ABSTRACT

QIU, FENG. Three Essays on Government Policy and Agricultural Markets. (Under the direction of Barry K. Goodwin.)

The dissertation includes three essays. The first essay investigates the impacts of decoupled and coupled government program payments on farmland rental contract choices for a subset of U.S. crop farms using a principal-agent model. The scope of the study includes cash and share contracts as well as hybrid contracts, which represent an increasingly prominent feature of U.S. agriculture. The conceptual framework suggests that restrictions on payments between contracting parties are ineffective and induce an offsetting contractual rearrangement. Empirical results from a multinomial logit model confirm that government support programs have large and significant effects on contract choices and these effects vary by types of programs.

The second essay proposes a modified Ricardian rent framework and evaluates to what extent farm programs affect farmland rental rates, taking risks and transaction costs into consideration. The empirical model corrects for potential selection biases that arise because of the influence of farmland leasing contract choices made by landlords and tenants. The empirical results suggest that government subsidies have significant effects on rental rates. The study finds that landlords capture around 41% of the aggregate subsidies under cash leases and 78% under share contracts. Farm program payments are found to have different impacts on rental rates depending on the types of programs and leasing arrangements.

The third essay investigates the effects of state-dependent policy interventions on price transmission. The empirical application is the price linkages between Ukrainian domestic wheat price and the world price. The empirical analysis follows the smooth transition conintegrating (STC) framework of Saikkonen and Choi (2004), Choi and Saikkonen (2004), and Choi and Saikkonen (2010), and follows the general procedure used to investigate long-run equilibrium and short-run error correction suggested by Engle and Granger (1987). The results indicate that there is regime-switching behavior for the long-run relationship between Ukrainian domestic and

world market prices, based on the world price. In particular, when the world price of wheat is below the threshold of \$185/ton, the transmission elasticity of domestic price with respect to the world price tends towards unity. However, when the world price is above the threshold level, the relationship between the two markets approaches another regime and the transmission elasticity drops to 0.7, which indicates that a 1% increase in the world market price results only in a 0.7% increase in the Ukrainian domestic price. The results suggest that Ukrainian market is well integrated into the world market. However, active government interventions in trade activities can cause great long-term losses for Ukrainian producers.

Three Essays on Government Policy and Agricultural Markets

by
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DEDICATION

To my parents and my husband.

BIOGRAPHY

Feng Qiu grew up in Zhejiang, China. She obtained a BA degree in International Economics and Trade from Nanjing University of Finance and Economics in May, 2003. Feng came to the United States for graduate study in the fall of 2004. After she earned her master's degree in Economics at North Carolina State University (NCSU) in 2006, she continued her Ph.D. study of Economics at NCSU. Feng is broadly interested in agricultural policy analysis, market and price analysis, agricultural production and supply, applied econometrics, and risk and insurance modeling. In addition to research, she has sharpened her teaching skills. Under the guidance and support of the excellent faculty at NCSU, she trains herself to become a researcher and a teacher in agricultural and resource economics. Feng has been recognized both as the University and the College Outstanding Graduate Student Teaching awards winner at NCSU. She is also the 2011 recipient of the North American Colleges and Teachers of Agriculture (NACTA) Graduate Student Teaching Award. Upon graduation, Feng will join the faculty of the Department of Resource Economics and Environmental Sociology at the University of Alberta in Edmonton, Alberta, Canada, and work as an assistant professor there.

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Chapter 1

Introduction

This study is comprised of two related parts. The first part has two essays that address an important issue in economics and agricultural policy; namely farm program benefit distribution between landowners and tenant farmers. The two essays evaluate the effects of U.S. farm programs on farmland leasing arrangements and rental rates. The second part is about the impacts of policy intervention on price transmission as it applies to the Ukrainian wheat market. Ukraine is an interesting case study as it is a typical transition country with active and frequent government intervention in its economy. It is also one of the world's top grain exporters. Appropriate investigations of the linkage between the domestic and world markets can provide valuable information for future policy recommendations regarding food security, market efficiency, and trade liberalization.

1.1 Benefit Distribution

The United States has a long history of providing generous support to the agricultural sector. The primary goals of these farm programs are to increase farm income and reduce income volatility for farmers. In this sector, land renting is a common practice, with about 40% of the farmland in operation rented from others. Contrary to conventional wisdom, most agricultural landlords are non-operator individuals that work in or are retired from non-farm-

related activities (Goodwin et al. 2010).

Farmland rental rates, land values, and land-use decisions are influenced by these farm programs. The basic economic theory of rent implies that these subsidies will be largely, if not completely, capitalized into land values and rents as they represent increases in the net returns of land use. Consequently, programs that aim to help poor farmers may actually benefit relatively wealthy and non-farming landowners. How landlords and tenant operators share program benefits is an important issue. The more benefits that are passed through to landowners, the less effective these programs are as a tool to help farmers. Furthermore, a large proportion of the capitalization of program benefits into land values and rents can also have significant impact on land-use decisions, which can result in other policy spillovers like reducing incentives for conservation and exacerbating deforestation.

1.1.1 Effects of Farm Program Payments on Farmland Contract Choices

Given this background, the first part of this study investigates how and to what extent landowners capture the program benefits that are supposedly intended for farmers. The most intuitive and direct way to measure the benefit division between landlords and tenants might be estimating the impact of subsidies on farmland rental rates, as the relevant literature usually does. However, this study argues that in addition to raising rents for a given leasing arrangement, landowners may choose to capture extra benefits through changing/switching leasing arrangements, especially when there are legislative restrictions on payment sharing between contracting parties.

For example under the 1996, 2002, and 2008 Farm Bills, legislation required that direct payments be shared between tenant operators and landlords on a "fair and equitable" basis. To receive payments, an individual had to share in the risk of production and had to be entitled to share in farm receipts. If a landlord received cash rents, then the tenant operator bore all the risks and should have received the entire amount of government subsidies. Naturally, landlords might indirectly capture a share of, if not all, the benefits by raising cash rents. In

contrast, a share lease involves a distribution of payments to both the landlord and the tenant operator directly from the government, according to the pre-determined share percentage as they share the crop. In such a case, landlords would obtain certain benefits from the government directly and might also capture extra benefits that have been distributed to the tenant through a monetary term added to the existing contract. Thus, changes in tenure arrangements, from a share contract to a hybrid contract, may actually reflect a benefit pass-through from tenants to landlords.[1]

In addition, one can observe a certain type of rental rate only if the farmer and the landowner first choose to use this type of contract. Thus, empirical studies of subsidy incidence, which directly evaluate the impacts of subsidies on rental rates under a certain type of leasing arrangement, may be biased because of this selection issue regarding contract choices.

Therefore, the first essay in Chapter 2 investigates the impacts of decoupled and coupled program payments on rental contract choices. The second essay in Chapter 3 evaluates subsidy incidence in rental rates, taking the selection issue into account.

The conceptual framework of the first essay utilizes a principle-agent model. It concludes that exogenous legal restrictions on the distribution of benefits between landowners and tenant operators are likely to be ineffective. This can cause an offsetting contractual rearrangement to restore the benefit sharing to an unrestricted market level, as long as the market and contracting activities are free and landowners have the private property rights of their land.

The main dataset used in this essay is from the 1999 Agricultural Economics and Land Ownership Survey (AELOS), which is an integrated survey of farm finance and land ownership. It includes comprehensive information collected from both tenants and landlords. Empirical results confirm that decoupled payments discourage the use of share contracts and increase the use of hybrid contracts. Other support programs, like loan deficiency payments (LDP) and disaster assistance payments, also have significant effects on contract choices that vary by types of programs.

[1] A hybrid contract consists of a predetermined share percentage plus a fixed cash payment.

This analysis leads to two important implications. First, legal restrictions on benefit sharing between landlords and tenant operators are ineffective and induce offsetting contractual rearrangements. The increasing use of hybrid contracts likely reflects a redistribution of program benefits. Second, the results indicate that it is important to account for the non-random nature of the sample results from contract choices when estimating the impacts of subsidies on rental rates. The selection concern calls for additional analysis and it is tackled in the second essay.

1.1.2 Effects of Farm Program Payments on Farmland Rental Rates

The second essay evaluates to what extent each of the main commodity programs affects farmland rental rates and investigates whether the subsidy incidence differs across different types of leasing arrangements. Intuitively, the work evaluates the extent to which each dollar of subsidy is captured by the landowner through higher rental rates.

The provision of taxpayer-funded support to a group that tends to be relatively wealthy has been the subject of considerable debate. Given the concern of benefit pass-through from tenant operators to landowners, relevant literature on examining the impacts of government subsidies on farmland rental rates has recently begun to emerge (e.g., Lence and Mishra 2003; Patton, et al. 2008; Kirwan 2009; Goodwin, Mishra, and Ortalo-Magne 2010; Qiu, Goodwin, and Gervais 2011). The existing empirical studies often have neglected two important issues related to rent determinants: uncertainty and transaction costs. Factors such as production lags, unstable prices, and bad weather validate that production in agriculture is subject to risks. Risk averse tenant operators need to be compensated for bearing revenue uncertainty. Chavas' (1993) uncertainty version of Ricardian rent theory indicates that equilibrium rent occurs at the maximum of the expected profit minus the risk premium. Under this conceptual framework, government programs can have two types of effects on rents: an income effect through raising expected net returns, and an insurance effect through reducing the uncertainty of net returns and the degree of risk aversion, if it applies.

Therefore, the expectation is that different programs will have different impacts on rental

rates, depending on how these two effects interact. For decoupled direct payments, the income effect can be expected to dominate, as they are designed to be independent of current production and thus have little to no effect on farming risks. For coupled program payments such as loan deficiency payments and counter-cyclical payments, both income and insurance effects can be expected. These program benefits will raise rental rates both through increasing the expected net return and reducing uncertainty. However, the insurance effect will differ by program as they have different capacities to reduce uncertainty.

In addition to uncertainty, this analysis extends previous work by introducing transaction costs into the model. While the notion of Ricardian rent, defined as the residual return to farmland after all other factors have been paid, does not incur transaction costs explicitly, econometricians are not able to observe this residual-based rent in reality. What is observed are usually self-reported rents provided by tenants and/or landowners. Under this condition, rent is a contractual agreement of income and risk sharing between two parties. To achieve an agreement or a contract always involves transaction costs, such as negotiation and enforcement costs. Any adjustments to the existing agreement to reflect changes in factors that determine the rent will incur transaction costs. These include information collection costs, renegotiation costs, switching partner costs, land-use conversion costs, and so on. With these costs, landlords will only seek to raise rents if the potential benefits netting transaction costs are positive.

Transaction cost is also one of the reasons that different programs can have different effects on rents, as the costs to negotiate a new contract are different. For example, direct payments are lump sum income transfers tied to farmland and are usually known prior to the lease contracts. Thus, the transaction costs for raising rents to reflect an increase in decoupled payments can be expected to be smaller than those required to capture coupled program payments. Transaction costs for landowners currently engaging in farming or having farm-related experiences can be expected to be smaller than those without any farming experience. Transaction costs can also explain why we do not observe a one-to-one, complete subsidy incidence in farmland rents.

This study accordingly adopts a modified Ricardian rent framework and evaluates to what

extent farm programs affect rental rates, taking risks and transaction costs into consideration. Utilizing farm- and county-level data from multiple sources, four commodity programs are evaluated. They are direct payments, counter-cyclical payments, loan deficiency payments (LDPs), and disaster assistance payments. In the empirical procedure, the selection issue raised from the first essay is corrected by utilizing the method developed by Dubin and McFadden (1984) and Bourguignon, Fournier, and Gurgand (2007).

The two essays together answer the question of how and to what extent landowners capture program benefits intended for tenant operators. By how, the two essays indicate that landowners obtain extra benefits through switching leasing arrangements and raising rental rates. Ignoring the former behavior and focusing on the latter gives only incomplete stories of the benefit distribution, and the results may also suffer from serious bias resulting from the selection issues. The results indicate that landlords capture around 41% of the aggregate subsidies under cash leases and 78% under share contracts. Farm program payments are found to have different effects on rental rates depending on the types of programs and leasing arrangements.

1.2 Impact of State-Dependent Policy Intervention on Spatial Price Transmission

Besides the domestic policy intervention, government restrictions also often exist in international trade, especially when dealing with agri-food markets and economies in transition. As to the measure of transaction costs, direct quantification of policy interventions is difficult. In reality, trade intervention often reflects a state-dependent reaction rather than constant behavior. For instance, if the objective of a policy active exporting country is to stabilize the domestic price, export controls might be triggered when the world price is "too high", and subsidies would be applied when the world price is "too low". This state-dependent feature indicates a nonlinear relationship between domestic and world market prices. Although the extension of the concept of cointegrating relationship to a nonlinear framework is not new (see Park and

Phillips 1999, 2001, Chang and Park 2003, Saikkonen and Choi 2004, Gonzalo and Pitarakis 2006, etc.), the procedure to test and estimate nonlinearity in cointegrating vectors is.

The objective of Chapter 4 is to provide an investigation of the effects of state-dependent policy intervention on spatial price transmission and its empirical application in the Ukrainian wheat market.

Ukraine is the second largest European country after Russia. It became independent when the Soviet Union dissolved in 1991. The economy experienced a large increase in GDP growth after an eight-year recession that immediately followed the dissolution. Ukraine is a globally important grain supplier largely due to its abundant endowment of arable land. Though grain production suffered from dramatic declines in the first decade following independence, output has considerably increased since then. In marketing year 2009/10, Ukraine was among the world's top three leading grain exporters (after Brazil and Russia).

On the other hand, although Ukraine is a large grain exporter, it is still plagued by food security issues. The transition was difficult and plunged the majority of the Ukrainian people into poverty. A large part of the population still cannot afford sufficient food, and some have to rely on a subsistence diet of bread and tea. Given the political sensitivity of food prices, combined with Ukraine's history as a planned economy, the Ukrainian government always reacts quickly to the global rise in grain prices. Both local and central governments try to control crop and food prices (Brümmer, von Cramon-Taubadel, and Zorya 2009). When world food prices soar, the government response is often populist in nature. The government often accuses traders of driving up grain prices. As a result, they introduce export controls to try to reduce the domestic food prices.

World food prices increased dramatically in 2007/08, creating what some have called a global food crisis. Price increases ranged from 37.5% (for sugar) to 224% (for rice) between January 2007 and June 2008. Wheat rose 118% and corn rose 77% in the same period. After the dramatic increases, the prices dropped suddenly in the third quarter of 2008. Food prices then started to rise sharply again in the middle of 2010, and the price index of food surpassed

the peak levels of 2007/08 (World Bank 2011). The Food and Agriculture Organization (FAO) Food Price Index increased by more than 30% between June 2010 and December 2011, while the price index for cereals jumped by more than 40% during the same period (FAO 2012).

The reasons for the dramatic increases in food prices are still under debate. On the demand side, income growth in many developing countries, especially those with large populations like China and India, has led to increased demand for livestock and dairy products. Since it takes significantly more than one a kilogram of feed to produce one kilogram of meat or a liter of milk, demand for grains and oilseeds has grown significantly. Demand for biofuel production in developed countries in recent years has also contributed to drive up the food prices. On the supply side, rising energy prices have influenced world food prices by increasing the costs of production (e.g., fertilizer) and transportation. Together with this, production has been lower than expected due to unfavorable weather in recent years in key exporting countries (e.g., severe droughts in Australia, United States, Russia, and Ukraine).

The FAO has estimated that the 2007/08 price spike increased the number of undernourished people from about 850 million in 2007 to about 1,023 million in 2009 (World Bank 2011). As a result, rising food prices have led to political unrest and violence. In many developing countries, rising food prices have caused protests and riots. Numerous government interventions have been triggered to protect vulnerable populations from the negative consequences of higher food prices. There is considerable heterogeneity across countries in terms of how governments response to food price increases. Three main interventions have been observed: 1) Increasing food availability to households in need through direct transfers of cash or goods; 2) Stimulating the domestic grain/food production and marketing responses, and increasing the food supply in the long term, by improving agricultural production systems including post-harvest management, marketing infrastructure, and access to finance and risk management instruments; and 3) Controlling domestic food prices through policy interventions, including the reduction of import tariffs and taxes, grain or bread subsidies, direct domestic price controls, state procurement and distribution, export taxes, and export restrictions/bans (World Bank 2011).

At both the 2007/08 and 2010/11 price peaks, the primary response of the Ukrainian government to rising food prices was to implement grain export controls, primarily by issuing export quotas. The argument made to support these market interventions was that they were needed to guarantee food security and protect domestic consumers from rising international food prices. At the same time, although direct government intervention in the grain markets is commonplace in Ukraine when the market price is "too high", it does not subsidize agricultural exports even when the world market price is low.

The conceptual framework builds upon the earlier efforts of Mundlak and Larson (1992). We expand their work by introducing the state-dependent nature of the policy interventions into the model and thus allowing the price linkage to exhibit regime-switching nonlinear behavior. A two-regime policy response is proposed: a free market with no active policy intervention and trade intervention based on world market prices, to model the relationship between domestic and world wheat prices.

This study uses weekly observations of the world market and Ukrainian domestic wheat prices from March 23, 2001 to September 9, 2011. The empirical analysis is based on the smooth transition cointegrating (STC) framework of Saikkonen and Choi (2004) and Choi and Saikkonen (2010). It follows the general procedure used to investigate long-run equilibrium and short-run error correction suggested by Engle and Granger (1987).

The STC regression results indicate that there is regime-switching behavior for the long run relationship between Ukrainian domestic and world market prices, based on the world price. The results also indicate that the world price does not respond to disequilibrium between the two markets. Therefore, price shocks in Ukrainian domestic markets do not push the world market price to make adjustments accordingly.

Chapter 2

An Empirical Investigation of the Linkages between Government Payments and Farmland Leasing Arrangements

2.1 Introduction

About 45% of United States (US) farmland was operated by a tenant in 1999 (USDA/NASS 2001). Historically, contractual arrangements between landlords and tenants mostly included either cash payments or sharecropping. More recently, a third form of leasing arrangement involving both forms of payments - an arrangement that we designate as a hybrid contract - has gained popularity. USDA/NASS (2001) defines a hybrid contract (also called a cash/share contract) as one under which the tenant pays part of the rent in cash and part as a share of crops or livestock products.[1]

[1] In what follows, we use a relatively narrower definition of the hybrid contract as one that consists of a predetermined share percentage plus a fixed cash payment.

The use of hybrid contracts is increasing in the US farmland leasing market. In 1999, about 11% of all US leased farmland was under hybrid contracts, compared to only 3% in 1988. The incidence of use of hybrid contracts was highest in the Corn Belt and the Northern Plains (USDA/NASS 2001). These two regions are mainly comprised of crop farms, which are also the primary beneficiaries of commodity and conservation program payments. In 1999, 26% of leased farmland in Indiana was rented under hybrid contracts, as compared to less than 2% in 1988. Similar situations can be observed in other important agricultural states, including Illinois, Ohio, Kansas, Nebraska, Missouri, and Iowa.

The literature on farmland contract choice is considerable. Marshall (1890) laid out the early foundations of the analysis of sharecropping and illustrated the source of inefficiency associated with sharecropping (in relation to a cash or wage contract). Sharecropping discourages the tenant's own input use because he/she receives only part of his/her marginal product. A number of studies challenged Marshall's conclusion. Cheung (1969) argued that sharecropping could be as efficient as other types of contracts if monitoring is costless. Stiglitz (1974), and Newberry and Stiglitz (1979) introduced land tenure choices into a principal-agent framework. The standard agency model suggests that contracts are designed to achieve a balance between efficient risk-sharing and appropriate incentives to discourage moral hazard.

Allen and Lueck (1992, 2002) argued that in developed countries, where insurance markets are well developed, risk-sharing should not be the primary determinant of contract choices. They argued that the benefit of a sharecropping contract is that it curbs the tenant's incentive to overuse the inputs (e.g., soil moisture and nutrients) supplied by the landlord. However, sharecropping requires the output to be divided between the landlord and the tenant and thus generates additional transactions and monitoring costs for the landlord.

More recently, Huffman and Just (2004) introduced a principal-agent model which allows for heterogeneity in the characteristics of principals and agents and relaxes the risk-neutrality assumption for landlords. They argued that the parameters of sharecropping vary across tenants and landlords because of tenants' heterogeneity (e.g., the agent-specific effort productivity).

Huffman and Fukunaga (2008) and Fukunaga and Huffman (2009) provided recent empirical evidence on the determinants of contract arrangements using a model in which agents choose between a share and cash-rent contract. They found that both risk sharing and transaction cost incentives are important determinants of the contract type. They also emphasized the role of the landlords' attributes in determining the optimal landlord-tenant contract choice.

The literature has neglected two main issues related to farm leasing arrangements. First, previous studies have largely ignored the existence of hybrid contracts, instead focusing on a binary decision rule that involves cash rentals versus share contracts (e.g., Allen and Lueck 2002; Fukunaga and Huffman 2009). As argued above, hybrid contracts capture a growing share of leasing arrangements. Second, most studies ignored the impacts of government support on contract choices.[2] Government support programs are especially important in US agriculture. From 2000 to 2009, more than 40% of US farms revived program payments annually. The average annual commodity and conservation program payments under the 2002 Farm Bill were $15.0 billion and the corresponding support payments are projected to be $10.9 billion per fiscal year under the 2008 Farm Bill (Monke and Johnson 2010). Previous studies have demonstrated that the optimal contract choice depends on farming risk, the tenant's and the landlord's risk preferences, and the expected returns from rented land. Income and price support programs will affect the landlord-tenant contract choice because they potentially have impacts on expected returns and income variability as well as on the individuals' degree of risk aversion.

The contribution of this study is twofold. First, we examine the effects of government programs on farmland rental contract choices. In an empirical model, we break down aggregate government support into five different program types and investigate to what extent each impacts the probabilities of selecting a given contract type. Second, we introduce hybrid contracts as a third alternative in the contract set available to landlords and tenants in order

[2] Bierlen et al. (2000) offer a notable exception, though it is not closely connected to our analysis. The authors used a 1997 survey of Arkansas farm operators and investigated the impacts of the 1996 FAIR Act on leasing arrangements. They investigated whether operators terminated or added farmland leases due to the FAIR Act. Their results indicate that the probability of adding leases due to the FAIR Act increased as operators' experience declined, financial position strengthened, and managerial independence increased.

to investigate the determinants of an increasingly popular form of rental arrangement in US agriculture. Individual contract-level data collected in the 1999 Agricultural and Economics Landlord Owner Survey (AELOS) and the 1999 Agricultural Resource Management Survey (ARMS) are combined to carry out the analysis.

The remainder of the paper is structured as follows. The next section introduces a principal-agent model to explain the landlord-tenant leasing arrangements. Section three discusses issues pertaining to data, model specification, and empirical strategies. Section four presents the results of the estimation procedure. Concluding remarks are offered in the last section.

2.2 Conceptual Framework

The model below builds upon earlier efforts of Huffman and Just (2004), and Huffman and Fukunaga (2008). We expand their work by introducing agricultural program payments into the model. We begin by addressing the role of decoupled payments (also known as direct payments) in leasing arrangements. Decoupled payments were introduced in the 1996 Farm Bill and were renewed in the 2002 and 2008 Farm Bills. They are annual lump-sum income transfers that are designed to be independent of current production and market prices. Decoupled payments are based on base acreage and historical yields, and the producer is not obligated to be currently growing any specific crops on the land. He/she may plant any crop (with the exception of fruits and vegetables) without losing benefits. Under the current legislation, decoupled payments are the largest payout among commodity programs. In the Congressional Budget Office (CBO) March 2010 baseline projection for FY2011-FY2020, decoupled payments total $49 billion (or $4.9 billion annually), which account for 77% of the total commodity payments under Title I (Monke and Johnson 2010). There are also restrictions on the distribution of decoupled payments. Legislation requires that the payments be shared among tenant farm operators and landlords subject to the contract on a fair and equitable basis. Under a cash rental arrangement, 100% of the decoupled payments are allocated to the farm operator. Under a share contract, the government distributes payments to both the landlord and the tenant operator according

to the share terms of lease.

