination	coefficient		t-value	coefficient		t-value	coefficient		t-value	coefficient		t-value
Spain	0.027367		1.5	0.000893		0.01	0.04267	**	2.05	0.736473		0.82
Netherland	0.254073	***	6.18	0.202959		0.43	0.216994	***	5.54	0.235281		0.5
England	0.27424	***	6.24	-0.86196	**	-2.04	0.238576	***	5.73	-0.35895		-0.55
U.S.				-0.41171		-1.41				1.155469	**	2.06
	Adj. R-squared=0.8997						Adj. R-squared=0.8528					
	# of observations=46						# of observations=44					
Paprika												
	Country effect			Exchange rate effect			Country effect			Exchange rate effect		
Destination	coefficient		t-value	coefficient		t-value	coefficient		t-value	coefficient		t-value
Mexico	-0.00395		-0.23	0.114352		0.37	-0.03142	*	-1.91	-0.42357	*	-1.87
Spain	-0.09793	***	-3.53	-0.25646		-0.62	-0.08096	***	-3.42	-0.94562		-1.62
U.S.				0.32474		1.45				0.131351		0.63
	Adj. R-squared=0.8803						Adj. R-squared=0.8487					
	# of observations=46						# of observations=44					
Artichoke												
	Country effect			Exchange rate effect			Country effect			Exchange rate effect		
Destination	coefficient		t-value	coefficient		t-value	coefficient		t-value	coefficient		t-value
France	0.281752	***	6.99	-0.32348		-0.48	0.246205	***	6.26	1.533187		1.57
Spain	0.170665	***	5.88	0.6089		0.98	0.064355	**	2.55	0.546275		0.9
U.S.				-0.33704	*	-1.81				0.180047		0.82
	Adj. R-squared=0.7298						Adj. R-squared=0.5901					
	# of observations=46						# of observations=44					

Level of significance: *** p<0.01, ** p<0.05, * p<0.1

The results obtained here(shown in table 2.4 and table 2.5) indicate that almost all the time when exchange rates effects were significant, prices were being stabilized in the importer's currency. This is consistent with PTM agricultural research findings, where the majority of the coefficients for importing countries had a negative sign in the exchange rate effect (Carew, 2000, Miljkovic, et al., 2003, Pall, et al., 2013).

2.6 Conclusion

This paper has examined the market power of Peruvian non-traditional vegetables in an international trade context. Using a PTM model, the results suggest that Peru is engaging in price discriminating behavior. On average, country markups through fixed effects were being applied and Peruvian exporters were stabilizing English pound prices in the asparagus market and Euro (Spain) prices in the paprika market. In addition, after the preferential free trade agreement between U.S. and Peru was enforced, Peruvian exporters were pricing-to-market the U.S. by amplifying the effects of movements in the exchange rate (U.S. dollar / Nuevo sol). However, the finding of incomplete pass-through should be cautiously interpreted because the assumption of constant marginal cost in the fresh asparagus market cannot be tested with available data.

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Chapter 3:

Hispanics and Citizenship Status: An Investigation of Wages Using Quantile Regression with Sample Selection Correction in the Southeastern United States

Abstract

It is undeniable that in the United States Hispanic people carry a greater stigma regarding their immigration status compared to other ethnicities. This study focuses on citizenship status' effect on wages using the model proposed by Albrecht et al. (2009) to estimate quantile wage regressions accounting for sample selection in the states of Alabama and Georgia. The sample selection variable is the household head living in a metropolitan area or not. Wages for two groups were analyzed: Hispanics and non-Hispanics (all other ethnicities). Results suggest that using an average estimation of the effect of citizenship on wages for non-Hispanics does not reflect what is really happening in the different quantiles. For the Hispanic group, all the quantiles showed a positive coefficient (of having citizenship) and a higher magnitude in the lower one. However, this higher magnitude was not economically significant compared to the other quantiles. It is concluded that Hispanic are negatively affected by not having the proper immigration status regardless of the quantile they are in, thus an average effect regression provides a good fit for testing the impact of citizenship status on Hispanic people.

