MEASURING MARKET INTEGRATION FOR APPLES ON THE SOUTH AFRICAN FRESH PRODUCE MARKET: A THRESHOLD ERROR CORRECTION MODEL

by

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ABSTRACT

Apples constitute the bulk of deciduous fruit produced in South Africa, i.e. in 2000, apples made up the largest percentage of the deciduous fruit crop (43%). From 1991/92 to 2002/03 production averaged 574 850 tons per annum with a standard deviation of 43 922 tons. The average distribution of the apple crop between the local market, exports and processing is more or less even. Because of its potential lucrative nature much emphasis in the apple industry is afforded to exports, but relatively little is known about how price transmission takes place on the domestic fresh produce markets (FPMs). Moreover, it is increasingly recognized that the formulation of market-enhancing policies to increase the performance of the local market requires a better understanding of how the market functions. Aggregate market performance is better understood by studying the level of market integration that exists, which in turn is affected by transaction costs in the value chain.

Hence, the primary objective of this study was to measure market integration for apples on the South African FPMs to determine the existence of long-run price relationships and spatial market linkages. Specific issues addressed in this study include, (i) determination of the effect of deregulation of the marketing of agricultural products in 1997 on average real market prices, price spread and volatility (risk), (ii) determination of how FPMs where apples are sold are linked and how prices are transmitted across these markets, (iii) determination of the threshold prices beyond which markets adjust and return to equilibrium, and (iv) establish the response of the FPMs to price shocks and how long it takes for shocks to be eliminated.

The FPMs included in this study are Johannesburg, Cape Town, Tswhane, Bloemfontein, Port Elizabeth, Durban, Kimberley and Pietermaritzburg. The criteria for selecting the FPMs were based on net market positions (surplus or deficit area), geographical distribution, the volume of trade and the importance of the market to the national apple trade flow.

The investigation revealed a statistically significant decline in real prices in six of the eight markets investigated, a statistically significant relation in prices (price spread) between the Johannesburg FPM and five other FPMs, as well as that the price spreads between these markets declined after deregulation, and that the variation in real apple prices declined for five of the eight markets after deregulation. Standard autoregressive (AR) and threshold autoregressive (TAR) error correction models were compared to determine whether transaction cost has significant effects in measuring market integration. Larger adjustment coefficients were found in the TAR model. This is an indication that price adjustments are faster in threshold autoregressive TAR models than in AR models. Also half-life deviations in the TAR model are much smaller than in the AR model. The TAR model requires less time for one-half of the deviation from equilibrium to be eliminated than the standard AR model. Therefore, it is better to use TAR models than AR models because TAR models give a more reliable result.

In addition, the parameter estimates of the threshold vector error correction model were analyzed. The results show that bidirectional and unidirectional causality exist between Johannesburg FPM prices and other markets. Regime switching estimates to investigate market integration in the selected markets show that no persistent deviation from equilibrium existed for all but one market pair and no clear evidence was found to support improved market integration after market deregulation in 1997.

A nonlinear impulse response function to investigate the impact of positive and negative price shocks in the Johannesburg FPM on other FPMs revealed that it takes about six to twelve months for positive and negative shocks to be completely eliminated in all the markets.

Generally, the results obtained confirmed strong market integration in terms of apples for selected FPMs.

Key words: Half-life, Threshold Error Correction, Cointegration, Market Integration, Autarky, Arbitrage, Impulse Response, Market Segmentation

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UITTREKSEL

Appels maak die grootste gedeelte uit van sagte vrugte wat in Suid-Afrika geproduseer word. In 2000 byvoorbeeld, het appels die grootste persentasie van die sagte vrugte opbrengs (43%) verteenwoordig. Vanaf 1991/92 tot 2002/03 is ‘n gemiddelde produksie van 574 850 ton per jaar behaal met ‘n standaard afwyking van 43 922 ton. Die gemiddelde verspreiding van die appel-oes tussen die plaaslike mark, uitvoere en prosessering is min of meer eweredig. Weens die potensiële winsgewendheid daarvan, word die aandag in die appelbedryf veral op die uitvoermark toegespits, maar relatief min is bekend oor prysoordrag op die plaaslike varsprodukte markte (VPM’e). Verder word al hoe meer erken dat die formulering van markopheffende beleidsrigtings om die doeltreffendheid van die plaaslike mark te verhoog, beter begrip vereis van hoe die mark bedryf word. ‘n Beter begrip van markprestasie in die geheel word verkry deur die bestudering van die bestaande vlak van markintegrasie, wat weer deur die transaksiekoste in die waardeketting beïnvloed word.

Die primêre doelwit van hierdie studie was dus om markintegrasie van appels op die Suid-Afrikaanse VPM’e te bepaal om vas te stel of daar bestaande langtermyn prysverwantskappe en ruimtelike markskakeling is. Bepaalde sake waaraan in hierdie studie aandag geskenk is, sluit in (i) die vasstelling van die uitwerking van deregulasie van die bemarking van landbouprodukte in 1996 op gemiddelde reële markpryse, prysverspreiding en onsekerhede (risiko), (ii) die vasstelling van hoe VPM’e waar appels verkoop word gekoppel is en hoe pryse tussen die markte oorgedra word, (iii) die vasstelling van die drempelpryse waarná

markte aanpassings maak en die balans herstel, en (iv) om vas te stel wat die reaksie van die VPM'e op prysskokke is en hoe lank dit neem om die skokke te verwerk.

Die VPM'e waarna in hierdie studie verwys word is in Johannesburg, Kaapstad, Tswhane, Bloemfontein, Port Elizabeth, Durban, Kimberley en Pietermaritzburg. Die maatstawwe waarvolgens die VPM'e gekies is, is gebaseer op netto mark posisies (surplus of te kort gebied), geografiese verspreiding, die omvang van die handel en hoe belangrik die mark vir die nasionale handel in appels is.

Die ondersoek het 'n statisties beduidende afname in reële pryse aan die lig gebring op ses van die agt markte wat ondersoek is, asook 'n statisties bepalende verwantskap in pryse (prysverspreiding) tussen die Johannesburg VPM en vyf ander VPM'e. Dit het ook daarop gedui dat die prysverspreiding tussen hierdie markte ná deregulasie afgeneem het en dat die variasie in reële appelpryse by vyf van die agt markte ná deregulasie afgeneem het. Standaard outoregressiewe (OR) en drempel outoregressiewe (DOR) foutregstellingsmodelle is vergelyk om te bepaal of transaksiekoste 'n beduidende uitwerking op die meting van markintegrasie het. Groter aanpassingskoëffisiënte is in die DOR model gevind. Dit dui daarop dat prysaanpassings vinniger plaasvind in DOR modelle as in OR modelle. Halflewe afwykings in die DOR-model is ook heelwat kleiner as in die OR-model. Die DOR-model kan die helfte van die balansafwyking gouer uitskakel as die OR-model. Dit is daarom beter om die DOR-modelle, eerder as die OR-modelle te gebruik, omdat die DOR-modelle meer betroubare resultate lewer.

Daarbenewens is die parameterskattings van die drempelvektor foutregstellingsmodel ontleed. Die resultate toon dat twee- en eenrigting oorsaaklike verbande tussen die pryse op Johannesburg se VPM en ander markte bestaan. Ramings ten opsigte van regime wisselings om markintegrasie in die verkose markte te ondersoek, het behalwe vir een markpaar, geen volgehoue afwyking van die ewewig aan die lig gebring nie. Daar is ook geen duidelike bewys gevind ter ondersteuning van verbeterde markintegrasie ná markderegulering in 1997 nie.

'n Nie-liniêre impuls respons funksie om die uitwerking van positiewe en negatiewe prysskokke op die Johannesburg VPM op ander VPM'e te ondersoek, het daarop gedui dat dit

op al die markte ongeveer ses tot twaalf maande neem voordat positiewe en negatiewe skokke geheel en al uitgeskakel is.

Die resultate behaal bevestig oor die algemeen sterk markintegrasie ten opsigte van appels vir uitgesoekte VPM'e.

Sleutelwoorde: Halflewe, Drempelfoutregstelling, Gelyktydige integrasie, Markintegrasie, Outarkie (selfversorgend), Arbitrage, Impuls Respons, Marksegmentasie.

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LIST OF ACRONYMS

ADF	Augmented Dickey-Fuller
AR	Autoregression
BLOM	Bloemfontein
CAPT	Cape Town
CPI	Consumer price index
CUSUM	Cumulative sum
DFPT	Deciduous fruit producers Trust
DURB	Durban
ECM	Error correction model
ESTAR	Exponential smooth transmission autoregression
FIPPT	Fruit industry plan project team
FPMA	Fresh produce marketing association
FPM	Fresh produce market
GCIS	Government communication and information system
JFPM	Johannesburg fresh produce market
JNB	Johannesburg
KIMB	Kimberley
LOOP	Law of one price.
NDA	National department of agriculture
OABS	Optimal agricultural business system
OLS	Ordinary least square
PACF	Partial autocorrelation function
PBM	Parity bound model.
PIETM	Pietermaritzburg
PE	Port Elizabeth
RESET	Residual error specification test.
RRA	Railroad association of South Africa
SABG	South African business guide
SAPO	South African plant improvement organization
STAR	Smooth transmission autoregression.
SURE	Seemingly unrelated regression estimation
TAR	Threshold autoregression

TSW	Tshwane
TVECM	Threshold vector error correction model

CHAPTER 1

INTRODUCTION

1.1 Background

Primary agriculture in South Africa contributed between 3% and 5% to the gross domestic product (GDP) of South Africa over the last decade, whilst the agro-industrial sector contributed approximately 15% to GDP (Government communication and information system (GCIS), 2002). Agriculture furthermore has strong forward and backward linkages with the rest of the economy and almost 9% of people that are formally employed are in this sector (GCIS, 2002).

At a sub-sector specific level the fruit industry plays a vital role in job creation and foreign exchange earnings. The deciduous fruit industry provides employment to approximately 210 000 people and supports approximately 1.13 million dependants (Jooste, Viljoen, Meyer, Kassier and Taljaard, 2001). The deciduous fruit industry had an annual turnover of R4.6 billion in 1999, of which approximately R3 billion was earned on the foreign market and 1.6 billion was generated on the domestic market (Jooste *et al.*, 2001).

During 2001/2002, deciduous fruit contributed approximately 29% to the gross value of horticultural products. In the same season about 363 768 tons of deciduous fruit were sold on the 16 major local fresh produce markets (FPM's) and to retailers. This represents a 2.1% increase compared to 356 169 tons sold during the 2000/2001 seasons (NDA, 2003).

Deciduous fruit includes grapes, apples, pears and stone fruits. Apples constitute the bulk of deciduous fruit produced in South Africa, i.e. in 2000, apples made up the largest percentage of the deciduous fruit crop (43%). The apple industry, in addition to its contribution to revenue generation, also provides employment to many people. According to Fruit Industry Plan (FIP) (2004) the apple industry provides employment to 28 068 people with 112 272 dependants.

According to Louw and Fourie (2003), the average distribution of the apple crop between the local market, exports and processing is more or less even. Because much emphasis in the apple industry is afforded to exports (FIP, 2004), and relatively little is known about how price transmission takes place on the domestic FPM's. Such information is important for apple producers and other apple value chain role players since it affects their marketing decisions (buying and selling), which in turn affects decisions related to logistical matters and eventually profits realized.