For simplicity, we assume that each landlord contracts with only one tenant. The principal is the landlord and the agent is the tenant operator.[3] Following Huffman and Just (2004), we allow heterogeneity in risk preferences of agents and principals. We also allow heterogeneity in the productivity of effort, cost of effort, and reservation utility.

2.2.1 A Principal-Agent Model with Decoupled Payments

The output of tenant operator i on one unit of leased land (or net revenue with appropriate normalization) is defined as:

$$y_i(e_i) = a_i e_i + \varepsilon_i,$$

where e_i is tenant i's effort/labor input and a_i is the tenant-specific productivity of labor. Differences in productivity may be related to human capital in the form of farming experience (Huffman and Just 2004). Output is also function of a stochastic term ε_i which is assumed to have zero mean and variance σ_i^2.

Following Huffman and Fukunaga (2008), we assume the landlord offers a linear incentive contract to the tenant operator. The tenant operator's compensation is:

$$I_i(e_i) = \alpha_i + \beta_i(a_i e_i + \varepsilon_i + g_d) - 0.5 k_i e_i^2,$$

where k_i is the tenant-specific effort cost parameter. A high (low) value of k_i indicates a steep (flat) marginal cost curve. The variable g_d represents decoupled government payments. The parameter α_i is the tenant-specific cash payment of the contract. A positive α_i represents the cash wage paid by the landlord to the tenant; a negative value for α_i means that cash rent payments are made to the landlord. The parameter $\beta_i (0 \leq \beta_i \leq 1)$ is an incentive rate representing a share of output. Hence, when $\beta_i = 1$ and $\alpha_i < 0$, the leasing arrangement is a cash contract as opposed to $0 < \beta_i < 1$ and $\alpha_i = 0$ which indicates a share contract. More

[3] Tenant operators include pure-tenant operators, who rent all of the farmland from others; and part-owner operators, who own part of the farmland and rent part of the land from others.

importantly in the context of this paper, $0 < \beta_i < 1$ and $\alpha_i < 0$ indicate that the leasing arrangement is a hybrid type contract.

Assume that the tenant has well-defined preferences over income summarized by the utility function $U_i(I_i)$. Expected income of the ith tenant operator is $E(I_i) = \alpha_i + \beta_i(a_i e_i + g_d) - 0.5 k_i e_i^2$. The variance of the tenant's income is $V(I_i) = \beta_i^2 \sigma_i^2$. Let $RP_i \equiv 0.5 r_i V(I_i)$ denote the risk premium where $r_i \equiv -U_i''/U_i'$ is the tenant's Arrow-Pratt coefficient of absolute risk aversion. Under the expected utility model, $EU_i(I_i) = U[E(I_i) - RP_i]$. Given that $U_i(I_i)$ is an increasing function of income, maximizing $EU_i(I_i)$ is equivalent to maximizing the expression $[E(I_i) - RP_i]$ (Chavas, 2004). Therefore, the tenant operator's optimal effort is determined by maximizing his/her certainty equivalent $CE_i = E(I_i) - RP_i$:

$$\max_{e_i} CE_i = \max_{e_i} \left[E(I_i) - 0.5 r_i V(I_i)\right] = \max_{e_i} \left[\alpha_i + \beta(a_i e_i + g_d) - 0.5 k_i e_i^2 - 0.5 r_i \beta_i^2 \sigma_i^2\right]. \quad (2.1)$$

The optimization problem defined in (2.1) solves the optimal effort $e_i^* = \beta_i a_i / k_i$.

Similarly, the lth landlord's expected return from ownership of the rented land equals $E(\pi_i) = E[(1 - \beta_i)(y_i + g_d) - \alpha_i]$ and its variance is $V(\pi_l) = (1 - \beta_i)^2 \sigma_i^2$. As in the case of the tenant operator, we write the landlord' optimization problem in terms of the certainty equivalent return:

$$\max_{\beta_i} CE_l = \max_{\beta_i} \left[E(\pi_l) - 0.5 r_l V(\pi_l)\right] = \max_{\beta_i} \left[(1 - \beta_i)(a_i e_i^* + g_d) - \alpha_i - 0.5 r_l (1 - \beta_i)^2 \sigma_i^2\right], \quad (2.2)$$

subject to the participation and incentive compatibility constraints: $\alpha_i + \beta_i(a_i e_i^* + g_d) - 0.5 k_i e_i^{*2} - 0.5 r_i \beta_i^2 \sigma_i^2 \geq \mu_i$ and $e_i^* = \arg\max_{e_i} [\alpha_i + \beta_i(a_i e_i + g_d) - 0.5 k_i e_i^2 - 0.5 r_i \beta_i^2 \sigma_i^2]$, where r_i is the Arrow-Pratt measure of the landlord's absolute risk aversion, and μ_i is tenant i's reservation utility.

The landlord's optimal choice of α_i will be determined by the binding participation constraint. Substituting e_i^* and α_i into (2.2) and optimizing over β_i yields the optimal incentive

rate offered to the ith tenant operator:

$$\beta_i^* = \frac{c_i + r_l\sigma_i^2}{c_i + (r_i + r_l)\sigma_i^2} = 1 - \frac{r_i\sigma_i^2}{c_i + (r_i + r_l)\sigma_i^2}, \tag{2.3}$$

where $c_i \equiv a_i^2/k_i$ is an index of tenant-specific effort productivity. Substitute the optimal share rate into the participation condition to obtain the optimal cash component of the contract:

$$\alpha_i^* = \mu_i - 0.5\beta_i^{*2}(c_i - r_i\sigma_i^2) - \beta_i^* g_d. \tag{2.4}$$

The optimal share rate β_i^* in (2.3) emphasizes the role of the landlord's and the tenant operator's degree of risk aversion. If a tenant operator is risk neutral ($r_i = 0$), the optimal share rate equals one and a cash contract is the optimal outcome. Similarly, the optimal share increases towards one as the landlord's coefficient of risk aversion goes to infinity (i.e., $r_l \to \infty$). Risk, represented by the variance of income, is negatively correlated with the optimal share rate. The higher is the variance of income, the smaller is the optimal share rate. Therefore, an increase in income volatility can have a negative impact on the choice of a cash contract, *ceteris paribus*. However, an increase in risk has an indeterminate impact on the optimal cash payments.

A quick look at (2.3) suggests that decoupled payments do not have a direct impact on the share rate. However, these payments may affect the contract choice indirectly, through their impact on the degree of risk aversion. If an individual has constant absolute risk aversion (CARA) preferences, decoupled payments will not have an impact on the solution in (2.3). However, if his/her risk preferences entail decreasing absolute risk aversion (DARA), decoupled payments will reduce the degree of risk aversion through their impact on wealth.

Decoupled payments however do have a direct effect on the optimal cash component a_i^* of the contract. It reflects a pass-through of program benefits from the tenant operator to the landlord. From (2.4), the optimal cash component with no decoupled payments (i.e., $g_d = 0$) is $\mu_i - 0.5(\beta_i^*)^2(c_i - r_i\sigma_i^2)$, which is greater than that with positive decoupled payments. The

difference is $\beta_i^* g_d$, which equals the share of decoupled payments going to the tenant operator under the current legislative environment for US farm programs. Hence, the landlord captures the benefits that go to the tenant by charging an extra cash amount of size $\beta_i^* g_d$, and restores the equilibrium that would have been attained under no governmental restriction, *ceteris paribus*.[4] Under the optimal leasing arrangement, the landlord is able to capture all of the benefits distributed to the tenant operator, given the conditions that payments are decoupled and the wealth effect is negligible. A governmental restriction on payment distribution does not influence the actual benefit distribution between landlords and tenant operators in the end. It merely results in offsetting contractual rearrangements.[5] This is consistent with Lence and Mishra (2003), and Goodwin, Mishra, and Ortalo-Magné (2010) who have found evidence that landlord capture 62%-86% of the benefits of decoupled payments by raising cash rents.

When referring to the equilibrium leasing arrangement, it is interesting to look at the comparison between the equilibrium contract choice with decoupled payments and the choice without the payments under three general circumstances. Assume both parties have CARA preferences and define β_{i0}^* and α_{i0}^* as the optimal share rate and cash component of the contract under no decoupled payments. First, if a share contract is optimal (i.e., $0 < \beta_{i0}^* < 1$, and $\alpha_{i0}^* = 0$), the introduction of decoupled payments will change the equilibrium to a hybrid contract, increasing the cash payment to the landlord and keeping unchanged the share rate ($0 < \beta_i^* < 1$ is constant, and $\alpha_i^* = 0 - \beta_i^* g_d < 0$). In a second case, if the equilibrium contract with no decoupled payments is a cash contract ($\beta_{i0}^* = 1$, and $a_{i0}^* < 0$), the introduction of decoupled payments would leave the share rate constant and the cash payment to the landlord would increase ($\beta_i^* = 1$, and $\alpha_i^* = \alpha_{i0}^* - \beta_i^* g_d < 0$). Therefore, the equilibrium contract choice will still be a cash contract; however the cash rent will increase. Finally, if the optimal leasing arrangement is a hybrid contract without decoupled payments, decoupled payments will not change the equilibrium contract type. The cash payments to the landlord will simply increase. In summary, under the CARA assumption, the introduction of decoupled payments increases

[4] We assume that transaction costs for renegotiating contracts are zero.

[5] Cheung (1969, chapter 5) reaches a similar conclusion.

the use of hybrid contracts and decreases the use of share contracts. Decoupled payments have no effect on the choice of cash contracts.

On the other hand, if both parties have DARA preferences, decoupled payments will lower the degree of risk aversion. From the solutions in (2.3) and (2.4), a cash contract will emerge in the case where the landlord's degree of risk aversion goes to infinity. Under more general conditions, the direct and indirect effects of decoupled payments on the cash component of the contract (α_i) can go in different directions. The net effect depends on the risk preferences of both contracting parties. Table 2.1 summarizes the effects of decoupled payments on optimal leasing arrangements by risk preferences. The ambiguous causal effect of decoupled payments on leasing arrangement can only be resolved empirically. However, we examine the effects of coupled payments on contract choices before considering the empirical investigation.

2.2.2 A Principal-Agent Model with Coupled Payments

Coupled payments are based on current production and/or market price. Many forms of coupled programs exist in the US. These include price and/or yield support mechanisms and disaster relief programs. For simplicity, we investigate a per-unit production subsidy in this section. As before, the landlord and the tenant share the program payments in the same proportion as they share output. The per-unit production subsidy rate is $\phi > 0$ and coupled support equals $g_c = \phi y_i$. The tenant operator and the landlord's payments are $\beta_i \phi y_i$ and $(1-\beta_i)\phi y_i$, respectively. Maximizing the objective function defined in (2.1) accounting for coupled support yields the optimal effort level $e_i^* = (1+\phi)\beta_i a_i / k_i$. The optimal share rate and the cash payment that maximize the landlord's objective function are:

$$\beta_i^* = \frac{(1+\phi)^2 c_i + r_l(1+\phi)^2\sigma_i^2}{(1+\phi)^2 c_i + (r_i + r_l)(1+\phi)^2\sigma_i^2} = 1 - \frac{r_i\sigma_i^2}{c_i + (r_i + r_l)\sigma_i^2},$$

$$\alpha_i^* = \mu_i - 0.5(1+\phi)^2\beta_i^{*2}(c_i - r_i\sigma_i^2)$$

Table 2.1: Effects of Decoupled Program Payments on Contract Choices

Program payments	**Risk preference**	**Effect**	**Optimal share rate β_i^***	**Optimal cash-part payments α_i^***	**Effects on Contract Choice**		
No program payments			$1 - \frac{r_i\sigma_i^2}{c_i+(r_i+r_l)\sigma_i^2}$	$\mu_i - \frac{1}{2}\beta_i^{*2}(c_i - r_i\sigma_i^2)$			
			$1 - \frac{r_i\sigma_i^2}{c_i+(r_i+r_l)\sigma_i^2}$	$\mu_i - \frac{1}{2}\beta_i^{*2}(c_i - r_i\sigma_i^2) - \beta_i^* g_d$	Cash $\beta_i^* = 1$ $\alpha_i^* < 0$	Hybrid $0 < \beta_i^* < 1$ $\alpha_i^* < 0$	Share $0 < \beta_i^* < 1$ $\alpha_i^* = 0$
Decoupled payments	Both CARA	direct	NO	–	NO	+	–
		indirect	NO	NO	NO	NO	NO
	Both DARA	direct	NO	–	NO	+	–
		indirect	+/–	+/–	+/–	+/–	+/–
	TO CARA LL DARA	direct	NO	–	NO	+	–
		indirect	–	+	–	+/–	+/–
	TO DARA LL DARA	direct	NO	–	NO	+	–
		indirect	+	–	+/–	+	–

Note: TO refers to the tenant operator and LL refers to the landlord.

The per-unit production subsidy has a direct impact on income variability and the marginal productivity of effort. These two effects however cancel each other when determining the optimal share rate, which remains constant if no wealth effects are present. Turning our attention to potential wealth effects, coupled payments have an ambiguous impact on the share rate. The effect depends on both the landlord and tenant's risk aversion. As discussed in the previous section, the optimal share increases towards one as the landlord's risk aversion increases to infinity. However, coupled support may decrease the landlord's risk aversion coefficient, which could entail a switch from a cash rental type to hybrid or sharecropping.

Production subsidy payments have a direct impact on optimal cash payments. If $\beta_{i0}^* = \beta_i^*$, the cash payments decrease from $\mu_i - 0.5(\beta_i^*)^2(c_i - r_i\sigma_i^2)$ to $\mu_i - 0.5(1+\phi)^2(\beta_i^*)^2[c_i - r_i\sigma_i^2]$, which suggests that cash payments increase as per-unit production payments increase. Note that a decrease in cash payments is possible if the wealth effect decreases the landlord's coefficient of absolute risk aversion.

The impacts of coupled payments on contract choice can differ substantially according to the support types. Coupled payments influence the optimal share rate and cash payments through one or more of the following factors: increases in expected returns, changes in income variability, changes in the (value of) marginal productivity of effort, and impacts on the contracting parties' degree of risk aversion. Programs that decrease income variability and/or decrease a tenant's effort decrease the optimal share rate, thus have a positive effect on the choice of a cash contract. While the types of programs have not been explicitly modeled, we use the insights of this section to state the hypotheses related to the causal relationship between US coupled farm payments and leasing arrangements.

2.3 Modeling Framework

2.3.1 Data and Empirical Model

The data used in this study come from five sources: the 1999 Agricultural Economics and Land Ownership Survey (AELOS), the 1999 Agricultural Resource Management Survey (ARMS), the Regional Economic Information Systems (REIS) dataset for the 1990-1999 period, the county level farm program payment data from the Farm Service Agency (FSA) of the USDA over the 1996-1999 period, and the county level farmland data from the 1997 Census of Agriculture. In contrast to other studies that only use one source of data (e.g., Huffman and Fukunaga 2008, 2009), we combine the above datasets in an effort to increase the explanatory power of the empirical model. The AELOS is an integrated survey of farm finance and land ownership. It includes comprehensive information collected from both tenants and landlords. Each observation in this dataset represents a unique contractual relationship between a landlord and a tenant operator.

The ARMS is a national survey that provides observations of farm-level production practices, economic attributes, and operator households' characteristics. We use this dataset to obtain individual farm level program payments as well as additional farm and operator characteristics that may impact the leasing arrangements. The REIS contains economic data and annual estimates of personal income for the residents of the entire nation as well as states, metropolitan areas, and counties. We obtain county level gross cash farm income (cash receipts from marketing and government payments) data from REIS and FSA, and county level farmland acres from the 1997 Agricultural Census.

We refine the combined dataset following these steps. First, we focus on the landlords who have only one renter. This accounts for about 90% of the entire dataset. Second, some outliers (less than 2% of the available sample) are excluded from the analysis because they represent atypical situations (for example, landlords reporting land rent exceeding $2,000 per acre). Third, because crop farm producers are the main recipients of farm program payments,

farms that reported livestock product sales that exceeded of 50% of their farm sales are excluded. Farms for which more than 50% of total sales were nursery products, fruits, or vegetables are also dropped from the sample. After this selection procedure, a total of 15,457 observations remain for the analysis. In the AELOS dataset, each landlord/operator observation has a different weight to represent their weight in the underlying population, as if a complete census had been carried out.[6] We present the weighted results in this article.

We address the choice of leasing arrangements using a multinomial Logit (MNL) model, appealing to the concept of random utility derived by individual n from a set of $j = 1, 2, \ldots, J$ different alternatives (Train 2003):

$$U_{nj} = V_{nj} + \varepsilon_{nj} \qquad \forall\, j, \tag{2.5}$$

where V_{nj} represents information that is known by researchers and ε_{nj} is the unobservable component of utility.

Let $\boldsymbol{x}$ be a vector of individual-specific characteristics and $\boldsymbol{\beta}$ a corresponding vector estimated coefficients. If ε_{nj} is unknown but follows a logistic distribution, the choice probability is (Long and Freese 2006):

$$P_{nj} = \frac{\exp(\alpha_{j|b} + \boldsymbol{x}\boldsymbol{\beta}_{j|b})}{\sum\limits_{j=1}^{J} \exp(\alpha_{j|b} + \boldsymbol{x}\boldsymbol{\beta}_{j|b})} \tag{2.6}$$

where b refers to the base alternative which is defined here as a "share contract". We normalize $\alpha_{b|b}$ and $\beta_{b|b}$ so that the log of the odds of an alternative compared with itself is always zero.

[6] For more information about the calculation of these weights, see the General Explanation for Agricultural Economics and Land Ownership Survey (1999). Online. Available at: http://www.agcensus.usda.gov/Publications/1997/Agricultural_Economics_and_Land_Ownership/appendix-a.pdf [Accessed June 2011.]

The log likelihood function for the MNL model is:

$$\ln L(\beta) = \sum_{n=1}^{N} \sum_{j=1}^{J} d_{nj} \ln P_{nj}. \tag{2.7}$$

The variable d_{nj} equals 1 if individual n choose alternative j, and equals zero otherwise.

2.3.2 Model Specification

In the following empirical investigation, we use a generalized MNL model with an alternative-specified constant (Train 2003). Each observation in the dataset constitutes a landlord and tenant operator pair (landlord-tenant hereafter) who is involved in a specific farmland contract. The landlord-tenant chooses a contract among three alternatives: a cash contract, a share contract, or a hybrid contract. The decision is made conditional on a set of independent variables which are specific to the landlord-tenant pair n and are included in the vector x_n. This vector can be decomposed into four different parts that include farm program payments, farming risk and risk preferences, tenant operator's effort productivity, and other factors, each of dimension I_G, I_R, I_P, and I_M, respectively. Alternatives are assumed to be mutually exclusive. The utility function can be written as:

$$V_{n,i|b} = \alpha_{i|b} + \sum_{g=1}^{I_G} \beta^R_{g|b} GovP_{n,g} + \sum_{r=1}^{I_R} \beta^R_{r|b} Risk_{n,r} + \sum_{p=1}^{I_P} \beta^P_{p|b} EffP_{n,p} + \sum_{m=1}^{I_M} \beta^M_{m|b} Other_{n,m}.$$

The subscript i refers to either the "cash" or the "hybrid" contract. The parameter $\alpha_{i|b}$ is the ith alternative specific constant which can be interpreted as the average effects of unobserved factors. The variable $GovP_{n,g}$ are payments (per acre) received from government program g. The variable $Risk_{n,r}$ are proxies to capture farming risk and both parties' risk preferences. One potential proxy candidate for risk is the coefficient of variation (CV) for gross income at the individual farm level. However, this may raise endogeneity concern if the individual CV is correlated with unobserved farm characteristics, such as land attributes. Therefore, we use

a CV of gross cash farm income per acre (which includes both cash receipts from market and government payments) in the county where the individual farm is located over the previous ten-year period. A tenant operator's risk preference is represented by farm net worth. On the landlord's side, we do not have data on net worth/wealth. We therefore use an indicator of whether the landlord purchased insurance for the farm business as a proxy to the landlord's risk preference.

The variable $EffP_{n,p}$ represent the tenant operator's productivity. We employ farming experience and the squared value of farming experience to proxy the tenant operator's effort productivity. Finally, the variables in $Other_{n,m}$ include the landlord's residence, the landlord's real estate taxes relative to his/her rent income, farm type, and the tenant operator's tenure status (whether the operator is a pure tenant or a part-owner tenant).[7]

Government program payments include six components. They are Production Flexibility Contract (PFC) payments, Market Loss Assistance (MLA) payments, Loan Deficiency Program (LDP) Payments (include marketing loan gains), Agricultural Disaster Payments (which include all market loss or disaster assistance payments, but exclude Federal Crop Insurance indemnity and other indemnity payments), Conservation Reserve Program (CRP) benefits, and a final category including all other minor program payments. As discussed in the conceptual framework section, decision makers use expectations of future payments to determine the contract type. Disaster, MLA, and LDP payments are not predetermined. Rather, they are triggered by market and production conditions. Measurement issues arise if actual reported payments are used to represent expectations, as is noted in Goodwin, Mishra, and Ortalo-Magné (2003). To control potential errors-in-variables problems, we follow their approach and use a four-year county average of payments per acre to proxy expected program payments.[8]

Future PFC payments are decoupled and known in advance of when a contract is signed.

[7] Following the suggestion of an anonymous reviewer, we investigated whether the landlord's and the tenant operator's relative real estate taxes had an impact on leasing arrangements. We found that the operator's real estate taxes were not statistically significant and thus excluded this variable from the final estimation.

[8] Market Loss Assistance payments were introduced in 1998 and we use the 1998-1999 average annual payments. For other programs, we use 1996-1999 average annual payments.

Therefore, we use self-reported, realized farm-level payments in the empirical model. Conservation Reserve Program pays farmers annual rents to place land in reserve. In order to be eligible for the payments, land must be erodible and environmentally fragile. Such payments could quite possibly be correlated with the unobserved factors which affect the contract choice (e.g., land attributes). Therefore, although CRP payments are usually known before signing the contract, our empirical investigation uses a four-year county average as a proxy for individual farms to avoid the potential endogeneity problem.

Table 2.2 presents the definition of key variables and summary statistics. In our crop farm sample, 57% of farmland contracts were on a cash basis while 18% were share contracts. The remaining 25% were hybrid contracts, making the latter form of leasing arrangement more popular than pure share contracts for crop farms. From 1996 to 1999, farms received on average $14.15 PFC payments per acre annually at the county level. The corresponding MLA, LDP, and Disaster payments were $10.03, $9.80, and $2.38 per acre on average. Finally, the annual county average CRP payments were $2.32 per acre. All monetary values were adjusted by the consumer price index to represent 2004 dollars. Tenant operators had 26.3 years farming experience on average. About 55% of landlords lived in a rural area and the remaining 45% of landlords were defined as absentee landowners and lived in a non-rural area. Principle crop farms - defined as grains, oilseed, dry beans, or peas farms- account for 63% of the crop farm sample. Around 83% of tenant operators are part-owner tenants who own some of the operating land and the remaining 17% are pure-tenant operators who rent the entire farmland from others.