3.1 Introduction

Regular regression models estimate the average effect of exogenous variables over an endogenous variable. However, the results (especially for policy analysis) may be misleading as they do not capture changes in the distributions for different segments of the data. Ordinary Least Squares (OLS) regressions using binary variables does not solve this problem as the technique only account for different intercepts and slopes (if including interactions) but not different inherent distributions. Is there a proper way to consider these differences in distributions without truncating the data? Yes, Quantile regression modeling.

Quantile regressions have been widely used in several fields. However, as in regular regressions, the inclusion of a predictor that is endogenous (or sample selection related) can bias the results. In labor economics, there are some studies applying quantile regression in addition to a correction for sample selection. Recently, some studies addressing the differences in wage salaries between men and woman while correcting for sample selection have been made (Albrecht, et al., 2009, Chzhen and Mumford, 2011, Garcia, et al., 2001). These studies have shown that the returns of different characteristics for each group (gender) have different impacts depending on what quantile is being analyzed.

In similarity to the studies mentioned, this study analyzes wages by using a quantile regression model with sample selection correction but instead of focusing on differences between men and women, it focuses on differences between Hispanic ethnicity and other ethnicities[8]. This distinction is made because it is generally accepted that not having a proper immigration status is a greater stigma for Hispanics than for all other ethnicities because of the huge amount

[8] For simplicity, it is assumed in that men and women groups can be pooled together allowing having only groups by ethnicity. In reality, first we would need to separate the groups by gender (as coefficients may be differ) or do a F chow-test (to prove that they can be pooled together) to then proceed to separate by ethnicity. The approach of separating by gender is not taken as it reduces the amount of observation (degrees of freedom) in specific groups thus limiting the power of the inferences. Interactions between gender and the other variables will help to reduce this concern.

of illegal immigration from Mexico and Central America to the United States. The study uses data from the states of Alabama and Georgia prior to the proposal of current immigration laws. The hypotheses of this study are that the impact of citizenship status in the wage equation is quite different for Hispanic people compared to the pool of all other ethnicities, and that these differences can be more pronounced depending of what quantile (especially low quantile) is being analyzed.

The rest of this paper is structured as follows: similar topics addressed by other authors are reviewed in section 2. The methodology and the data are described in section 3; section 4 shows and comments on the results obtained; conclusions are presented at the end.

3.2 Literature Review

Latino or Hispanic ethnicity is the fastest growing minority in the United States, so it is not surprising that there are numerous studies addressing the effect of being Latino or Latino descendent on wages. The easiest and simplest way to address this topic would be to include a binary variable in a wage regression model. One example is the paper by Goodrum (2004) who estimates the effect of being Hispanic in the construction industry. However, it is implausible to think that the groups being analyzed (white and Hispanic ethnicities) can be pooled together (unless there is a Chow test stating it). Better econometric methods are needed.

A couple of studies as early as in the 80's (King, et al., 1986, Reimers, 1984) addressed the differences in wages for several different types of Hispanic men (several groups according to country of origin) compared to black non-Hispanic and white non-Hispanic men by using simple OLS regressions. Note that these early studies already distinguished between ethnicity and race (e.g. a person can be Latino and of black race). The most interesting findings by Reimers (1984) were that Hispanic groups had lower returns to education than Anglos and that the race among Hispanics has no significant impact on wages; meanwhile the most interesting finding by King et al. (1986) was that as Hispanic migration becomes larger, it positively impacts the wages of U.S.-born Hispanics.

Latino wage studies in the U.S. became more specific and started to account for further differences between Latinos, especially separating by gender. A study made by Torres and McQuillan (2007) using OLS and the Integrated Public Use Microdata Series (IPUMS) data found that, after separating by country of origin subgroups, new economy characteristics (such as high tech working ability) are most important for non-Hispanics men compared to other groups. They also found that Mexican women have the lowest estimated average earnings. McCall (2001) also used OLS and found similar results in her study. Significant differences were found in the sources of wage inequality across race, ethnicity, and gender and that these differences are greater between ethnic groups not much between men and women. She also found a large negative effect of immigration on the relative wages of Latinos. In conclusion, these two studies found that ethnicities should not be pooled together as estimated coefficient are likely to be different, that human capital characteristics contribute significantly to earning gaps, and that these gaps are greater among ethnic groups rather than gender.

