Time Series Analysis of a Principal-Agent Model to Assess Risk Shifting and Bargaining Power in Commodity Marketing Channels

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Abstract

We apply the classic agency model to investigate risk shifting in an agricultural marketing channel, using time series analysis. We show that if the principal is risk-neutral and the agent is risk-averse instead of risk-neutral, then a linear contract can still be optimal if the fixed payment is negative. Empirical results for the Dutch potato marketing channel indicate that while fixed payments to farmers (agents) have decreased over time, even to negative levels, the incentive intensity has approximately doubled, and the risk premium the farmers ask for has remained considerable. These results imply that risk has shifted from wholesalers, processors, and retailers to farmers; we argue that this shift could be the consequence of chain reversal, i.e., the transformation of the traditional supply chain into a demand-oriented chain.

Key words: risk shifting, agency theory, commodity marketing channel, chain reversal, time series analysis.

Introduction

Marketing firms that convert raw farm products into finished consumer goods by performing a set of marketing services, such as collection, cleaning, processing, transportation, and retailing (see Helmberger and Chavas, p. 134) have become much larger than farms. Hence, risk shifting has become an important topic of study for agricultural economists and policy makers. In spite of marketing co-operatives, concern is growing that the increasingly large processors and supermarket chains will be able to dictate the terms of trade and transfer the market-level risk to farmers (e.g., Weaver and Kim).

However, if marketing firms can dictate the terms of trade, they will do so to maximize profit. Transferring risk to farmers, who have fewer opportunities to spread risk compared with marketing firms and therefore find it more costly to bear, simply reduces the gains from trade. In contrast, marketing firms would prefer to bear the risk themselves (reducing risk-bearing costs) and extract the gains from this by lowering the price they pay to farmers. Consequently, if marketing firms do transfer market-level risk to farmers, there must be another reason for doing so than mere risk aversion. In this paper we argue that the classic agency model (e.g., Gibbons; Furubotn and Richter; Milgrom and Roberts; Valimaki) provides such another possible reason. Using sector-level, time-series data we outline and empirically illustrate how the usefulness of this model for indicating risk shifting in a food supply chain can be tested.

Originating in economics literature, agency theory has been the backbone of research on corporate governance (Jensen and Meckling; Fama and Jensen; Schleifer and Vishny). It has been applied to, amongst others, budget control in business research (Demski and Feltham), domestic franchising (Rubin; Mathewson and Winter; Brickley and Dark), retail sales compensation (Eisenhardt), and supplier–distributor relationships (Lassar and Kerr). In this paper we apply the agency theory to assess risk shifting in a commodity marketing channel. By using sector-level, time-series data we take a more indirect approach than Knoeber and Thurman, who also applied the agency model to assess risk shifting, but used contract-specific information instead of the widely available data we use here.

The classic model in agency theory is based on the concept of the principal–agent relationship. The agent performs a task for the principal, and the principal values the agent’s output and pays him compensation, as specified in a contract. To generate the output required and/or desired by the principal, the agent has to put in effort. As well as depending on the effort invested, an agent's output also depends on a random component: unexpected events that are beyond his control. While the principal is observing the agent's output, he does not usually have access to the know-how necessary to be able to make the agent’s effort; but even if the principal does get hold of the necessary know-how, he does not have the ability to interpret it. This information asymmetry in the principal–agent relationship is not a problem *per se*. However, it does become a problem when principal and agent have or develop different goals, creating a moral hazard on the part of the agent in the supply of effort. Therefore, if an agent is risk-averse, preferring a certain reward over an uncertain one, to obtain an optimal relationship with the agent the principal might consider a contract that allows for a trade-off between incentives and insurance.

Receiving a fixed salary, independent of the output realized, would provide the agent with full insurance but no incentive. Receiving a percentage of the output value obtained by the principal would give the agent full incentive, yet no insurance. We may hypothesize that the optimal contract lies somewhere between these extremes, consisting of a fixed payment plus a bonus rate of the value received by the principal for the agent's output. Such a *mixed share-wage contract* or *share contract*, is consistent with Stiglitz’s theory from tenancy literature, in which the distribution of the output in a sharecropping context is based on the trade-off between the landlord’s (i.e., principal’s) need to provide both incentives and insurance to his tenants (i.e., agents). This trade-off is the core of the principal–agent problem and provides a useful framework from which Knoeber reviews the literature on agricultural contracting.

The agency model offers a possible explanation for why marketing firms (i.e., the principal) wish to transfer risk to farmers (i.e., the agent), in spite of the higher risk-bearing costs. These higher risk-bearing costs might not outweigh the higher profits the supply chain achieves when farmers are given more incentives to meet the delivery conditions that enable marketing firms to increasingly produce high value-added products in addition to the mainstream homogeneous products. This phenomenon, whereby traditional supply-oriented chains are transformed into demand-oriented chains, can be denoted as "chain reversal" (cf. Boehlje's "industrialization of agriculture"). Chain reversal has been growing in importance now that consumer food markets in the western world have become saturated, international competition is growing by the day, and agri-food companies must concomitantly meet the rising demand for product differentiation and deal with the stiffer competition in their markets. On top of this, consumers and governments expect improvements in production quality and environmental care.

Given that the marketing firms are eclipsing the farmers because of the need to produce more products with greater added value, it is important to note that although the fixed payment can be thought of as equivalent to the reservation wage (i.e., the wage that an agent receives for an alternative job without risk), the classic agency model shows that a Pareto-optimal solution is not inevitable (e.g. Valimaki, p. 35). Upon reflection, solutions with a negative fixed payment can be Pareto optimal. In such cases, the agent's degree of risk aversion allows for a *mixed share-rent contract*. This entails the agent paying a fixed amount to the principal for the opportunity to perform for the principal, in exchange for a percentage of the total value that the principal receives for the agent's actual output. In these cases, the agent has no insurance, despite his risk aversion. Such a contract implies shifting the risk from the marketing firm to the farmer, to increase the latter's incentive — possibly to involve the farmer more in the investments of the marketing firm that has to develop products that better satisfy consumer needs.