2.3.3 Expected Impacts of Key Factors on Contract Choices

The PFC payments are decoupled payments which are independent of current production and market price. The impacts of decoupled payments on leasing arrangements are summarized in Table 2.1. More specifically, when the wealth effects are small or negligible, the PFC payments will entice agents to move from a share contract to a hybrid contract and thus will redistribute the benefits between contracting parties. The MLA, LDP, and disaster assistance programs are

Table 2.2: Summary Statistics (N=15,457)

Choices		**Frequency**	**Percentage**
Cash contract		8,806	56.97
Hybrid contract		2,817	24.80
Share contract		3,834	18.23
Variable	**Definition**	**Mean**	**Std. Dev.**
1996-1999 County average program payments ($/acre)			
PFC	Production flexibility contract payments	14.15	8.73
MLA	Market loss assistance payments	10.03	6.21
LDP	Loan deficiency payments (including marketing loan gains)	9.80	6.76
Disaster	Disaster payments	2.38	2.96
CRP	Conservation reserve program payments	2.32	2.40
Other	Other payments	0.12	0.16
Risks and risk preferences			
CV	10-year county level coefficient of variation of cash receipts from market and government payments (per acre)	0.13	0.07
NetWorth	Net Worth of the farm	183032.70	70011.40
Insurance_l	1 if landlord's purchase insurance for the rented farm	0.34	0.47
Tenant operator's effort productivity			
FarmingExp	Tenant operator's farming experience	26.30	11.96
FarmingExp^2	Tenant operator's farming experience squared	821.50	732.74
Other factors			
Rural_l	1 if landlord lives in a rural area	0.55	0.50
Ft_main	1 if the farm type is grains, oilseed, dry beans, or peas.	0.63	0.48
RealTax_l	Landlord's real estate tax expenditure relative to total rent received (100%)	0.62	4.64
PartOwner	1 if the tenant is a part-owner and 0 if he or she is a pure-tenant	0.83	0.38

coupled and are associated with current production and/or market conditions.

When wealth effects are negligible, we can expect the following impacts of program payments on the contract choice. The coupled programs (MLA, LDP, and disaster) lower income variation and have a positive effect on the optimal share lease rate. Thus, they raise the probability of selecting a cash contract. However, if a wealth effect influences the degree of risk aversion of both parties, the effects of program payments can have opposite impacts. In general, government programs can shift incentives to use a particular type of contractual arrangement and can redistribute income and risk between the landlord and tenant. The CRP is a special type of program when considering the impacts of government payments on leasing arrangements. In most cases, payments are not related to the leased land. Tenant operators receive payments from their own land. The CRP pays land owners annual rents to set their land aside under a ten to fifteen year lease agreement. Land committed to CRP must be removed from production. Because the CRP payments usually do not involve rented land, they may not affect the landlord's incentives. However, they may have an impact on the contract choice by affecting the tenant's degree of risk aversion (through wealth effects). According to the optimal share rate derived in (2.3), risk is expected to have a negative impact on the optimal share rate. However, an increase in generates conflicting effects on the optimal cash payments and it makes it impossible to unambiguously sign the net impact of risk on cash payments.

One concern at the empirical stage is the possibility that a particular type of principal contracts with certain types of agents, a phenomenon dubbed endogenous matching by Ackerberg (2002). Ackerberg argues that if: 1) there exist incentives for particular parties to contract with a specific subset of the other parties (e.g., a risk-averse tenant being more likely to contract with a risk neutral landlord); and 2) some characteristics (e.g., landlord's true risk preference) of contracting parties are not observable, explaining the outcome may involve a possible bias if the endogeneity is not addressed.

To investigate this possibility, we carried out a two-stage regression procedure that involves in the first stage, regressing the tenant operator's risk preference (represented by net

worth of the farm) on the landlord's risk preference (proxied by purchase of insurance) and other exogenous factors may that may have an impact on matching (e.g., contracting parties' ages and education). We found no significant correlation between the contracting parties' risk preferences. In a second stage, we use the predicted value of the tenant operator's risk preference proxy and estimate the multinomial logit model. The results from the second stage are quite similar to the uncorrected MNL estimation results which does not control for endogenous matching. Intuitively, the similarity between the results is consistent with prior studies (e.g., Sherrick and Barry 2003; Allen and Lueck 2002) that emphasize how contracts emerge from long-run business relationships due to close ties between the landlord and the tenant. Therefore, it is not unreasonable to treat the matching of contracting parties as exogenous to the leasing arrangements in the US farmland market.

A tenant operator normally contracts with several different landlords (on average, one tenant operator contracted with four landlords in 1999). Some correlation among observations from the same tenant operator may exist. Therefore, clustered robust standard errors are used and based on the tenant operator's id number in this analysis. The logit model implicitly imposes the Independence of Irrelevant Alternatives (IIA) assumption which states that the probability of choosing among two alternatives is unaffected by the presence of additional alternatives. We test the IIA using the Chi-Square test statistic proposed by Hausman and McFadden (1984). We are not able to reject the null hypothesis that the IIA assumption is valid at a high level of significance. Tests for combining alternatives (Long and Freese 2006) are also computed to examine if hybrid contracts are distinguishable from share and cash contracts. The Wald tests reject the hypothesis that any two of the alternative contracts are indistinguishable at a 0.01 level.

2.4 Results

Table 2.3 reports the estimates of the coefficients in the three-alternative MNL model while Table 2.4 reports the marginal or discrete changes in predicted probabilities for each alternative

derived from the estimates in Table 2.3.

2.4.1 Government Program Payments

Recall that program payments are measured in 2004 dollars. Not surprisingly, the change in the predicted probability following a dollar increase is small. Therefore, we report the effects of a standard-deviation change in Table 2.4. We define a standard deviation increase (centered on the mean) as one unit change when we refer to the marginal/discrete effects. Table 2.4 shows evidence that the PFC payments have a positive impact on the selection of hybrid contracts and a negative effect on share contracts. When a PFC payment increases by one unit ($8.73), the probability of choosing a hybrid contract increases by 1.1% and the probability of choosing a share contract decreases by 4.5%. This is consistent with the theoretical explanation that landlords are more likely to capture the program benefits through a hybrid contract. The impact of decoupled payments on choosing a cash contract is positive. Direct payments have an impact on wealth and decrease risk aversion under DARA-type preferences, and thus increase the probability to choose a cash contract.

Both the disaster payments and the loan deficiency payments encourage the choice of a cash contract by reducing the income volatility. If a tenant operator receives an additional unit ($6.76) of loan deficiency payments, the probability of choosing a cash contract increases by 2.1%. The predicted probability of choosing a cash contract is 5.0% higher following a one unit ($2.96) increase in the tenant operator's disaster payments. Meanwhile, both the LDP and the disaster payments decrease the probabilities of choosing a share contract. In contrast, the MLA payments have negative impacts on the cash contract choice. The marginal effects of the MLA payments on both the cash and share contracts are the largest among all government programs. Getting an additional unit of MLA payments decreases the probability of choosing a cash contract by 6.8%. The extent of LDP and disaster payments was determined by the 1996 Farm Bill. However, the MLA was determined outside of the Farm Bill. In 1998, the prices of many crops declined significantly. Congress authorized $2.86 billion as emergency MLA

Table 2.3: Maximum Likelihood Estimation of MNL Models of Contract Choice

Explanatory Variable	**Choice**	**Coefficient**	**Standard Error**
1996-1999 County average program payments $/acre			
PFC	Cash	0.02***	4.02E-3
PFC	Hybrid	0.02***	4.22E-3
MLA	Cash	-0.07***	0.01
MLA	Hybrid	-0.04**	0.02
LDP	Cash	0.03**	0.01
LDP	Hybrid	0.03**	0.01
Disaster	Cash	0.08**	0.03
Disaster	Hybrid	0.02	0.04
CRP	Cash	-0.02	0.02
CRP	Hybrid	-0.03	0.03
Other	Cash	0.20***	0.07
Other	Hybrid	0.07	0.08
Risks and risk preferences			
CV	Cash	-1.97**	0.84
CV	Hybrid	-1.76*	0.91
NetWorth	Cash	2.12E-07	1.50E-07
NetWorth	Hybrid	1.42E-07	1.66E-07
Insurance_l	Cash	-0.93***	0.09
Insurance_l	Hybrid	-0.61***	0.09
Tenant operator's effort productivity			
FarmingExp	Cash	0.01	0.02
FarmingExp	Hybrid	0.04	0.02
FarmingExp^2	Cash	-1.85E-04	3.05E-04
FarmingExp^2	Hybrid	-3.30E-04**	9.90E-05
Other factors may affect the contract choice			
Rural_l	Cash	0.30***	0.09
Rural_l	Hybrid	0.02	0.09
Ft_main	Cash	-0.37**	0.18
Ft_main	Hybrid	0.33	0.21
RealTax_l	Cash	0.17***	0.06
RealTax_l	Hybrid	-0.01	0.06
PartOwner	Cash	0.32*	0.16
PartOwner	Hybrid	0.36**	0.18
Constant	Cash	1.31***	0.31
Constant	Hybrid	-0.38	0.41
Log pseudo-likelihood = -863750.98, Wald chi^2 = 294.99, Prob > chi^2 = 0.00			

Note: Asterisks (*,**,***) represent significance at the 10%, 5%, and 1% levels.

Table 2.4: Marginal and Discrete Changes on the Predicted Probabilities

	Unit change in the variable	**Change in predicated probability (100%)**		
		Cash	**Hybrid**	**Share**
Receipt of program payments in 1999				
PFC	$8.73	3.42	1.08	-4.50
		[0.64, 6.21]	[-0.97, 3.12]	[-6.61, -2.40]
MLA	$6.21	-6.77	1.15	5.62
		[-10.43, -3.10]	[-1.66, 3.96]	[3.12, 8.12]
LDP	$6.76	2.10	1.09	-3.19
		[-1.16, 5.35]	[-1.44, 3.62]	[-5.50, -0.87]
Disaster	$2.96	5.01	-1.97	-3.04
		[1.44, 8.59]	[-4.91, 0.97]	[-5.85, -0.23]
CRP	$2.40	-0.41	-0.41	0.82
		[-3.07, 2.25]	[-2.55, 1.74]	[-0.75, 2.38]
Other	$0.16	3.93	-1.33	-2.60
		[1.27, 6.58]	[-3.35, 0.70]	[-4.67, -0.53]
Risks and risk preferences				
CV	0.07	-1.77	-0.32	2.10
		[-4.41, 0.86]	[-2.36, 1.71]	[0.49, 3.71]
NetWorth	$70011.40	2.32	-0.21	-2.11
		[-1.91, 6.55]	[-3.50, 3.08]	[-5.07, 0.85]
Insurance_l	$0 \rightarrow 1$	-15.25	0.88	14.36
		[-18.71, -11.78]	[-1.41, 3.18]	[11.39, 17.34]
Tenant operator's effort productivity				
FarmingExp	11.96	-3.85	6.28	-2.44
		[-12.79, 5.10]	[-1.36, 13.93]	[-8.83, 3.96]
NetWorth	732.74	5.20	-9.45	4.25
		[-4.13, 14.54]	[-17.60, -1.30]	[-2.12, 10.62]
Other factors				
Rural_1	$0 \rightarrow 1$	7.02	-3.47	-3.55
		[3.98, 10.06]	[-5.67, -1.27]	[-6.03, -1.07]
Ft_main	$0 \rightarrow 1$	-12.59	9.61	2.98
		[-19.51, -5.67]	[4.37, 14.85]	[-1.88, 7.84]
RealTax_l	1 %	4.27	-2.30	-1.97
		[1.89, 6.66]	[-3.88, -0.73]	[-3.82, -0.12]
PartOwner	$0 \rightarrow 1$	3.37	2.08	-5.45
		[-3.46, 10.21]	[-2.84, 7.00]	[-10.60, -0.31]

Note: For continuous independent variables (program payments, CV, and farming experience), a unit change equals a standard deviation around the mean, holding other variables at their sample mean. Numbers between brackets provide 95% confidence intervals for changes in predicted probabilities (Long and Freese, 2006).

payments (trigged by low market price, but based on historic base acreages) to help farmers deal with income losses. Therefore, MLA actually targeted higher risk farms/crops. This would in turn make MLA correlated with higher risk (high CV) and thus have negative impacts on the optimal share rate (i.e., a decrease in the probability of choosing a cash contract). The impacts of CRP payments on the landlord-tenant contract choices are found to be insignificant. The other payment category reveals a positive impact on the choice of a cash contract and a negative impact on a hybrid or a share contract.

In conclusion, the results indicate that decoupled payments encourage the use of cash and hybrid contracts relative to share contracts. Benefits from most of the program payments (with the exception of MLA) have positive effects on the choice of a cash contract. Also, the impacts of payments on the probability of selecting a hybrid contract are positive (except for disaster payments). The payment effects on share contracts differ depending on the specifics of the program. Most programs have negative impacts on the probability to observe a share contract, with the exception of MLA payments. The impacts of program payments on contract choices show that risk-sharing and benefit distribution are important determinants of farmland leasing arrangements.

2.4.2 Risk and Risk Preferences

The role of risk sharing in the determination of leasing arrangements is somewhat controversial in the literature. Some empirical studies (e.g., Fukunaga and Huffman 2008, 2009) find it an important determinant of leasing arrangements. Others (e.g., Allen and Lueck 1992, 2002) disagree. Our results provide evidence that risk has significant impacts on leasing arrangements. Risk (as proxied by the income CV variable) has a negative effect on the choice of a cash contract. A standard deviation increase in the coefficient of variation will reduce the probability of choosing a cash contract by 1.8%, and increase the probability of choosing a share contract by 2.1%. The landlord's purchase of insurance is found to be a significant determinant of contract choices. The results show that if a landlord purchases insurance for the target farm

business (denoting possible risk aversion), he/she is less likely to choose a cash contract. This is not consistent with the intuition summarized in Table 2.1. We expected that a risk-averse landlord would be more likely to choose a cash contract. One possible explanation is that the purchase of insurance indicates a more risky farming activity (is large) which deters the use of cash contracts. However, the farm's net worth is found to be insignificant.

2.4.3 Productivity of Effort and Other Attributes

Table 2.4 reports that farming experience is not a statistically significant determinant of contract choices. However, the squared value is found to be significant. It has a positive impact on the probability of selecting cash and share contracts and a negative effect on hybrid contracts. The results indicate that a landlord living in a rural area is more likely to choose a cash contract than those who live in an urban area. The evidence supports the transaction cost hypothesis proposed by Allen and Lueck (2002) which states that an absentee landlord is more likely to choose a share contract, under which the tenants' incentive to overuse the land is smaller than under a cash contract. It does not lend support to the alternative transaction cost hypothesis that an absentee landlord is less likely to choose a share contract since the cost of monitoring is relatively high (e.g., Cheung 1969). The results show that the farm type significantly affects contract choices as well. If the target crop farm belongs to a principle crop farm type (i.e., oilseed and grain farms), the probability of choosing a hybrid contract increases 9.6%. The landlord's ratio of real estate taxes to total rent income is found to be a statistically significant variable. The higher is the ratio, the higher is the probability of cash lease. The tenure status of the tenant operator is found to be a statistically significant determinant of contract choices. Tenant operators who are part-owners of the land are found to be more likely to choose cash or hybrid contracts and less likely to use share contracts.

2.5 Conclusion

This paper provides a simple conceptual model to evaluate the impacts of government programs on contract choices in agriculture. The theoretical model shows that exogenous legal restrictions on the distribution of program benefits between contracting parties, such as the restriction on the direct payments distribution between landlords and share tenants under the 1996, 2002, and 2008 Farm Bills, can cause an offsetting contractual rearrangement in order to restore the benefit distribution to the unrestricted level. The increasingly common use of hybrid contracts (and decreasing use of share contracts) on crop farms may be a form of this contractual rearrangement. We use data from a variety of sources to empirically analyze the determinants of contract choices using a multinomial logit (MNL) model with alternative specified constants. The results confirm that different policy mechanisms have different effects on the farmland contract choices. More specifically, we find that a one standard deviation increase in the PFC (decoupled) payments increases the probability of using a hybrid contract 1.1% and decreases the probability of selecting a share contract by 4.5%. Other farm programs are also found to be significant determinants of leasing arrangements. Their effects vary by the types of programs. Risk-sharing incentives are important determinant of contract choices.

This study generates two important implications. First, it illustrates the potential biases that may arise when restricting the set of potential leasing arrangements to only cash and share contracts. Introducing hybrid contracts into the analysis is especially important to understanding the impact of program payments on leasing arrangements. Second, the analysis suggests that governmental and legal restrictions on benefit sharing between contracting parties are ineffective and induce offsetting contractual rearrangements. The increasing use of hybrid contracts likely reflects a redistribution of program benefits between contracting parties. Most existing empirical research that analyzes the distribution of program benefits between landlords and tenants effects focuses on the cash rental contracts (e.g., Lence and Mishra 2003). Only a few studies examine the benefit distribution under share contracts (e.g., Goodwin, Mishra, and Ortalo-Magné 2010). Future studies may find it helpful to consider different types of contracts,

especially hybrid contracts. Future research endeavors could also use panel data to investigate the impact of policy changes on leasing arrangements.

Chapter 3

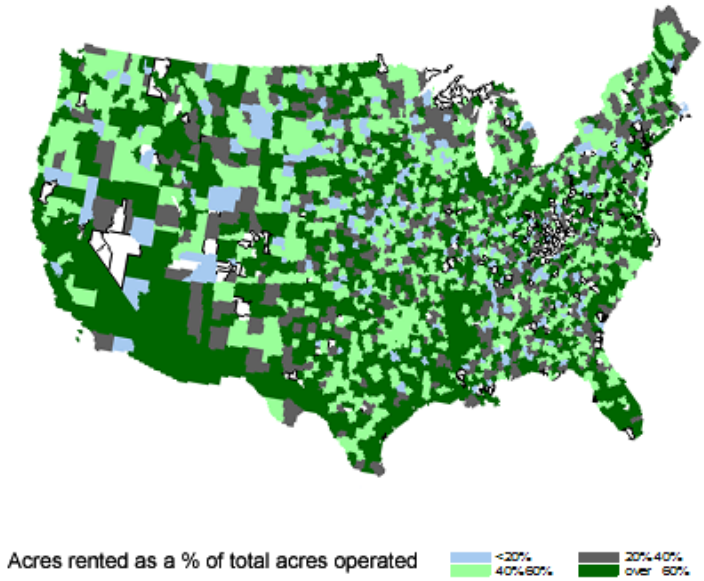
An Empirical Investigation of the Agricultural Subsidy Incidence in Farmland Rental Rates

3.1 Introduction

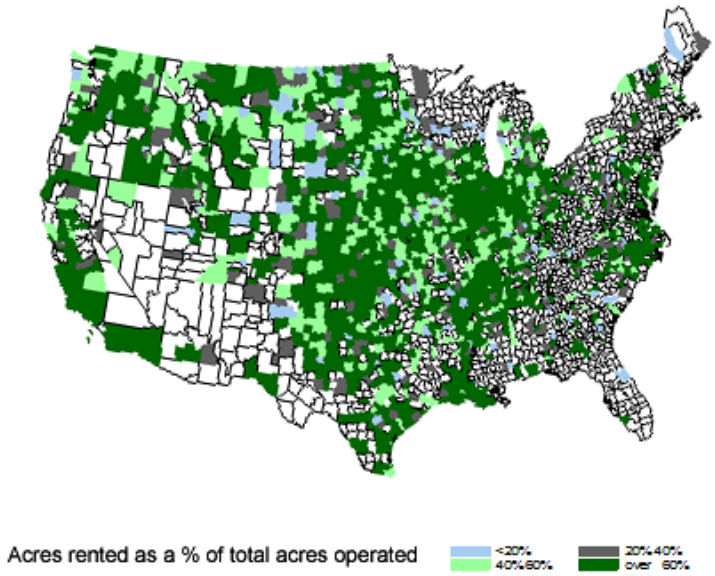
Land is an essential input for farming and ranching and agricultural real estate accounted for around 70% on average of all farm assets between 1960 and 2004 (Mishra and Dedah 2008). Farmland renting is a common practice in U.S. agriculture. Nationwide, about 40% of farmland is leased by the operator. This percentage is even larger in major crop regions such as the Corn Belt (USDA/NASS 2001) and under share leasing arrangements, as is indicated in Figure 3.1.

Agricultural subsidies have become an important source of farm income in the United States over the years. Annual average government payments made directly to all recipients in the farm sector (including landlords) accounted for 20% of total net farm income between 2007 and 2011 (USDA/ERS 2011).[1] A wide variety of income support programs exist that are explicitly

[1] "Farm Income and Costs: Farm Sector Income Forecast." For more details, visit http://www.ers.usda.gov/briefing/farmincome/data/nf_t2.htm [Accessed July 2011.]



(a)



(b)

Source: Authors' compilation from weighted averages of ARMS farm level data: 2002-2007

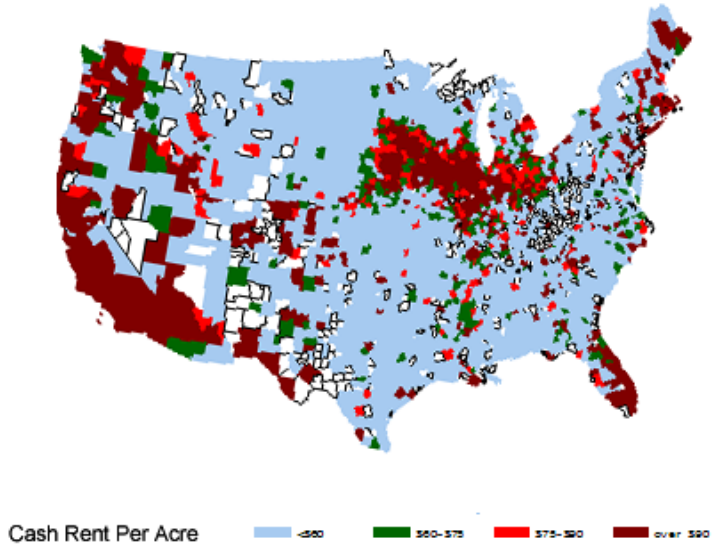
Figure 3.1: Farmland Rented from Others as a Percent of Total Land Operated under Cash Contracts (Above) and Share Contracts (Below)

intended to raise and stabilize farm incomes. In the 2002 and 2008 Farm Bills, commodity program payments which were intended to provide eligible farmers with income support and to address price and revenue risks were estimated to be about $9.14 billion (Monke and Johnson 2010). The eligibility of program payments is often tied to land, or to historical base acreage (in the case of direct payments and counter-cyclical payments), and can directly impact farmland rental rates (Goodwin, Mishra, and Ortalo-Magné 2010). Government payments may also influence rental rates through their impacts on production decisions, income variability, and wealth (Qiu, Goodwin, and Gervais 2011).

Figure 3.3 presents the 2002-2007 county average rental rates under cash and share leases, respectively. Figure **??** shows corresponding county level government subsidies. Correlation between high rental rates and program payments appears to be significant in many areas, particularly in major crop regions such as the Corn Belt. This correlation suggests it is important to evaluate how landlords and tenant operators share government subsidies. The greater is the share of the payments that goes to landlords (e.g. through increases in rental rates), the less effective are subsidies to support farm income (Patton, et al. 2008). It is often argued (see for example, Sherrick and Barry 2003) that the benefits of agricultural programs accrue entirely or almost entirely to operators who own all or part of the land and to non-operator landlords. This line of research argues that tenants gain little from programs since they have to pay higher rents as a result of the subsidies. In addition, tenants do not capture the potential capital gains generated by farmland appreciation.

Other researchers (e.g., Kirwan 2009) suggest that tenant operators are the main beneficiary of government subsidies. Goodwin, Mishra, and Ortalo-Magné (2003b, 2010), and Patton et al. (2008) question this "general statement". They argue that the effects of program benefits vary substantially across the types of programs because the benefits and risk associated with these payments are different. Generalizations with regard to the overall effects of program benefits on land values or rental rates may therefore be misleading.