There are a few studies addressing the issue of immigration status' effect on wages in the U.S. labor market. Bratsberg et al. (2002) estimated the effect of naturalization on adult males using cross-sectional and panel data. The authors used regression model with a sample selection correction for the job being in the public sector (Heckman model). The first stage of the Heckman model included a naturalization exogenous variable. The authors found that naturalization has a positive effect on wages on all models. Zhou and Lee (2013) used cross-sectional data to explore the effects of U.S. citizenship on wages of Asian immigrant women. They found that citizenship has a positive effect of on wages and also that higher education significantly boosts wages for Asian women who have citizenship but not for those who do not have it. Examination of the U.S. nurse labor market and the effect of citizenship was made by Schumacher (2011). The author's regression models showed a 4.5% lower wage for non-citizen nurses. They also showed that this disadvantage is greater when the nurse is new to the U.S. One particularity of these studies is that they have used regular regressions models (generally OLS) to make their inferences, so, an average effect was estimated. A quantile regression would be more informative as it allows for different distributions for different segments of the data.

Not much studies using quantile regression in the labor market have been made. Most of the recent ones tend to focus on gender differences. However there are very few studies of this type using U.S. data and moreover none of these have addressed the effect of U.S. citizenship on wages. Kim and Min (2006)examined the effect of technology adoption on the wage dispersion in the U.S. manufacturing sector using a quantile regression. They found that high-wage quantiles have adopted technologies more actively than the others and that this adoption has contributed to widening the gap between quantiles. Another study made using U.S. data was made by Martinez-Sanchis et al(2012). Here it was found that age (measuring a proxy for ability) had a negative impact on wages only in low quantiles.

As it was mentioned before, the base methodology for the present study is the one proposed by Albrecht et al(2009) where they used quantile regression with sample selection correction. In this study (Albretch et al, 2009), the authors found that, in the Netherlands, there is a 15-20% gap between men and women wages and that this gap increases as we move up the distribution. A recent study using Albrecht et al's (2009) methodology was made by Chzhen and Mumford

(2011) to assess any differences on wages between genders in Great Britain. This study found that the raw gender wage gap shows a tendency to increase across the distribution and when sample selection was included a larger gender earning gap was found.

This study makes two contributions to the literature. First, it measures the effect of citizenship status on U.S. wages in a sample selection corrected quantile regression context. Second, it focuses on ethnic differences especially on Latino population.

3.3 Methodology and Data

Using the Integrated Public Use Microdata Series (IPUMS), which is the world's largest individual-level population database, a dataset of 56,740 observations was created. The dataset contains information regarding the household head of homes in the states of Alabama and Georgia for the year 2010. This year was chosen because it was before the Alabama HB56 immigration law was proposed, allowing to have a smaller bias through lying about immigration status. With this information, a semi-log (log-linear) wage equation was built. The wage measurement is the natural logarithm of the annual earning. The human capital model, as the basis for the earning function (Becker, 1962), is adopted here. This model assumes that wages increase with measurements of accumulated skills such as education and experience. Education is measured by binary variables depending on the highest level of education achieved and experience is measured by age. Other binary variables such as if the household head lives/works in a metropolitan area, gender, if the person is of black race[9] , if the person has citizenship, and if the person lives in the state of Georgia are also included as they will affect earnings[10]. In our

[9] Take note that Hispanic is not a race but an ethnicity. There are Hispanics that can be white, black, and even Asian in race.

[10] As the model is using a lot of dummy variables, it is important to state the baseline comparison. The baseline is: no high school degree, living in a non-metropolitan area, female, white, non-citizen, and lives in the state of

model, the metropolitan variable is sample selection related and should be corrected for[11]. Other variables have also the potential to be endogenous (especially education or citizenship), however, the use of a Hausmann endogeneity test assumed this concern away. The model in its general form is as follows:

Log(wage) = f(age, education, metropolitan, male, black, Georgia, citizenship) (3.1)

The variable "age", which reflects experience, is continuous, thus allowing the regression to have predictive power and not being just an ANOVA using qualitative variables. A quadratic specification in "age" was considered as it has been extensively shown in the literature that after a certain point there will be an inflection point as too much age affects negatively productivity. Table 3.1 shows the descriptive statistics of the data. It can be seen that, on average, Hispanic household heads have lower annual income compared to other ethnicities. Hispanics also have lower level of education and black percentage of people compared to other ethnicities. Hispanic people tend to be younger, live more in metropolitan areas, have a greater male/female ratio, and live more in Georgia than Alabama compared to other ethnicities. One very important variable in our model is citizenship status (having citizenship). The general statistics show that only half of the Hispanic household heads have United States citizenship compared to 98% of other ethnicities household heads.