In line with the classic agency model, we have chosen a linear contract, because it corresponds to real-world settings. Holmstrom and Milgrom have shown that the optimal compensation scheme for providing incentives over time to an agent with a constant absolute risk aversion is a linear function of the end-of-period results, such as revenues, costs, or profits. This result is based on the fact that a linear contract provides more uniform incentives. In contrast, if, for instance, we consider the annual output as the result of many small daily actions performed by the agent, a non-linear contract may create unintended or non-uniform incentives for the agent in the course of the year, depending on the agent's performance so far (Gibbons).

Below, we will outline the classic agency model and its consequences for risk shifting and incentive transfer. We will then explain how the model can be applied to time-series data. Subsequently we will present an empirical application of the time-series-based principal–agent model, using data from the Dutch supply chain for ware potatoes. Finally, we will discuss the main conclusions of our analysis and propose an avenue for future research.

The Classic Agency Model

Performance in the classic model of principal and agent is assumed to satisfy

(1) $x = h + \varepsilon = E(p|I)e + \varepsilon$

where x is the value obtained by the principal for the agent's actual performance, h the actual amount of effort of the agent (e) valued at the output price (p) for which both principal and agent form the same rational expectation conditional on their common knowledge (I), and ε are the events in the performance process that are beyond the agent's control (i.e., "noise"). The random term ε is normally distributed, with zero mean and variance σ^2.

The costs incurred by the agent when performing for the principal are described by a cost function $C(e)$, such that $dC/de > 0$ and $d^2C/de^2 > 0$, i.e., cost is a convex function of e. For ease of demonstration, but without loss of generality for the main conclusions we have yet to draw, we adopt the following specification

(2) $C(e) = 0.5ce^2$

where c is a positive parameter.

The principal pays the agent a compensation w according to the linear function

(3) $w = \alpha x + \beta$

where αx and β are the variable (i.e., uncertain) and fixed (i.e., certain) compensation components, respectively, and α represents the output-value sharing rate, such that $0 \leq \alpha \leq 1$. The function in (3) is referred to as a linear incentive contract if $\alpha > 0$. The magnitude of α measures the strength of the incentives. Absence of incentives, i.e., $\alpha = 0$, reduces (3) to a fixed-wage contract. A mixed share-wage contract is obtained if $0 < \alpha < 1$ and $\beta > 0$.

In the classic agency model, the principal is assumed to be risk-neutral, while the agent is risk-averse. This assumption is based on the observation that the principal can usually diversify, while the agent cannot. The agent's utility function is

$$(4) \qquad U(w, e) = -\exp\{-r[w - C(e)]\}$$

where $r > 0$ is the agent's coefficient of constant absolute risk aversion (henceforth CARA and implying $r = -[d^2U/de^2]/[dU/de]$). Consequently, a principal trying to maximize his expected payoff will solve

$$(5) \qquad \max_{e,\alpha,\beta} E(x - w)$$

subject to

$$(5a) \qquad E(-\exp\{-r[w - C(e)]\}) \geq U(\overline{w})$$

and

$$(5b) \qquad e \in \arg\max_{e} E(-\exp\{-r[w - C(e)]\})$$

where $\overline{w}$ is the certain monetary equivalent, so that (5a) represents the agent's participation constraint and (5b) reflects the agent's incentive compatibility constraint.

Let us first consider (5b). If we assume that the agent's net payoff $w - C(e)$ is a normally distributed random variable, then the certainty equivalent $\hat{w}$ of $w - C(e)$ under CARA preferences, i.e.

(6) $$U(\hat{w}) = E\{U[w - C(e)]\}$$

has a particularly simple form, namely

(7) $$\hat{w} = E[w - C(e)] - 0.5r\text{var}[w - C(e)]$$

where the difference between the mean of the random net payoff, i.e. $E[w - C(e)]$, and its certain equivalent $\hat{w}$ is referred to as the risk premium: $0.5r\text{var}[w - C(e)] = E[w - C(e)] - \hat{w}$. Substituting (7) in (6), the resulting expression for (6) in (5b), and working out $E[w - C(e)]$ and $\text{var}[w - C(e)]$, shows that the optimization problem of the agent is equivalent to

(8) $$\max_{e} \{\alpha E(p|I)e + \beta - 0.5ce^2 - 0.5r\alpha^2\sigma^2\}$$

which yields

(9) $$\alpha E(p|I) = ce$$

Equation (9) is called the incentive constraint and must be satisfied by any feasible contract. It says that the agent will select the amount of effort he inputs in such a way that his marginal gains from more effort, i.e., $\alpha E(p|I)$, equal his marginal personal cost of effort, i.e., ce.

Inserting (9) into the participation constraint (5a) yields

$$(10)\qquad [\alpha E(p|I)]^2/c + \beta - 0.5[\alpha E(p|I)]^2/c - 0.5r\alpha^2\sigma^2 = \overline{w}$$

from which the following expression for the fixed compensation β results

$$(11)\qquad \beta = \overline{w} + 0.5r\alpha^2\sigma^2 - 0.5[\alpha E(p|I)]^2/c$$

Substituting the expressions for e, see (9), and β, see (11), into (5), where $E(x - w) = E(p|I)e - \alpha E(p|I)e - \beta$, as can be derived from (1) and (3), the principal solves

$$(12)\qquad \max_{\alpha} \{\alpha[E(p|I)]^2/c - [\alpha E(p|I)]^2/c - (\overline{w} + 0.5r\alpha^2\sigma^2 - 0.5[\alpha E(p|I)]^2/c)\}$$

of which the first-order condition yields

$$(13)\qquad \alpha = 1/(1 + rc[E(p|I)]^{-2}\sigma^2)$$

Equation (13) can be referred to as the incentive intensity principle and shows that since r, c, $E(p|I)$ and σ^2 are positive, the optimal incentive parameter α is between zero (full insurance) and one (full incentive). Furthermore, α is smaller if the agent is more risk-averse (r is higher), if the marginal cost of effort increases more quickly (c is higher), if the marginal gains of effort increases less quickly ($E(p|I)$ is lower), or if there is more uncertainty in production (σ^2 is higher).