The distribution of subsidy benefits is complicated by the existence of different leasing



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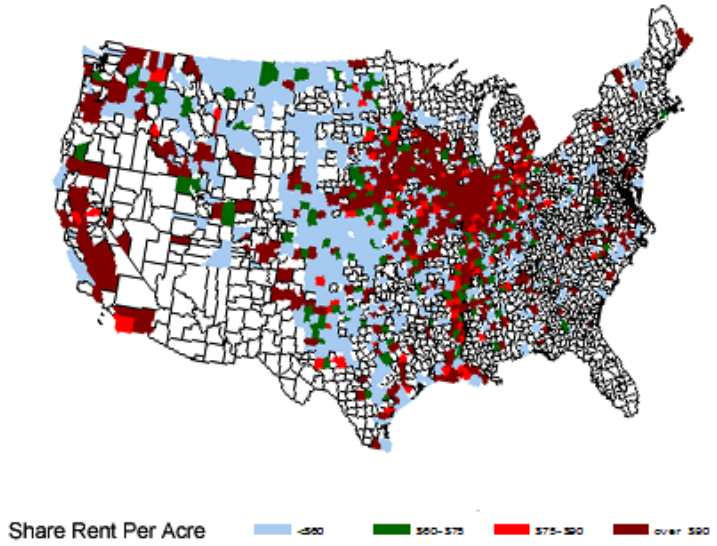
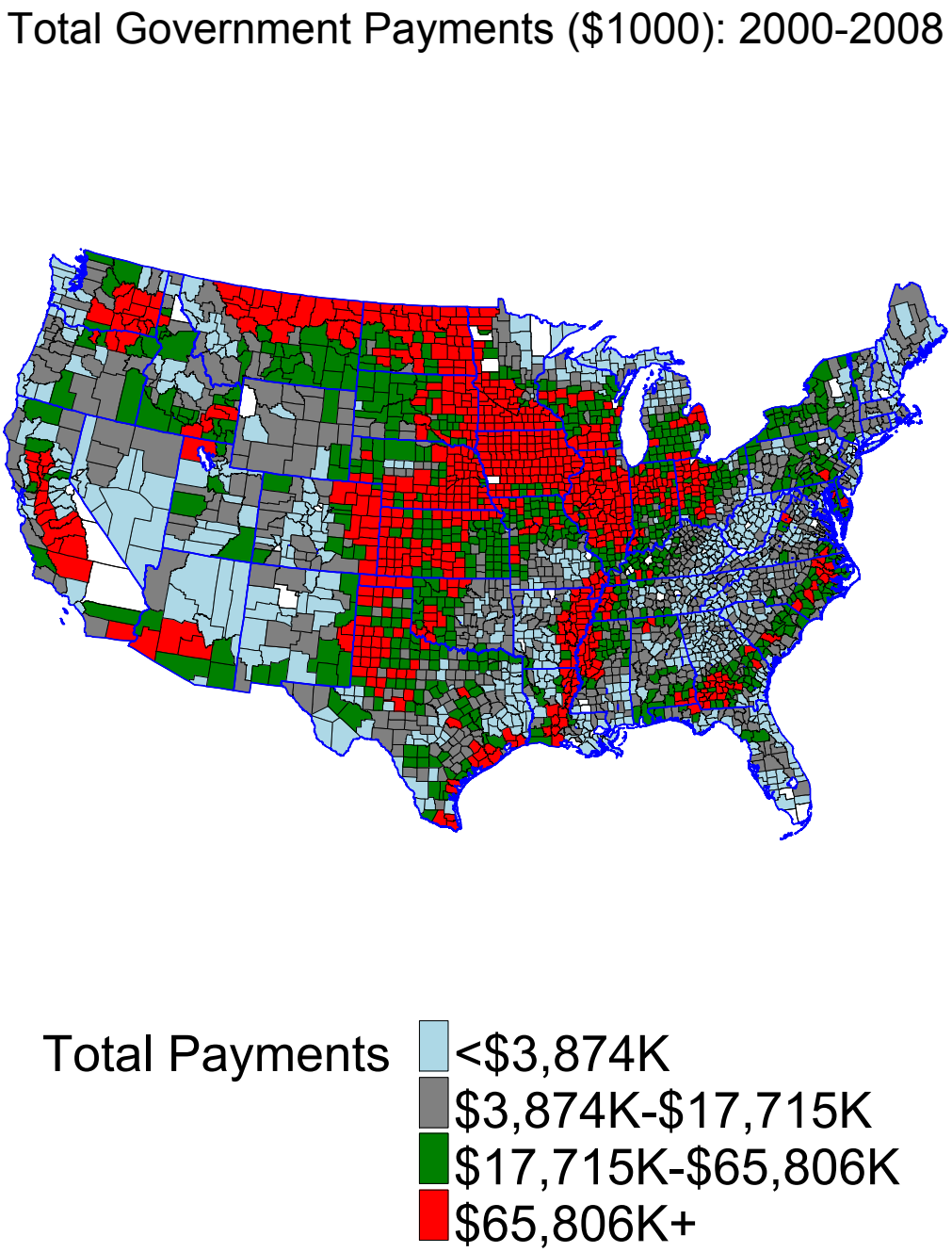


(b)

Source: Authors' compilation from weighted averages of ARMS farm level data: 2002-2007

Figure 3.2: Calculated 2002-2007 County Average Rental Rates under Cash Leases (Above) and Share (include Hybrid) Leases (Below)



Source: Authors' compilation from the county level farm program payment data from the Farm Service Agency (FSA) of the USDA

Figure 3.3: County Level Total Government Subsidies: 2000-2008

arrangements. Legislation also involves restrictions and constraints on the allocation of program benefits. For example, under the 2002 and 2008 Farm Bills, legislation required that direct payments and counter-cyclical payments be shared between tenant producers and landlords on a "fair and equitable" basis. To receive payments, an individual must share in the risk of production and must be entitled to share in farm receipts. If a landlord receives cash rents, then the tenant operator bears all the risk and should receive the entire amount of government subsidies (see Sec 1101 (13A) of the 2008 Farm Bill). Naturally, landlords may indirectly capture a share of those payments by raising cash rental rates. In contrast, a share lease involves a distribution of payments to both the landlord and the tenant operator according to the predetermined share arrangement as they share the crop. In such a case, program benefits for the landlord can come directly from the government as well as through the monetary terms of the leasing contract. Changes in tenure arrangements, for example, from a share contract to a hybrid contract, may also reflect a redistribution of benefits between tenants and landlords (Qiu, Goodwin, and Gervais 2011).[2]

The subsidy incidence literature has generally focused on cash leasing arrangements, although there are exceptions (e.g., Goodwin, Mishra, and Ortalo-Magné 2010). Other land tenure contracts are commonly used in the U.S. In 1999, about 59% of all leased farmland was rented under cash contracts. The remaining 41% was mainly leased under share or hybrid contracts (USDA/NASS 2001). The percentage of share or hybrid contracts is even larger in major crop regions.

The objectives of this paper are to evaluate to what extent main commodity farm programs affect farmland rental rates and to investigate whether the subsidy incidence differs across different types of leasing arrangements. We make two important contributions to the literature. First, we adopt a modified Ricardian rent framework and evaluates to what extent farm programs affect rental rates, taking uncertainty and transaction costs into consideration. Second, our estimation procedure corrects for potential selection issues. Qiu, Goodwin, and Gervais

[2] A hybrid contract combines elements of sharecropping and cash rental agreements such that the tenant pays part of the rent in cash and part as a share of crops or livestock products.

(2011) have shown that government payments have a significant impact on the form of the contractual arrangement used by landlords and tenants. This line of sample selection issue is referred to as incidental truncation (Heckman 1979 and Wooldridge 2002) for its strong self-selection component: farmers self-select into certain types of contracts, so whether or not we observe a certain type of rent depends on an farmer's contract choice first. It is important to account for the nonrandom nature of the sample we have for estimating the rental rate equation. Different methods are investigated to detect and correct any potential selection biases.

3.2 Conceptual Framework

In this section, we develop a behavioral model of a tenant farmer facing revenue uncertainty. The model builds upon earlier efforts of Chavas (1993). We expand their work by introducing program payments and transaction costs into the analysis.

For simplicity, we focus on a single crop year where all input decisions are made at the beginning of the period. The tenant pays the landlord a fixed cash rental rate for the use of the land. The contract is signed before harvest, which implies that the tenant bears the price and yield uncertainty. Assume land is the only fixed input that earns rent. The tenant faces a production technology represented by the production frontier $y = \varepsilon f(x, L)$, where y is farm output and ε is a stochastic term distributed with mean one and finite variance. The stochastic term ε represents multiplicative production uncertainty reflecting the influence of exogenous shocks (e.g., bad weather and inappropriate management) on yields, and $\bar{y} = f(x, L)$ denotes expected production. Assume that farming is a competitive industry, output y and input x being traded on competitive markets. Let q be the output price and r the vector of input prices for x. Denote the land rental rate as s. The output price q is treated as a random variable reflecting price uncertainty. Tenants face both production and price uncertainty. Following Chavas (1993), this uncertainty is characterized by a subjective probability distribution of the random variables ε and q. Let $p = q\varepsilon$ represent uncertain revenue per unit of expected output. Also, assume uncertainty can be represented by a location and a spread parameters, that is, $p = \bar{p} + \sigma e$,

where $\bar{p} = E(p)$ is the location parameter,σ the spread parameter, and e is a random variable with mean zero. Assume there are two types of government subsidies: decoupled and coupled. Decoupled payments are lump-sum income transfers which are independent of current market and production conditions. There are different types of coupled programs. The eligibility for the coupled payments can be triggered by market price (e.g., the LDPs), or current production (e.g., disaster assistance payments), or current revenue (a product of price and production, e.g., the ACRE payments). Denote the net farm income from agricultural-related activities by:

$$\pi = pf(x, L) - r'x - sL + G_1 + G_2(p),$$

where $pf(x, L)$ is agricultural revenue, $r'x$ is production costs, and sL is the total rent paid to the landlord. G_1 and $G_2(p)$ denote decoupled and coupled government payments, respectively. Without loss of generality, we set G_2 as a function of uncertain revenue p . Let I be exogenous non-farm income. Then the objective function of the tenant operator can be written as:

$$EU(\pi) = EU[I + pf(x, L) - r'x - sL + G_1 + G_2(p)], \tag{3.1}$$

where E is the expectation operator, U is a Von Neumann-Morgenstern utility function that satisfy the conditions $U' = \partial U/\partial\pi > 0$ and $U'' = \partial^2 U/\partial\pi^2 < 0$.

We are interested in analyzing the determination of the rental rate of land. This can be done by investigating the bid price for land. The marginal analysis indicates that the optimal rental rate of land is the value that makes the tenant operator indifferent between farming and not farming after the uncertainty has been compensated. This indicates that the optimal rent s will satisfy the following equation:

$$EU(\pi) = EU[I + pf(x, L) - r'x - sL + G_1 + G_2(p)] = U(I) \tag{3.2}$$

where $EU(\pi)$ represents the situation where the tenant operator farm the land L, while the

$U(I)$ corresponds to the situation where he does not participate farming. Let $s^*(I, \bar{p}, \sigma, r, x, L)$ be the implicit solution of (3.2) for s. The Ricardian theory indicates that rent is the highest price that the tenant can afford for the land use (Chavas 1993). The Ricardian rent to land L can thus be defined as:

$$s^e(I, \bar{p}, \sigma, r) = \max_{x \geq 0, L \geq 0} [s^*(I, \bar{p}, \sigma, r, x, L)]. \tag{3.3}$$

Chavas (1993) has demonstrated that s^e is also the market equilibrium rental rate of land under free entry and exit. The optimization problem of (3.3) can generate the optimal choice for $x^*(I, \bar{p}, \sigma, r)$ and $L^*(I, \bar{p}, \sigma, r)$. This study focuses on the optimal rental rate and we do not address the properties of x^* and L^* here in details. However, future studies interested in investigating the optimal input, farm size, production reaction, and effects of government subsidies on these factors under uncertainty can further explore the properties of x^*, L^* and the optimal output given $x*$, L^*.

As argued by Pratt (1964), risk-averse tenant farmer must be compensated by receiving a risk premium for bearing privately revenue uncertainty. This can be seen by defining the Arrow-Pratt risk premium as the amount of money RP which satisfies $EU(\pi) = U(E\pi - RP)$, implying that risk premium RP can be expressed as $RP(x, L) = [\bar{p}f(x, L) - r'x - sL + G_1 + G_2(\bar{p}) - U^{-1}EU(\pi)]$, and $RP > 0$ under risk aversion. The optimal rental rate can then be generated implicitly from Equation (3.2):

$$s^* = [\bar{p}f(x, L) - r'x + G_1 + G_2(\bar{p}) - RP(x, L)]/L, \tag{3.4}$$

where $RP(x, L)$ measures the cost of private risk bearing to the tenant farmer. Accordingly, the Ricardian rental rate defined in (3.3) becomes:

$$\begin{aligned} s^e &= \max_{x \geq 0, L \geq 0} [\bar{p}f(x, L) - r'x + G_1 + G_2(\bar{p})]/L \\ &= \max_{x \geq 0, L \geq 0} [Mkt(x, L) + g_1 + g_2(\bar{p}) - rp(x, L)] \end{aligned} \tag{3.5}$$

Where $Mkt(x, L) = [\bar{p}f(x, L) - r'x]/L$ is the expected net market return per acre, g_1 and g_2 are, respectively, decoupled and coupled program payments per acre, and $rp(x, L)$ represents the average risk premium per acre. It follows that the equilibrium occurs at the maximum of the expected net return per acre (both from farming and government subsidies) minus the average risk premium per acre. The result here indicates that higher revenue uncertainty reduces the Ricardian rent. This conclusion is quite general as it is independent of the pattern of absolute or relative risk aversion. Revenue uncertainty makes the risk-averse tenant worse-off, thus he needs to be compensated for the implicit cost of his private risk bearing. For example, compensate the tenant by offering a lower rental payment for land use. Under such circumstance, the landlord would benefit from a stabilization program not only through increasing the expected net return, but also reducing revenue uncertainty. This can be one reason that a dollars subsidy of coupled payments often results in more than one dollars rise in rental rates. In contrast, the landlord benefits from a decoupled direct program only through increasing the expected net income, with no risk reduction. Equation (3.5) indicate that if the land market is competitive and no transaction cost needed, and if farmland is the only fixed input that gains the rent (residuals after all variable costs have been controlled) then the coefficient of decoupled payments can be expected to be one and the coefficients of coupled payments might be greater than one as they represent both increases in expected net returns and decreases in uncertainty. The coefficients associated with uncertainty shall be expected to be negative for risk-averse tenants.

In addition to uncertainty, the rental rate can also be influenced by transaction costs. Although the notion of Ricardian rent, defined as the residual return to farmland after all other factors have been paid, does not incur transaction costs explicitly, econometricians are not able to observe this residual-based rent in reality. What we can observe are usually self-reported rents provided by tenants and/or landowners. Under this condition, rent is a contractual agreement of income and risk sharing between two parties. To achieve an agreement or a contract always involves transaction costs, such as negotiation and enforcement costs. Any adjustments to the existing agreement to reflect changes in factors that determine the rent will incur transaction

costs. These include information collection costs, renegotiation costs, switching partner costs, land-use conversion costs, and so on. With these costs, landlords will only seek to raise rents if the potential benefits netting transaction costs are positive.

There are many ways to specify transaction costs. We adopt here one of the most widely used specifications—the so-called iceberg transaction costs. In a system involving iceberg transaction costs, an individual receives the fraction k when he sells/rents out one unit of a good, or receives k dollar when he sells/rents goods with value of one dollar, where $k \in (0, 1)$. In our case, recontracting always involves some costs for both parties and the landlord then receives only the fraction k of each dollar of increases in net returns. This specification implies that the fraction $1 - k$ of their value disappears because of recontracting. Transaction costs can explain why we do not observe a one-to-one, complete subsidy incidence in rents, even for the decoupled payments. The corresponding Ricardian rent with iceberg transaction costs can thus be expressed as:

$$s^e = \max_{x \geq 0, L \geq 0} [k_0 Mkt(x, L) + k_1 g_1 + k_2 g_2(\bar{p}) - rp(x, L)] \tag{3.6}$$

The nature of different programs indicates that the k associated with each program would probably be different as the costs to negotiate a new contract are different. For example, direct payments are lump sum income transfers tied to farmland and are usually known prior to the lease contracts. Thus, the transaction costs for raising rents to reflect an increase in decoupled payments can be expected to be smaller than those required to capture the coupled program payments. That said, a larger k can then be expected for decoupled payments. Transaction costs for landowners are currently engaging in farming or have farm-related experiences can be expected to be smaller than those who have no farming experience or are not currently engaged in farming.

Given the above conceptual analysis, a corresponding econometric specification for the rent

Equation (3.6) can be written as:

$$s = \alpha + \beta E(Mkt) + \sum_{j=1}^{J} \gamma_j E(g_j) + \sum_{k=1}^{K} \varphi_k rp_k, \tag{3.7}$$

where α is a constant term that can be interpreted as an overall effect of a set of unobserved factors affecting the rental rate, g_j is government subsidy from program j, rp is a vector of proxies presents the uncertainty/risk premium, which includes the coupled program payments that reduce the uncertainty and the 10-year county average coefficient of variation (CV) of cash receipts from market.[3]

U.S. agricultural commodity programs have been characterized by three main program instruments since the 1996 Farm Bill. The first is fixed, direct payments (FDPs), which are decoupled payments, providing annual lump-sum income transfer that are independent of production and market condition. A second category of programs operate by supporting market prices through loan deficiency payments (LDPs) and marketing loans, which entail subsidies whenever market prices are lower than the loan rate. A third category of support is conveyed through countercyclical payments (CCPs), which provide payments that are independent of production whenever aggregate market prices fall beneath a target level. The 2008 Farm Bill introduced the average crop revenue program (ACRE) as an alternative to the CCP program. ACRE is designed to protect eligible farmers against revenue losses, regardless of the cause: price decline, yield loss, or a combination of the two. However, program participation has been fairly low to date. For the 2009 crop year, only about 8% of the total number of eligible farms participated in the ACRE program (Shields, Monke, and Schnepf 2010).

In the empirical stage, four types of government subsidies have been investigated: the above three major commodity program payments and the disaster assistance payments. The next section discusses the construction of the data and the empirical challenges involved in the estimation phase.

[3]We assume that for a give coupled program payments, income effect and risk effect are separable.

3.3 Data and Empirical Strategy

We use the 2002-2006 farm-level Agricultural Resource Management Survey (ARMS) data and the 1993-2006 county level government program payment data from the Farm Service Agency (FSA) of the USDA. We also utilize data from the 1988-2006 Regional Economic Information System (REIS) and county level data from the 1997 and 2002 *Censuses of Agriculture*.

The ARMS is a national survey that provides observations of farm-level production practices, economic attributes, and operator households' characteristics. We utilize this dataset to calculate farm-specific rental rates, as well as to obtain information about additional farm and operator characteristics that may impact leasing choices. The REIS contains annual estimates of personal income at the national, state, metropolitan area, and county levels. We computed county-level, net agricultural market returns using marketing receipts and production costs from the REIS.

We selected farms in the ARMS that rented at least part of their land. The ARMS data also include information about farmland leasing arrangements. Farmland contracts are divided into three categories: cash rental arrangements, share leases (include both pure sharecropping and hybrid contracts), and land rented for free. This study focuses on farms using cash and share leases (we refer to these farms as target farms) which accounted for about 95% of the total number of farms. Tenant operators reported acreage rented from landlords under each leasing arrangement. They also reported total cash rents paid and the shares of production that went to the landlords. We categorize target farms according to one of three leasing arrangements: cash-only, share-only, and both. In our sample, 68% of farms are cash-only farms (i.e., farms only use cash leases to contract with one or more landlords), 12% are share-only farms (similarly, farms only use share leases to contract with their landlords), and the remaining 20% use both arrangements. We exclude some observations (less than 2% of the available sample) from the analysis because they represent atypical situations (for example, farms reporting rents exceeding $800 per acre). After this procedure, a total of 48,886 observations are available for the empirical analysis. The data from ARMS are merged with other county-level variables. Table 3.1 presents

the definitions of key variables and summary statistics. All monetary values were adjusted by the consumer price index to represent 2004 dollars.

3.3.1 Measurement Errors in Dependent Variables

While total acres and total rents are available, the true per-acre rental rates are usually unobserved; hence the first step of the empirical analysis is to compute a proxy for this dependent variable. Analyses that have relied on a constructed measure of the rental rate may suffer from measurement errors that could ultimately cause important biases in the estimation stage. Consider the *Census of Agriculture* data as an example. The *Census* does not directly record per-acre rental rates. It only records the total cash rent paid as well as the total acreage rented. The latter variable includes land rented under cash, share, and all other types of leasing arrangements. Previous studies (e.g., Kirwan 2009) computed a per-acre cash rental rate by dividing total cash rent by total acres rented. Because the dependent variable also included acres rented under share arrangements (and other forms of leasing arrangements) and because acreage or contract choices could correlate with exploratory variables like government subsidies (Goodwin and Mishra 2006; and Qiu, Goodwin, and Gervais, 2011), there would be an important endogeneity bias in the estimation stage.

In the case of the ARMS data, the 2002-2006 surveys report rented acres under cash and share contracts separately. The latter category however also includes hybrid contracts which have rental payments based on a fixed cash payment along with shared production. Given the prominence of hybrid contracts, it may be misleading to compute the cash rental rate directly as the total cash amount (paid as rent) divided by total acres (rented under cash contracts) without specifically differentiating the contract type. The cash rental rate would be overestimated while the share rental rate would be underestimated since part of the cash actually comes from hybrid contracts. Such a proxy for the rental rate may therefore suffer from measurement error.

We solve this problem by distinguishing hybrid contracts from pure sharecropping contracts, and then by calculating the rental rates under different types of contracts separately. First

consider cash-only farms. There are no reporting issues for these farms because the cash rental rate can be calculated as total cash rent divided by total acreage under contracts, as long as we exclude all other types of contracts from calculation. We then turn to farms in our database which are classified as share-only (i.e., only use share contracts). Recall that in the ARMS data, the farms that report share contracts include both pure sharecropping and hybrid contracts. For this share-only category of farms, we can conclude that pure share contracts were used if reported cash payments are zero. The per-acre share rent can then be calculated as the total value of the landlords' share of production divided by the acres under share contracts, and these observations will fall under the pure share category. Alternatively, a farm in this category can report share contracts as well as positive cash payments. This indicates that hybrid contracts have been used. The number of farms that use only hybrid contracts in the sample is small (a little over 1% of the target farms); and thus our investigation of subsidy incidence under different leasing arrangements focuses on pure cash and pure share leasing arrangements.[4]

Meanwhile, most empirical studies directly use the computed rental rate the one like we just calculated as the dependent variable. This may be problematic because it is quite common that landlords share part of the cost of productive inputs (other than land) with tenants, especially under share leasing arrangements (Allen and Lueck 2002). In our sample, more than 30% of the landlords reported sharing production costs with their tenants. In that case, the computed rents reflect the payments made to landlords in order to compensate for the costs of the non-land inputs. When these payments are correlated with one or more of the regressors (e.g., market returns), measurement problems which may result in biases in the estimation stage are introduced. To account for this, we subtract the landlords' share of variable costs from the calculated rental rates, thereby providing a more reliable measure of the dependent variable - pure rental rates - in the empirical models.

[4]Although the number of farms using only hybrid contracts or hybrid contracts with pure share contracts is small, it does not follow that hybrid leasing arrangements are uncommon. In 1999, about 11% of the leased farmland were under hybrid contracts in the United States. Farms are generally more likely to use other leasing arrangements (e.g., cash leases) together with hybrid contracts. In the U.S., a typical tenant farmer contracts with more than 4 landlords on average (USDA/NASS 2001).

Table 3.1: Summary Statistics (N=48,886)

Variable	**Definition**	**Frequency**	**Percentage**
Cash-only	Farms using only cash leases	41,583	68.19
Share-only	Farms using only share leases	7,225	11.85
Both	Farms using both cash and share leases	12,173	19.96
Variable	**Definition**	**Mean**	**Std. Dev.**
Cash rent	Cash rental rate (exclude the LL's non-land cost share)	72.44	91.73
Share rent	Share rental rate (exclude the LL's non-land cost share)	92.03	100.31
	Historical county average payments		
Total Payments	Total government subsidies received by tenants and landlords ($/acre)	31.25	5.56
LDPs	Loan deficiency payments (include other marketing loan benefits) ($/acre)	7.39	7.98
CCPs	Counter-cyclical payments ($/acre)	10.3	10.14
Disaster	Disaster payments ($/acre)	2.77	3.28
Other	All other government payments ($/acre)	4.41	5.56
Market returns	Net returns from market sales ($/acre)	20.34	134.73
CV	10-year county average coefficient of variation of cash receipts from market ($/acre)	0.17	0.07
FDPs	Farm operators' self-reported direct decoupled payments ($/acre)	13.09	21.49

Prior to the 2002 Farm Act, production flexibility contract payments functioned as direct decoupled payments. Counter-cyclical payments are represented by market loss assistance payments between 1998 and 2001.