Alabama.

[11] It would have been better to use an explicit distinction between living in the city or rural area, however, the metropolitan area variable was the closer proxy found.

Table3.1 Descriptive statistics

	All other ethnicities			Hispanics		
Variable	**N**	**Mean**	**Std Dev**	**N**	**Mean**	**Std Dev**
wage	54794	29256.06	44120.76	1946	25694.88	35022.92
age	54794	52.32	16.75	1946	40.49	13.48
metropolitan	54794	0.65	0.47	1946	0.71	0.44
male	54794	0.51	0.49	1946	0.61	0.48
bachelor	54794	0.17	0.38	1946	0.12	0.33
gradschool	54794	0.11	0.31	1946	0.07	0.25
highschool	54794	0.36	0.48	1946	0.28	0.45
black	54794	0.25	0.43	1946	0.05	0.23
citizenship	54794	0.98	0.13	1946	0.52	0.49
georgia	54794	0.64	0.47	1946	0.83	0.37

Following Albrecht et al (2009), the quantile regression accounting for sample selection model is as follows: first we have group "a" and group "b". "a" denotes all workers and "b" denotes those people who select to work in a metropolitan area. Y is the random dependent variable (natural log of wage), X is a stochastic vector of predictors measuring the different characteristics, and x is the realization of this vector. Let $Y_a$be the counterfactual random variable representing the log wage that a randomly selected person would earn were he/she to work in a metropolitan area. The quantiles of Y_a conditional on x_a are given by

$$quant_u(Y_a|X_a = x_a) = x_a\beta^a(u) \quad u\epsilon[0,1] \tag{3.2}$$

Where $\beta^a(u)$ is the value of the coefficient correcting for selection at the "u" quantile. Equation Then using a semi-parametric estimator we obtain

$$quant_u(Y_b|Z_b = z_b) = x_b\beta^a(u) + h_u(z_b\gamma) \quad u\epsilon[0,1] \tag{3.3}$$

where Z is a set of observable characteristics that influence the probability that a person working in a metropolitan area, and the term $h_u(z_b\gamma)$ corrects for selection at the uth quantile. Regarding the selection equation, Z can use variables that were also included in the wage regression but needs at least one variable that is not included in the main model. In this application Z includes: age, education, and the instrument is if the person was single or not. The person being single or not is used as an instrument because being married is generally strongly correlated with having kids. It is also generally recognized that when a family is starting to grow, a bigger place that has backyard and where not too much traffic is present is usually preferred. Usually you cannot find this in the city (metropolitan area). As for an easy way to understand the other term in equation (3.3) Albrecht et al (2009) is quoted: “$h_u(z_b\gamma)$ plays the role that the mills ratio in the usual Heckman procedure…”. As a final step, Buchinsky (1998) suggested a series estimator for $h_u(z_b\gamma)$ as follows

$$h_u(z_b\gamma)hat = \delta_0(u) + \delta_1(u)\lambda(z_b\gamma) + \delta_2(u)\lambda(z_b\gamma)^2 + \ldots \tag{3.4}$$

where λ is the inverse mills ratio. The function in equation (3.4) is a power series approximation of $h_u(z_b\gamma)$. For appropriate values of the δ’s, $h_u(z_b\gamma)hat$ tends to approach $h_u(z_b\gamma)$as the number of terms goes to infinity.

For comparison purposes, the results will also include the following models: OLS without sample correction, OLS with sample correction (Heckman), and quantile regression without sample correction. The quantile regression without sample correction used the resampling option to allow for bootstrapping standard errors and probability values. The main model was estimated using the command “mmsel” in STATA 11.0. The other models were estimated using SAS 9.2.

3.4 Results and Discussion

3.4.1 All Other (non-Hispanic) Ethnicities

For the first part, the results for other ethnicities are presented. Table 3.2 shows the results for all other ethnicities using uncorrected and corrected OLS (thus average effect). All the coefficients were significant at a 10% level (except for high school in the uncorrected version). It can be seen that when the sample correction is accounted for, all of the coefficients change. The effects of age, male, gradschool degree, and citizenship become greater; meanwhile the opposite happens for the effect of bachelor degree, black, and if the person lives in the state of Georgia. Regarding our variable of interest, citizenship status makes the person earn on average 44% more compared to those who do not have it. The interaction effect between age and male is positive in both specifications, meaning that experience in males is more important compared to experience in females for determining wages.