1.2 Problem statement

Since the 1980s several policy changes that affected the production and marketing of agricultural products in South Africa characterized the agricultural landscape. There were changes in the fiscal treatment of agriculture including the removal of tax concessions, reduction in budgetary allocation, land and institutional reforms including the deregulation of the marketing of agricultural products. Marketing of agricultural products in South Africa was deregulated in 1996 following many years of a distorting agricultural marketing policy. Prior to deregulation there were state controlled monopolies in the agricultural marketing system, which, amongst others, controlled the movement of goods, information flow and prices. With deregulation all state controlled marketing boards were abolished (Jooste *et al.*, 2001).

As far as the deciduous fruit industry is concerned the emphasis was mainly on supply-side economics, while fruit exports were controlled by Unifruco. After deregulation many changes took place in the deciduous fruit industry (not all a direct consequence of deregulation, but nevertheless important for the industry as a whole) which include, amongst others, increased competition internationally, higher and stricter standards to access both the local and international market, new legislation affecting different levels of the supply chain, etc. In this changing environment much attention was afforded to the export market and one could argue that the domestic market was to a large extent neglected in so far as understanding local market forces are concerned (FIP, 2004). As stated, the marketing of apples to the international and local markets is more or less even, but little is known about the price transmission effects between different markets in the deregulated environment.

Moreover, it is increasingly recognized that the formulation of market-enhancing policies to increase the performance of the local market requires a better understanding of how the market functions. Aggregate market performance is better understood by studying the level of market integration that exists, which in turn is affected by transaction costs in the value chain.

In the context of this study market performance will be investigated by studying the impact of deregulation on average market prices, price spreads and market volatility in the domestic apple industry. A review of the literature showed that no study has been specifically conducted to measure the extent of market integration and price transmission in this industry. The model developed for purposes of this study can also be used to investigate the extent of market integration in other fruit sub-sectors, and for that matter all other agricultural products.

1.3 Objectives of the study

The primary objective of this study is to measure market integration for apples on the South African FPMs to determine the existence of long-run price relationships and spatial market linkages. In order to meet the primary objective the following secondary objectives must be met:

- Determine the effect of reform on average market price, price spread and volatility (risk).
- Determine how FPMs where apples are sold are linked and how prices are transmitted across these markets.
- Determine the threshold prices (transaction cost) beyond which markets adjust and return to equilibrium.
- Establish the response of the FPMs to price shocks and how long it takes for shocks to be eliminated.

1.4 Data and methodology

Using time series apple price data from 1991 to 2004 the study analyzes price transmission in eight selected South African FPMs. The FPMs included in this study are Johannesburg, Cape Town, Tswhane, Bloemfontein, Port Elizabeth, Durban, Kimberley and Pietermaritzburg. The criterion for selecting these FPMs is based on net market positions (surplus or deficit area), geographical distribution and the volume of trade or the importance of the market to the national apple trade flow.

1.5 Outline of the study

Chapter 2 is the literature review that provides a general overview of the theoretical background of co-integration analysis of spatially separated markets. Chapter 3 provides an industry overview that uses the structure, conduct and performance method. In Chapter 4 the empirical method used in this study is discussed, while Chapter 5 presents the results of the study. In Chapter 6 conclusions and recommendations are provided.

CHAPTER 2

LITERATURE REVIEW

2.1 Introduction

Various approaches that have used to study market integration have been criticized because transaction costs were not considered. Recent improvements have been the introduction of the parity bound and threshold cointegration models that account for transaction cost. The objectives of this chapter are to assess the development of these concepts and to decide on an appropriate modeling framework based on the theory and findings in the literature reviewed.

2.2 The concept of market and price analysis

In a market driven economy, the marketing system serves at both the micro and macro levels as mechanism to transmit to market participants' information that is useful in decision making. Transparent, accurate and timely price signals play a significant role in the conduct and performance of an efficient marketing system. In a competitive economy the pricing mechanism is expected to transmit orders and directions to determine the flow of market activities. Pricing signals guide and regulate production, consumption and marketing decisions over time, form and place (Kohls and Uhl, 1998). Identifying the causes of differences in prices in interregional or spatial markets has therefore become an important economic analytical tool to understand markets better. The next section explores this concept.

2.2.1 Spatial arbitrage price analysis

According to Negassa, Myers and Gabre-Maldhin (2003), the price relationships between spatially separated markets are generally analyzed within the framework of spatial price equilibrium theory developed by Enke (1951), Samuelson (1964) and Takayama and Judge (1964). The key assumption underpinning spatial price equilibrium theory is that price relationships between spatially separated competitive markets depend on the size of transaction costs. This implies that transaction costs play a key role in the study of spatial price relationships and should not be ignored (Faminow and Benson, 1990). The principle underlying the differences between regions in a competitive market structure with homogeneous commodities is that price differences between any two regional markets that trade with each other should equal transaction cost, while in a situation of autarky price differences will be less than or equal to transaction costs (Tomek and Robinson, 1990). These are called spatial arbitrage[1] conditions (Faminow and Benson, 1990).

When the price difference between different markets exceeds transaction costs, arbitrage opportunities will be created and profit seeking merchants will seek to exploit such opportunities, by purchasing commodities from a low-price surplus market and transferring them to a higher-priced deficit market. Arbitrage opportunities occur only when the deviation in prices is substantial enough for potential profit to exceed the cost of trading. This will raise prices in the surplus region and reduce them in the deficit region (Tomek and Robinson, 1990).

This concept of constant market price and arbitrage are consistent with the "law of one price". The "law of one price" states that, under competitive market conditions, all prices within a market are uniform after taking into consideration the cost of adding place, time and form utility to the products within the market (Kohls and Uhl, 1998). The "law of one price" (LOOP) is useful in determining the size of a market, predicting price changes within a market and evaluating the pricing efficiency of a market (Kohls and Uhl, 1998). Failure of one or more regions to adhere to the LOOP means that the regions may not be linked by arbitrage, certain factors act as impediments to efficient arbitrage, e.g trade barriers, government intervention policy, imperfect information, or risk aversion. The LOOP has been investigated by several researchers. Baffles (1991) and Ardeni (1989) found the LOOP to be a short-run phenomenon. They found no evidence to support the LOOP as a long-run relationship and suggest institutional factors, transaction cost, price and time–specific problems as the main reasons for this failure rather than the concept itself.

Point-space price relationships have been used by Takayama and Judge (1992) to model perfectly competitive markets characterized by distinct regions or centers. The idea is that all transactions within a region occur at one point; in other words, no regional transaction costs apply and all buyers and sellers converge to a single point. Geographical regions are divided into a discrete number of regions within which transaction costs is assumed to be zero. Intra-regional competition and trade, they argued, are characterized by the perfect competition

[1] Arbitrage is a widely used concept defined as a riskless profit without investment. It is used in spatial market efficiency studies to signify exploitation of profit opportunity created by market inefficiency.

model and regional boundaries are assumed fixed. Faminow and Benson (1990) argue that this is an oversimplification of reality due to (i) the existence of spatial market price interdependence even within distinct regions; (ii) the fact that market participants are spatially separated, and (iii) intra-regional transaction costs. They also state that there are no fixed geographical boundaries; markets are spatially integrated and transaction costs affect the net price received or paid.

Moreover, transaction costs and location play a vital role in marketing decisions. Market participants will prefer least cost transactions, choosing a nearby market over another across a geographically separated area due to transaction costs. This tendency creates a viable justification for the importance of transaction costs in spatial price analysis. Transaction costs limit the scope of arbitrage or area relevant for competition among firms to occur. In a non-competitive market, the presence of transaction costs and other non-competitive institutional structures, such as organized oligopoly arrangements, either through price leadership or collusion, may bring about a pricing system based on a specific point[2]. This situation affects market integration and efficiency.

2.2.2 Spatial market integration

Spatial market integration refers to co-movement or a long run relationship of prices. It is defined as the smooth transmission of price signals and information across spatially separated markets (Golletti, Ahmed and Farid, 1995). Two trading markets are assumed integrated if price changes in one market are manifested to an identical price response in the other market (Goletti *et al.,* 1995; Barrett, 1996). Market integration can also be defined as a measure of the extent to which demand and supply shocks in one location are transmitted to other locations (Negassa *et al.,* 2003). Barrett (1996) distinguished market integration into (i) vertical market integration involving different stages in marketing and processing channels, spatial integration relating to spatially distinct markets, and (ii) inter-temporal market integration which refers to arbitrage across periods of time. This study intends to examine spatial market integration.

[2] Base-point pricing was in fact described more than a quarter of century ago by E.A.G. Robinson in his book "Monopoly".

Based on the definition by Gonzalez-Revera and Helfand (2001), a market within a distinct location will be considered integrated if physical flows of goods and services exist among the locations and there is evidence of a long run relationship. These criteria are important in identifying the sets of locations that are directly or indirectly spatially linked by trade.

2.2.3 Necessary and sufficient conditions for efficient market arbitrage

Since market integration is interwoven with the concept of arbitrage, efficiency in market performance requires that markets must be linked by perfect arbitrage, i.e. the spatial conditions for arbitrage must be fulfilled. The necessary and sufficient conditions for efficient market arbitrage are summarized as follows. The necessary conditions are fulfilled if there are physical flows of goods and information across space, time and form between any two trading partners. The sufficient conditions are fulfilled when the price differential between any two regional markets that trade with each is equal to or less than transaction costs. That is, market prices in the two regions adhere to the law of one price. When price spreads exceed transfer cost, spatial arbitrage conditions are violated irrespective of whether or not trade occurs. The violation of spatial arbitrage conditions indicates a lack of market integration.

Efficient market arbitrage therefore provides a strong driving force for market equilibrium conditions. Efficiency in spatial arbitrage leads to market price stability and encourages risk sharing practices among markets. For example, crop failure in one region will result in increased prices in all regions. Efficiency in spatial market arbitrage will cause the price risk to be spread among markets, a higher price in a deficit region will be transmitted to surplus markets and lower prices from a surplus region will be transmitted to deficit markets, causing prices in the different markets to equalize through arbitrage.

2.2.4 Economic justification for market integration

Most studies on spatial price linkages focus on market integration and price domination (Kuiper, Lutz and Tilburg, 2003). Measurement of market integration can be viewed as a basic tool for an understanding how markets work (Ravallion, 1986). If markets are integrated the effects of policy intervention in one market would be transmitted to other markets. Duplication of interventions to spatially separated markets will otherwise be undertaken at a higher cost (Goletti *et al.,* 1995). By giving a more detailed picture of the process of transmission of incentives across the marketing chain, knowledge of market integration is relevant to the success of policies such as market liberalization, price stabilization programs and food security programs (Amha, 1999). Market integration ensures that a regional balance occurs among food deficit, surplus and non-cash crop producing regions (Goletti *et al.,* 1995). According to Barrett (1996), studies on market integration provide information on market performance which is necessary for proper policy formulation and macroeconomic modeling.

If markets are not spatially or inter-temporally integrated it could be indicative that market inefficiencies exist as a result of, amongst others, collusion and market concentration which result in price fixing and distortions in the market. In such cases cross-sectional or inter-temporal aggregation of demand and supply loses its logical foundation (Barrett, 1996). The result is that agricultural producers will fail to specialize according to long run comparative advantages and gains from trade will not be realized (Baulch, 1997). This implies that if the assumptions of marketing integration hold, optimal allocation of scarce resources could be attained. However Nebery and Stiglitz (1984), quoted in Amha (1999), argue that the existence of a free market alone does not necessarily guarantee optimal allocation of resources. According to Wyeth (1992), market integration deals with one aspect of market performance and a perfectly competitive market will probably be integrated, but an integrated market may not be perfectly competitive.