3.3.2 Measurement Errors in Independent Variables

Farmland rental rates are functions of expected cash flows from various sources. Certain payments (such as direct decoupled payments) are known with certainty, at least in the short-run, prior to signing a contract. However, payments such as disaster assistance and price supports are not predetermined. They are triggered by market and/or production conditions. Measurement issues arise if actual realized payments are used to represent expectations, a point noted in Goodwin, Mishra, and Ortalo-Magné (2003a). In addition, realized individual payment data may be correlated with unobserved factors, such as land productivity. This can cause endogeneity problems and may result in biases as well. To control for the potential errors-in-variables and endogeneity problems, we follow Goodwin, Mishra, and Ortalo-Magné (2010) and use a historical five-year county average payments instead of the realized individual data reported in the ARMS data.[5] As such, expected net market earnings are measured by a historical five-year county average value per acre. We also include a measure of "farming risk" by computing a ten year coefficient of variation of the county-level agricultural market net returns.

3.3.3 Selection Bias

One can observe a specific type of rent only if the farmer initially chooses this type of leasing arrangement. Sample selection bias occurs if unobservable characteristics (e.g., land productivity) that affect the rental rates are correlated with factors that affect the contract choice (e.g., farm level government subsidies and production risk). There exists an extensive literature on the detection and correction of selection issues. Bourguignon, Fournier, and Gurgand (2007, BFG hereafter) provide an overview of the methods available to account for selection issues in the context of the multinomial logit model. BFG conduct a set of Monte Carlo experiments and find that in most cases, the approach introduced by Dubin and McFadden (1984, DMF1 hereafter) is the preferred method in comparison to the most commonly used procedure proposed

[5] Because they depend upon market prices. Loan Deficiency Payments can vary substantially according to the time period chosen. We thus use a 10-year historical average.

by Lee (1983). BFG also develop an estimator (which they label DMF2) that circumvents the specific linearity restriction of the DMF1 estimator and which is more robust. Our empirical analysis mainly focuses on the DMF2 approach.[6]

3.3.4 General Description of The Selection Model

Consider a two-equation, censored regression model. Let the subscript j define a categorical variable that describes the choice of a decision-maker among M alternatives, based on the level of utility associated with each alternative y_j^*. The variable of interest y_1 is observed if and only if alternative 1 is chosen:

$$\begin{aligned} y_1 &= \boldsymbol{x}\boldsymbol{\theta}_1 + u_1, \\ y_j^* &= \boldsymbol{z}\boldsymbol{\lambda}_j + \eta_j, \qquad j = 1, 2, \ldots M, \end{aligned} \tag{3.8}$$

where vectors $\boldsymbol{x}$ and $\boldsymbol{z}$ are exogenous variables and u_1 is a disturbance term for which $E(u_1 \mid x, z) = 0$ and $V(u_1 \mid x, z) = \sigma^2$. If alternative 1 is selected, we know that $\varepsilon_1 = \max_{j \neq 1}(y_j^*) - y_1^* < 0$. Assume that η_j is independent and identically distributed according to a Gumbel distribution and define $\Gamma \equiv \{z\lambda_1, \ldots, z\lambda_M\}$. After some algebraic manipulations, a consistent estimate of the vector $\boldsymbol{\theta}_1$ can be obtained based on the regression:

$$y_1 = \boldsymbol{x}_1\boldsymbol{\theta}_1 + \mu(\Gamma) + \omega_1 = \boldsymbol{x}_1\boldsymbol{\theta}_1 + \tilde{\mu}(P_1, P_2, \ldots, P_M) + \omega_1, \tag{3.9}$$

where $\mu(\Gamma) = \tilde{\mu}(P_1, P_2, \ldots, P_M) = E[u_1 \mid \varepsilon_1 < 0, \Gamma]$ is the conditional mean of the disturbance in the equation of interest, P_j is the probability that alternative j is chosen, and ω_1 is a residual which is mean-independent of regressors.[7] Restrictions on $\tilde{\mu}(\cdot)$, or equivalently, $\mu(\Gamma)$ are required in the estimation stage. In what follows, we briefly review two broad approaches

[6] DMF1 results are quite similar to DMF2 results in this study. DMF1 results are available from authors on request.

[7] More detailed information about the calculation can be found in Bourguignon, Fournier, and Gurgand (2007). An overview of sample selection models can also be found in Shadish, Cook, and Campbell (2001); Stolzenberg and Relles (1997); and Winship and Morgan (1999).

to estimate (3.9) based on the insights of BGF.

3.3.5 Lee's Model

Lee (1983) proposes a generalization of the two-step selection bias correction method introduced by Heckman (1979). Following Lee, let $F_{\varepsilon_1}(. \mid \Gamma)$ be the marginal distribution function of ε_1, which can be transformed into a standard normal random distribution: $J_{\varepsilon_1}(. \mid \Gamma) = \Phi^{-1}(F_{\varepsilon_1}(. \mid \Gamma))$, where Φ is the standard normal cumulative distribution function. We assume that u_1 and $J_{\varepsilon_1}(. \mid \Gamma)$ are jointly distributed under the hypothesis that $E(u_1 \mid \varepsilon_1, \Gamma) = \sigma\rho_1 J_{\varepsilon_1}(\varepsilon_1 \mid \Gamma)$ and that the joint distribution is independent of Γ. The expected value of the disturbance term u_1, conditional on alternative 1 being chosen, is then expressed as:

$$E(u_1 \mid \varepsilon_1 < 0, \Gamma) = -\sigma\rho_1 \frac{\phi(J_{\varepsilon_1}(0 \mid \Gamma))}{F_{\varepsilon_1}(0 \mid \Gamma)}$$

where ϕ is the standard normal density function. In this case, Eq. (3.9) then can be consistently estimated by a two-stage method on the basis of:

$$y_1 = \boldsymbol{x}_1\boldsymbol{\theta}_1 - -\sigma\rho_1 \frac{\phi(J_{\varepsilon_1}(0 \mid \Gamma))}{F_{\varepsilon_1}(0 \mid \Gamma)} + \omega_1.$$

Schmertmann (1992) points out that this method fully specifies the joint distribution, which can be very restrictive and may therefore have strong practical limitations. In general, it implies that the correlation between u_1 and $(\eta_j - \eta_1)$ must be of the same sign for all j alternatives. In other words, unobservable determinants of the choice of alternative 1 against any other alternative should be correlated in the same direction with unobservable determinants of the outcome y_1. In the special case here, in which the selection choice is based on an *i.i.d.* multinomial logit model, Lee's (1983) implicit assumption is even stronger. It implies that correlations between u_1 and $(\eta_j - \eta_1)$ are all identical, which is likely to be unrealistic in practice. For example, let land productivity be represented by the unobserved disturbance u_1 and η_j be farm level government subsidies distributed to the landlord associated with contract choice j. Correla-

tion between land productivity and government subsidies could be positive or negative under a share contract. However, there could be no correlation between the two under a cash contract because all subsidies are distributed to the tenant under current legislation.

3.3.6 Dubin and McFadden's Model and Its Variation

Dubin and McFadden (1984) adopt a different approach that introduces a linearity assumption with regard to the original shock η_j instead of ε_1. They assume:

$$E(u_1 \mid \eta_1, \eta_2, \ldots, \eta_M) = \sigma \frac{\sqrt{6}}{\pi} \sum_{j=2,\ldots,M} r_j(\eta_j - E(\eta_j)) \text{, and} \sum_{j=1,\ldots,M} r_j = 0$$

where r_j is a correlation coefficient between u_1 and $(\eta_j - \eta_1)$. The estimated equation (3.9) thus becomes:

$$y_1 = \boldsymbol{x}_1\boldsymbol{\theta}_1 + \sigma \frac{\sqrt{6}}{\pi} \sum_{j=2,\ldots,M} r_j \Big[\frac{P_j \ln(P_j)}{1 - P_j} + \ln(P_j) \Big] + \omega_1$$

BFG propose to make the shock u_1 a linear function of normal variates:

$$E(u_1 \mid \eta_1, \eta_2, \ldots, \eta_M) = \sigma \sum_{j=1,\ldots,M} r_j^* \eta_j^* \text{, given} \sum_{j=1,\ldots,M} r_j = 0$$

where η_j^* is a standard normal variable with pdf $\eta_j^* = \Phi^{-1}(G(\eta_j))$ and r_j^* is the correlation between u_1 and η_j^*. The corresponding estimation equation (3.9), conditional on choosing $j = 1$, becomes:

$$y_1 = \boldsymbol{x}\boldsymbol{\theta}_1 + \sigma \Big[r_1^* m(P_1) + \sum_{j=2,\ldots,M} r_j^* m(P_j) \frac{P_j}{P_j - 1} \Big] + \omega_1 \tag{3.10}$$

where $m(P_1) = E(\eta_1^* \mid y_1^* > \max_{s \neq 1}(y_s^*), \Gamma)$ and $m(P_j)\frac{P_j}{P_j - 1} = E(\eta_j^* \mid y_1^* > \max_{s \neq 1}(y_s^*), \Gamma)$.

Correspondingly, we define $y_1 \equiv r_{t,c}$ as cash rental rates and y_j^* as the level of utility associated with farmland contract choice j, where j is a choice among: cash-only, pure-share only, or both cash and share contracts. The cash rental rates are observed if and only if cash-

only contracts are chosen. The empirical model implies that rental rates $r_{t,c}$ are conditional on expected market earnings and expected government payments as well as selection correction terms obtained from a multinomial logit model that explains leasing arrangements from among three alternatives: cash-only, pure-share only, or both cash and share contracts. The leasing arrangement outcome is conditional on expected market returns and government payments as well as other independent variables such as farm type, total asset value, and age of the farm operator.[8]

The model specification for pure-share rental rates can be defined similarly. Then consistent estimation of subsidy incidence under cash and pure-share contracts could be obtained based on the two-step regression in equation (3.10) we just discussed above.

Several other aspects about the empirical procedure are important to discuss before presenting the results. First, the ARMS applies complex stratified and probability-weighted sampling methods. The individual strata used in collecting the data are not identifiable, making possible efficiency gains in the estimation stage irrelevant. We can observe population weights for each individual farm and we would be able in theory to incorporate these weights into the estimation procedure. However, this study utilizes both farm-level ARMS data (e.g., rental rates) and county-level payment and returns information (e.g., government payments and net market returns). ARMS weights are created at the state or, sometimes, multi-state level. Even though the sample farms may be in a target county, the farms represented by the weights are likely to be located elsewhere in that state or, in many cases, other parts of the United States. This makes the value of weighted estimation procedures dubious at best. Dubman (2000) suggests county-level estimates are best left unweighted because of their potential to introduce distortions in the estimation.[9] For these reasons, we present and focus on unweighted regression results in the following section. However, the unweighted results need to be interpreted within

[8]When reporting the results, we present the second-stage regression equation only. Results from the multinomial logit estimated in the first-stage are available upon request.

[9]Wooldridge (2001) provides a comparison of systematic treatments of weighted and unweighted M-estimators under variable probability stratified sampling. Provided the underlying feature of the conditional distribution is correctly specified (in a linear regression model this means that the error must have zero conditional mean), the unweighted estimator is consistent.

the context of this particular sample of farms, and should not be directly extended to the entire population.

Second, the selection equation implicitly imposes the Independence of Irrelevant Alternatives (IIA) assumption, which states that the probability of choosing among two alternatives is not affected by the presence of additional alternatives. We used both the Chi-Square test proposed by Hausman and McFadden (1984) and the likelihood ratio test proposed by Small and Hsiao (1985) to test the IIA. Both tests failed to reject the null hypothesis that the IIA assumption is valid at conventional levels of significance.

Also, correlation among observations from the same county or same type of farm may exist. Therefore, clustered robust standard errors, based on the farm type and county, were implemented. However, the clustered and unclustered results are quite similar and thus we only present the unclustered results.

Meanwhile, as this is a two-stage estimator, conventional OLS standard errors will be incorrect. We apply weighted least squares in the second stage to account for heteroskedasticity present in the model due to selection issues. The estimator variances for all methods are also bootstrapped (200 replications) to derive valid standard errors in light of the two-stage estimation.

3.4 Results

The empirical analysis consists of two parts. First, we investigate the incidence of aggregate government subsidies on farmland rental rates under both cash and share arrangements. This provides a general idea of the extent of capitalization into farmland rental rates of an extra dollar of government payments. Second, we disaggregate government payments into four distinct programs and evaluate the impacts of each type of program payment on rental rates. In each part, we report the results from three different models: 1) generalized linear regression (GLR); 2) Lee's selection procedure; and 3) the Dubin-MacFadden flexible selection procedure (DMF2).

3.4.1 Aggregate Program Payments

We first consider the effects of total subsidies, aggregated across all programs, on rental rates. Table 3.2 presents estimates of the impacts of aggregate subsidies on cash rental rates obtained from a generalized linear regression (GLR) and selection regression models (Lee and DMF2). The results from the generalized linear regression show that each additional farm subsidy dollar raises cash rents by \$0.50 per acre. In other words, landlords capture 50% of the total benefits and leave 50% to tenant producers. The same empirical model predicts that each dollar obtained from market returns increases rental rates by \$0.04. The estimated coefficient of the variable CV is 103.13. Given the mean and standard deviation of CV reported in Table 3.1, an increase of one standard deviation in CV will decrease cash rental rates by \$7.22 on average. The significant and negative effect of CV on cash rental rates suggest that risk is an important determinant of rents. The more uncertainty are farming activities, the lower are the rental payments needed to compensate the landlords.

Table 3.2: Effects of Aggregate Subsidies on Cash Rental Rates Models (N=48,886)

Cash rent	**GLR**		**Lee**		**DMF2**	
Variable	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**
Total Payments	0.50*	0.42	0.30*	0.04	0.41*	0.03
Market Returns	0.04*	0.01	0.10*	0.01	0.08*	0.01
CV	-103.13*	15.15	-199.01*	24.26	-106.86*	19.2
Constant	54.94*	3.38	31.07*	7.81	30.25*	9.18
σ^2	–	–	3230.34*	97.7	12323.95*	2501.17
ρ_1	–	–	-1.61*	0.41	-0.13	0.22
ρ_2	–	–	–	–	0.73**	0.41
ρ_3	–	–	–	–	-1.59*	0.27

*Note:*Asterisks (*) and (**) denote significance at the 5% and 10% level, respectively.

We argued in the previous section that a generalized linear regression framework that does not account for the role of government payments and other related factors on the types of leasing arrangements may result in estimation bias because of selection issues. Both selection models reveal statistically significant correlation coefficients between the error terms from the cash rental rate equation and the leasing arrangement choices. This suggests that selection issues in the context of government payments and leasing arrangements are very relevant. As we argued in the previous section, there exists evidence in the literature that the DMF2 model is superior to the most commonly used method of Lee (1983). In what follows, the discussion of results largely focuses on the DMF2 approach.

The DMF2 results suggest that an additional dollar of government subsidies tends to increase cash rental rates by \$0.41, *ceteris paribus*. The incidence of government payments on cash rents is smaller than that under the GLR model, and can be explained by the selection issues not being captured by the GLR model. The results also show that an extra dollar of market returns raises rental rates by \$0.08 on average. This estimate is higher than that obtained by the GLR model. The impacts of the CV of market returns on cash rents are similar under both the GLR and DMF2 models. The results from the selection models confirm that a government subsidy has a significant positive effect on cash rental rates. However, the rates of benefit pass-through are not as large as the one predicted by a simple linear framework, which does not account for selection issues.

Table 3.3 presents the corresponding estimates of the GLR, Lee and DMF2 models under share leasing arrangements. Recall that when analyzing the impact of aggregate subsidies, share rents include both the value of the share of production that went to landlords (according to lease terms) and government payments directly distributed to landlords. In contrast to cash rents, the results suggest a much higher benefit pass-through from tenants to landlords. The DMF2 estimates suggest that each additional subsidy dollar raises share rental rates by \$0.78.

In the U.S., 50-50 share arrangements are common (Huffman and Just 2004). Allen and Lueck (2002) point out that 50-50, 60-40 (the tenant obtains 60% of the crop and leaves 40%

Table 3.3: Effects of Aggregate Subsidies on Share Rental Rates Models (N=48,886)

Share rent	**GLR**		**Lee**		**DMF2**	
Variable	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**
Total Payments	0.77*	0.05	0.77*	0.02	0.78*	0.04
Market Returns	0.08*	0.03	0.08*	0.01	0.08*	0.03
CV	-129.91*	16.76	-133.79*	18.27	-117.17*	24.51
Constant	76.52*	3.93	81.38*	1.95*	38.16*	35.33
σ^2	–	–	6782.15*	363.74	7111.89*	1797.75
ρ_1	–	–	-0.02*	0.01	-0.74**	0.42
ρ_2	–	–	–	–		
ρ_3	–	–	–	–		

*Note:*Asterisks (*) and (**) denote significance at the 5% and 10% level, respectively.

to the landlord), and 67-33 are the three most common sharecropping rules in their sample of Nebraska and South Dakota farmers. Our results indicate that landlords are able to capture benefits that are substantially higher than what is dictated by the legislative guidelines for payment distribution. Hence, the results offer support to the argument that landlords are capturing program benefits from both direct program payments and from increased rental rates. In other words, legal restrictions on benefit distribution between contracting parties are largely ineffective since benefits can be redistributed by other means such as adjusting/renegotiating the rental rates. The results also indicate that landlords extract 8% of the market returns under share leases. When considering the impact of risk on rental rates, the results suggest that rents will decrease by $8.20 per acre following a one standard deviation increase in the CV.

In summary, the results confirm that aggregate government subsidies have statistically significant and economically substantial effects on rental rates. These effects vary across leasing arrangements. The incidence of program payments on share rental rates is larger than that for cash rental rates. One explanation is that transaction costs for landowners under sharecropping (i.e., are still actively engaged in farming) can be expected to be smaller than those using cash contracts (are not participating in farming). Given that more than 50% of leased farmland is

under share contracts(including hybrid contracts) in primary crop regions (see Figure 3.1), it must be noted that results obtained from cash rental arrangements may not be used to draw inferences about the entire distribution of benefits between tenant producers and landlords. The significant and negative effect of CV on rental rates confirms that uncertainty is an important determinant of rents. The more risky are farming activities, the lower are the rental payments needed to compensate the landlords.

3.4.2 Disaggregate Program Payments

Table 3.4 presents the estimation results for the cash rental equation when payments are disaggregated according to the different programs. The results from the GLR model show that all four program payments have significant impacts on cash rental rates. In particular, an increase of one dollar in LDP and FDP payments increases the cash rental rates by \$2.16 and \$0.26, respectively. In contrast, disaster payments are found to have a negative effect on cash rents. The results also show that the impact of CCPs on cash rents is very small; an additional dollar obtained from the CCPs only raises cash rental rates by 8 cents. Yet, results from the linear regression framework may suffer from selection biases.

The results from the DMF2 model are quite different. A one dollar increase in LDPs adds \$1.19 per-acre to cash rental rates. Meanwhile, an increase of \$1 in FDPs yields a benefit of \$0.26 to the landlords. In contrast, CCPs are found to exert a significant effect on cash rental rates. An additional dollar of CCPs increases cash rents by and \$0.43 per acre. Disaster payments are also found to have a negative impact on rental rates. This is consistent with many prior studies. For example, Lence and Mishra (2003), and Goodwin, Mishra, and Ortalo-Magné (2010) reported similar findings. This negative relationship may reflect the fact that this particular type of payments is directed at riskier and/or less productive land. The results also indicate that landlords claim \$0.08 in benefits for each additional dollar in returns from the market. This number is relatively stable under both types of leasing arrangements; given risk and landlords' share of production costs have been controlled for.

Table 3.4: Effects of Disaggregate Subsidies on Cash Rental Rates Models (N=48,886)

Cash rent	**GLR**		**Lee**		**DMF2**	
Variable	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**
LDPs	2.16*	0.19	1.32*	0.12	1.19*	0.08
FDPs	0.26*	0.02	0.25*	0.03	0.26*	0.02
CCPs	0.08*	0.15	0.40*	0.06	0.43*	0.07
Disaster	-0.56*	0.21	-0.64*	0.09	-0.57*	0.09
Market Returns	0.06*	0.01	0.08*	<0.01	0.08*	¡0.01
CV	-100.49*	16.25	-156.78*	7.78	-151.37*	5.31
Constant	56.03*	3.8	47.59*	1.42	47.01*	2.53
σ^2	–	–	2890.98*	47.64	3226.43*	189.03
ρ_1	–	–	-0.88*	0.03	0.56*	0.22
ρ_2	–	–	–	–	-0.55*	0.25
ρ_3	–	–	–	–	-0.56**	0.32

*Note:*Asterisks (*) and (**) denote significance at the 5% and 10% level, respectively.

Table 3.5 repeats the previous analysis for share rental rates. The DMF2 estimates indicate that share rents increase by $1.54 for each dollar increase in LDPs. The results show that an increase of one dollar in FDP payments under share contracts increases the share rental rates by $0.52. Meanwhile, an additional dollar in CCPs raises share rental rates by $0.60. It again indicates that legislative restrictions on benefit distribution between contractual parties may be ineffective since landlords capture additional program benefits through rental rates. Both coefficients are larger than the results under cash contracts. It indicates that landlords capture more FDP and CCP benefits under share leasing arrangements than cash contracts, *ceteris paribus*. As in the case of cash contracts, the impacts of disaster payments are also found to be negative.

When referring to net market returns, we find landlords claim $0.09 in benefits for each additional dollar return from the market under share leasing arrangements. In general, the proportion is smaller than estimates reported in the literature (e.g., Goodwin, Mishra, and Ortalo-Magné 2010). The difference likely reflects a compensation for non-land inputs that

Table 3.5: Effects of Disaggregate Subsidies on Share Rental Rates Models (N=48,886)

Share rent	**GLR**		**Lee**		**DMF2**	
Variable	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**
LDPs	1.64*	0.37	1.64*	0.16	1.54*	0.22
FDPs	0.54*	0.08	0.54*	0.06	0.52*	0.06
CCPs	0.52	0.33	0.53*	0.18	0.60*	0.19
Disaster	-0.55	0.53	-0.57**	0.33	-0.43*	0.34
Market Returns	0.09*	0.04	0.09*	0.03	0.09*	0.03
CV	-110.95*	29.75	-108.90*	19.98	-102.54*	22.29
Constant	80.07*	6.39	77.73*	10.3	44.03*	25.62
σ^2	–	–	5582.15*	263.74	6077.55*	709.22
ρ_1	–	–	-0.02	0.07	-0.32	0.31
ρ_2	–	–	–	–	0.07	0.08
ρ_3	–	–	–	–	-0.50**	0.29

*Note:*Asterisks (*) and (**) denote significance at the 5% and 10% level, respectively.

landlords share with tenant farmers, explicit control of risk associated with farming activities, and the recognition of selection biases undertaken in this analysis.

In summary, the results confirm the transaction costs and uncertainty hypotheses, and in particular confirm three effects. First, government subsidies do have large significant effects on rental rates that vary by program and tenure. Second, the LDPs raise the rental rates the most, and then the counter-cyclical payments and direct payments, respectively. This confirms that uncertainty is an important determinant in farmland rental rates. The LDPs are paid based on the current market price and individual yields; the counter-cyclical payments are triggered by current market price but independent of current production; the direct payments are lump-sum income transfer that is independent of market price and yields. Therefore, the LDPs are expected to have the largest risk effect among all three payments, while the direct payments have the smallest, if any. Third, for each program payment, landowners utilizing share contracts capture more benefits than those using cash leasing arrangements.