Table3.2 Estimation results for other ethnicities using OLS.

	WITHOUT CORRECTION			**WITH CORRECTION**		
Parameter	**Estimate**	**Standard**	**Pr> |t|**	**Estimate**	**Standard**	**Pr> |t|**
Intercept	3.41	0.21	<.0001	3.16	0.24	<.0001
age	0.19	0.006	<.0001	0.23	0.007	<.0001
age2	-0.003	0.00005	<.0001	-0.003	0.00006	<.0001
metropolitan	-0.03	0.002	<.0001			
male	0.63	0.03	<.0001	2.32	0.14	<.0001
Age*male	2.64	0.11	<.0001	-0.02	0.002	<.0001
bachelor	1.45	0.05	<.0001	1.43	0.06	<.0001
gradschool	1.97	0.06	<.0001	1.98	0.07	<.0001
highschool	0.01	0.04	0.8079	-0.09	0.05	0.0860
black	-0.42	0.04	<.0001	-0.33	0.04	<.0001
citizenship	0.41	0.13	0.0022	0.44	0.14	0.0019
georgia	0.31	0.03	<.0001	0.28	0.04	<.0001

Turning to the quantile regression estimation, table 3.3 shows the result for the quantile regression using the 0.25, 0.50, and 0.75 quantiles. Again it can be seen that the coefficients change depending on if the correction was applied or not.

The corrected quantile regression model shows that in the lower level quantile, the variable age has a negative significant coefficient meanwhile in the other quantiles it has the expected positive effect. This may be explained that in lower level jobs, the age variable is reflecting more an incapability through getting old rather than work experience. These low level jobs may be requiring physical activities thus these results are plausible. The interaction between age and male is negative in the 0.25 and 0.50 quantile but positive in the upper quantile meaning perhaps that in upper level jobs, work experience reflected by age is more important for men than for women. Leaving in a metropolitan area increases the wage for lower and medium quantiles and decreases it for the upper quantile. Being male gets you a higher wage only in lower and middle quantiles. Having a bachelor or a graduate level degree gets you higher wages in all quantiles compared to elementary education. Being black gets you a lower wage in all quantiles. Talking about our variable of interest, citizenship, having this status affects positively wages in the mid and high quantiles. It is more important for those in the middle quantile making them to earn 34% more compared to those who do not have it. However, this status affects negatively those in the lower quantile. The last part could be perhaps explained by that lower jobs may require physical challenges and American may be perceived as “lazier” compared to immigrants more motivated by achieving the “American dream”. Lastly, living in the state of Georgia earns you a higher salary compared to Alabama in all quantiles.

Table3.3 Quantile regression for other ethnicities.