The distinction between integrated and non-integrated markets has obvious importance for the formulation of empirical models of trade in general. Markets which are independent must be modeled in a disaggregated manner, while markets which are integrated may be amenable to aggregate analysis. Hick's composite commodity theorem guarantees that commodities can be treated as a single aggregate when their relative prices remain constant (Diakosavva, 1995). In light of these assertions several studies have attempted to provide a better understanding of how specific markets work. In the next section, price transmission and measures of market integration will be illustrated using different approaches.

2.3 Measures of market integration

2.3.1 Bivariate correlation approach

Early studies on the degree of spatial market integration utilized simple bivariate price correlation between two price series in two competing markets (Negassa *et al.,* 2003). In general, this notion is intuitively related to the idea that prices in integrated markets move together (Goletti *et al.,* 1995). Price correlation is an easy and simple way to measure co-movement of prices. Jones (1972), Thodey (1969) and Lele (1967) used bivariate correlation and regression coefficients to measure the adequacy of infrastructure, level of competition, the impact of legal barriers to the movement of agricultural products and negotiated transactions per time period (quoted in Amha, 1999).

Several researchers question the usability of bivariate price correlation to investigate the degree of market integration, for example:

- Golletti *et al.,* (1995) state that bivariate analysis masks the presence of certain factors such as general price inflation, effects of government policies, etc.
- Barret (1996) argues that the bivariate approach is weak because it produces high correlation results even for markets with no physical linkage. In addition, price data of reasonably high frequency are often synonymous with the heteroscedasticity problem and a simple pair wise price correlation statistic will fail to recognize the presence of heteroscedasticity inherent in such price data series.
- According to Delgado (1986), bivariate correlation coefficients have presented a distorted picture by indicating relatively low price correlation between the markets even in cases where evidence suggests competitive and rational behavior by a large number of market participants. He argues that bivariate correlation analysis is a pair-wise analysis of two markets; difficulties arise when more than two markets are to be analyzed and compared.
- Ravallion (1986) and Sexton *et al.* (1991) state that bivariate price correlation assumes instantaneous price adjustment and cannot capture the dynamic nature of a marketing system. There is a high tendency of spurious market integration because the prices may tend to move together even though markets are not integrated.

- Barrett (1996) and Negassa *et al.* (2003) found that bivariate price correlation may overstate the lack of price integration if a lag in market information produces a lag in the price response between markets.

2.3.2 Variance decomposition approach

Since the mid 1980s several attempts have been made to improve the much criticized bivariate price correlation approach to market integration analysis. Based on the assumption of trend stationary economic series, Delgado (1986) criticized most bivariate correlation analysis for undetrended market time series price data. Under this presumption Delgado (1986), in his variance component approach to food grain marketing in northern Nigeria disaggregated his analysis by seasons and controlled for heteroscedasticity and autocorrelation and first removed common trend and seasonality present in the price series before testing for market integration. In his approach, Delgado (1986) jointly evaluated information from a series of markets to assess the integration of the market system and then jointly tested for equality of seasonal trends across all markets. The approach is aimed at decomposing the variance of food grain prices into components. This instance precipitated assumptions of constant variance of prices, transportation and transaction costs for various components and between any two markets within the system over the season. Then the spatial integration between pairs of markets for a given season is indicated by the equality between the spatial price spread and the constant transport and transaction costs during that season, subject to random noise (Delgado, 1986, Negassa *et al.,* 2003). This means that if the price spread between any pair of market is random, then inter-market price differentials are equal due to the presence of transportation and transaction cost, with deviation from the constant being random noise. In other words, if price divergent trends are non-random over seasons between a pair of markets, this supports the hypothesis of a constant relationship between the series.

However, the variance decomposition approach is based on tests of contemporaneous price relationships, and thus, does not allow for dynamic relationships between prices in different markets. It assumes constant inter-market transfer cost (Negassa *et al.,* 2003). This model was applied to eighteen months of weekly grain prices for twenty-two villages in northern Nigeria. Results suggest that markets are not well integrated in the six months covering the harvest period.

2.3.3 Radial market integration approach

Another approach to the concept of market integration is the radial market integration system proposed by Ravallion (1986). This approach also aims at improving the static bivariate correlation procedure. According to Ravallion (1986), in order to improve the static bivariate model it must be extended into a dynamic model of spatial price differentials. Price series will have their own dynamic structure incorporating both correlated local seasonality and interlinkage with other markets. To achieve this Ravallion (1986) assumes a radial spatial market structure between a group of local markets and a single central market whose prices are weakly exogenous from those of other markets. While there may be some trade among the local markets, it is trade with the central market which determines local price formation (Ravallion, 1986 and Negassa *et al.,* 2003). The central market price can also be influenced by local prices, depending on their size and number. In order to instrument the price in the central market, the local market prices were lagged. By adopting the above assumptions, an implicit binary relation can be obtained between each local market and the central market.

This method allows testing a set of hypothesis regarding spatial market integration between local and central markets after controlling for seasonality, common trend and autocorrelation. The first hypotheses testing assume market segmentation in that central market prices do not influence prices in the local market if the price coefficient is zero. The second hypothesis assumes short-run market integration. A price increase in the central market will be immediately passed on to the local market if the coefficient is equal to one. The third hypothesis assumes long-run market integration; if market prices are constant over time, undisturbed by any local stochastic effects the sum of the coefficients of the lagged market prices must be one. The acceptance of the short-run restrictions implies long-run market integration, but the reverse is not attainable.

However, this approach also has limitations; firstly, the assumption of radial market structure does not always hold due to inter-seasonal flow reversals and direct trade links between regions (Barrett, 1996). Secondly, the assumptions of constant inter-market transfer cost will introduce bias in the test for market integration if transfer costs are time variant (Barrett, 1996). Thirdly, aggregated price differential produces inferential difficulty when one investigates linkage location of any impediments to trade (Ravallion, 1986). Fourthly, multicollinearity may occur among the regressors. In that case, a high standard error on the coefficient for central market price may be due to its high correlation with lagged local prices rather than weak market integration. Fifthly, the method does not distinguish market integration due to non- competitive behavior such as collusion (Faminow and Benson, 1990)

An application of this approach to monthly rice price data for Bangladesh suggests some quite significant departures from the conditions for both short and long-run market integration. Accordingly these findings are not revealed by test using static bivariate correlations (Ravallion, 1996).

2.3.4 Structural determinants of market integration

Studies in market integration have mostly concentrated on efforts to relate price transmission to market performance. Market integration has largely been characterized as co-movement of prices, and several researchers have made conclusions on the basis of information conveyed by price signals in the marketing system. But markets are complex institutions and their performance and integration are the result of numerous factors. The various approaches used to measure integration (as highlighted in this text so far) have largely neglected analysis of structural factors affecting market integration.

Goletti, Ahmed and Farid (1995) systematically relate market integration to structural factors. Marketing infrastructure, volatility of government intervention and the degree of self-sufficiency of production were identified as major determinants of market integration.

Market infrastructure, transportation, communication, credit and storage facilitate smooth functioning of markets. For example, Penzhorn and Arndt (2002), testing maize market integration in Mozambique, found that a lack of adequate infrastructure and the existence of structural, institutional and political impediments have inhibited free movement of information, capital, investment, goods and services, resulting in periodic segmentation and deviations from the law of one price. Government policies, price stabilization, trade restrictions, availability of credit and transport regulations affect marketing systems in various ways. Government intervention destabilizes a market economy. Levels of production, government policies and infrastructural development determine the level of self-sufficiency of a country. There are different categories of self-sufficiency, and this may be linked to the likelihood of market integration (Goletti *et al.,* 1995).

However, Goletti *et al.,* (1995) based their findings on the hypothesis that a marketing infrastructure makes a positive contribution to integration. The degree of similarity of production per capita affects markets positively; that is, the more dissimilar the markets, the greater the incentives to trade with each other. Government intervention affects integration positively or negatively. Price stabilization policy can lead to co-movement of market prices, and may inhibit transmission of price signals across spatial markets. Transportation infrastructure plays a key role in determining integration. The greater the distance between two markets, the higher the transaction cost.

Analysis of the role of structural factors in determining market integration of the rice market in Bangladesh by Goletti *et al.,* (1995) indicates a moderate degree of market integration. Segmented markets made up less than 10 percent of all conceivable links in the network of the 64 markets in the data set used in the analysis. Different measures of market integration (e.g. speed of adjustments, co-integration coefficient, and correlation of prices) respond differently to structural factors. The weak congruence of the effects of various structural factors suggests that market integration is affected negatively by distance between markets.

2.4 Statistical properties of the time series variables

Econometric analysis of time series data is usually preceded by investigation of the characteristics of the data series. Enders (2004) suggests that a visual plot of the correlogram will give a vital clue. The correlogram of a stationary series drops off as the lag becomes large, but that of a non-stationary series does not. The objective is often to establish whether the series is trended. This will help highlight the dynamics of movement of slow long-run evolution of time series in the levels. Understanding the trend process is important in economic analysis. An economic time series is ether stationary or non-stationary. A stationary stochastic series has a constant mean, variance and covariance. It is time invariant, mean reverting, and fluctuations around its mean have constant amplitude. Non-stationary stochastic series have varying mean, or time varying variance. A non-stationary process exhibits random walk and has unit roots

Unit root test has since 1980's attracted interest in econometrics and statistical literature. Nelson and Plosser (1982) argued that most macroeconomics variables have a univariate time series structure. Much more work has been done following Nelson and Plosser's (1982) expository research. Analysts are apparently uncertain on the issue of the unit root test because of the difficulty in ascertaining the data generating process of most economics series. Varied opinions exist in the literature as to the most appropriate technique of determining the statistical or stochastic mechanism of a data series to avoid spurious analysis.

Before the early 1980s economic time series were generally assumed to be characterized as stationary fluctuations around a deterministic trend. Many studies on the measurement of business cycles were based on this assumption. Nelson and Plosser (1982) criticized Bodkin (1996), Lucas (1973), Barro (1978), Sargent (1978), Taylor (1979) Hall (1980), and Kydland and Prescott (1980) for their studies which were based on linear detrending of time trends for the measurement of a business cycle. In their own study Nelson and Plosser (1982) investigated whether macroeconomic time series are better characterized as stationary fluctuations around a deterministic trend or as a non-stationary process that have the tendency to return to a deterministic path. Using 14 historical macroeconomic time series for the U.S., they could not reject the hypothesis that the series are non-stationary stochastic processes with no tendency to return to the trend line. They observed that time series may contain both a secular or growth component and a cyclical component. A cyclical component is assumed to be transitory (stationary) in nature. They argued that since cyclical fluctuations are assumed to dissipate over time, any long run or permanent movement (non-stationary) is attributed to the secular component. Secular component exhibits stochastic trends with a random walk and does not follow a deterministic path. Therefore, a linear detrending of a stochastic time series characterized by a random walk is inappropriate and the model based on a time trend residual will be seriously misspecified.