Now that the optimal incentive parameter has been determined in (13), the fixed part of the agent's compensation can be derived by substituting (13) into the participation constraint (11), giving

$$(14)\qquad \beta = \overline{w} + 0.5(r\sigma^2 - [E(p|I)]^2/c)/([1 + rc[E(p|I)]^{-2}\sigma^2]^2)$$

Equation (14) reveals that β should not necessarily be positive since $r\sigma^2 - [E(p|I)]^2/c$ can be smaller than zero, such that $|0.5(r\sigma^2 - [E(p|I)]^2/c)/([1 + rc[E(p|I)]^{-2}\sigma^2]^2)| > \overline{w}$. Moreover, this situation may occur while still having $r\sigma^2 > 0$. In other words, the classic agency model allows for a contract in which the principal obtains $x - w = (1 - \alpha)x - \beta$, where a negative β represents the lump sum of x (i.e., rent) received by the principal and $(1 - \alpha)x$ is the variable amount assigned to the principal, leaving the agent with a variable compensation of αx minus the lump sum taken by the principal. Such a contract is called a mixed share-rent contract and provides the agent with no insurance, even though the agent is still risk-averse. Why do marketing firms shift risk to the farmers instead of profiting from bearing the risk themselves? A plausible explanation is that chain reversal becomes necessary if farmers obtain more incentives to accommodate investments that enable marketing firms to react promptly to the increasingly varying consumer demands in a saturated market with increasing worldwide competition. Our objective in this article is therefore to find out how the classic model in agency theory can be applied to time-series data in order to find empirical evidence of risk shifting in the marketing channel as a possible result of the purported chain reversal.

Econometric considerations

The solutions of the game theory model in the previous section are given by the expressions for α in (13) and β in (14). The unknowns in the expression for α are r, c, $E(p|I)$ and σ^2. If we consider these unknowns as constant parameters over time, then α is a constant as well. In order to impose a minimum of time invariance restrictions, let us ignore that $E(p|I)$ varies over time. According to the incentive constraint given by (9), α

is equal to $ce/E(p|I)$. Although we may consider c as time-invariant, this cannot be imposed on e. Hence, in terms of time-series variables, the incentive constraint implies that α varies with time:

(9') $$\alpha_t = ce_t/E(p_t|I_{t-1})$$

where the index $t = 1, \ldots , n$ refers to observations through time.

For annual data, as used in the empirical part of this research, it can typically be assumed that σ^2, i.e., $\mathrm{var}(x - h)$, is constant in the food supply chain where the farmers are the agents and the marketing firms the principals[1]. Consequently, in order to comply with the time-varying behavior of α, the other time-varying coefficient in (13) must be r:

(13') $$\alpha_t = 1/(1 + r_t c[E(p_t|I_{t-1})]^{-2}\sigma^2)$$

From this and the fact that $\overline{w}$ can be considered to vary with time as well, it can also be expected that β varies with time:

(14') $$\beta_t = \overline{w}_t + 0.5(r_t\sigma^2 - [E(p_t|I_{t-1})]^2/c)/(1 + r_t c[E(p_t|I_{t-1})]^{-2}\sigma^2)^2$$

Now given that w and x are also time-varying variables, substituting (13') and (14') into (3) and using

(15) $$r_t = ([E(p_t|I_{t-1})]^2/c\sigma^2)(\{[E(p_t|I_{t-1})]^2 - ch_t\}/ch_t)$$

as can be derived from (1), (9') and (13'), we obtain the following equation

(16) $$(w_t - \overline{w}_t - 0.5h_t) = ch_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2}$$

in which c is the single unknown parameter. Before c, as parameter of interest, can be estimated, it should first be identified (cf. Ackerberg and Botticini). If $(w_t - \overline{w}_t - 0.5h_t)$ and $h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2}$ are stationary, then the estimation model

$$(16') \quad (w_t - \overline{w}_t - 0.5h_t) = ch_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2} + u_t$$

in which u_t is an unobserved component, does not typically allow for simple OLS estimation, because $h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2}$ and u_t could well be correlated, in particular with h_t included on both sides of (16'). This problem, however, vanishes when $(w_t - \overline{w}_t - 0.5h_t)$ and $h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2}$ are co-integrated (Engle and Granger). But if these variables, as well as $(w_t - \overline{w}_t)$ and h_t, are stationary, then we may test for the absence of simultaneity bias by performing the omitted variable version of the Hausman test, as in

$$(16'') \quad (w_t - \overline{w}_t) = \lambda_1 h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2} + \lambda_2 h_t + \gamma_1 \hat{v}_{1t} + \gamma_2 \hat{v}_{2t} + u_t^*$$

to first test the null hypothesis $\gamma_1 = \gamma_2 = 0$, i.e., $h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2}$ and h_t are exogenous, by an F test, where $\hat{v}_{1t}$ and $\hat{v}_{2t}$ are the residuals of a bivariate VAR(k) for $h_t(x_t - h_t) \times [E(p_t|I_{t-1})]^{-2}$ and h_t, with k being much smaller than the sample size. If the null hypothesis cannot be rejected, we can test the restriction $\lambda_2 = 0.5$ by testing for the absence of h_t in the regression of $(w_t - \overline{w}_t - 0.5h_t)$ on $h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2}$ and h_t.