	Quantile=0.25			Quantile=0.50			Quantile=0.75		
Parameter	**Estimate**	**Standard**	**Pr> |t|**	**Estimate**	**Standard**	**Pr> |t|**	**Estimate**	**Standard**	**Pr> |t|**
				UNCORRECTED					
Intercept	15.55	0.32	<.0001	1.63	0.11	<.0001	2.67	0.09	<.0001
age	-0.40	0.005	<.0001	0.40	0.004	<.0001	0.37	0.004	<.0001
age2	0.002	0.00	<.0001	-0.005	0.00	<.0001	-0.004	0.00	<.0001
Age*male	-0.06	0.002	<.0001	-0.008	0.001	<.0001	0.01	0.001	<.0001
metropolitan	0.32	0.02	<.0001	0.39	0.02	<.0001	0.25	0.01	<.0001
male	4.63	0.16	<.0001	0.95	0.06	<.0001	-0.05	0.05	0.3678
bachelor	0.94	0.05	<.0001	0.77	0.02	<.0001	0.58	0.01	<.0001
gradschool	1.17	0.09	<.0001	1.08	0.02	<.0001	0.89	0.03	<.0001
highschool	-0.00	0.02	1.0000	-0.03	0.02	0.1742	-0.06	0.01	0.0002
black	-0.26	0.03	<.0001	-0.29	0.02	<.0001	-0.30	0.01	<.0001
citizenship	-0.57	0.22	0.0089	0.39	0.05	<.0001	0.32	0.04	<.0001
georgia	0.14	0.02	<.0001	0.21	0.02	<.0001	0.12	0.01	<.0001
				CORRECTED					
Intercept	11.10	0.35	<.0001	1.40	0.11	<.0001	2.99	0.08	<.0001
age	-0.46	0.005	<.0001	0.39	0.004	<.0001	0.37	0.002	<.0001
age2	0.002	0.00	<.0001	-0.005	0.00	<.0001	-0.004	0.00	<.0001
Age*male	-0.05	0.002	<.0001	-0.008	0.001	<.0001	0.01	0.001	<.0001
metropolitan	11.39	0.44	<.0001	1.48	0.19	<.0001	-0.29	0.14	0.0415
male	4.18	0.15	<.0001	0.94	0.05	<.0001	-0.05	0.05	0.3206
bachelor	3.89	0.14	<.0001	1.19	0.04	<.0001	0.55	0.03	<.0001
gradschool	4.42	0.16	<.0001	1.50	0.05	<.0001	0.83	0.05	<.0001
highschool	-0.47	0.03	<.0001	-0.13	0.03	<.0001	-0.08	0.01	<.0001
black	-0.07	0.03	0.0203	-0.18	0.02	<.0001	-0.27	0.01	<.0001
citizenship	-0.53	0.17	0.0021	0.34	0.04	<.0001	0.30	0.04	<.0001
georgia	0.15	0.02	<.0001	0.18	0.01	<.0001	0.11	0.01	<.0001

3.4.2Hispanic Ethnicity

For the second part, the results for Hispanic ethnicity are presented. Table 3.4 shows the results for the average coefficients for Hispanic wages. Again, it is seen that the coefficients change if the correction is applied or not. Regarding our variable of interest, those people who have citizenship earn on average 37% more compared to those who do not have it.

Table 3.4 Estimation results for Hispanics using OLS.

	WITHOUT CORRECTION			WITH CORRECTION		
Parameter	**Estimate**	**Standard**	**Pr> |t|**	**Estimate**	**Standard**	**Pr> |t|**
Intercept	-0.84	0.85	0.3257	8.14	0.69	<.0001
age	0.36	0.03	<.0001	0.10	0.02	0.0001
age2	-0.004	0.0003	<.0001	-0.001	0.0002	<.0001
metropolitan	0.06	0.19	0.7467			
male	3.94	0.56	<.0001	0.77	0.33	0.0198
Age*male	-0.04	0.01	0.0005	-0.003	0.007	0.6817
bachelor	1.22	0.28	<.0001	-0.14	0.36	0.6948
gradschool	2.05	0.36	<.0001	0.84	0.46	0.0727
highschool	0.12	0.20	0.5373	-0.77	0.27	0.0050
black	0.09	0.39	0.8135	-0.05	0.20	0.7729
citizenship	0.65	0.19	0.0008	0.37	0.11	0.0011
georgia	-0.35	0.23	0.1385	0.01	0.14	0.8973

Now we turn the attention to the quantile regression. Table 3.5 shows the results for the quantile regression with and without sample correction. Again it can be seen that the coefficients change depending on if the correction was applied or not.

Regarding our variable of interest, citizenship is important for all quantiles. The quantile where this variable impacts the most is the lower one. This means that if a person has citizenship in this quantile, his wage will increase by 53%. However, as all the quantiles show a positive effect and not that much difference in magnitude, the use of an average coefficient would have been enough (remember that a coefficient of 0.37 was obtained using corrected OLS)[12].

The corrected quantile regression model shows that age and its squared form have the correct sign for all quantiles; however, the coefficient with largest magnitude (for its linear form) is present in the lower quantile meaning that having experience have greater returns in low paying jobs. The interaction between male and age is negative and significant in the lower quantile, meaning that an older male tends to earn less in this quantile. This makes sense for low level jobs as for example in the construction industry, an employer rather have a young male employee (usually correlated with good health) that does not have much experience but can work on very physical demanding conditions. Another way to see this is taking a look at the other side of the coin. In the house cleaning industry, an employer rather have an experienced woman to do the cleaning as know-how is very important in this case. Living in a metropolitan area has no significant impact on any quantile. Being male will make the household head to earn more specially in low level jobs. Having bachelor and graduate school degrees also help to earn a higher wage compared to elementary school in all quantiles. These two degrees have the biggest impact on the lower quantile. Finally, being a Hispanic of black race or living in the state of Georgia does not affect the wage received in any of the quantiles.