In contrast to the findings of Nelson and Plosser (1982), Perron (1989) used the same data set as Nelson and Plosser (1982) but his findings were different. Perron (1989) reported that most macroeconomics time series are not non-stationary stochastic processes and are not characterized by the presence of a unit root. Fluctuations are indeed stationary around a deterministic trend function. According to Perron (1989), 11 out of the 14 series he analyzed were rejected under the null hypothesis of unit root as compared to only one series (money stock) that was rejected under the null hypothesis of unit root in the Nelson and Plosser (1982) study. This is a matter of misspecification as Perron (1989) suggested. In another study, Alemu, Oosthuizen and Van Schalkwyk (2004) found that Agricultural GDP in Ethiopia is a trend stationary process. This implies that fluctuations in the agricultural GDP series in Ethiopia are temporary and dissipate in a short period of time.

Given the above findings, statistical testing of unit root is crucial in the evaluation of the non-stationarity that most time series data exhibit. It will be particularly useful in determining whether the trend component is stochastic through the presence of unit root or deterministic through the presence of a polynomial time trend (Perron, 1989; Nelson and Plosser, 1982). While a stochastic trend may exhibit systematic variations which are hardly predictable, a deterministic tend does not vary and is completely predictable (Gujarati, 2003). For plausibility in the economic analysis, the removal of the trend component is important. If the process that generated the series contains outliers or is autocorrelated, a linear combination of two non-stationary series will produce residual error terms that are also correlated. If the residuals are correlated, the OLS estimators will underestimate residual variance. This will result in wrong conclusions about the statistical properties of the time series.

Two approaches have been commonly used in the literature to remove trend and seasonal components (Maddala and Kim, 1998). The methods are, either by regressing the variables against time (Nelson and Plosser, 1982; Gujarati, 2003; Maddala and Kim, 1998) or taking the first difference of the series (Gujarati 2003; Maddala and Kim, 1998). If simply taking the first difference of the series can eliminate the non-stationarity, it is known as a difference stationary process (DSP). Nelson and Plosser (1982) argued that the implication of stochastic trend (unit root) hypothesis is that under this hypothesis random shocks have a permanent effect on the system. Fluctuations are not transitory. A series with a deterministic time trend is rendered stationary by removing the time trend. In a trend stationary process, fluctuations are dominated by temporary deviations and have the tendency of dissipating in a short period of time (Alemu, Oosthuizen and van Schalkwyk, 2004).

Notably, errors could arise if the data generating process (DGP) of a series is misspecified. If the time series is DSP but treated as a trend stationary process (TSP), the process is underdifferenced. Chan, Hayya and Ord (1977) found that when the true model of a time series is a random walk, the use of linear ordinary least square (OLS) regression to eliminate a suspected trend will create spurious positive autocorrelations in the first few lags. Nelson and Kang (1981) found that this will imply strong pseudo-periodic behavior in the detrended series. On the other hand, if the series is a TSP and treated as a DSP we have a case of overdifferencing of the series (Gujarati, 2003 and Maddala and Kim, 1998). The use of first differences to eliminate a linear trend will result in a residual that may be stationary but which is not white noise with a first lag negative autocorrelation. It is argued in the literature that overdifferencing is a less serious problem than underdifferencing. Under or overdifferencing may not have serious consequences if serial correlation in the error term is taken care of (Plosser and Schwert, 1978, and Nelson and Kang, 1991, 1984). Nelson and Kang (1991) argue that if the true process is DSP[3]and treated as TSP, then the detrended series will exhibit spurious periodicity. Contrary to the Nelson and Plosser (1982) argument, Maddala and Kim (1998) suggest that under or over differencing will not constitute a problem if the serial correlation structure of DSP and TSP models is taken into account. In other words, Maddala and Kim (1998) recommend estimating the equations in both levels and first difference and choosing the one that requires the smaller amount of correction to remove autocorrelation of residuals.

The situation in the last two decades has been to derive test statistics that approximate a plausible stationarity test. Dickey and Fuller (1981) developed the Dickey-Fuller test (DF) for unit roots. The DF test was constructed on the assumption of an independently normally distributed error term. Based on the criticism of the DF test, Dickey and Fuller augmented the test by adjusting it to take care of possible serial correlation in the error term. The augmented Dickey-Fuller test (ADF) includes lagged difference terms of the regressand (Dickey and Fuller, 1979, 1981; and Gujarati, 2003). Engle and Granger (1987) proposed the use of several test statistics among which include the DF and ADF tests (these residual-based tests statistics are discussed in details in Appendix A). The residual-based test involves procedures that are designed to detect the presence of a unit root in the residual of co-integrating regressors among the levels of economic time series (Phillips and Ouliaris, 1990). Comparatively, residual–based tests have gained much recognition by many empirical researchers. Preference for residual-based testing is due to its ease, convenience and clarity of objective under the null

[3] The study is based on the periodic properties of detrended series. Error terms are assumed to be non-cyclical.

hypothesis. According to Asche, Gordon and Hannesson (2004), two problems can be identified with the residual-based test. Firstly, it is subject to the normalization problem. Secondly, normal statistics inference and test for the law of one price (LOP) are not valid, although co-integration testing for between two commodities is possible. The Johansen-Juselius non-residual-based maximum likelihood approach is used in Vector autoregressive frame work to test for co-integration.

However, Engle and Granger (1987) recommend ADF test statistics based on its performance and showed that the ADF test and other proposed tests are similar when the data set follows a vector random walk by independently and identically normally distributed (iid) innovations. This phenomenon has been criticized, based on assumptions of error being independently and normally distributed with zero mean and constant variance $(0, \sigma^2)$. Intuitively, the assumptions of independence and homoscedasticity will result in aggregate time series characterized by random walk (Phillips 1987). Phillips however, proposed new test statistics that do not depend on these assumptions. These are Z_α and Z_t tests, which are based on the limiting distribution theory. In addition, Phillips and Ouliaris extended the ADF, Z_α and Z_t tests to include a variance ratio test and a multivariate trace test statistic. ADF, Z_α and Z_t tests all have limiting distribution often expressed as a stochastic integral. In essence, Z_α and Z_t are found to be asymptotically equivalent. Phillips and Perron (1988) proposed a non-parametric test. The test accommodates models with a fitted drift and a time trend so that they may be used to discriminate between unit root non-stationary and stationarity about a deterministic trend (Phillips and Perron, 1988). The asymptotic distribution of Phillips-Perron (PP) and ADF are similar (Gujarati, 2003). Therefore the Phillips-Perron test is often used in conjunction with ADF because they re- enforce each other (Bamba and Reed, 2004). Nevertheless the Phillips-Perron (1988) test has less restrictive assumptions compared to the ADF test and the possibility of heteroscedasticity is more accommodated in the Phillips-Perron test (Bamba and Reed, 2004).

The Range Unit Root (RUR) test statistics, developed by Aparicio, Escribano and Garcia (2004) constitute another type of test statistics. This is a non-parametric range unit root test. It is superior to the standard unit root test because it does not impose severe restrictions on the data generating process of the series. It is invariant to non-linear monotonic transformation and to the distribution of the model errors. It is robust against structural breaks, parameter shifters and outliers. It does not depend on the variance of any stationary alternative; hence it outperforms standard tests in terms of power on near-unit root stationary time series. Finally, it is not affected by the presence of additive noise on the series (Aparicio *et al.,* 2004). In general, the usefulness of these various tests depends on their performance. The size distortions and the power properties of the commonly used test will be discussed in the next section.

2.4.1 Size distortion and power property of unit root test

The size distortion and the power property of the commonly used unit root test, DF, ADF, and PP, have been criticized. Gujarati (2003) and Maddala and Kim (1998) observed that the tests suffer from size distortion if the underlying distribution contains a moving average (MA) component. Schwert (1989), quoted in Maddala and Kim 1998, suggest that the PP test suffers from size distortions when the (MA) parameter is large[4]. The ADF test display size distortions in the presence of negatively correlated (MA) or error structures (Schwert, 1989 and Dejong, Nankervis and Whiteman, 1992b). Gujarati (2003) suggests that power property of the various tests depends on many factors. For a given sample size, the power is greater when the span is larger. For example, if the random walk component can have arbitrarily small variance, test for unit root or trend stationarity will have low power in small samples (Cochrane, 1991). If the parameter coefficient is closer to one, the test may fail to reject the null hypothesis due to lack of power. Researchers usually assume the order of integration to be 1(1); if the series is integrated of the order 1(2) the traditional unit root test will perform poorly. Although DF, ADF, PP, tests are often used, Gujarati (2003) suggests that no uniformly powerful test for unit root hypothesis exists so far in the literature. Despite insufficiencies in the tests procedures discussed, ADF has been recommended as showing better performance (Engle and Granger, 1987, and Gujarati, 2003). Hence, we shall use the ADF test for stationarity tests in this study because it performs better among the available tests and also takes into account that serial correlation is prevalent in most time series.

[4] Schwert (1989) suggest correct specification of the autoregressive integrated moving average ARIMA process before testing for the presence of unit root.

2.4.2 Null and alternative hypothesis under unit root test

Traditionally, analyses of economic time series have been based on the null hypothesis of unit root or absence of co-integration (i.e. $\rho = 1$, where ρ is the correlation coefficient of the co-integrating variables). The null hypothesis of unit root in the residual is tested against the alternative that the root is less than one, 1(0). Under this dispensation, if the null of unit root I(1) is rejected then the alternative of co-integration 1(0) is accepted. In some cases, the commonly expressed view is that the null hypothesis of stationarity should be used. In view of this, Phillips and Ouliaris (1990), Kwiatkowski, Phillips, Schmidt and Shin (1992), acknowledge that the power properties of many standard tests depend critically on their method of construction.

However, many researchers argue that the hypothesis of co-integration (stationarity) is a better choice (Engle and Granger, 1987; Kwiatkowski *et al.,* 1992; and Phillips and Ouliaris, 1990). The supportive argument for the null hypothesis of stationarity is expressed based on the facts that most standard unit root tests fail to reject the null hypothesis of unit root for most economic time series. Kwiatkowski *et al.* (1992) suggest that the failure to reject the null hypothesis of no co-integration is due to the fact that most economic time series are either informative about whether or not there is a unit root, or that standard tests are not very powerful against the relevant alternative of the null hypothesis. Kwiatkowski *et al.* (1992) suggest simultaneously performing tests of the null hypothesis of stationarity as well as unit root. Using the same U.S. time series data set used by Nelson and Plosser (1982), they test the null hypothesis of stationary around a deterministic trend, thus expressing the series as the sum of deterministic trend, random walk and stationary errors. They presented a statistical test of the hypothesis of stationarity, either around a level or around a linear trend. For all the series the null hypothesis of level stationarity is rejected, but could not be rejected for the null hypothesis of trend stationarity. The null hypothesis of trend stationarity corresponds to the modified version of Lagrange multiplier hypotheses that variance of the random walk equals zero. The random walk is assumed normal and errors are white noise. By testing both unit root and stationary hypothesis they distinguished among series that are stationary, have unit root, and series for which the data set do not posses sufficient information to determine whether they are stationary or not.

Kwiatkowski *et al.* (1992) found the unemployment series to be trend stationary while consumer price, real wages, velocity and stock prices all have unit root. GNP, nominal GNP, and the interest rate also have unit root. Real per capita GNP, employment, unemployment rate, GNP deflator, wages, and money could not be rejected under both null hypothesis of unit root and trend stationary; hence the researchers concluded that the data series had insufficient information to distinguish between these hypotheses. The results of Kwiatkowski *et al.* (1992) and other researchers suggest a better performance under the null hypothesis of stationarity. In spite of these suggestions few residual-based statistical tests of co-integration proceed practically along these lines in the literature.