Suppose that we have been able to estimate c. Then, from (9'), we obtain the estimate of α_t. Next, we can derive r_t from (15), and then β_t from (14'). Finally, substituting α_t and β_t in (3), w_t can be estimated as

(3') $\tilde{w}_t = \alpha_t x_t + \beta_t$

and compared with the actual values of w_t. This comparison evaluates the validity of the model. If it is valid and the empirical model shows a situation in which β_t has been decreasing over time from a positive to a negative value, whereas r_t has always remained positive, we can conclude that although farmers are risk-averse, marketing firms still find it optimal to increase farmers' rent instead of reducing the risk farmers have to be compensated for. This allows us to hypothesize that marketing firms need farmers in the marketing channel for more than just supplying the primary produce: as sales and profit tend to become a responsibility of the chain as a whole in reversed chains, marketing firms also need farmers to finance some of the activities they want to initiate (or they want farmers to initiate) to successfully process and market the final consumer goods. By way of example, the empirical case of the Dutch ware potato chain outlined in the next section shows that farmers have increasingly become involved in storing the raw potatoes they produce.

Empirical application

Every year, some eight million tons of potatoes are produced in the Netherlands, mainly on family farms. About half are ware potatoes, approximately 20 percent are seed potatoes, while the remaining 30 percent are potatoes grown for starch. Most ware potatoes are sold to wholesalers. A negligible amount is sold directly by the farmer to the processor or retailer (De Graaf; Smidts). The basic marketing problem facing wholesalers is how to optimize the supply of potatoes in terms of time (storage), quantity and quality (assembly and sorting), and place (transport), so as to meet the requirements of the different users.

Most of the wholesale trade has become concentrated in relatively few hands, as the major users, particularly the large retailers, processors and export markets, demand large quantities with tight specifications which only the larger wholesalers can meet. Because of this development in the market, the need has arisen to procure potatoes before harvest, and hence a number of different arrangements to do so have emerged. The most important include fixed-price contracts and pooling contracts (e.g., Young; Smidts).

The fixed-price contract involves selling a net amount of potatoes at a fixed contract price. This marketing strategy entails transferring the entire price risk from the farmer to the wholesale company. In the pooling-contract system, the potatoes delivered by the farmers are sold by wholesalers throughout the season. The resulting gross returns from these sales, minus the wholesalers' expenses, are distributed across the producers, proportional to the amount of potatoes delivered. The reason non-fixed price arrangements have been adopted is because wholesalers wish to retain their core suppliers by offering them contracts that bear some relation to the market price. Note that this complies with the concept of chain reversal. Our empirical results will shed light on the growing importance of the non-fixed price contracts in the Dutch chain for ware potatoes.

For our empirical analysis of the Dutch ware potato marketing system, Statistics Netherlands provided us with annual data over the period 1946 – 1996, for the following variables: the farm and retail prices (Euro/kg) of ware potatoes, both deflated by the consumer price index (1990 = 1.00), the area planted (1000 ha), the yield per hectare (100 kg/ha), and the rent price of land (Euro/ha), deflated by the consumer price index.

From these variables, we derive the following variables of interest. First, the output value at consumer prices (billion Euro), i.e., x_t, is computed as the retail price times the yield per hectare times the area planted (divided by 10^4). Next, to compute h_t, i.e., the expected output value at consumer prices, we fit the retail price by a univariate AR(3) model and consider this fit to be the expected retail price, i.e., $E(p_t|I_{t-1})$, under bounded rationality (e.g., Pesaran; Roumasset, Boussard and Singh)[2]. The yield per hectare clearly shows a positive linear trend, so we use the fit of the linear trend as a proxy for the expected yield per hectare. The expected retail price times the expected yield per hectare times the area planted (divided by 10^4) gives h_t (billion Euro). Lastly, w_t (billion Euro) is computed as the farm price times the yield per hectare times the area planted (divided by 10^4), and for $\overline{w}_t$ (billion Euro), we take the rent price of land times the area planted (divided by 10^6). In the computation of x_t, h_t and w_t ($\overline{w}_t$), we divide by 10^4 (10^6) each time, to ensure uniformity in the units of measurements of the components that made up each of these variables.

Before estimating c in (16'), we first investigate the order of integration of the time series of $(w_t - \overline{w}_t - 0.5h_t)$ and $h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2}$. The graphs of these two time series display a downward trend from which it is difficult to decide whether or not these series are trend stationary. However, Johansen's co-integration test (Johansen and Juselius; Osterwald-Lenum) rejects all hypotheses according to which the rank of matrix Π is not full in the model

$$(17) \qquad \Delta X_t = \Pi X_{t-1} + \sum_{j=1}^{k-1} \Gamma_j \, \Delta X_{t-j} + \mu_t + \varepsilon_t$$

where $X_t = [(w_t - \overline{w}_t - 0.5h_t), h_t(x_t - h_t)\{E(p_t|I_{t-1})\}^{-2}]'$, μ_t captures the deterministic terms, and $\{\varepsilon_t\}$ is Gaussian white noise[3]. The test results are presented in Table 1, where the *trace* statistic has been computed for the case where the linear trend is restricted to be included only in the cointegrating space and $k = 1$, as selected by the Akaike Information Criterion (AIC) for a VAR in levels with a linear trend and pre-specified upper bound of order six. We also tested for the absence of the linear trend in case rank(Π) = 2 and found a value of 25.98 for the likelihood ratio test statistic. The asymptotic distribution of the test statistic is $\chi^2(2)$, and its 95% quantile equals 5.99. Thus, the value of the test statistic is highly significant. Based on these results, we conclude that $(w_t - \overline{w}_t - 0.5h_t)$ and $h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2}$ are trend stationary.