Finally, figure 3.1 allows an easy comparison of the different coefficients along the quantiles between the two groups. It can be seen that other ethnicities' confidence interval coefficients tend to be narrower (less variation) compared to Hispanic's ones.

[12]Recall that this OLS results would be closer to match up to a simple average using the results from the different quantiles only If we were using continuous and not discrete values for quantiles

Table 3.5 Quantile regression for Hispanics.

	Quantile=0.25			Quantile=0.50			Quantile=0.75		
Parameter	**Estimate**	**Standard**	**Pr> |t|**	**Estimate**	**Standard**	**Pr> |t|**	**Estimate**	**Standard**	**Pr> |t|**
					UNCORRECTED				
Intercept	-8.02	0.94	<.0001	1.19	0.61	0.0523	5.99	0.50	<.0001
age	0.42	0.04	<.0001	0.39	0.02	<.0001	0.18	0.02	<.0001
age2	-0.004	0.0005	<.0001	-0.004	0.0003	<.0001	-0.002	0.0004	<.0001
Age*male	-0.11	0.01	<.0001	-0.01	0.01	0.1791	0.00	0.005	0.4253
metropolitan	0.04	0.15	0.7637	0.11	0.07	0.1114	0.08	0.05	0.1449
male	12.10	0.60	<.0001	1.44	0.47	0.0021	0.31	0.20	0.1239
bachelor	1.16	0.20	<.0001	0.68	0.11	<.0001	0.60	0.07	<.0001
gradschool	1.79	0.28	<.0001	1.20	0.16	<.0001	1.14	0.09	<.0001
highschool	0.17	0.18	0.3415	0.07	0.06	0.2978	0.03	0.05	0.4806
black	-0.09	0.47	0.8494	0.06	0.17	0.7050	0.10	0.12	0.4042
citizenship	0.55	0.15	0.0005	0.53	0.07	<.0001	0.46	0.05	<.0001
georgia	0.07	0.19	0.7206	0.006	0.10	0.9521	0.01	0.06	0.8064
					CORRECTED				
Intercept	-8.13	1.15	<.0001	1.14	0.68	0.0963	5.72	0.49	<.0001
age	0.42	0.04	<.0001	0.39	0.02	<.0001	0.18	0.02	<.0001
age2	-0.004	0.0005	<.0001	-0.004	0.0003	<.0001	-0.002	0.0004	<.0001
Age*male	-0.11	0.01	<.0001	-0.01	0.01	0.2609	0.001	0.005	0.8151
metropolitan	0.23	1.57	0.8795	0.36	0.87	0.6779	0.77	0.57	0.1752
male	12.01	0.54	<.0001	1.31	0.45	0.0037	0.44	0.21	0.0376
bachelor	1.24	0.41	0.0030	0.82	0.23	0.0004	0.84	0.16	<.0001
gradschool	1.82	0.39	<.0001	1.30	0.21	<.0001	1.31	0.14	<.0001
highschool	0.22	0.25	0.3859	0.10	0.11	0.3294	0.13	0.08	0.1087
black	-0.12	0.47	0.7875	0.07	0.19	0.7010	0.08	0.13	0.4989
citizenship	0.53	0.17	0.0025	0.51	0.07	<.0001	0.44	0.05	<.0001
georgia	0.06	0.17	0.7298	-0.003	0.09	0.9679	0.03	0.07	0.6730