2.4.3 Unit root test in the presence of structural break

The influence of structural break on the plausibility of the unit root test has recently been recognized by researchers. Co-integration analysis involves the use of long span time series data, which are most likely to have structural breaks. Failure to detect and account for structural shifts in parameters would be characterized by misspecifications, biased inference and poor performance (Gabriel *et al.,* 2001). The unit root test will exhibit low power, and fail to distinguish between I (1) series and I (0) process in the presence of a structural break (Gabriel *et al.,* 2001).

Structural breaks affect the null hypothesis of unit root differently (Lee *et al.,* 1997). Leybourne *et al.,* (1998) showed that using standard DF test can lead to spurious rejection of the null hypothesis of unit root if a structural break occurs early in the series for a finite sample.

Lee (2000) pointed out that the spurious rejection of the null hypothesis of unit root under the DF test is due to efficiency losses, given that DF is based on the conditional distribution of data discarding the first observation. For example, Lagrange multiplier unit root tests conducted by Schmidt and Phillips (1992), quoted in Lee (2000), utilizing the unconditional distribution of data, show no spurious rejection, even if the break occurs early in the series.

Non-constancy of error variance can contribute to spurious stationarity tests (Gabriel *et al.,* 2001). According to Hamori and Tokohisa (1997), spurious stationarity can arise if a DF test is applied to a process that suffered an upward break in variance. An early shift in a finite data series will contribute to increase the size distortion of the parameter estimates. According to Gabriel *et al.,* (2001), a similar study by Nellson, Piger and Zivolt, (2001), and Psaradaki, (2001) with series that exhibit decrease in error variance show several spurious rejections. Spuriousity can also result from the presence of large influential additive outliers (Franses and Haldrup, 1994), quoted in Gabriel *et al.,* (2001).

Conversely, Lee, Huang and Shin (1997) show that ignoring structural shifts will bias a test towards rejecting the null hypothesis of stationarity in favor of a false null hypothesis of a unit root. The distribution of stationary test is asymptotically invariant to the exclusion of the existing shift if the alternative hypothesis of a unit root is true, therefore there will be no power or efficiency losses (Lee *et al.,* 2000). Because of this Gabriel et al (2001) suggest that rejection of the null hypothesis is relatively dependent on the construction of the null hypothesis. Some stationary tests may have power against structural change due to the way they are constructed. In the test characterized by structural change, the unit root test tends to under reject the null hypothesis of unit root. According to Maddala and Kim (1998), the poor performance of the co-integration test under structural influence indicates that the assumption of co-integration relationship based on the constant co-integration (i.e., $\rho = 1$) is not appropriate.

Analysis of the effect of breaks has been preceded in the literature by the statistical testing for the existence of a break including tests for the number of breaks. Allen and Fildes (2004) suggest that structural change only has relevance in the context of a particular model. An approach suggested by Allen and Fildes (2004) is to detect the break point by standard structural test, using all the data set and treat a single break as exogenous (because it is hard to reject the null of a unit root when the break is considered endogenous). Multiple breaks can be specified and tested for. Several break tests have been criticized for testing for existence of a break(s) without testing for the break points (Maddala and Kim, 1998). For example, Perron (1989) showed that the standard test of unit root hypothesis against trend stationary alternatives cannot reject the unit root hypothesis, if the true data generating mechanism is a stationary fluctuation around a trend function which contains a one-time break. Accordingly, this test procedure was conditional on a change occurring at a fixed known date, and the breaks were assumed exogenous[5]. In a study on Agricultural GDP in Ethiopia, Alemu, Oosthuizen and van Schalkwyk (2004) measured the extent of the sustained impact of a break or a shock on agricultural GDP. Their approach differs from that of Perron (1989) in that the date of structural break was treated as unknown a priori. Their recursive regression test procedure identified one significant break point out of the three events considered in their investigation.

2.4.4 Unit root and Outliers

Structural breaks are a kind of outlier. A structural break causes a break on the trend while outliers cause a break on the mean (Perron, 1998). Outliers are unusual observations that are away from the rest of the data points. Outliers manifest themselves as unusually large residuals, i.e. the residual representing positives or negatives between the actual values of the regressand and its value estimated from the regression model. Outliers may be due to recording error, policy changes or major events. Outliers may also be caused by misspecification of the relationship between variables. For example, a non-linear relationship may be misspecified as linear, or there may be omission of important variables. Just like structural breaks, outliers cause size distortion in unit root test. Under rejection or over rejection of the null hypothesis of unit root introduces inferential bias. The solution to the problem of outliers is to detect and possibly discard them. Automatic rejection of outliers is however not a wise procedure, because the outlier sometimes provides information other data points cannot. Outliers may arise from an unusual combination of circumstances which may be of vital interest and require further investigation rather than rejection (Gujarati, 2003). Model transformation or change may be recommended in a severe outlier's problem. But in most cases additive outliers are taken care of through the use of dummy variables. Alternatively, unit root tests that are robust to moving average (MA) errors correct for additive outliers (Perron and Ng, 1996).

[5] Assumptions of endogenous breakpoints have been found by many to be totally unacceptable. According to Maddala and Kim (1998) if a search is conducted it should be around the events.

2.5 Co-integration analysis

Co-integration analysis is an alternative procedure for evaluating spatial market linkage by taking the presence of stochastic trends in the price series into account. It was developed and applied in earlier work by Engle and Granger (1987) and also Engle and Yoo (1987). Co-integration analysis ensures that deviations from equilibrium conditions between two economic variables which are individually stationary in the short-run should be stationary in the long-run. Intuitively, the concept of co-integration implies that economic forces should prohibit persistent long-run deviations from equilibrium, even though short-run deviations may be observed (Goodwin and Schoeder, 1991 and Negassa *et al.,* 2003). An important implication of this is that, while individual economic variables such as price drift apart, certain pairs of such variables should not diverge from one another in the long-run (Goodwin and Schroeder, 1991). Co-integration thereby admits instability in market margins, which need only be stable in the long-run, but not fixed (Barrett, 1996). Therefore, if market prices are co-integrated, thus, the markets concerned are integrated (Alexander and Wyeth, 1994; Goletti and Babu, 1994; Goletti, Ahmed and Farid; 1995, Dercon, 1995; Goodwin and Schroeder, 1991; Negassa *et al.,* 2003). In other words, market integration is an indication of interdependence. Co-integration has been regarded by many researchers as not absolute but a measure of degree of market integration (Gonzalez-Rivera and Helfand, 2001, and Goodwin and Schroeder, 1991). Spatial market prices that diverge from each other for a long time would have a weak long run relationship while two prices that co-move are likely to be co-integrated. According to Goodwin and Schroeder (1991) various factors affect co-integration, e.g. transaction cost, risk associated in transacting business and influence of volume of trade. Low-volume markets have the tendency of large price variability and the distance between markets has a great influence on transaction costs. This idea is consistent with the findings of Goletti *et al.,* (2005) in describing the influence of structural factors in determining market integration.

According to Barrett (1996), co-integration is unfortunately not a sufficient tool for spatial market analysis. Negassa *et al.,* (2003) pointed out that, if transaction costs are non-stationary, a failure to find co-integration between two market price series may be completely consistent with market integration (Barrett 1996; Negassa *et al.,* 2003). In other words, co-integration may be assumed unnecessary, because price can be co-integrated without the market being integrated or efficient (Baulch, 1997; Negassa *et al.,* 2003; Fackler, 1996, and Barrett, 1996). Barrett (1996) asserts that the insufficiency of co-integration as a tool for spatial market analysis stems from the fact that, if the coefficient of prices in the central market is negative, a negative relationship is observed, implying that prices move in the opposite direction rather than co-movement as indicated by the concept of market integration. The magnitude of the price co-integration coefficient may diverge far from one, contradicting the intuition behind market integration hypothesis. For example, the study by Asche, Bremnes and Wessells (1999) on market integration of Salmon prices was criticized based on the estimation of negative co-integrating vectors, implying that equilibrium Salmon prices move in an opposite direction.

In an imperfect market where competition is marred by market failures and government policy intervention, market segmentation can result from a market price spread greater than transfer cost. If discontinuity in trade flows exists, a co-integration test imposes a linear approximation to a non-linear conditional expectation function. The greater the transfer costs between markets and the more frequent the discontinuities, the more pronounced this non-linearity and the more unreliable the findings of cointegration regression.

According to Baulch (1997), a co-integration test may be deemed as being both an unnecessary and insufficient condition for a measure of spatial market integration. It is an unnecessary condition because if transfer costs are non-stationary, arbitrage between two markets may be efficient even when their price series are not co-integrated. It is an insufficient condition, because the price series may be co-integrated but their price differentials may be too small to offset the transfer cost. The practical importance of co-integration is not as a test for market integration but as a pre-test for other tests of market integration (Alexander and Wyeth, 1994).

2.5.1 Estimating Co-integration relationship

In economic estimation of co-integration relationships, it is often important to integrate short-run economic behaviour with long-run equilibrium relationships. That is, if the short-run relationship is known the long-run relationship can also be identified. But economic theories do not provide much information about short-run relationships. Engle and Granger's (1987) co-integration theory addresses this issue of integrating short-run dynamics with long-run equilibrium.

Equilibrium theories involving non-stationary variables require the existence of a combination of the variables that is stationary. An equilibrium relationship, among a set of non-stationary variables, implies that their stochastic trends be linked and do not drift apart as time goes on. Since the variables cannot move independently, their linkage necessitates that the variables be co-integrated. Therefore, estimating the co-integration relationship is necessary to reveal the existence of a long-run relationship.

2.5.1.1 Order of Integration

Assuming that each series of two economic variables P_t^1 and P_t^2 are non-stationary and require a single differencing to remove trend and make them stationary, the residual approach relies *a priori* on the premise that if the two series are co-integrated of the order I(1,1) then the residual from the co-integrating regression should be I(0). A series is integrated of the order (d) if it must be differenced d times to be stationary. Two series are integrated of the order (d,b) if individually they are integrated of the order d and b and their linear combination is integrated of the order (d-b), where b>0 (Engle and Granger, 1987). Practically, differencing series has been found to remove the long-run relationship which co-integration estimation intends to find. Co-integration is thus confirmed by testing the residuals using the various test statistics.

Granger's representation theorem recommends that if a co-integration test is accepted (i.e., residuals are white nose) then error correction models can be developed to study the short-run price relationship (Engle and Granger, 1987). Conversely, the theorem also applies if the series are individually I (1) since their linear combination will be I (0). This does mean that all variables which are integrated of the same order are co-integrated. A set of variables integrated of different orders are not co-integrated (Maddala, 1992 and Enders, 2004). It is nevertheless possible to find relationships among groups of variables that are co-integrated. This can be found in a multicointegration situation. Several methods have been used to estimate co-integration in the literature.

2.5.1.2 Engle and Granger two-step estimation procedure

Engle and Granger (1987) suggest two-step estimation procedures. The co-integration regression is firstly estimated by simple ordinary least squares, obtaining the residual of the co-integrating relationship and applying a residual-based unit root test for co-integration. The second procedure is the estimation of the error correction model (ECM). This is based on the representation theorem of Engle and Granger (1987; which states that, if a set of variables are co-integrated of the order (1,1), then there exists a valid error correction representation of the data.