[INSERT TABLE 1 ABOUT HERE]

We now estimate c in (16') by the following regression model

(16''') $(w_t - \overline{w}_t - 0.5h_t) = c_0 + c_1 t + ch_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2} + u_t$

where the deterministic component $c_0 + c_1 t$ is considered as an extension of the cost function specification in (2):

(2') $C(e_t) = 0.5c[h_t/E(p_t|I_{t-1})]^2 + c_0 + c_1 t$

However, before we are allowed to use the estimate of c, obtained from applying OLS to (16'''), we first have to find out whether $h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2}$ and u_t are uncorrelated. For this, we apply the Johansen test to check for the trend stationarity of $(w_t - \overline{w}_t)$, $h_t(x_t -$

$h_t)[E(p_t|I_{t-1})]^{-2}$ and h_t, as is required when applying the omitted variable version of the Hausman test, as in

$$(16'''') \quad (w_t - \overline{w}_t) = \kappa_0 + \kappa_1 t + \lambda_1 h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2} + \lambda_2 h_t + \gamma_1 \hat{v}_{1t} + \gamma_2 \hat{v}_{2t} + u_t^*$$

where $\hat{v}_{1t}$ and $\hat{v}_{2t}$ are the residuals of a bivariate VAR(...) for $h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2}$ and h_t, as selected by the AIC.[4] The results in Table 2 are for $k = 1$ and show that $(w_t - \overline{w}_t)$ and h_t can also be considered to be trend-reverting. So, next, we estimate the parameters in (16'''') and test the restrictions $\lambda_1 = \lambda_2 = 0$. The p value of the F test is 0.82 and hence, we conclude that $h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2}$ and h_t are exogenous. Moreover, after omitting $\hat{v}_{1t}$ and $\hat{v}_{2t}$, we cannot reject the restriction $\lambda_2 = 0.5$ either (p value = 0.07). Therefore, we now arrive at estimating (16''').

[INSERT TABLE 2 ABOUT HERE]

The OLS estimates in (16''') are $c_0 = 0.172$ (t value = 2.83; p value = 0.01), $c_1 = -0.013$ (t value = –5.65; p value = 0.00), and $c = 0.056$ (t value = 4.39; p value = 0.00). The negative coefficient of the linear trend complies with the cost-reducing technological advances in agriculture. Furthermore, the $R^2 = 0.61$, the Jarque-Bera statistic testing for normality of the residuals has a p value of 0.08, and the F version of the LM statistic testing for the absence of first-order (fourth-order) autocorrelation in the residuals has a p value of 0.054 (0.085) and the CUSUM test does not find parameter instability. From these diagnostic test results and the results of the specification tests with regard to (16''''), we conclude that $\{u_t\}$ is Gaussian white noise and uncorrelated with $\{h_t(x_t$ –

$h_t)[E(p_t|I_{t-1})]^{-2}\}$. In what follows, we graphically show and discuss the relevant variables in the model.

[INSERT FIGURES 1- 4 ABOUT HERE]

Using the estimate of c we obtain the following graph of α_t from (9'), see Figure 1. The graph shows a negative trending pattern between 1949 to 1965, according to which α_t decreases from 0.56 to 0.30. After that, α_t slightly rises to 0.48 in 1980. Thereafter, α_t shows a much more positive trend and increases to 0.86 in 1996. This sharp rise in α_t implies a decrease in r_t, see (13'), as shown in Figure 2, i.e., less risk aversion among farmers. Moreover, at the same time, the risk premium $0.5r_t\alpha_t^2\sigma^2$ decreases from an average of about 0.40 billion Euro in the 1970s to 0.15 billion Euro in 1996, while $C(e_t) + 0.5r_t\alpha_t^2\sigma^2$ seems to perform reasonably well as an expectation of w_t, conditional on the information set available at time $t - 1$, see Figure 3. Nevertheless, in spite of the result that the farmers are still asking for a positive risk premium – one which, compared with the total production costs $C(e_t)$ and compensation w_t, is considerable –, the fixed compensation β_t, computed as

(14") $$\beta_t = \overline{w}_t + 0.5(r_t\sigma^2 - [E(p_t|I_{t-1})]^2/c)/(1 + r_tc[E(p_t|I_{t-1})]^{-2}\sigma^2)^2 - 0.013t + 0.172$$

where $-0.013t + 0.172$ originates from the extended cost function and estimated in (16'''), declines steadily, becoming negative during the 1970s. In Figure 4 its decline thereafter is clearly shown. The Figure also reveals that the model explains w_t quite well for many of the years studied. Conditional on this, we conclude that risk has been shifted to the potato growers, such that from receiving a lump-sum payment, they have now lost all this payment to the marketing firms, even though the risk premium they asked for is still considerable. It is the marketing firms, however, who have been able to compensate for some of their expenses without risk. They have done so by steadily increasing the proportion of output value at consumer prices: from –20 percent in the early 1950s to 60 percent in the mid 1990s, with an average annual increase of 2.9 percent since 1975 (see Figure 5).

[INSERT FIGURE 5 ABOUT HERE]

Given that the marketing firms can be assumed to be risk-neutral, we might have expected them to behave differently and bear all the risk themselves, so as to reduce the risk-bearing costs of the farmers.[5] This would also be in their own interest, since it would allow them to lower the price they pay to the farmers. The above results, however, suggest that farmers play a crucial role in the process of chain reversal, as they seem to be the ones who have to finance some of the activities wanted by marketing firms in order to meet consumers' needs and demands in the increasingly saturated consumer food market, amidst growing competition and globalization. The fact that growers have become more involved in storing potatoes is a clear example of this development.