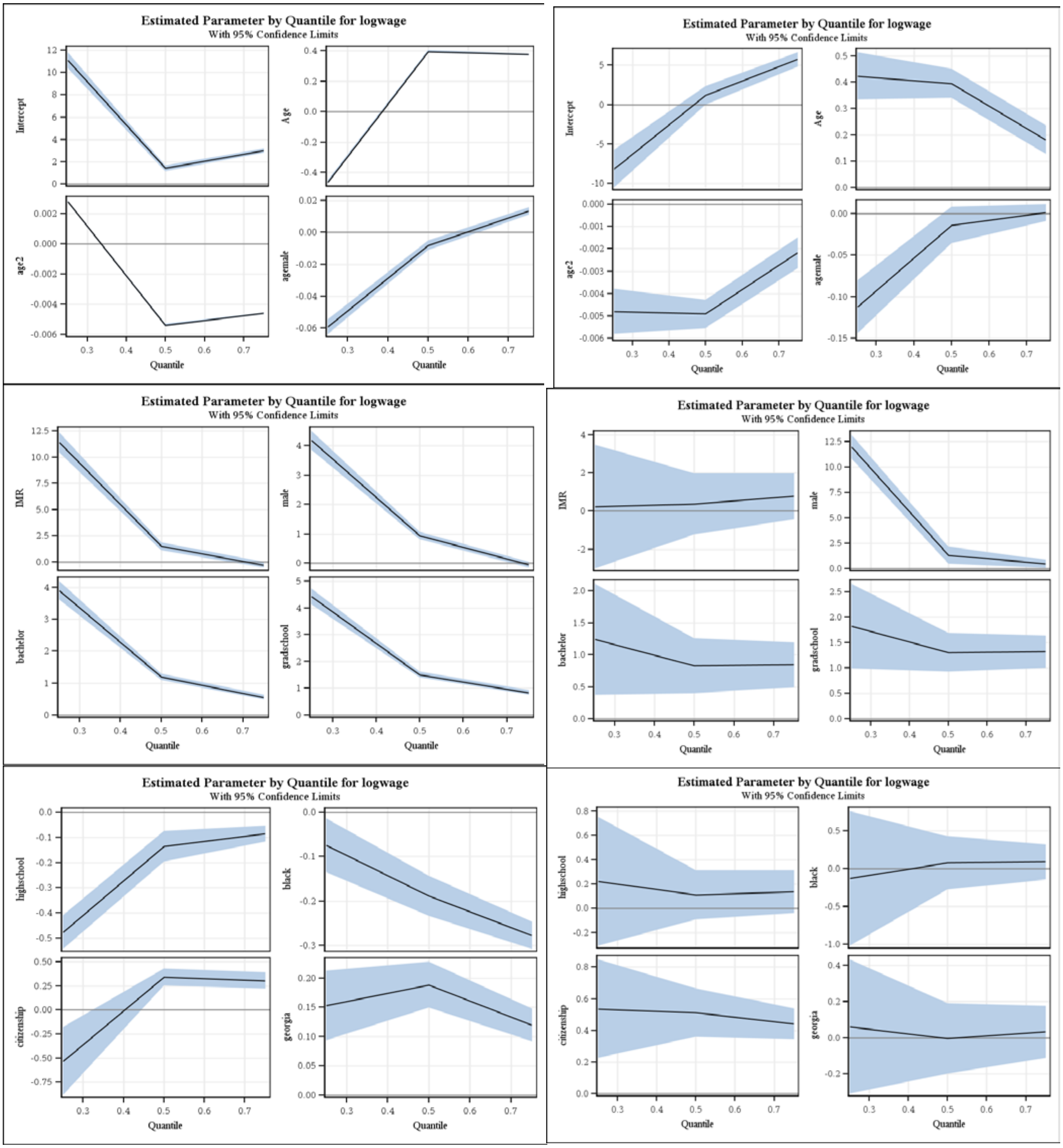

Figure 3.1 Comparison between estimated coefficients for all ethnicities and Hispanics using corrected quantile regressions. The left side is for other ethnicities and the right side is for the Hispanic group

3.5 Conclusion

It can be concluded that U.S. citizenship has a strong effect on wages in the states of Alabama and Georgia. There are differences in the magnitude of this effect depending on different ethnicities and the different quantiles being analyzed. The average effect for non-Hispanics was not really reflecting the importance of this variable on wages as different distributions had different coefficient's signs being negative in the lower quantile and positive in the two other quantiles. In the case of Hispanic people, the average effect reflected an acceptable approximation of what happened in the majority of the quantiles as the difference between the effect on wages in the 25th and 75th quantile was only of 9%. This result suggests that Latinos face a similar citizenship stigma on wages regardless of the quantile in which they are located.

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