Engle and Granger's (1987) approach has been criticized. Diakosavvas (1995) argues that the use of OLS is inconsistent and inappropriate because both variables P_t^1 and P_t^2 are non-stationary. Besides, a long-run relationship is not guaranteed, given the value of co-integrating parameter β. He suggests using the autoregressive distributed lag model (ADL) developed by Phillips and Loretan (1991). Phillips and Loretan (1991) estimated β in a non-linear least square (LS) method compared to Engle and Granger's linear estimation. Lack of a lag term in the single equation approach of Engle and Ganger may lead to bias in the estimation of β in finite samples (Maddala and Kim, 1998). Therefore, unrestricted ECM and the use of a large number of lag length is recommended. Engle and Granger were criticized for arbitrarily imposing normalization with respect to the dependent variable. According to Ng and Perron (1997), least square estimators will have poor finite sample properties when normalized in one direction, but perform better in a two-way normalization. In other words, it is better to use a less integrated regressand (Ng and Perron, 1997).

2.5.1.3 The system estimation method

In the Engle and Granger (1987) estimation procedure discussed above, only one co-integrating parameter (vector) is expected, subject to the normalization problem. The system method avoids the problem of normalization. The number of co-integrating vectors is not fixed a priori but is determined through estimation. Even though known vectors can be used, Balke and Fomby (1997) recommend estimating them.

The Johansen procedure is the most commonly used system estimation method. Other methods include the Box-Tiao, and Principal Component methods. The Johansen method is based on canonical correlation analysis. The procedure utilizes two tests, the trace test and the maximum eigenvalue test. The trace test is used to test the hypothesis of **r = 0** against the alternative of **r+1** co-integrating vector. The maximum eigenvalue test is used to test the hypothesis of **r = 0** against the alternative of **r = 1** co-integrating vector. The asymptotic distribution of the test is given by the maximum eigenvalue of the stochastic matrix, while that of the trace test is given by the trace of the stochastic matrix. The Box-Tiao test differs from Johansen's only in the specification of the order of integration of the dependent variable. Johansen uses I (0) while Box-Tiao uses I (1). According to Johansen, a linear combination of correlated ΔP_t I (0) and P_{t-1} I (1) series will yield residuals that are uncorrelated. However, the Johansen test is criticized for being based on the assumption of iid in error term. The test is susceptible to a type I error with the null hypothesis of no-co-integration when the errors are not iid (Huang and Yang, 1996). The test can detect spurious co-integration which does not exist. It has been found to have a high variance and to be more likely to produce outliers than other tests (Maddala and Kim, 1998, and Phillips, 1994). In the principal component method, co-integrating vectors are identified as components with maximum variance which corresponds to eigenvalue of the covariance matrix $\sum$**.** The principal components corresponding to the smaller eigenvalues give the co-integrating vectors and those corresponding to the larger eigenvalue give the common stochastic trend.

2.5.2 Empirical application of the co-integration test

This section reviews studies that applied the concept of cointegration with the aim to compare various ways the concept is used in relation to spatial market analysis.

Several researchers have applied the co-integration test to study the spatial interdependence of markets as measured by price relationships. Diakosavvas (1995) used co-integration analysis based on the ordinary least square approach and the time varying parameter estimation approach to examine the market integration between Australian and United States (US) beef prices at the farm gate. Using monthly time series data from 1972:1 to 1993:2 co-integration was found between Australian and U.S. beef prices. The time varying convergence analysis indicates the degree of convergence between the various price pairs has not increased over time. The result implies that Australian prices cannot be adopted as the world price in empirical analysis.

Asche, Bremnes and Wessells (1999) used the concept of co-integration and the law of one price tested for product aggregation, market integration and relationships between prices in the world salmon market from 1986 to 1996. The aim was to investigate whether different species of salmon compete in the same market. The Johansen multivariate test was used to conduct tests for law of on price LOOP and co-integration. Their result indicates, firstly, that the markets are co-integrated. This implies that five species of salmon used in the analysis compete in the same market. Secondly, the LOOP and the composite commodity theorem hold. Their findings have implications for trade restraining policy measures. The analysis supports the claims that all salmon prices move together in the long-run (co-integrate) with some short term deviations. This calls for relaxation on restrictions on trade measures in the long-run. The results also have significant implications for European Union regulations regarding minimum prices on imported Norwegian salmon and the plan by the Norwegian government to reduce supply to drive up price.

However, Miljkovic and Paul (2001) criticized the study by Asche, Bremnes and Wessells (1999) on the grounds that the analysis ignored the concept of market tradability. Positive trade flow is a strong indication of market integration (Barrett, 1996), and markets can not be integrated if there is no trade. In addition, transaction costs were completely overlooked. When prices and transaction costs are non-stationary, co-integration is not necessary for markets to be in equilibrium (McNew and Fackler, 1997). Miljkovic and Paul (2001) pointed out that co-integration is not sufficient to conclude integration, noting that the signs and magnitude of changes in long-run equilibrium values of salmon prices run contrary to the equilibrium concept for substitute commodities.

Goodwin and Schroeder (1991) used a co-integration test of regional price series to evaluate spatial linkages in regional cattle markets in the U.S. The aim was to determine the impacts of co-integration of several regional markets characteristics. Weekly price series were used from January 1980 through September 1987. Co-integration tests were conducted on spatial price relationships among eleven regional markets. The result was that several markets were not integrated over 1980 through 1987. Market volume, industrial concentration ratio, and distance between markets were found to significantly influence co-integration.

Liu and Wang (2003) used Johansen's multivariate co-integration to test for egg market integration of six Pacific states in the US (Washington, Idaho, Oregon, California, Nevada, and Arizona). Annual price data from 1960 to 1996 were used. There are physical flows among the six pacific states which indicate integration. However, the law of one price was rejected by testing linear combinations of co-integrating vectors. Due to transaction costs, the LOOP has also been rejected in many other studies using the integration test involving non-stationary series (Baffles 1991; Ardeni, 1989, and Asche *et al.,* 2004). The authors conclude that eggs from these states substitute for each other or to some degree, and arbitrage opportunities through trade bind the egg prices.

McNew and Fackler (1997) dispute the notion that co-integration methods to test for market efficiency and integration. They distinguish between the concepts of efficiency and market integration. Using a simple spatial equilibrium model to stimulate price behavior, they show that in a well integrated market, prices need not be co-integrated. The number of co-integrating relationships among prices is furthermore not a good indicator of the degree of market integration. Two models, the point-space model of Enke (1951), Samuelson (1964) and Takayama and Judge (1964), and the agent on links models of Hotelling (1929) and Smithies (1941), as quoted in McNew and Fackler (1997), were used in their analysis.

2.6 Parity bound model approach to market integration

The various methods used in analyses of efficient testing for spatial market linkage discussed so far have mostly been based on assessments of co-movement of price series or long-run relationships between prices. Many of these assessments have used time series econometrics price data and have assumed or failed to recognize the crucial role played by transaction costs. The parity bound model approach to spatial market analysis, first developed and applied by Spiller and Haung (1986), and Spiller and Wood (1988), and further developed by other researchers (Sexton *et al.,* 1991, Fafchamps and Gavian, 1996; Baulch, 1997; Barrett and Li, 2002; Park *et al.,* 2002, and Penzhorn and Arndt, 2002), extends the scope of price series analysis by using explicit information on transaction costs at a single point in time. In addition to price data it acknowledges the continuity in trade flows between markets.

According to Baulch (1997) and Penzhorn and Arndt (2002), conventional analysts make erroneous assumptions about the continuity of trade flows between markets and the nature of price formation in a multi-market system. The conventional analyst assumes that transaction costs and trade flows are unobservable phenomena and hence are constant. Advocates of parity bound models criticize the estimation of transaction costs based on inter-market differentials, especially when trade flows between two markets are infrequent but occur regularly between each of the two and a third market (Baulch, 1997). In this circumstance price differential will not reflect the cost of inter-market transfer of commodities.

The parity bound model developed by Baulch (1997) is an extension of earlier studies on stochastic frontier and switching regression models. It allows for transfer cost to vary, makes no implicit assumptions about the nature of marketing margins, and may be estimated using time series data that are incomplete in contrast to assumptions made under traditional analysis. Baulch (1997) used a parity bound model, and estimated transaction costs using structure, conduct and performance studies by physically examining transportation, interviewing traders, tracking shipments, and looking for unexploited arbitrage opportunities. Because transaction costs is time variant, an extrapolation from observed transaction costs in one period is used as an estimate over the entire time series.

The parity bound model has detected violations of the spatial arbitrage conditions with a high degree of accuracy when estimated with sample sizes that are typical of short food price series in most developing countries (Baulch, 1997). It allows for transaction cost, trade reversal, and autarky. It measures the probability of being in different spatial market efficiency regimes over the sample period. The parity bound model can be indicative not only whether markets integrate but also the extent to which they do so. The model specification of the parity bound model is discussed in Appendix B.

2.6.1 Modifications of the Parity Bound Model (PBM)

Earlier work on PBM by Spiller and Huang (1986) and Spiller and Wood (1988) attempted to resolve some of the methodological problems encountered in arbitrage models that were based on contemporaneous relationships and point-space trading characterizations of markets. In the Spiller and Huang (1986) approach, prices in one region are not regarded as predetermined and transaction costs are endogenously determined. Market integration is not viewed as an absolute preposition in all regions. That is, regions often may rather be linked by arbitrage in some periods, but not in others, depending on the supply-demand market equilibrium in each region at a particular time as well as the magnitude of transaction cost. Therefore, the probability that markets are integrated is allowed to vary continuously. However, the methodology of Spiller and Huang (1986) has come under various criticisms. Sexton, Kling, and Carman (1991) suggest that production of many agricultural products is concentrated in few regions (supplies are from surplus to deficit regions). Therefore, this does not conform to the concept of regional autarky markets implied in the Spiller and Huang (1986) model. Baulch (1997) and Penzhorn and Anrdt (2002) extrapolated transaction costs exogenously from known transaction costs in a single time periods rather than an endogenously estimated parameters as adopted by Spiller and Huang (1986), and Sexton *et al.,* (1991). This facilitates accurate sample separation between regimes and avoids the difficulty with estimating transaction costs based on inter-market price spreads.

Negassa, Myers, and Gabre-Madhin (2003) suggest statistical tests of structural changes in the probabilities of different trade regimes due to marketing policy change. It will help to determine the length of time required before the full effects of the marketing policy change is realized. Barrett and Li (2002) introduced a new method of spatial price analysis based on maximum likelihood estimations of a mixture of distributional models incorporating price, transaction costs and trade flow. They distinguished between market integration from competitive spatial equilibrium and derived four measures of inter-market trade, perfect integration, segmented equilibrium, imperfect integration, and segmented disequilibrium. They suggest that one cannot observe all possible transaction costs values; therefore trade flow information can offer indirect evidence on the effect of omitted transaction costs, thereby providing full information with which to analyze market relationships. The PBM was extended to include trade flow data and a joint probability distribution of inter-market margin (rent) and transaction costs is estimated to include the four market conditions. According to the researchers one can also use these estimates to derive semi-parametric measures of time-varying regime probabilities to track changing market conditions. In this case six regime probabilities where used to define the four market conditions.