Conclusion and Discussion

In this paper we apply the classic agency model to shed light on risk shifting and chain reversal in a commodity marketing channel. The model involves a mixed share-wage/rent contract with a time-varying fixed wage/rent and output value sharing rate. It can be tested on sector-level time series data that are widely available. To perform this test, we have outlined how to take the time-series properties of the data into account, in relation with the simultaneity problem regarding the parameter of interest to be estimated. If the model complies with the data, it can be used to detect risk shifting as a possible indication of a marketing channel changing from a traditional supply-oriented chain into a demand-oriented chain. The estimates may then reveal a situation where the fixed wage eventually becomes a rent, while the risk premium the agents demand remains considerable.

Our empirical application to the Dutch marketing channel of ware potatoes has shown that risk has been shifted from the purchasers of potatoes to the potato growers. Having received 20 percent of the retail sales as a fixed payment in the early 1950s, the average decline of 2.9 percent per annum since 1975 means that potato growers now have to pay a rent equivalent to 60 percent of the retail sales to their purchasers. This, despite the fact that the growers are still demanding a hefty risk premium. The rise in the output-value sharing rate implies that farmers' attitudes to risk have changed over time, i.e. they have become less risk-averse. This finding contributes to the debate on whether risk attitude is a stable concept (e.g., Pennings and Garcia).

The method used in this paper differs from the procedure in Knoeber and Thurman. Knoeber and Thurman already knew which contracts were used in the course of time. Using simulation methods along with production and payment data from a panel of individual farmers, they measured the risk shift between principal and agent, based on these contracts. By estimating the parameter of interest, our method is also able to reveal how the contracts have changed over time. However, for this purpose it uses only sector-level data on prices and quantities that are widely available.

Knoeber and Thurman applied their method to the U.S. broiler industry, where the agents are the growers and the principals are the integrator firms. They concluded that risk had shifted from the agents to the principals. In contrast to their study, our application to the Dutch marketing channel of ware potatoes includes the retail sector among the principals. Our results show risk shifting from principals to agents. This is consistent with the fact that retailers have become more powerful than upstream stages in the channel (e.g., Kuiper and Meulenberg). As a result, they can force processors and wholesalers to better fit the needs and wants of the consumer which, in turn, processors and wholesalers can only do with the farmers' support. Nevertheless, the difference in the results shows the importance of extending the classic agency model to more than two stages in the marketing channel. It also indicates a future avenue of research: the possibility of testing for different strategic interactions between these stages.

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Table 1. Testing Rank(Π) When $\mathbf{X}_t = [(w_t - \overline{w}_t - 0.5h_t), h_t(x_t - h_t)\{E(p_t|I_{t-1})\}^{-2}]'$

Rank(Π)	*Trace* Statistic	5 % Critical Value[a]	1% Critical Value[a]
= 0	98.02**	25.32	30.45
≤ 1	36.80**	12.25	16.26

[a] Critical values are from Osterwald-Lenum, Table 2*
** denotes significant at the one percent level

Table 2. Testing Rank(Π) When $\mathbf{X}_t = [(w_t - \overline{w}_t), h_t(x_t - h_t)\{E(p_t|I_{t-1})\}^{-2}, h_t]'$

Rank(Π)	*Trace* Statistic	5 % Critical Value[a]	1% Critical Value[a]
= 0	149.88**	42.44	48.45
≤ 1	74.93**	25.32	30.45
≤ 2	30.65**	12.25	16.26

[a] Critical values are from Osterwald-Lenum, Table 2*
** denotes significant at the one percent level

Figure 1. The Output Value-Sharing Rate (α_t)

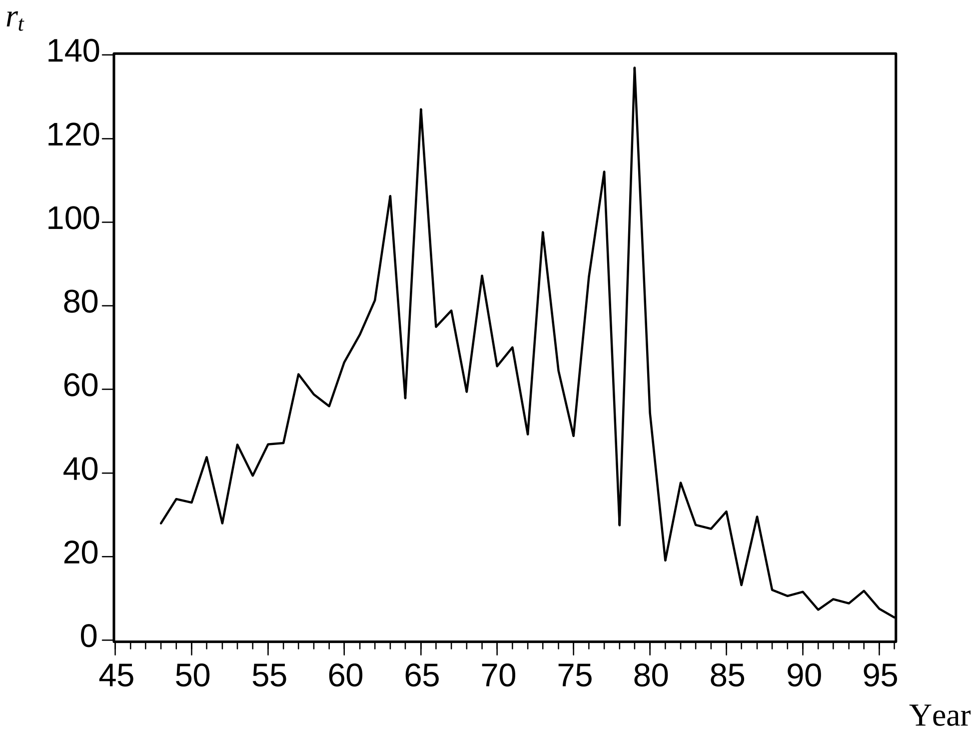

Figure 2. Constant Absolute Risk Aversion Coefficient (r_t) (1/billion Euro)