The explanation for the third effect is that transaction costs for re-contracting are lower

for those landowners utilizing share contracts as compared to those using cash arrangements. Those landowners who are sharecropping with tenants are usually still actively involved in farming. They usually have more information about the government policies and programs. Furthermore, according to the AELOS (1999) survey, compared to those landlords under cash contracts, more sharecropping landlords are holding farm-related occupations or are retired from farming. Thus, landowners under share contracts are able to capture more benefits compared to their cash leasing fellow landowners.

3.5 Conclusion

This study incorporates uncertainty and transaction costs into the Ricardian rent theory and provides new insights beyond the rent determination and subsidy incidence literature. In addition, whether one observes a specific rental rate for an individual farm depends on the leasing arrangements between the tenant and landlord for that farm. We account for non-random selection in the data and analyze the degree of capitalization of government payments using selection bias correction models. The results show that government payments have statistically significant impacts on farmland rental rates. These impacts vary substantially across different leasing arrangements. In particular, we find that landlords capture around 41 percent of aggregate subsidies under cash leases and about 78 percent under share contracts. Disaggregate farm program benefits are also found to have different impacts on rental rates according to the type of program. All three commodity-related program payments (LDPs, FDPs, and CCPs) are found to have positive impacts on rental rates. The results confirm that risk-sharing and transaction costs are important determinants of farmland rents.

Given the increased reliance on contracting in agriculture and the complex mix of leasing arrangements that is emerging in U.S. agriculture, this study should appeal to policy makers who attempt to understand the impacts of government programs under different institutional organizations. We also illustrate the potential biases that may arise from measurement errors in rents and when restricting the subsidy incidence to apply to only cash contracts. Introducing

share contracts (as well as other types of leases) into the analysis is especially important in order to understand the impact of program payments on rental rates. Most existing empirical research that analyzes the distribution of program benefits between landlords and tenants focuses on the cash rental contracts alone. Future studies may find it helpful to consider different types of leasing arrangements.

Chapter 4

The Law of One Price under State-Dependent Policy Intervention: An Application to the Ukrainian Wheat Market

4.1 Introduction

The law of one price (LOP) is one of the fundamental principles of trade theory. It states that homogeneous goods sold in different regions will sell at the same price when expressed in the same currency. The LOP has been considered as an important indicator of market efficiency because it illustrates to what extent markets are linked across space. When referring to economies in transition, the LOP is also an important index of the market liberalization.

A rich body of the empirical economics literature has investigated the LOP among spatially separated markets. Early studies use correlation coefficients and regression techniques to directly test the equality of prices in different regions (e.g., Isard 1977, Richardson 1978, Protopapadakis and Stoll 1986). The results usually do not support the LOP. Some economists

blamed transaction costs (primarily transportation costs) for the failures of prices to converge. They thus proposed a modified version of the LOP which stated that prices of homogeneous products in any two locations should not differ by more than the costs of transferring those goods from one location to the other. As long as trade can take place freely, price shocks in one region can be buffered and price co-movement between regions will be observed. Several empirical studies (e.g., Goodwin 1992 and Michael et al. 1994) have incorporated transaction costs into the analysis and found some supportive evidence.

Modern empirical studies have noticed the nonstationary attribute of the price data and proposed a different framework for testing the LOP. Engle and Granger (1987) point out that, given a pair of first-order integrated series, if there is a linear combination between them which is stationary, the two processes are said to have a long-run equilibrium or simply are said to be cointegrated. Their approach has provided researchers of price transmission (spatial and vertical) with valuable tools for jointly modeling and drawing inferences about the long-run price relationship, together with the short-run adjustments toward the equilibrium.

Some economists (e.g., Goodwin and Piggott 2001) suggest that given cointegration, the short-run adjustments to the equilibrium may not be linear because of the transaction costs associated with arbitrage. Deviations from long-run equilibrium within the transaction cost band will not trigger any adjustment simply because it is not profitable to do so; but deviations that fall outside of the band will trigger trade activities and thus should be mean reverting. This validates the introduction of nonlinear regime-switching autoregressive models and the corresponding (vector) error correction (EC) models into the analysis. Following this idea, an extensive literature has investigated price transmission accounting for nonlinear adjustments by using various versions of regime-switching EC models (i.e., threshold EC, smooth transition EC, and Markov-switching EC). Under this framework, supportive evidence for the LOP have been reported by Lo and Zivot (2001), Sephton (2003), Balcombe, Bailey, and Brooks (2007), Park, Mjelde, and Bessler (2007), and Goodwin, Holt, and Prestemon (2011).

An assumption underlying this transaction-cost version of the LOP, and therefore the use

of error correction models, is that trade is free and open (i.e., without barriers, such as tariffs, quotas, or regional arbitrage interventions). However, trade restrictions do often exist, especially when dealing with agri-food markets. Policy interventions may not only affect short-run dynamic adjustments, but may also alter or even eliminate any long-run market integration under certain conditions. Import/export taxes, subsidies, quotas, certificates, and direct bans, can create a considerable wedge between world and domestic prices, and thus lead to incomplete transmission in prices.

As a measure of transaction costs, direct quantification of policy intervention is difficult. Policy intervention often reflects a state-dependent reaction rather than a constant behavior. For instance, if the objective of policy active exporting country is to stabilize the domestic price, export controls might be triggered when the world price is too high, and subsidies would be applied when the world price is too low. This state-dependent feature indicates a nonlinear relationship between prices. Although the extension of the concept of cointegrating relationship to a nonlinear framework is not new (see Park and Phillips 1999, 2001, Chang and Park 2003, Saikkonen and Choi 2004, Gonzalo and Pitarakis 2006, among others), the procedure to test and estimate nonlinearity in cointegrating vectors is. The policy effects therefore are often investigated indirectly by, on one hand, adding dummies or conducting investigations in different time periods (e.g., Thompson, Sul, and Bohl 2002 and Baffes and Ajwad 2001) and adding a constant term (sometimes together with a proportional term) in the price transmission equations, to account for a fixed policy effect (e.g., Mundlak and Larson 1992).

The objective of this paper is to provide an investigation of the effects of state-dependent policy intervention on spatial price transmission. In pursuing this objective, this study contributes to the literature in three ways. First, we relax the linear cointegrating restriction and allow the long-run equilibrium to be nonlinear based on the state of intervention. Second, we also allow the short-run error correction processes to differ by state, conditional on nonlinearity in the long-run price relationship. Third, we propose an empirical application related to the Ukrainian wheat market. We investigate the price linkages between the Ukrainian and the

world markets. Ukraine is an interesting case study, as it is a typical transition country with active and frequent government intervention. It is also one of the world's top grain exporters. Appropriate investigations of integration between the market and the world market (if any) will provide valuable information for future policy recommendations regarding food security, market efficiency, and trade liberalization.

The remainder of the paper is structured as follows. Section 2 outlines the conceptual framework which introduces the state-dependent policy intervention into the price transmission analysis and develop a simple regime-switching LOP framework. Section 3 provides a brief background on the Ukrainian wheat market and relevant trade policies used over the sample period. Section 4 is dedicated to the empirical procedure, followed by a presentation of the results. Section 5 discusses the policy implication and Section 6 concludes the study.

4.2 Conceptual Framework

The model below builds upon earlier efforts of Mundlak and Larson (1992). We expand their work by introducing the state-dependent feature of policy intervention into the model. It thus allows price linkages to exhibit regime-switching behavior. To empirically investigate the relationship between the domestic and world prices in the presence of policy intervention, Mundlak and Larson (1992) propose the following model

$$P_{it} = P^*_{it} E_t S_{it} \tag{4.1}$$

where P_{it} denotes the domestic price of commodity i at time t. According to the LOP, it can be expressed as a product of the world price P^*_{it} , the nominal exchange rate E_t, and the policy intervention S_{it}. This study does not investigate exchange rate transmission issues and focuses on the linkages between the two prices that are measured in the same currency (US$ in our case), which is a common feature of internationally traded commodities. When rewriting the

price relation equation in the logarithmic form, we obtain

$$p_{it} = p^*_{it} + s_{it} \tag{4.2}$$

where $p^*_{it} = \ln(P^*_{it}E_t)$. Assuming policy depends on world market conditions, Mundlak and Larson (1992) propose the following policy reaction relationship

$$s_{it} = \phi_i + \pi_i p^*_{it} \tag{4.3}$$

where π is a policy reaction index which reflects to what extent the government reacts to world market price. Combining (4.2) and (4.3), for a given homogenous commodity i, the domestic and world price relationship can be expressed in logarithmic form

$$p_{it} = \phi_i + (1 + \pi_i)p^*_{it} \tag{4.4}$$

We expand Mundlak and Larson's (1992) work by letting the policy intervention equation in (4.3) be a state-dependent reaction function which itself is induced by world market conditions

$$s_{it} = \begin{cases} 0 & \text{if } \theta_1 < p^*_{it} < \theta_2, \\ \phi_1 + \pi_{i1}p^*_{it} & \text{if } p^*_{it} \leq \theta_1, \\ \phi_2 + \pi_{i2}p^*_{it} & \text{if } p^*_{it} \geq \theta_2. \end{cases} \tag{4.5}$$

Substituting (4.5) into (4.4), we then obtain the corresponding state-dependent price linkage as

$$p_{it} = \begin{cases} k_i + p^*_{it} & \text{if } \theta_1 < p^*_{it} < \theta_2, \\ \phi_{i1} + (1 + \pi_{i1})p^*_{it} & \text{if } p^*_{it} \leq \theta_1, \\ \phi_{i2} + (1 + \pi_{i2})p^*_{it} & \text{if } p^*_{it} \geq \theta_2. \end{cases} \tag{4.6}$$

In logarithmic form, the econometric specification can be written as

$$p_{it} = (c_1 + \rho_1 p^*_{it})I_1 + (c_2 + \rho_2 p^*_{it})I_2 + (c_3 + \rho_3 p^*_{it})I_3 + \varepsilon_{it} \tag{4.7}$$

where c_i, $i = 1, 2, 3$ are the constant terms that can be interpreted as an overall effect of a set of factors affecting price signals, including transportation costs, the degree of product homogeneity, changes of the consumer or producer price indexes, and the fixed part of policy effects as shown in (4.5), and so on. The term ε_{it} is a stationary disturbance and I_i, $i = 1, 2, 3$ are indicator functions which satisfy the conditions that the world price is within a certain range, or below or beyond a certain threshold. Again, assume the purpose of government intervention is to stabilize the domestic price, as long as the world price is staying within a certain range, let's say, a commodity-specified "open trade band", the government will not (actively) intervene in trade and the open trade assumption holds. However, if the world market price goes outside the band, either by becoming too low or too high, the government will intervene. Under these circumstances, as long as the world price is still within the band, any fluctuations of the world price would not trigger government intervention, thus one can expect a close-to-unity price transmission elasticity from the (conintegrating) regression in (4.7). However, if the world price goes below the lower threshold, export subsidies might be introduced to maintain a relatively high and stable domestic price and to support the domestic producers. In this case, a positive π_1 is expected, thus a greater-than-unity, price transmission elasticity (i.e., $\rho_2 > 1$) is also expected, if the LOP holds. Conversely, when the world price is "too high" and beyond the upper threshold value θ_2 , the government intervenes through the introduction of export taxes, bans, and/or quotas to lower the domestic price. A less-than-unity coefficient ρ_3 would be expected.[1]

[1]It is worthwhile to mention that, in reality, direct government interventions to domestic markets/prices may not occur in developed counties, but are not rare in those less developed counties and economies in transition. Some reasons are: lack of trade and economics knowledge, traditions of mixed and/or planned economy, poor infrastructure system, tight budgets, and less developed social welfare supportive programs.

4.3 Ukrainian Wheat Market and World Food Crisis

Ukraine is the second largest European country after Russia. It became independent when the Soviet Union dissolved in 1991. The economy experienced a large increase in GDP growth after an eight-year recession that immediately followed the dissolution. It is a globally important grain supplier largely due to its endowment arable land. Ukraine has more than 100 million acres of cropland and permanent pasture with fertile soils–approximately 40% of the world's black soils, year-round ice-free ports, and proximity to key import markets in the Middle East, Northern Africa, and the European Union (von Cramon-Taubadel and Zorya 2001). Though grain production suffered from dramatic declines in the first decade following independence, output has considerably increased since then. In marketing year (MY) 2009/10, Ukraine was easily among the world's top three leading grain exporters (after Brazil and Russia). Between 2008 and 2010, Ukraine, together with Russia, exported an average of 29 million tons of wheat annually. This accounted for 21.3% of world wheat exports and was greater than the exports of any of the other major exporters US, Canada, EU-27, and Australia (Goychuk and Meyers 2011).

Although Ukraine is a large grain exporter, it is still plagued by food security issues. As pointed out by von Cramon-Taubadel and Zorva (2001), food consumption of an individual (or a country) does not just depend on its production ability, but more importantly, on his/her endowments, working capacity, and exchange entitlements (i.e., the ability to exchange these endowments for food). Even if a country is a net exporter of food, its vulnerable, low-income groups can still suffer from hunger. In Soviet times, the economy of Ukraine was the second largest in the Union and was an important industrial and agricultural component of the country's planned economy. With the dissolution of the Soviet system, the country moved from a planned economy to a market economy. The transition was difficult, and plunged the majority of the Ukrainian people into poverty. A large part of the population could not afford food, and some had to rely on a subsistence diet of bread and tea (von Cramon-Taubadel and Zorya 2001). As a result, rising food prices are most likely to incite political unrest and violence. Given the

political sensitivity of food prices, combined with Ukraine's history of a planned economy, the Ukrainian government always reacts quickly to the global rise in grain prices. In Ukraine, grain markets are often considered as a "political tool". Both local and central governments control crop and food prices (Brmmer, von Cramon-Taubadel, and Zorya 2009). When world wheat prices soar, the response of Ukrainian government is often populist in nature. The government often accuses traders/speculators driving up wheat prices. As a result, they introduce export certification, export quotas, and fixed bread prices to try to control the market prices.

World food prices increased dramatically in MY 2007/2008, creating a global food crisis. In 2008, U.S. wheat export prices rose from $375/ton in January to $440/ton in March, and Thai rice export prices increased from $365/ton to $562/ton. This came on the heels of a 181% increase in global wheat prices over the 36 months preceding February 2008, and an 83% increase in overall global food prices over the same period (Revenga 2011). Similarly, since July 2010, prices of many crops have risen significantly. World food prices reached a historic peak in January 2011, exceeding prices reached during the food crisis of MY 2007/08. Corn increased by 74%; wheat prices went up by 84%; and sugar prices by 77% (Oxfam online 2011).

Food price crisis caused political and economic instability in Ukraine. In both periods, the initial response of the Ukrainian government to rising food prices was to implement grain export controls, primarily by issuing export quotas. The argument behind these market interventions is that they are needed to guarantee food security and protect domestic consumers from rising international food prices. The first export quotas were introduced in late September 2006. The quota volumes set for the MY 2007/08 were especially low. They virtually banned exports over a certain time period.[2] In July 2008, export quotas were cancelled due to the gradual decreases of world market prices and a large domestic grain harvest in MY 2008/2009. In addition, Ukraine had an obligation to cancel the export restrictions as part of its WTO commitments.[3] In October, 2010, the Ukrainian government again enacted a resolution requiring quotas and

[2]The total export quota in MY2007/08 is 1.2 million tons, compare to a 12.9 million tons net export in MY2008/2009.

[3]Ukraine became the WTO's 152nd member on May 16, 2008.

licenses for exporting grain. While the protectionist policy came under attack from both foreign and domestic observers, the government extended the export grain quotas until June 30, 2011. Moreover, the government used corrupt practices of allocating export quotas and licenses wherein an unknown company Khlib Investbud received the majority share and gained market power in the grain export industry. In place of the quota, contract price export duties of 9% of the contract price were introduced on July 1, 2011 and remained in effect until January 1, 2012. At the same time, although direct government intervention in the grain markets is common in Ukraine when the market price is "too high," the government does not subsidize grain exports when the world market price is low.

Following the above description, we thus propose to use a two-regime policy response model—free trade and policy intervention—based on the world market price–to evaluate the relationship between domestic and world wheat prices. In particular, when the world market price falls below a certain threshold, no significant export controls are triggered. We thus expect a near-unitary price transmission elasticity of the domestic price respect to the world market price, if the LOP holds. However, if the world market price reaches and exceeds the threshold, export controls would be triggered. Accordingly, increases of the world market price would not fully pass-along to the Ukrainian domestic price and a less-than-unity transmission elasticity can be expected.

Based on this information, a two-regime threshold cointegrating regression model is appropriate to model the price linkages. However, the threshold models are based upon the assumption that the transition from one regime to another is abrupt and discontinuous. If threshold models are used to capture the policy-switching behavior, the break between regimes can only be sharp and discontinuous if any policies can be fully carried out instantly without any delay. However, both policy intervention and market adjustment take time and would probably develop gradually for a while before any changes can be made. Therefore the regime-switching behavior of the price transmission is likely to be smooth. A smooth transition conintegrating regression model is thus utilized in the empirical stage. It is important to note that such a

model specification also allows for rapid adjustment, such as that imposed by discrete threshold models.

4.4 Data and Empirical Procedure

This study uses weekly observations for the world market and Ukrainian wheat prices from March 23, 2001 to September 9, 2011. Ukrainian domestic wheat price is measured as ex warehouse price of milling wheat of class III (obtained from Information Agency APK-Inform). The FOB price of wheat (classification other wheats) in Rouen, France (obtained from consulting company HGCA 2009) is used as the world market price for Ukraine. World prices and Ukrainian ex warehouse prices are converted based on the daily exchange rates provided by the European Central Bank into US$ per ton. Figure 4.1 shows the Ukrainian domestic and world wheat price series. Figure 4.2 presents plots of relationship between these two prices. We also plot the relationships between U.S. and German domestic wheat prices and their corresponding world reference prices as a comparison (there was no export/import controls in grain trade activities by these countries during the two food price crisis periods). Visual inspection leads us to suspect a regime-switching pattern in the relationship between Ukraine and world wheat prices. When the prices are low, the correlation coefficient of Ukraine's wheat price with respect to world reference price is larger than when both prices are high. However, we do not observe such switching behaviors for the U.S. and German situations. This suggests an impact on price linkages resulting from government intervention.

We begin by assessing the time series properties of price series using the standard Augmented Dickey-Fuller (ADF) test (Dickey and Fuller 1979) and the KPSS test of Kwiatkowski, Phillips, Schmidt, and Shin (1992). Table 4.1 presents the test results. The ADF tests fail to reject the unit root hypothesis for both price series and the KSPP tests reject the stationarity null for the two series. Meanwhile, test results reject the unit root hypothesis and are not able to reject stationarity for the first difference of price series. Hence, the price series may be considered as

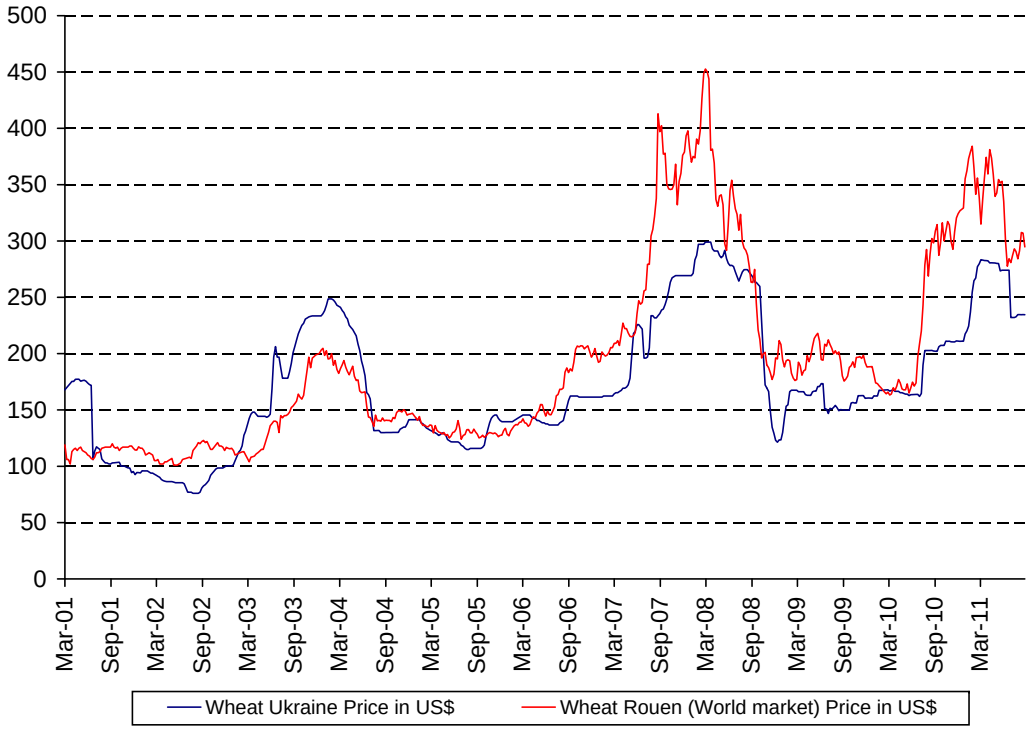
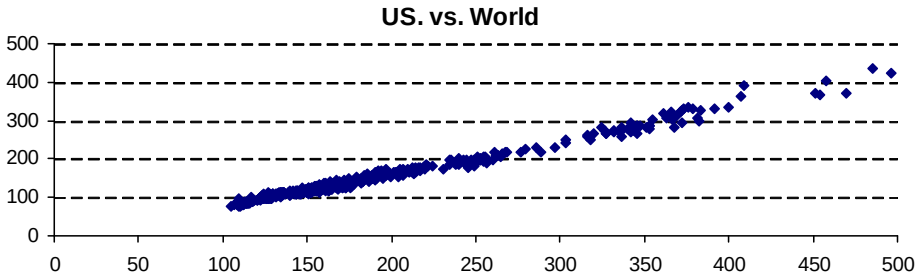
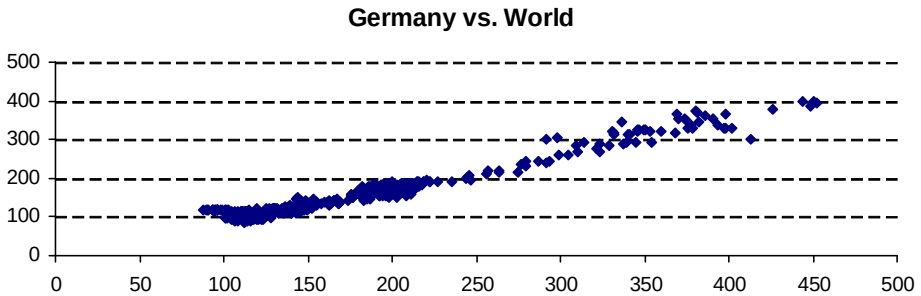


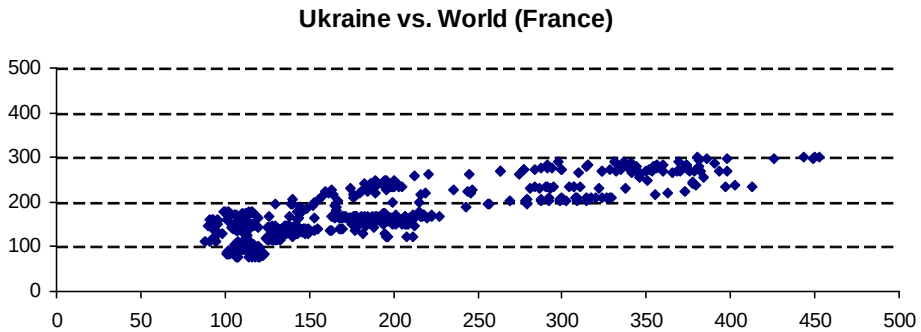
Figure 4.1: Ukrainian and World Market Prices (US$/ton)



(a)



(b)



(c)

Note: For the U.S., the FOB price of hard red winter wheat at the USA Gulf port (HGCA 2009) has been utilized as the relevant world market price for the USA; and for the Germany and Ukraine, the world reference price is the FOB price of wheat (classification other wheats) in Rouen, France (HGCA 2009).