Apart from modifications, the overall performance of PBM has been criticized on many grounds. Fackler (1996) argues that there is no link between economic theory and the distributional assumptions used in the switching regime model. Therefore, to assess the statistical reliability of PBM and respond to the argument by Fackler (1996), a Monte Carlo experiment has been used to test for the sensitivity of the result to the distributional assumptions made (Baulch, 1997; Penzhorn and Arndt, 2002; Barrett and Li, 2002). A PBM handles only a limited number of markets. It may be misleading because it considers short-run deviations from equilibrium as inefficiency, whereas it may actually represent traders' rational response to a lag in information and trade flows.

2.7 Linear versus non-linear time series modeling

The aim of this section is to identify the need for non-linear modeling of stochastic processes. Several non-linear models are compared and the reason for the choice of model used in analysis is given.

Several stochastic processes used in modeling time series are usually assumed to be linear. Perhaps, these assumptions can be useful in describing the actual time-path or behaviour of economic variables. But errors could be made if policy makers are guided by these assumptions when some (or all) of the time-paths are actually non-linear. In that case economic adjustment processes will follow a non-linear pattern. For example, it has been established that downturns in the business cycle are sharper than recoveries because key macroeconomic variables, like output and unemployment, fall more sharply than they rise (Enders, 2004).

Recent economic theory suggests that a number of important time-series variables exhibit non-linear behaviour. Non-linear models have been developed to address the problem. Examples of some non-linear models are the generalized autoregressive (GAR) model, the bilinear (BL) model and the threshold autoregressive (TAR) models. The GAR is used to model series with unknown functional form. It approximates an unknown functional form by using differentiable functional forms with polynomial terms not more than the order of three. The GAR model extends the standard AR model by including various powers of the lagged AR process. The BL model is an extension of the autoregressive moving average (ARMA) model by adding cross products. It utilizes the interaction of the ARMA terms to approximate a high-order GAR. The weaknesses of these two models are that while GAR model can easily be estimated with OLS, BL models are more complicated and are estimated using maximum likelihood functions. Both models are useful only when the functional form of the non-linear process is unknown. To compare the speed of adjustment and performance of AR, GAR, BL, and TAR processes, Enders (2004) used a simulated time-path of the four models and found that the TAR process displays a substantial degree of mean reversion than the other models. GAR returns to its mean far more slowly than others. As measured by the residual sum of squares, the TAR model is a better fit than AR (2), GAR but not than the BL model. The weakness of the TAR model is the assumption of equal residual variance and it is criticized for being overparametized (Enders, 2004).

The TAR model has however been applied in the literature to study market integration and other related studies. Goodwin and Piggott (2001) compared the speed of adjustments of AR and TAR models and found that threshold models indicate adjustments that are generally about twice as fast as those implied by the standard AR model. Obstfeld and Taylor (1997) found that TAR models yield empirical estimates of threshold and convergence speed that are more reasonable than that of the AR process. In this study TAR models will be used to study the price transmission mechanism and adjustment processes in South African apple markets. The TAR process will be discussed in detail in the next section.

2.8 Threshold Co-integration

Traditionally, economic time series models require the co-integrating relationship underlying the price transmission process to be linear. Recent studies have however recognized the potential for non-linear threshold-type adjustment in error correction models. Time series non-linear models were originally introduced by Tong (1978). Tsay (1989) developed a method to test for threshold effect in autoregressive models and to use it to model threshold autoregression. Balke and Fomby (1997) applied the threshold autoregression to co-integration framework which can be used in market integration analysis. Balke and Fomby (1997) noted that threshold co-integration can be used to explain the fact that economic time series exhibit both local and global characteristics. While they may locally have a unit root, they are stationary globally (co-integrated). According to Balke and Fomby (1997), co-integration is a global characteristic of the time series while the threshold regimes are the local characteristics. Because it is difficult to test the null hypothesis of threshold co-integration, Balke and Fomby (1997) suggest a two-step approach in which co-integration and threshold behaviour are tested differently. They suggest Engle and Granger's single equation approach or Johansen's system approach where threshold co-integration implies a threshold vector error correction model. However, the concept of transaction costs is embedded in threshold co-integration analysis.

Contemporary economic time series analysis has recognized the important role of transaction costs in spatial market analysis. This has resulted in a new approach to the econometric study of spatial market linkages. The co-integration-based test of market integration has been criticized on the premise that results are inconclusively drawn due to the omission of transaction costs (McNew and Fackler, 1997; Baulch, 1997; and Barrett, 1996). Spiller and Wood (1988), Sexton, Kling and Carman (1991), Baulch (1997) and Penzhorn and Arndt (2002), with direct observation on transaction costs, used endogenous switching models which account for multiple regimes due to transaction costs to study the degree of spatial market linkages. Recently, the influence of transaction costs has been used in a threshold autoregression and co-integration framework (Balke and Fomby, 1997; Goodwin and Piggott, 2001).

Threshold co-integration is based on the assumption that the presence of transaction costs which is unobservable creates a "neutral band" within which prices in different markets are not linked (Balke and Fomby, 1997, Goodwin and Piggott, 2001 and Goodwin and Harper, 2000). Price equalizing arbitrage activities are triggered only when localized shocks result in price differences which exceed the "neutral band". The aim is thus to estimate a threshold caused by transaction costs that must be exceeded by deviation in prices before provoking equilibrating price adjustment which leads to market integration. Additionally, Goodwin and Piggott (2001) observed that threshold models imply faster adjustment to deviations from equilibrium conditions than in the case where thresholds are ignored. Co-integration tests have often been used to characterize long-run equilibrium relationships between economic variables. Usually, in a co-integrating system the error correction model is used as an adjustment process through which the long-run equilibrium between co-integrating variables is maintained[6]. Balke and Fomby (1997) describe this equilibrating system as discrete, and add that the presence of a fixed transaction costs may prevent economic agents from adjusting continuously.

Therefore, threshold effects occur when shocks above a certain threshold bring about a symmetric or asymmetric response (Goodwin and Piggott, 1997; and Goodwin and Harper, 2000). Due to transaction costs arbitrage trading will cause mean–reversion to equilibrium when the deviation from equilibrium is large. In this case, the benefit of adjustment exceeds the cost, and therefore economic agents will act to move the system back to equilibrium. On the other hand, smaller deviations will reduce the incentive to trade and market players will fail to converge. Infrequent trading will in the long-run restore the system to equilibrium. However, the impact of these mean reversions will depend on the size of the deviation from the no-arbitrage condition (Martens and Vorst, 1998). A shock may have to be of a particular size to provoke a particular response. If the equilibrium error is less than the threshold value, there will be no tendency for mean reversion to equilibrium, there will be no co-integration, and the series will exhibit unit root. If the equilibrium error exceeds the threshold, arbitrage transactions will cause mean-reversion back to equilibrium; the series becomes stationary with zero mean. However, the equilibrium error will not change very much when deviations from zero are small. In other words, in the no-arbitrage regime, where the "neutral band" occurs the AR (1) coefficient will tend to be one when deviations are small. In the outer regimes the coefficients will be much smaller when the equilibrium error is large as to reflect mean-reversion towards zero. That is, the series are co-integrated if they diverge too far away from the equilibrium relationship but are not co-integrated if they are close to equilibrium (Balke and Fomby, 1997). Examining the AR (1) in the outer regime is not enough to determine whether the series is stationary because even if the AR (1) coefficient is equal to one in all regimes, the series may still be stationary. As long as the drift parameters act to

[6] In market integration models, the error correction model (ECM) specification has gained popularity because of its intuitively appealing interpretation (Meyer, 2004)

push the series back towards the equilibrium band the series is stationary even though it may exhibit random walk (Balke and Fomby, 1997).

Balke and Fomby (1997) identified three types of threshold regime models. Firstly, the "Equilibrium-Threshold autoregressive model (EQ-TAR)", shows that equilibrium reversion is towards the center of the band and not the edge. In this model the strength of the error correction effect depends, in part, on how far the variable is from the equilibrium relationship. A second alternative is the "Band-TAR", which represents the case where the process returns to an equilibrium band rather than to an equilibrium point. Thirdly, the "Returning–TAR" model is where reversion is in form of a random walk in every regime with drift parameters that move the process back to equilibrium when the process is outside the threshold. The RD-TAR model also has the property of returning back towards an equilibrium band rather than an equilibrium point. Balke and Fomby (1997) showed that all of these processes are globally stationary (that is, they are co-integrated in the long-run), but they have different short-run dynamics. The three models are compared according to the persistence of their constant parameters in the outer regime. Balke and Fomby (1997) suggest that Band-TAR models are more persistent that others. Obstfeld and Taylor (1997) suggest that the "RD-TAR" model is the most persistent. In this study we will use the Band–TAR model. The adjustment process in the Band–TAR model returns to an equilibrium band instead of an equilibrium point. Secondly, there is continuity of the process at the threshold, unlike in EQ-TAR were there is no continuity at the threshold (Obstfeld and Taylor, 1997).

2.8.1 Estimation of threshold

This section discusses methods of estimation of threshold lags, threshold values and test procedure for threshold behaviour.

Tsay's (1989) procedure involves the estimation of threshold lags (p, k, d), testing for the existence of threshold behaviour in the estimated equilibrium error from the co-integration relationship followed by the location of a threshold value(s). A search algorithm is embarked on to locate the threshold values that define the non-linear series into regimes characterized by transaction costs.

Firstly, the autoregressive order p, and other threshold lags are arbitrarily selected. The AR order (p) is selected using a partial autocorrelation function (PACF). PACF measures correlations between time series observations that are k time periods apart after controlling for correlation at intermediate lags (Enders, 2003). If the ADF test conducted on the residual of the OLS levels regression is lagged k times, the number of p can be determined from the number of AR coefficients that are significant from the t-ratio. The AR order may differ from regime to regime (Tsay, 1989), and from one linear combination to another. Akaike information criteria (AIC) can be used to select p (Tong and Lim, 1980). Tsay (1989) prefers PACF because according to him, AIC best defines linear models and could be misleading if the process is non-linear. PACF provides good approximation for non-linear models, besides it can be refined if desired.

The delay parameter d defines the number of lags appropriate to the error correction term in the threshold autoregression. It gives an indication of the speed by which markets respond to the deviation from the no-arbitrage relation (Martens, Kofman and Vorst, 1998; Goodwin and Piggott, 2001).The equilibrium error determines the regimes. Because it could take more than one period for regime switch to occur, regime switch is alloyed to occur according to the value of equilibrium error lagged d times where d = 1, 2, 3…. Tsay (1989) suggests that the choice of d is optional. Many researchers prefer d to be one. Other parameters can be used to select the value of d .The value of d can be chosen based on the minimum of Alaike information criteria (AIC) or Bayesian information criteria (BIC), especially when the optimal value of p and the threshold depend on the value of d . Tong and Lim (1980) used AIC. Tsay (1989) proposed a procedure that is based on the performance of the F- statistics. If the AR order p is known, Tsay (1989) suggests selecting the value of d that maximizes the F-statistics. In this study d is chosen based on PACF.

2.8.1.1 Testing for non-linearity

The aim of this section is to examine various non-linearity test procedures. The choice of method used will also be discussed.