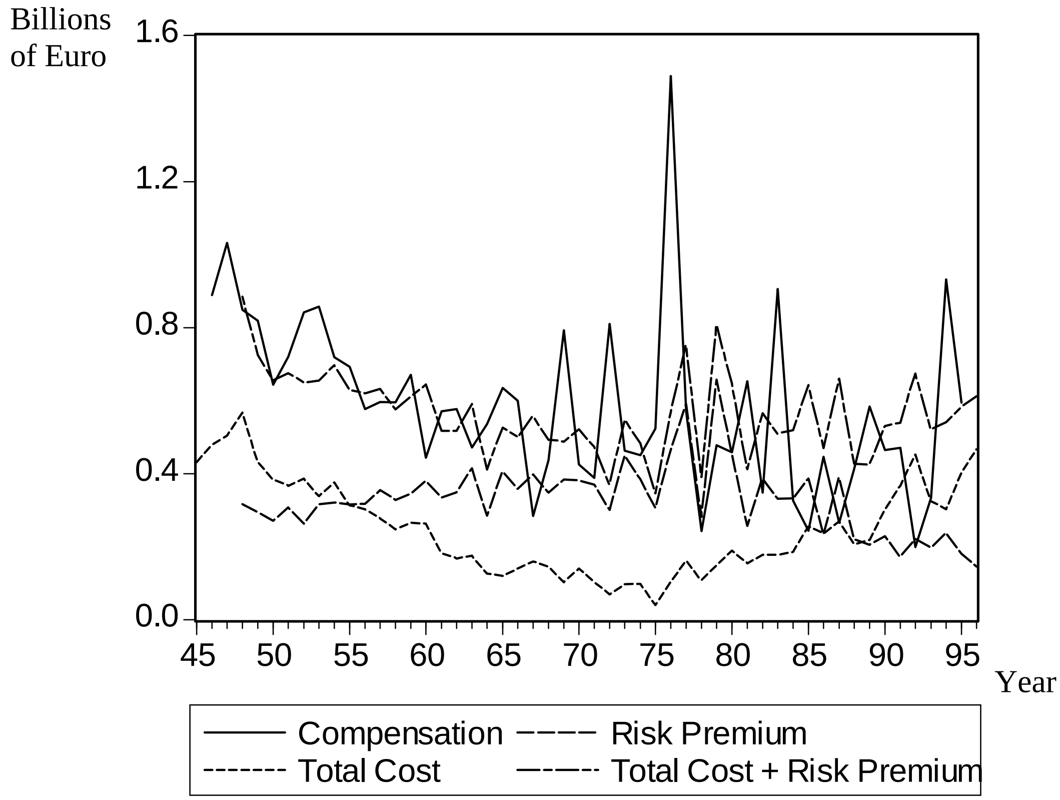

Figure 3. Compensation for Farmers (w_t), Farmers' Total Cost (C_t), Risk Premium ($0.5 r_t \alpha_t^2 \sigma^2$), and the Sum of the Farmers' Total Cost and Risk Premium

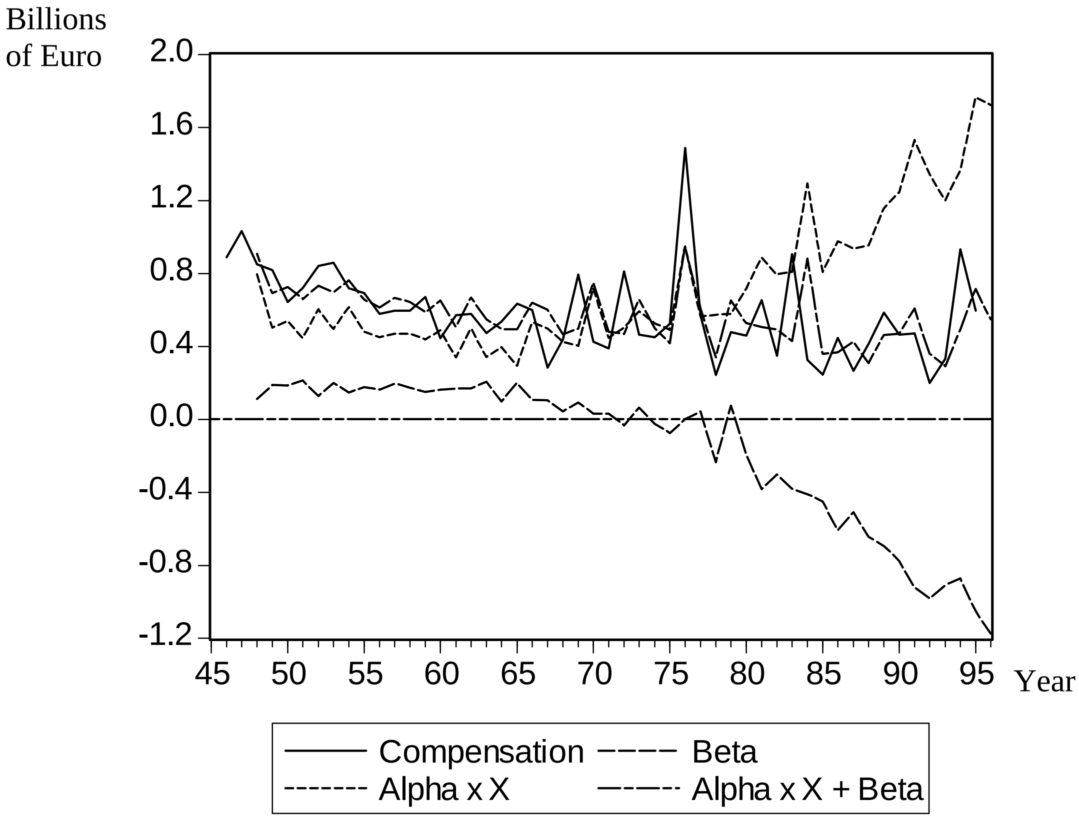

Figure 4. Compensation for Farmers (w_t), Estimated Compensation ($\alpha_t x_t + \beta_t$), the Variable (Uncertain) Compensation ($\alpha_t x_t$), and the Fixed (Certain) Compensation (β_t)