Figure 4.2: Domestic and Its Corresponding World Market Prices.

Table 4.1: Unit Root Tests for Price Data (in natural logarithms)

		World price		Ukraine price		ΔWorld price		ΔUkraine price
Dickey-Fuller								
Single Mean	Lags		Lags		Lags		Lags	
ρ	3	-3.15	3	-6.81	3	-319.74	3	-252.802
Pr $< \rho$		0.64		0.29		(< 0.001)		(< 0.001)
τ_μ	3	-1.19	3	-1.75	3	-9.95	3	-9.23
Pr $< \tau_\mu$		0.68		0.4		(< 0.001)		(< 0.001)
Trend								
ρ	6	-8.54	3	-12.98	6	-319.78	6	-8.54
Pr $< \rho$		0.54		0.26		(< 0.001)		(< 0.001)
τ_μ	6	-2.05	3	-2.66	6	-9.94	6	-253.78
Pr $< \tau_\mu$		0.58		0.25		(< 0.001)		(< 0.001)
KPSS								
Single Mean	6	4.81	6	2.93	6	0.07	6	0.1
Trend	6	0.3	6	0.26	6	0.07	6	0.08

Note: The 10%, 5%, and 1% critical values for KPSS-single mean test are 0.35, 0.46, and 0.74, respectively; and for KPSS-trend test are 0.12, 0.15, and 0.22 respectively.

$I(1)$ processes.[4]

The next step in the empirical investigation is to estimate the relationship between the Ukrainian and world prices. Introduced by Engle and Granger (1987), the concept of cointegration has become a popular tool in the analysis of nonstationary time series. The premise is that, for two nonstationary $I(1)$ series, if there is a linear combination of them which is stationary, then these two series are said to have a long-run equilibrium and thus are said to be cointegrated. This definition leads to interesting interpretations in the price transmission analysis as the prices can then be interpreted to have a stable long-run relationship and can be represented in a vector error-correction framework.

[4]In practice, the ADF and KPSS tests (other unit root and stationary tests alike) often have low power. It is almost impossible to differentiate a difference-stationary series from a highly autoregressive one. Similarly, the differences between a trended series and a difference-stationary series may be extremely difficult to see in small samples. That said, results from these unit root tests are not necessarily to reflect the true properties of the true data generating process. So we need to think of unit root tests as useful but not definitive information. Statistical tests are best used together with economic theory and an understanding of the economy in question.

Empirical implementation involves a two-step procedure for jointly modeling and conducting inferences about the long-run equilibrium together with the short-run adjustment processes towards the equilibrium: 1) estimate the linear equilibrium relationship and test for cointegration; 2) conditional on rejecting the null hypothesis of no-cointegration, test the nonlinearity of residuals, estimate the error correction model (ECM), and investigate how short-run dynamics in the system are influenced by the level, or the sign, of deviations from equilibrium.

Though both economic theories (e.g., market power in supply chain and sticky wage rates in labor markets) and practical economic conditions (e.g., in our case, the state-dependent policy intervention) often imply a nonlinear equilibrium, empirical studies typically only attempt to detect nonlinearity in the adjustment process to the equilibrium while the equilibrium relationship itself has been taken to be represented by a linear regression model.

The development and application of nonlinear cointegrating techniques are still young. Enders and Siklos (2001) propose to test nonlinearity in the residuals of the linear cointegrating vector using a threshold behavior as the alternative hypothesis. The drawback of this approach is that there are no workable approaches to derive a general limiting distribution of this test because the threshold parameters are not identified under the null. Seo (2006) proposes a sup-Wald statistic in the spirit of Davies (1987) to solve the problem, but the procedure is strictly valid only under the assumption that the cointegrating relation is known. Gonzalo and Pitarakis (2006) introduce threshold type nonlinearities within a single equation cointegrating regression model and propose a procedure for testing the null hypothesis of linear cointegration versus cointegration with threshold effects. Krishnakumar and Neto (2009) generalize the estimation and inference procedures of Gonzalo and Pitarakis (2006). However, their threshold cointegrating model requires the threshold/forcing variable to be stationary and ergodic, which may be too restrictive when applying the model to price series, as most commodity data are usually $I(1)$ (Wang and Tomek 2007). For example, in our case, the domestic and world price relationship depends on the world market price, which is a nonstationary series.

Saikkonen and Choi (2004) propose a smooth transition conintegrating (STC) regression

model where regressors are $I(1)$ and errors are $I(0)$. The regressors and errors are allowed to be dependent both serially and contemporaneously. Our approach is based on the STC framework of Saikkonen and Choi (2004), Saikkonen and Choi (2004), and Choi and Saikkonen (2010), and follows the procedure suggested by Engle and Granger (1987). The empirical procedures for analysis of the regime-switching price transmission can be described as follows:

1. Test linear versus STC long-run relationship using the method developed by Choi and Saikkonen (2004);

2. Estimate the STC regression model if linearity is rejected in favor of STC (as in our case), using the method proposed by Saikkonen and Choi (2004); [5]

3. Test stationarity using the residuals obtained from the estimated STC model (Choi and Saikkonen 2010);

4. Test linearity versus nonlinearity for error correction procedures, again using residuals from the estimated STC regression model;

5. Estimate the error correction models, based on the test results from (4), to investigate the dynamic adjustments in the relationship between two prices.

4.4.1 Test Linear Versus STC Long-Run Relationship

Consider a smooth transition cointegrating (STC) model[6]

$$y_t = (\alpha_1 + \beta_1 x_t) + (\alpha_2 + \beta_2 x_t) g(x_t - c; \gamma) + z_t, \quad t = 1, 2, \ldots, T \tag{4.8}$$

where y_t denotes the (logarithmic) Ukrainian wheat price and x_t represents the (logarithmic) world reference price; z_t is a zero-mean stationary error term, α_1 and α_2 are constant terms,

[5] If not, then follow the common practice and estimate the linear cointegration.

[6] Theoretically, the STC model we adopt here requires the independent variable x to be an $I(1)$ series. But for the dependent variable y, it is not necessarily to be $I(1)$. In general, it could be $I(1)$, a higher order of integration, or not integrated at all, depends on the type of the smooth transition function.

β_1 and β_2 are parameters that measure the price transmission elasticity, and $g(x_t - c; \gamma)$ is a smooth transition function of the process x_t , with smoothness parameter γ and threshold value c . The non-linear nature of the model is determined by the transition function. Like other smooth transition autoregressive (STAR) models, the STC can be thought of as a regime-switching model that allows for two regimes, associated with extreme values of the transition function, $g(x_t - c; \gamma) = 1$ and $g(x_t - c; \gamma) = 0$, and where the transition from one regime to the other is smooth. The regime that occurs at time t is determined by the observable variable x_t and the associated value $g(x_t - c; \gamma)$. Different choices for the transition function give rise to different types of regime-switching behaviors. In our study, we use a first-order logistic function as the transition function

$$g(x_t - c; \gamma) = \left[1 + \exp(-\gamma(x_t - c))\right]^{-1} \tag{4.9}$$

The parameter c can be interpreted as the threshold between the two regimes, in the sense that the logistic function changes monotonically from 0 to 1 as x_t increases. When x_t is small (relative to the threshold c), g approaches 0 and the behavior of y_t is given by $\alpha_1 + \beta_1 x_t + z_t$. Similarly, as x_t becomes large, g goes to 1 and the behavior of y_t is then given by $(\alpha_1 + \alpha_2) + (\beta_1 + \beta_2)x_t + z_t$. The parameter γ determines the smoothness of the change in the value of the logistic function and, thus, the smoothness of the transition from one regime to the other. As $\gamma \to 0$, the STC model becomes an AR(p) model. When $\gamma \to \infty$, the regime-switching from 0 to 1 becomes instantaneous at $x_t = c$. Hence, the STC model in (4.8) includes a two-regime threshold autoregressive (TAR) model as a special case. In the Logistic STC model, the two regimes are distinguished by small and large values of the transition variable x_t (relative to c). This type of regime-switching is appropriate in our case, as the relationship pertains to the active or inactive state of policy intervention, which itself is triggered by the level of world market prices. For detailed discussions on the choice of transition functions, the reader is referred to van Dijk, Teräsvirta, and Franses (2002) and Teräsvirta, Tjøstheim, and Granger

(2010).

Testing linearity against the STC specification constitutes a first step towards building the STC models. The null hypothesis of linearity can be expressed as equality of the autoregressive parameters in the two regimes of the STC model in (4.8). That is, $H_0 : \alpha_2 = \beta_2 = 0$, whereas under the alternative hypothesis of H_1: at least one of α_2 and $\beta_2 \neq 0$. The testing problem is complicated by the presence of unidentified nuisance parameters under the null hypothesis. Informally, the STC model constrains parameters which are not restricted by the null hypothesis, but about which nothing can be learned from the data when the null hypothesis holds. The null does not restrict the parameters in the transition function γ and c, but when H_0 holds, the likelihood is unaffected by the values of γ and c. Another attractive alternative might be testing the null hypothesis $H_0' : \gamma = 0$ directly from Equation (4.9). However, under H_0' , the magnitudes of α_2 and β_2 are completely irrelevant. In other words, the values of α_2 and β_2 are unidentified under the null hypothesis when the model is linear. In this case, it is impossible to perform an LM linearity test. Luukkonen et al. (1988) and Granger and Teräsvirta (1993) develop tests that circumvent the problem associated with the presence of nuisance parameters by replacing the transition function with a Taylor series approximation. However, since we are working with cointegrating regressions, and thus with $I(1)$ data, this brings about notable new challenges to the testing problem.

Choi and Saikkonen (2004) develop a nonlinearity test that extends the approaches developed by Luukkonen et al. (1988) and Granger and Teräsvirta (1993), and that can be applied in the context of STC. In particular, their test relaxes the exogeneity requirement for the regressors and follows the common practice in cointegrating regressions and permits both serial and contemporaneous correlations between the regressors and the error term of the model. In order to allow for this feature, the test uses the leads-and-lags approach proposed by Saikkonen (1991) and Stock and Watson (1993) for linear cointegrating regressions.

Following Luukkonen, Saikkonen, and Teräsvirta (1988), Choi and Saikkonen (2004) propose a set of tests based on the first- and third-order Taylor series approximation of the transition

function . The authors argue that a third-order Taylor expansion is superior to a first-order version, since it has more power when β_2 in (4.8) is small. We thus adopt the third-order Taylor approximation and rewrite the transition function as

$$g(x_t - c; \gamma) \approx b\gamma(x_t - c) + d[\gamma(x_t - c)]^2 + h[\gamma(x_t - c)]^3 \tag{4.10}$$

The testing procedure involves estimating the corresponding auxiliary regression using OLS[7]

$$\begin{aligned} y_t &= \alpha_1 + \alpha_2 \Big\{ b\gamma(x_t - c) + d\big[\gamma(x_t - c)\big]^2 + h\big[\gamma(x_t - c)\big]^3 \Big\} \\ &\quad + \beta_1 x_t + \beta_2 x_t b\gamma(x_t - c) + \sum_{j=-K}^{K} \pi_j \Delta x_{t-j} \\ &= \omega + \phi_1 x_t + \phi_2 x_t^2 + \phi_3 x_t^3 + \sum_{j=-K}^{K} \pi_j \Delta x_{t-j} + \eta, \quad t = K+1, \ldots, T-K \end{aligned} \tag{4.11}$$

The null hypothesis of linearity is $\phi_2 = \phi_3 = 0$. The LM statistic is $\tau = \hat{\Phi}'[\hat{\sigma}_\varepsilon^2 (M^{-1})_{xx}]^{-1}\hat{\Phi}$, where $\hat{\Phi} = [\hat{\phi}_2 \;\; \hat{\phi}_3]'$ are the OLS estimates of $[\phi_2 \, \phi_3]$, $\hat{\sigma}_\varepsilon^2$ is the variance estimator based on the residuals of the corresponding OLS estimation constrained by $\phi_2 = \phi_3 = 0$, M is the sample moment matrix for the auxiliary regression, and thus $(M^{-1})_{xx}$ is the element of the inverse of the sample moment matrix associated with $[x_t^2 \, x_t^3]'$. Under the null hypothesis, $\tau \xrightarrow{d} \chi^2(p+1)$, where p (1 in our case) is the dimension of the model. Test results are presented in Table 4.2.

Under all levels of lags and leads (K), the test rejects the null of linearity in favor of the STC framework. We thus use the STC for modeling the long-run relationship for Ukraine and world wheat prices. As a comparison, we also test the linearity of the U.S. and Germany wheat prices with their corresponding world price relationships. Neither of the tests is able to reject the linearity assumption, which suggests STC is not appropriate for the U.S. and German wheat markets. This is consistent with our prior expectation since these two countries

[7]Choi and Saikkonen (2004) argue that because the motivation for using the third-order instead of the first-order approximation is to improve the power of test statistics, they thus suggest using a third-order approximation only for the transition of the intercept term and using the first-order approximation for the transition involving the regressors.

Table 4.2: Linear vs. Smooth Transition Cointegrating Vector Tests

	Ukrainian vs. world market price	United States vs. world market price	German vs. world market price
Lags and Leads		Statistic τ	
$\sum_{j=-K}^{K} \alpha_j \Delta p_j^{wd}$		(3rd order Taylor approx.)	
K=1	12.83	0.88	1.13
K=2	11.99	0.39	1.05
K=3	12.17	0.54	0.87

Note: The tau statistic follows a chi-square distribution with two degree of freedom. The null hypothesis is linear cointegrating vector and the alternative is STC. The critical value is $\chi(2)_{0.05} = 5.99$.

have not implement trade restrictions during the food crisis. In our next step, we estimate the STC relationship for the Ukrainian case. Of course, as always, before we can draw any formal conclusion about the long-run equilibrium, we will need to test the stationarity of the residuals to decide if indeed these prices are cointegrated.

4.4.2 Estimation of the STC Long-Run Relationship

Given that the null hypothesis of linearity has been rejected, our next step is to estimate the STC model. Previous studies (for example, van Dijk, Tersvirta, and Franses 2002 and Enders 2010) usually suggest using a nonlinear least square (NLLS) technique to obtain the estimates of the parameters in (4.8).[8] The estimate of the parameter vector $\theta = [\gamma\, c\, \alpha_1\, \alpha_2\, \beta_1\, \beta_2]$ will satisfy

$$\hat{\theta} = \arg\min_{\theta} Q_T(\theta) = \sum_{t=1}^{T} \left[\tilde{y}_t - y_t(x_t;\theta)\right]^2 \tag{4.12}$$

where $\tilde{y}_t$ is sample observations and $y_t(x_t;\theta)$ is the so-called skeleton of the model given in (4.8). As before, we are working with the STC model where regressors are $I(1)$ and errors are $I(0)$, and the regressors and errors may be dependent both serially and contemporaneously. Saikkonen

[8]Many empirical studies may utilize maximum likelihood methods in application. Under the additional assumption that the errors of Equation (5) are normally distributed, NLLS is equivalent to maximum likelihood. Otherwise, the NLLS estimates can be interpreted as quasi-maximum likelihood estimates.

and Choi (2004) point out that, although the nonlinear least squares estimator from (4.12) is consistent, the asymptotic distribution involves a bias if regressors and error are dependent, which makes the above NLLS estimator inefficient and unsuitable for use in hypothesis testing. They thus propose a GaussNewton (G-N) type estimator that utilizes the NLLS estimator obtained from (4.12) as an initial estimator and expands the model by including leads and lags as extra regressors. Using leads and lags enables the G-N estimator to eliminate the bias and have a mixture of normals distribution in the limit, thereby making it more efficient than the NLLS estimator and thereby suitable for use in hypothesis testing. That said, the estimation procedure is comprised of two steps: to compute the NLLS estimator $\hat{\theta} = [\hat{\gamma}\,\hat{c}\,\hat{\alpha}_1\,\hat{\alpha}_2\,\hat{\beta}_1\,\hat{\beta}_2]$ for equation (4.12) and then to use $\hat{\theta}$ as the initial value and estimate the following augmented STC model

$$y_t = (\alpha_1 + \beta_1 x_t) + (\alpha_2 + \beta_2 x_t) g(x_t - c; \gamma) + \sum_{j=-K}^{K} \pi_j \Delta x_{t-j} + \eta, \quad t = K+1, \ldots, T-K \quad (4.13)$$

The Saikkonen and Choi (2004) approach has provided us with valuable suggestion for obtaining a consistent and unbiased estimates for the STC models. Actually, all methods for nonlinear optimization are iterative: from a starting point θ_0 the method produces a series of vectors $\theta_1, \theta_2, \ldots$ which (hopefully) should converge to θ^*, a global minimum for the given function. If the given function has several (local) minima, the result will depend on the starting point θ_0. Thus, the starting point for estimation is important in the empirical procedure. The Saikkonen and Choi (2004) approach provides a suitable starting point for the second stage G-N estimation. Given that the estimate from the first NLLS stage is the true θ^* for the first NLLS estimation, the second G-N approach supplies the better estimates. We adopt their iterative estimation procedure and utilize a damped G-N method—known as the Levenberg-Marquardt (L-M) method. Given the initial values of the parameters are close to the final optimal values, the L-M method has proved to be more efficient and can almost always guarantee quadratic final convergence.

Also, as just discussed, the estimate results could be sensitive to the initial values of γ and c. van Dijk, Tersvirta, and Franses (2002) thus suggest normalizing the transition function by dividing γ by the sample standard deviation of the transition variable x_t to make γ approximately scale free. We thus replace the transition function 4.9 with a normalized version

$$g(x_t - c; \gamma) = \left[1 + \exp(-\frac{\gamma}{\hat{\sigma}_x}(x_t - c))\right]^{-1} \tag{4.14}$$

Table 4.3 presents the (iterated) L-M estimates of the cointegration models for the linkages between Ukrainian and world wheat markets. Before discussing the results, we need to test the stationarity of the residuals first. We thus conduct a stationarity test utilizing the residuals obtained from the above STC regression in the next session.

4.4.3 Tests for Nonlinear Cointegrating Relationship

In order to test for the existence of this STC relationship, we need to test for the stationarity of the error process z_t. Under the linear cointegration situation, there are two approaches generally adopted to test the stationarity. One approach is to test the null hypothesis of cointegration against no cointegration, whereas the other approach reverses the roles of the null and alternative hypotheses. In the present nonlinear context, the former approach is more convenient. In the latter case, one would need to establish the asymptotic properties of the estimators $\hat{\theta}$ obtained from the STC regression when the error term z_t is $I(1)$. This is difficult as it involves solving the problem of spurious regression in a nonlinear context. Therefore, we adopt the approach developed by Choi and Saikkonen (2010), which tests the null hypothesis that z is a stationary process through the KPSS tests.

However, Choi and Saikkonen (2010) show that the KPSS test statistics depend on the limiting distributions of the estimators and unknown nuisance parameters of the STC model if one uses the full-sample residuals. Tabulating these limiting distributions is therefore impractical. To solve this problem, they develop a test using subsamples of the STC regression residuals.

That is, the KPSS tests are applied using sub-residuals of size b. They have demonstrated that as long as $b/T \to 0$ as $T \to \infty$, where T is the sample size, the tests have limiting distributions that are not affected by the limiting distributions of the full-sample estimators and the parameters of the STC model.

The formula for the test statistic of the sub-residual KPSS test is:

$$C_{NLLS}^{b,i} = b^{-2} s_i^{-2} \sum_{t=i}^{i+b-1} \left[\sum_{j=i}^{t} \tilde{z} \right]^2 \tag{4.15}$$

where s^2 is a consistent estimator of the long-run variance σ_z^2 based on the sub-residuals and $\tilde{z}$ is the residuals from the STC regression. The test statistic formula has the same functional form as the usual KPSS test (Kwiatkowski, Phillips, Schmidt, and Shin 1992). The $\{\tilde{z}\}_{t=i}^{i+b-1}$ represents a block of sub-residuals. The index i denotes the starting point of the sub-residuals and b denotes the size of sub-residuals (i.e., block size).

Choi and Saikkonen (2010) have shown that although the test statistic $C_{NLLS}^{b,i}$ is free of nuisance parameters in the limit, it is likely to have low powers compared to those utilizing full residuals. Thus they propose to do the test along with the Bonferroni procedure. For this, select M tests C_{NLLS}^{b,i_1}, C_{NLLS}^{b,i_2},...,C_{NLLS}^{b,i_M} and defines $C_{NLLS}^{b,\max} = \max\left[C_{NLLS}^{b,i_1}, \ldots, C_{NLLS}^{b,i_M}\right]$. These M tests have the same block size but use different starting points $i_1, i_2, \ldots, i_M$.

For a given block size b, the number of sub-residual-based tests M and the starting points $i_1, i_2, \ldots, i_M$ are:

1. $M = [T/b]^*$, where $[\cdot]^*$ denotes the smallest integer greater than or equal to x;

2. The starting points $i_1 = 1, i_2 = T - b + 1, i_3 = b + 1, i_4 = T - 2b + 1$, and the block size b is chosen by using the minimum volatility rule.

More specifically, the procedure of the minimum volatility method can be summarized in four steps as follows:

Step 1: Calculate the integers b_{small} and b_{big} with a restriction $b_{small} < b_{big}$. In this study, we

follow Choi and Saikkonen (2010) and use the rule $b_{small} = [T^{0.7}]$ and $b_{big} = [T^{0.9}]$, with $[x]$ denoting the integer part of x.

Step 2: Calculate test statistics $C_{NLLS}^{b_{i-m},\max}, \ldots, C_{NLLS}^{b_{i+m},\max}$ for each integer b_i in the interval $[b_{small}, b_{big}]$. Here m is a small positive integer. We again follow Choi and Saikkonen (2010) and set $m = 2$.

Step 3: Calculate the standard deviation of $C_{NLLS}^{b_{i-m},\max}, \ldots, C_{NLLS}^{b_{i+m},\max}$ obtained in step 2 and denote it as SD_i.

Step 4: Choose the block size that gives the minimum of SD_i over $b_i = b_{small}, \ldots, b_{big}$.

The test statistic in Equation (4.15) can then be used to test the stationarity. Table 4.4 reports the statistics calculated based on the above procedures. The results in Table 4.4 indicate that the null of stationarity is not rejected at the 5% level. We therefore conclude the Ukraine and world market wheat prices are cointegrated via a smooth transition mechanism.

Table 4.3: Estimates of the Smooth Transition Cointegrating Models.

	STC, no lags and leads			STC, with lags and leads		
Parameter	Estimate	Approx Std Err	Approx $Pr > \|t\|$	Estimate	Approx Std Err	Approx $Pr > \|t\|$
γ	3.87	1.73	0.03	3.23	1.18	< 0.01
c	5.21 ($185)	0.05	< 0.01	5.17	0.05	< 0.01
α_1	-0.86	0.49	0.08	-1.45	0.5	< 0.01
α_2	2.13	0.67	< 0.01	2.77	0.69	< 0.01
β_1	1.14	0.1	< 0.01	1.19	0.1	< 0.01
β_2	-0.44	0.13	< 0.01	-0.57	0.13	< 0.01
π_{t+1}^0				-0.44	0.45	0.34
π_{t+1}^1				0.59	0.54	0.27
π_t^0				-0.48	0.45	0.29
π_t^1				-0.19	0.54	0.73
π_{t-1}^0				-0.22	0.46	0.64
π_{t-1}^1				-0.54	0.56	0.33
$\sum(y_t - \hat{y}_t)^2$	8.21			7.54		

Table 4.4: Sub-residual KPSS Stationarity Test Results

$C_{NLLS}^{b,\max}$			
Statistics	Standard Deviation	Block Size	M
0.456	0.0004	276	2

Note: The 10%, 5%, and 1% critical values for KPSS-single mean test are 0.35, 0.46, and 0.74, respectively. The lag length used to calculate the s^2 in this table is 8. We have tried different lag length based on different selection criteria. In general, for $lags \geq 6$ (i.e., $4(T/100^{0.25})$), the results do not reject the stationarity hypothesis at a 10% level.

After the test of stationarity, now we go back to the discussion of the results from STC regression. The results presented in Table 4.3 are consistent with the institutional background and with our conceptual framework. When comparing the results from STC models with and without lags and leads, we find no significant difference. This may indicate that regressor-error dependence is not an issue in our sample set. Equation (4.16) is based on the STC with no lags and leads. It reveals the STC long-run equilibrium relationship for the two prices.

$$\begin{aligned} \hat{y}_t &= \begin{cases} -0.86 + 1.14x_t & \text{if } g = 0, \\ 1.27 + 0.70x_t & \text{if } g = 1, \end{cases} \\ \text{and} \quad g(x_t - c; \gamma) &= 1/\big\{1 + \exp\big[-3.87(x_t - 5.21)/0.16\big]\big\}. \end{aligned} \tag{4.16}$$

The results confirms a regime-switching behavior in the long-run relationship between Ukrainian and world prices, based on the level of world market prices. The estimated threshold value for the transition variable is 5.2 in logarithms, or \$185. When the world price is below the threshold of \$185/ton, the transmission elasticity of domestic price with respect to the world price tends to be 1.1. The two markets are closely integrated. This provides evidence that when the world price is not "too high", no active export control has been triggered, and thus that price changes or shocks in the world market can be fully transmitted to the Ukrainian market. At the same time, when the world market is "too high" (from the perspective of the

Ukrainian government), and exceeds the threshold level of $185, the relationship between the two markets gradually switch to another regime and the transmission elasticity decreases to 0.70. This reflects the effects of trade interventions on price transmission. The two food crisis periods, with strict export controls, belonging to this regime. The fitted price relationship is also presented in Figure 4.3. Finally, it is quite interesting to see what happens when the wheat price is between two regimes. In that case, an increase of one unit in the world market price will only partially be passed along to the domestic market while a similar decrease in the world price will fully be transmitted to the domestic market. The domestic growers under such a situation are thus worse off from price increases as compared to the potential benefit they might gain from the same price increase in the world market, all else being equal.

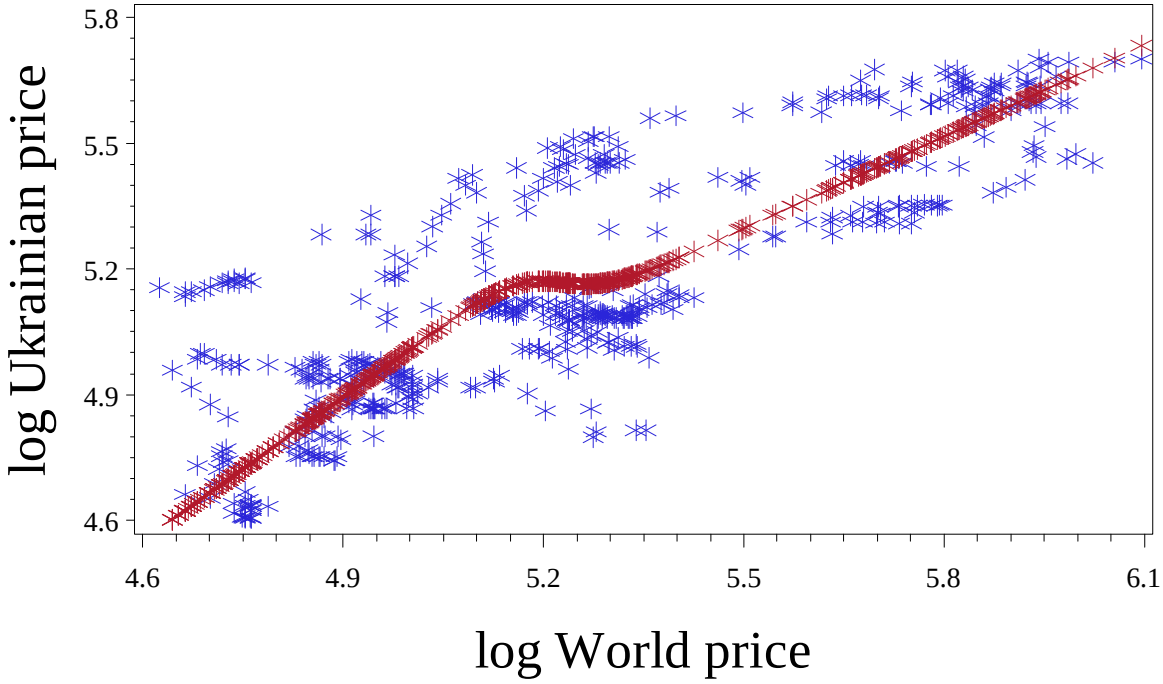


Figure 4.3: Smooth Transition Cointegrating Model Fit

4.4.4 Short-run Dynamic Adjustment

The transaction cost version of the LOP provides justification for using the momentum threshold autoregressive (M-TAR) or Exponential STAR types of regime-switching models which allow the adjustment behavior to be asymmetric inside and outside the transaction cost band. A standard two-parameter and three-regime M-TAR model when applied to the deviations from equilibrium, can be expressed as

$$\Delta z_t = \begin{cases} \phi_1 z_{t-1} + \varepsilon_1 & \text{if } z_{t-1} < \theta_1, \\ \phi_2 z_{t-1} + \varepsilon_2 & \text{if } \theta_1 < z_{t-1} < \theta_2, \\ \phi_3 z_{t-1} + \varepsilon_3 & \text{if } z_{t-1} > \theta_2. \end{cases} \tag{4.17}$$

where z_{t-1} is the previous deviation from long-run equilibrium. An equivalent vector error correction representation of (4.17) can be written as

$$\Delta y_t = \begin{cases} \sum_{i=1} \alpha_{1i} \Delta y_{t-i} + \sum_{j=1} \beta_{1j} \Delta x_{t-j} + \varphi_1 z_{t-1} + e_1 & \text{if } z_{t-1} < \theta_1, \\ \sum_{i=1} \alpha_{2i} \Delta y_{t-i} + \sum_{j=1} \beta_{2j} \Delta x_{t-j} + \varphi_2 z_{t-1} + e_2 & \text{if } \theta_1 < z_{t-1} < \theta_2, \\ \sum_{i=1} \alpha_{3i} \Delta y_{t-i} + \sum_{j=1} \beta_{3j} \Delta x_{t-j} + \varphi_3 z_{t-1} + e_3 & \text{if } z_{t-1} > \theta_2. \end{cases} \tag{4.18}$$

In (4.17) and (4.18), the interval $[\theta_1, \theta_2]$ defines an asymmetric transaction cost band within which arbitrage is not profitable. The ϕ_i can be interpreted as the speed-of-adjustment parameter. In this specification, deviations from the long-run cointegrating relation trigger error correcting movements in prices when the deviations fall outside of the band. If $z_{t-1} < \theta_1$ or $z_{t-1} > \theta_2$, then error correction follows a stationary AR(1) process and trade or arbitrage between markets is profitable. However, we are investigating a situation which is one-sided because of the nature of policy interventions. There is no transaction cost band, only one-sided transaction costs for trade from the domestic market to world market, it is thus more appropriate to utilize a two-regime threshold model to investigate the error correction process.[9]

[9]Due to severe winter-kill, the smallest harvest in more than 45 years was produced in marketing year (MY)

We begin by conducting a linearity test for the residuals which is based on Hansen's (1999) self-exciting threshold autoregressive (SETAR) approach. SETAR models with one regime (which shrinks to a linear AR model) and two regimes are

$$z_t = \alpha_1 z_{t-1} + e_t, \text{ and} \tag{4.19}$$

$$z_t = \alpha_1 z_{t-1} I_1(\gamma) + \alpha_2 z_{t-1} I_2(\gamma) + e_t, \text{ respectively.} \tag{4.20}$$

where z_t here is the predicted residuals from STC regression, $I(\gamma)$ is an indicator that $I_i(\gamma) = 1$ when ith regime occurs and γ is the threshold. The estimates of α_1 and α_2 are obtained from OLS along with the sum of squared residuals, denoted as SSR_2 . The threshold has been chosen when the estimation of (4.20) gives the minimum sum of squared residuals ($SSR_2^{\min}$), alternatively, $\hat{\gamma} = \arg\min SSR_2(\hat{\gamma})$. The search over all possible values of the threshold is restricted to the values of z_{t-1} that lie between the 15th and 85th percentiles. Let SSR_1 denote the sum of squared residuals from (4.19) and $SSR_2^{\min}$ denote the minimum sum of squared residuals from (4.20), which is the chosen threshold model, and the F-statistic can be constructed as

$$F_{12} = n(SSR_1 - SSR_2^{\min})/SSR_2^{\min} \tag{4.21}$$

where n is the observations associated with the values of z_{t-1} that lies between 15th and 85th percentiles (i.e., $n = 0.7(T-1)$). The F statistic has a non-standard asymptotic distribution under the SETAR hypothesis, so conventional critical values are not appropriate. Hansen (1999) showed how to obtain the appropriate critical value F_{12}^* using a bootstrapping procedure. The method involves resampling the data utilizing the residuals obtained from the above threshold model and for each bootstrap sample, searching the optimal threshold as we did before and calculating the test statistic F_{12}^*. This is repeated a large number of times (1000 in our case) to find the bootstrap distribution and thus the p-value for that representing the percentage of

2003/2004 in Ukraine, which made Ukraine a wheat importer in that year. This one exception aside, Ukraine is a pure wheat exporter in our sample time period.

test statistics for which the test taken from the estimation sample exceeds the observed test statistics. This method will be applied to the full sample residuals obtained from the STC regression.

Before we proceed with the error correction procedures, an issue is worth discussing. As we are dealing with an economy in transition with government intervention, it might be inappropriate to use a very short time period as a unit of reaction time span when investigating the error correction procedures. The model identification should reflect the reality that market reactions and adjustments may occur with a lag, especially for a transition economy. We therefore also consider multi-week differentials as a unit change in the "first-order difference" identification. That is, we identify the first-order of the error term as $\Delta_k z_t = z_t - z_{t-k}$ and its corresponding short-run response $\Delta_k y_t = y_t - y_{t-k}$, $k = 1, 2, \ldots, k_{\max}$ where k is the number of weeks that define a unit change, with $k = 1$ as the special case usually applied in the literature. We then use the same SETAR method to test linearity utilizing the following equation,

$$z_t = \alpha_1 z_{t-k} + e_t \text{ vesus } z_t = \alpha_1 z_{t-k} I_1(\gamma) + \alpha_2 z_{t-k} I_2(\gamma) + e_t. \tag{4.22}$$

We test linearity of the residuals, with $k_{\max} = 4$. When we estimate models using k greater than one as a unit change, some observations are lost. To accurately compare the alternative models with different k value, the sample time period should be kept fixed (at $T - k_{\max} - \text{lags}$). Otherwise, we would be comparing the performance of the models over different sample periods. The results are presented in Table 4.5. Model selection is based on AIC and SBC.

The Hansen tests do not reject the linearity hypothesis for all values of k. We then estimate the corresponding linear error correction models $\Delta_k y_t = \sum_{i=1} \alpha_i \Delta y_{t-i} + \sum_{j=1} \beta_j \Delta x_{t-j} + \lambda z_{t-k} + \varepsilon_t$ with k from 1 to 4. Both AIC and SBC indicate that for each group of residuals, the case $k = 1$ fits the best. We thus conclude the domestic price does respond to a deviation in a short time period. But as we will see, domestic price adjustments under both open trade and the active intervention regimes also respond to lagged price changes.

Table 4.5: Residual-based Tests of Linearity, Hansen F test

	Bootstrap P-value for Hansen 1999 test		
	Full sample residuals	Residuals from STC regime 1 (world price <= \$185)	Residuals from STC regime 2 (world price > \$185)
$\Delta_k z_t = z_t - z_{t-k}$			
$k = 1$	0.93	0.49	0.43
$k = 2$	0.9	0.45	0.42
$k = 3$	0.9	0.44	0.44
$k = 4$	0.92	0.46	0.42

Table 4.6: Estimates for Linear ECMs

	Residuals from STC Estimation	
Variable	Coefficient	Std. Err.
z_{t-1}	-0.04	0.009
Δy_{t-1}	0.23	0.042
Δy_{t-2}	0.21	0.066
Δy_{t-3}	0.21	0.067
Δx_{t-1}		
Δx_{t-2}	-0.14	0.053
Δx_{t-3}	-0.13	0.054
Half-life	17.7wks	
AIC	-294.92	
SBC	-252.42	
Observation	542	

The results of error correction models when $k = 1$ are presented in Table 4.6. We exclude the statistically insignificant regressors. The results indicate that the adjustment of Ukrainian domestic price responds to the deviation from equilibrium and the lagged own price shocks and the world market price shocks. The results also suggest that short-run dynamics of the Ukrainian prices react to the shocks from the world market with a lag of two and three weeks, but do not respond to shocks that occurred in the prior week. This was expected for an economy like Ukraine which has less developed market infrastructure and potentially high adjustment costs. To provide a little more intuition on the adjustment processes, we present the deviation half-lives for each group in Table 4.6.[10] Adjustment towards the long-run equilibrium–takes place through changes in Ukrainian domestic wheat price alone–with half of the deviation from the equilibrium being corrected requiring nearly 18 weeks. The slow adjustment speed again may be a reflection of the institutional and economic characteristics of Ukrainian grain markets.

Meanwhile, Ukraine is a major grain exporter. With intense world competition for commodities such as wheat, there is a legitimate concern that Ukraine may have some control over world market prices, at least in the short run. Some researchers and policy makers suggest that the export control in Ukraine is not only harming domestic markets and producers, but is also creating negative impacts on world grain markets and thus exacerbating the food crisis. We thus investigate whether world market prices also respond to deviations. We simultaneously estimate the error correction models for domestic and world prices using a seemingly unrelated regression technique. The results indicate that both under the full sample and the subsample situations, the world price does not respond to disequilibrium between the two markets. We also find that lagged changes in Ukrainian prices have no effect on adjustments of the world price. The results thus indicate that adjustments toward the long-run equilibrium take place through changes in Ukrainian prices alone. The result is consistent with the idea that the world market is large relative to Ukraine. This is also consistent with the 2008 World Bank report

[10] Deviation half-lives, given by $\ln(0.5)/\ln(1+\lambda)$, where λ is the OLS estimate of $\Delta y_t = \sum \alpha_i \Delta x_i + \sum \alpha_i \Delta y_i + \lambda z_{t-1} + \varepsilon_t$, represent the period of time (in weeks) required for one-half of a deviation from equilibrium to be eliminated.

suggesting that Ukraine's market power alone is limited in the long run and Ukraine would be ill-advised to attempt to exercise this influence by deliberately reducing exports in the long run in an effort to drive up world market prices and thus export revenues. However, our finding should not be interpreted as evidence that Ukraine has absolutely no effect on the world market price, but price shocks in Ukrainian domestic markets alone do not push the world market prices to make adjustments accordingly. Further investigation of the influence on the supply side would be helpful to understand the effects of Ukrainian trade interventions on world grain markets.

4.5 Policy Implications

This paper uses a more flexible STC model to investigate the price relationship between Ukraine and world markets, taking the state-dependent trade intervention into account. We find that a long-run equilibrium relationship exists and varies according to the world price. When the world price is below a certain threshold, Ukrainian and world markets are well integrated. However, when the world price exceeds the threshold level, it triggers active interventions, and the two markets are less integrated. In particular, only 70% of changes of the world price would be transmitted to Ukrainian price. In other words, 30% of potential export revenues are lost, other things being equal. The regime-switching long-run equilibrium provides a framework to estimate and predict the potential domestic export loss under certain scenarios. For example, consider the average world price during January 2010 to September 2011, \$213.8/ton. Assume further that the reduced export quantity is 10 million tons. Then, a 50% increase in world price will result in a \$320.7 ($213.8 \times 10 \times 0.5 \times 0.3$) million revenue loss for the domestic growers. What makes the domestic producers lose even more is that on the input side, rising energy prices in recent years have influenced the costs of production and trade. Production revenues have been further reduced for Ukrainian producers as their production depends on importing energy from Russia and fertilizers from international markets.

In summary, the two-regime long-run price transmission results indicate that the Ukrainian

market itself is well integrated with the world market. However, continuous government interventions in trade activities can cause significant losses for the domestic producers in the long run.

To give a more complete story of the impacts of export controls on Ukrainian domestic economy, we briefly discuss some important findings from other studies, in the hope of offering some suggestions for future policy recommendations. According to recent studies (e.g., von-Cramon and Raiser 2006 and Brmmer, von Cramon-Taubadel, and Zorya 2009), although the stated purpose of these export controls is to help those low income consumers, these are the people who actually benefit the least from the quota. First, wheat prices contribute only a certain percentage to the final bread price. The impact of lower wheat prices on the prices of meat and dairy is quite limited. Second, though wheat prices have been somewhat controlled, prices for flour and bread have actually risen since the introduction of the quote in 2006. Instead of the poor consumers, flour millers and animal feed producers, whose profit margins increase as a result of falling grain prices on the domestic market, are the main beneficiaries of the quota.

The quota system has also imposed big losses on international agribusiness companies and traders that have invested billions of dollars in farming, trading, storage, processing and export facilities. Furthermore, some have argued that the government used corrupt practices of allocating export quotas and licenses which resulted in unfair and nontransparent competitions in the trade market which hurt the majority of traders.

The future policy implications, in this paper are in accordance with von-Cramon and Raiser report (2006) which argues:

> "The quota system is ineffective (does not reach the poor), inefficient (imposes large cost for very limited gain), and led to corruption. The suggestion is therefore to abolish the quota system as soon as possible .... Alternative measures including the use of means tested cash transfers need to be considered to protect the poor from rising food prices."

4.6 Conclusion

The extent and magnitude of policy intervention on price transmission, when allowing for state-dependent attributes, on price transmission offer valuable information on price linkages and market integration. More generally, state-dependent or regime-switching long-run price equilibrium can result from other factors, such as state-dependent exchange rate pass-through, market power, and/or asymmetric information. It is thus a useful extension and generalization of linear cointegration approaches for modeling price transmission that has appeared in the literature. However, the development of nonlinear cointegration techniques and their application to price transmission are both novel and deserve more attention.

Chapter 5

Conclusions

This study investigates the effectiveness (reaching the goals) and efficiency (how much it costs to reach the goals) of government intervention in agricultural markets. Two specific focuses are on: 1) farm program payments as a way to help farmers; and 2) government intervention as a way to reduce domestic food prices, protecting the poor domestic food consumers.

This study evaluates the effectiveness and efficacy of farm program payments as a tool to help farmers and considers how and to what extent landowners capture any program benefits intended to tenant operators. The first essay investigates the effects of program payments on farmland contract choices. Farmland leasing arrangements are seen not only as a balance between efficient risk-sharing and appropriate incentives to discourage moral hazard, but also as a reflection of program benefit redistribution. The theoretical results confirm that governmental and legal restrictions on benefit sharing between contracting parties are largely ineffective and induce offsetting contractual rearrangements. However, changing/switching contracts is only one form of offsetting contractual rearrangements. There are other possible non-monetary offsetting activities such as increasing the tenant supplied inputs, decreasing other non-land landowner inputs, and/or tenant's provision of farmland maintenance services. These may be applied as compensations to landowners' shares of benefits, when legal restrictions exist on landowners' eligibility for direct receipt of payments. Future work may find it helpful to

investigate these other forms of compensation rearrangements to present a complete story of benefit distributions between landowners and tenant operators.

The second essay evaluates the subsidy incidence in farmland rental rates. From a theoretical aspect, this work extends previous work by introducing uncertainty and transaction costs simultaneously into the traditional Ricardian rent theory. The uncertainty and transaction costs version of Ricardian theory is able to explain the three puzzles arise from the findings of existing studies. In the empirical procedure, the selection issue is corrected by utilizing the method developed by Dubin and McFadden (1984) and Bourguignon, Fournier, and Gurgand (2007). The results confirm the transaction costs and uncertainty hypotheses.

This analysis incorporates uncertainty and transaction costs into the Ricardian rent theory and should help provide new insights in the rent determination and subsidy incidence literature. The results shall also provide policy makers some insight about the effectiveness relative to targeted to non-farming landowners instead of farm operators. Additional insight is offered about the efficiency. The results indicate that current support programs are inefficient as they cause extra costs for the government to distribute the program payments, and for landowners and tenants to negotiate the redistribution of benefits.

This study on the Ricardian rent theory is focused on the rent determinants. Future work may find it helpful to extend the current work to the analysis of optimal inputs, production, and farm size, and the impacts of government subsidies on these aspects under uncertainty and transaction costs.

The third essay investigates state-dependent government intervention on price transmission between domestic and international markets, with the empirical application to the Ukrainian wheat market. Price transmission has been regarded as an important indicator of market efficiency. If two markets are integrated, the price differential of homogeneous goods sold in these two regions shall not be greater than the transaction costs. When referring to transition economies, price transmission is also an index of market liberalization, which indicates whether or not the domestic market is connected with the rest of the world.

Prior studies in the literature on price transmission mainly utilized the linear cointegration technique associated with non-linear error correction adjustments. These studies usually test linear long-run price relationship versus no existing long-run relationship. As explained in the third essay, the justification for assuming a linear long-run relationship is free and open trade/arbitrage. However, open trade assumptions often do not hold when dealing with economies in transition, where government intervention is a common practice.

This essay relaxes the open trade assumption and introduces state-dependent government intervention into the analysis. In the empirical stage, it adopts a more flexible approach of a smooth transition cointegrating regression to allow the price relationship to be nonlinear. The results indicate that there is regime-switching behavior in the long-run relationship between the Ukrainian and world markets, conditional on the world price. The threshold value is USD $185/ton. When the world price of wheat is below the threshold, the transmission elasticity of the domestic price with respect to the world price is approximately unity. However when the world price is above the threshold, the transmission elasticity drops to 0.7. The results suggest that the Ukrainian wheat market is well integrated into the world market. However, government intervention causes significant long-term losses for the Ukrainian producers. The results also indicate that adjustments toward the long-run equilibrium take place through changes in the Ukrainian domestic price alone; the world wheat price does not react to the Ukrainian price shocks.

The world food crises and price peaks both posed challenges and presented opportunities for Ukraine, a net grain exporter with a significant exploitable yield gap, and one of the few countries in the world in a position to significantly increase net exports and make up for emerging deficits elsewhere (World Bank 2008). With appropriate policies and investments, Ukraine could significantly increase its grain production and gain global market share in an environment of rising global demand. However, the Ukraine government seems to fail in seizing the opportunity. Measures to reduce domestic prices are sometimes easier to implement and budget, but they have serious disadvantages as they reduce the export revenues and production incentives

for domestic producers. They also impose big losses on international agribusiness companies and traders that have invested billions of dollars in farming, trading, storage, processing and export facilities.

Seizing this opportunity would require a shift in policies and corresponding increases in private and public investments. A transparent, predictable, and market-oriented policy framework would both increase the effectiveness and efficiency of public expenditures and reduce uncertainty, and thereby increase private investments. A key priority is to reduce direct interference in agricultural markets and eliminate the sudden and unpredictable steps to regulate quantities and prices that make investment and producing in Ukrainian agriculture much less attractive for investors than it could be.

Finally, when referring to the technique adopted in this study, regime-switching, long-run price relationships can result from other factors besides government intervention, such as less developed transportation infrastructure, market power, state-dependent exchange rate pass-through, and asymmetric information. It is thus a useful extension and generalization of linear cointegration approaches for modeling price transmission that has previously appeared in the literature. Future studies may find it helpful to develop and/or adopt more flexible forms of nonlinear equilibrium and structural analysis.

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