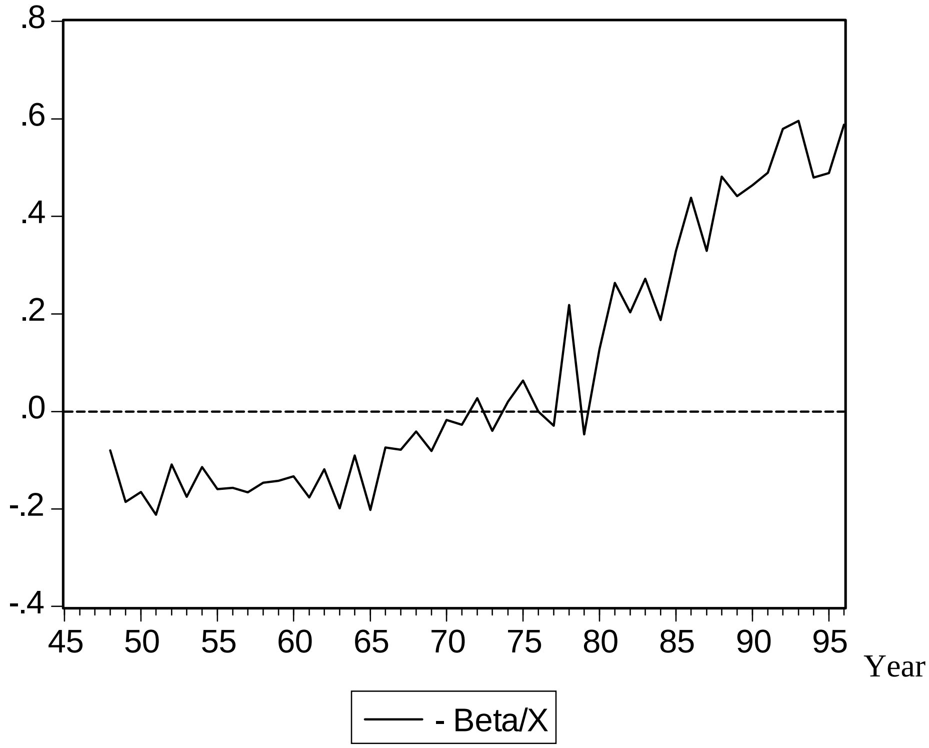

Figure 5. Part of Output Value of Potatoes at Consumer Prices Received by Marketing Firms without Risk (-β_t/x_t)

[1] There is a negative relationship between the frequency of data and the constancy of the variance of the distribution of those data over time. Thus, for example, we expect annual data to have a more constant variance over time compared with daily, weekly, or monthly data.

[2] Contrary to rationality, bounded rationality does not require the economic actors to know the structural equilibrium relations. Instead, it assumes that they use some reduced form of the model that is much easier to specify than the structural model.

[3] Suppose that X_t is a $(d \times 1)$ vector of variables. Given that ΔX_t is stationary, then, if the matrix Π is full rank, X_t is already stationary and has no unit root. If Π is the null matrix, then there are d unit roots, and hence the proper specification of (17) is one without the term ΠX_t. If there are g cointegrating relations $(0 < g < d)$, we can decompose Π into $\rho\delta'$, where ρ and δ are $(d \times g)$ matrices of full column rank, such that $\delta' X_t$ are g linearly independent combinations of variables in X_t that are stationary in spite of the non-stationarity of X_t. Johansen's trace statistic tests for the rank of Π.

[4] This VAR includes a linear trend, but the adjusted R^2s after detrending all variables are still 0.39 and 0.62 for the equations of the de-trended $h_t(x_t - h_t)[E(p_t|I_{t-1})]^{-2}$ and h_t, respectively.

[5] Of course, risk shifting from marketing firms to growers could also be understood by a model in which the agent is risk-neutral and the principal is risk-averse. In that case it can be derived that $\alpha = 1$. We then allow that c varies with time: $c_t = E(p_t|I_{t-1})/e_t$. Furthermore, $\beta_t = \overline{w}_t - 0.5E(p_t|I_{t-1})e_t$ and the fit of the regression, denoted $\hat{d}_t$, of $w_t - 0.5c_te_t^2$ on a constant and a linear trend, is added to $0.5c_te_t^2$ to form the total production costs. To evaluate this model, w_t is compared with $x_t + \beta_t + \hat{d}_t$. This comparison yields a mean absolute error of 0.172 billion Euro. To compare, the mean absolute error obtained from Figure 4 is 0.147 billion Euro. Moreover, if the prediction of w_t, say $\tilde{w}_t$, is optimal in terms of the information used to construct it, then we would expect $(\tau_0, \tau_1) = (0, 1)$ in the "Mincer-Zarnowitz regression" $w_t = \tau_0 + \tau_1 \tilde{w}_t + v_t$ (e.g., Diebold, p. 342), where v_t are nid$(0, \sigma_v^2)$ unexpected events. If the agent is risk-averse (i.e., Figure 4) the restrictions $(\tau_0, \tau_1) = (0, 1)$ are not rejected (p value = 0.50), but for the model in which the principal is risk-averse these restrictions are strongly rejected (p value << 0.01). These results comply with our observation that agents are risk-averse and the principals are risk-neutral.