Many procedures have been developed to test non-linearity. Most tests are based on examining the residual of the linear combinations for the presence of autocorrelation. Some examples of the residual-based portmanteau test are the autocorrelation function test (ACF), the McLeod-Li Test, the Regression error specification test (RESET), the Brock, Dechert, Scheinkman and Lebaron (1996) Test (BDS), quoted in Maddala and Kim (1998) and the Tsay (1989) F-Test. A non- residual based test is the Lagrange Multiplier Test (LM).

The ACF test is based on examining the ACF of the squared or cubed values of the time series data. The test is carried out using simulated samples and checking the statistical significance of the correlation coefficients. The ACF test may fail to detect non-linear relationships because it is ideally suited for linear models. The McLeod-Li Test is used to determine if there are significant autocorrelations in the squared residuals from a linear regression using any standard linear model example OLS. The Ljung-Box statistic is used to determine whether the squared residuals exhibit serial correlation. The Ljung-Box statistic has an asymptotic chi square distribution with n degrees of freedom if the residual sequence is uncorrelated. If the null hypothesis of linearity is rejected, the series is non-linear. An alternative way of conducting the McLeod-Li Test is by regressing the estimated residual on its squared and lagged values. The resulting coefficient will be zero if the series is linear. In a large sample, if the non-linearity is not found the test statistics will converge to a χ^2distribution with n degrees of freedom. In a small sample, F-statistics can be used to test the null hypothesis that joint coefficients are zero. The RESET test is based on the assumption that if the residual from a linear model are not correlated, they will not be correlated with the explanatory variables used in the series that generated them. The test is therefore run by regressing the residual from an OLS regression on the constant and explanatory variables. If the model is linear the sample value of the F-statistics should be small. The null hypothesis of linearity is rejected if the sample value of the F-statistics for the null of joint equal coefficients exceeds the critical value from a standard F-table.

The BDS test examines the distance between different parts of the residual series to determine if they are correlated. If all the values of the residual series are independent, then the probability that the distance between any pair of the residual series is less than a given distance should be the same for all combinations of the variables. The BDS test can detect serial correlation, parameter instability, neglected non-linearity and structural breaks. The

BDS test does not however have good small sample performance. Other methods for the non-linearity test include the cumulative sum (CUSUM) test statistics developed by Petruccelli and Davis (1986), and the sup-Wald test statistics. The CUSUM test is based on cumulative sums of standardized residuals from autoregressive fits to the data. The significant levels for the test under the null hypothesis of linearity are obtained from the asymptotic distribution of the cumulative sum as Brownian motion (Petruccelli and Davis, 1986). The maximum sup-Wald for structural breaks can also be used to test for non-linearity. In the Sup-Wald test, the chow-type test for structural breaks is applied on an arranged regression under the null hypothesis of no-structural break against the alternative of structural break.

The general weakness of all the residual-based tests is that they have general alternative hypotheses. They can be used to determine whether non-linearity is present in a series but can not detect the form of non-linearity preset. The Lagrange Multiplier test has been found to detect non-linearity as well as the form of the non-linearity present. But the parameters of the Lagrange function must be identified under the null hypothesis of linearity. In the Lagrange Multiplier test, auxiliary regression is run with estimated residual and partial derivatives of the function. The value of the test statistics has a χ^2 distribution with degrees of freedom equal to the number of regressors used. If the calculated value of the test statistics exceeds the critical value from the χ^2distribution table, the null hypothesis of linearity is rejected. One of the popular residual based tests in the literature is the Tsay's (1989) F-test. The test performs well for both small and large samples. Since the number and location of thresholds are unknown Tsay (1989) suggests there is no most powerful test for threshold non-linearity. Tsay's (1989) F-test has power, it is simple and it is feasible. Therefore this study uses it as the method to test for non-linearity. The test is based on arranged autoregression and predictive residuals. An arranged autoregression is an autoregression with residuals re-arranged, based on the values of a particular regressor. The residual from the recursive (LS) autoregression of the arranged residual is used to test for non-linearity by regressing the predictive (or standardized) residual on the arranged autoregressor of the stationary series using (OLS). A significant F-statistic of the resulting regression rejects the null hypothesis of linearity and confirms the presence of non-linearity in the series. Non-linearity in the series indicates a threshold co-integration process. Following the confirmation of threshold behaviours in the series, the TAR analysis will then be extended towards establishing the number and location of the threshold values.

2.8.1.2 Locating the value and position of threshold

The aim is to locate the number and location of potential threshold(s) values. Tsay (1989) suggests using internal estimate or point sample percentiles as point estimates to locate thresholds. Accordingly, thresholds are not found at the extreme points on the percentile because of a lack of enough observations to provide efficient estimates (Tsay, 1989). Following this, Tsay (1989) and CUSUM test statistics suggest discarding 10 percent of the data (i.e., the first 10 and the last 10 observations.). For the one-break sup-Wald test Balke and Fomby (1997) suggest using the interior 80 percent of the arranged sample. For the two-break sup-Wald test sample the 5^{th} to 30^{th} percentiles of the arranged residual sample are examined to locate the lower threshold (first regime). For the upper threshold (third regime) the search is from 70^{th} to 95^{th} percentile. The 40^{th} percentage of the sample falls in the middle regime. The number of observations in the algorithm search is expected to be reasonable enough to increase the chances of locating a potential threshold. Thus, the threshold must lie between maximum and minimum values of the data. At minimum of at least 20 observations is recommended for the search in each regime. Each data point within the chosen band is treated as a potential threshold. The search is conducted on arranged residuals. The order of the arranged residual has been debated. Balke and Fomby (1997) suggest arranging the residuals in ascending order (from low to high). Obstfeld and Taylor (1997) suggest arranging them in both ascending and descending order. The reason is that in small samples the case data may not fall in all the regimes delineated by every threshold value. For the TAR model, arranged autoregression orders the data according to the value of the potential threshold variable rather than by time (Tsay, 1989; and Balke and Fomby, 1997). Tsay (1989) points out specifically that arranged autoregression provides a means by which the data points are grouped so that all of the observations in one group follow the same linear AR model. However, arranged regression allows more power in discerning thresholds for which data are concentrated in a particular regime at either end of the arranged series (Goodwin and Piggott, 2001). It does not change the dynamic relationship between the dependent variable and its lag but if the data follow a threshold model, the threshold translates to structural breaks in the arranged series (Trenkler and Wolf, 2003). An approach recommended by Balke and Fomby (1997) to locate threshold value is the "grid search" method. Two alternative grid search techniques are proposed. Obstfeld and Taylor (1997) use grid search to obtain thresholds that maximize a likelihood function, while Balke and Fomby (1997) used grid search that minimizes the sum of squared error criterion. Tsay (1998) used grid search that minimize Akaike information criteria. For each data point the TAR model specified in TAR model equations is estimated until all the observations within the band are exhausted. According to Balke and Fomby (1997), the regression containing the smallest residual sum of squares contains the estimate of the threshold.

The use of graphic device to locate the number and position of thresholds has been proposed by Tsay (1989). It involves a scatter plot of various t-statistics against the specified threshold variable. The scatter plot of t-values of the constant and AR-coefficients are plotted against the threshold variable. The number and location of the potential threshold values are determined from the scatter plot by simply identifying the points where the t-ratio turns and changes direction as estimates of the potential threshold is reached. Two or more thresholds can be identified in the process, either from the negative or positive sides of the scatter plot. To motivate use of the scatter plot, Tsay (1989) asserts that the t-values indicate the significance of the coefficient of a linear model. For a significant coefficient, the t-ratio converges gradually to a fixed value as the recursion continues. The scatter plot provides information for locating the number and position of threshold. This approach has been criticized for inconsistency and subjective interpretations. Tsay (1998) used another type of graphic device developed by Chan (1993) to find the consistent estimate of a threshold. The sum of squares residuals from a grid search in an arranged autoregression were plotted in a graph. The idea is that the sum of squared residual (SSR) will form local minima at a threshold. The closer to the true threshold, the smaller should the SSR be. Hence, the SSR should be minimized at the true value of the threshold. If there is a single threshold, there should be a single trough; if multiple thresholds exist the SSR will have several local minima.

2.8.1.3 Threshold error correction mechanism

If the threshold is located, the data points are divided according to whether equilibrium error is above or below the threshold and OLS regression is estimated. Dummy variables are incorporated and ECM is specified using TAR models. The coefficient of the inner band (the neutral band) is not expected to be statistically significant because there is no mean reversion in this regime and adjustment is expected to be zero. The outer band is expected to adjust faster than the inner band. Therefore, its coefficient is expected a priori to be larger and to have a significant t-ratio.

2.8.2 Empirical application of threshold co-integration

The aim of this section is to examine the scope of application of the concept of threshold co-integration in the literature.

Goodwin and Piggott (2001) evaluated price linkages in spatially separated markets using threshold autoregressive and co-integration models to account for neutral band representing transaction costs. Using daily prices from January 1992 to March 1999 price linkages among four corn and soybean markets in North Carolina were analyzed. Impulse response functions were used to investigate dynamic patterns of adjustment to exogenous, localized price shocks. The results confirm the presence of thresholds. Impulse response is in agreement with a priori expectations and provides strong support for market integration. Price equilibrium arbitrages were found to occur in response to localized shocks that exceed the thresholds of the neutral band. The significance of transaction costs is shown by the faster adjustment in response to deviations from equilibrium than when threshold behaviour is ignored. According to the study adjustments were complete after 15 days.

Tsay (1989) applied the concept of threshold co-integration in testing and modeling threshold autoregressive processes. Based on an arranged AR and predictive residual, a single F-statistic test proposed by Tsay (1989) was used for threshold non-linearity test. Following specification of the threshold variables a scatter plot graphic device was used to identify the number and position of threshold values. Four-step procedures for modeling TAR models were proposed. Tsay (1989) applied these procedures and test statistics to three real data sets. Firstly, much analyzed sunspot data from 1700 to 1979 with 280 observations and reduced sunspot data from 1700 to 1920 were used for the analysis. Results obtained for both sets of data were similar with those of the previous researchers who used the same data. This is the evidence that the proposed Tsay (1989) procedures can perform well in multiple threshold frameworks. With similar procedures on logged Canadian lynx data with 114 observations and hourly attic temperature data sets with 251 numbers of observations, Tsay (1989) obtained result that agreed with those of previous authors.

The second threshold obtained with attic temperature data was less pronounced. Tsay (1989) proposed an iterative procedure of discarding data cases in the first regime after locating the first threshold and carrying out the same recursion estimate of arranged AR using the remaining cases to identify the second threshold (if present). The result was significant at the 5% level but there were apparent differences in variance.

Much of the work on threshold co-integration in the literature is based on the pioneering work of Balke and Fomby (1997), Tsay (1989), Tong (1978) and a few others. Balke and Fomby (1997) suggest two-step procedures for examining threshold co-integration. They suggested examining both local and global characteristics of the time series, by testing for co-integration and threshold behaviour. They found that standard time series methods developed for testing co-integration in the linear case perform well when threshold co-integration is present.

Threshold autoregression models have been applied in derivative markets to study the role of arbitrage transaction in Index-futures returns. Martens and Vorst, (1998) used 390 price data constructed on 15 seconds basis per day on the New York stock exchange market for two contract months May/June and November/December 1993 to estimate the band around futures prices within which arbitrage is not profitable. Adequate test procedures for threshold co-integration as proposed by Tsay (1989) were used. Applying threshold autoregressive models they found that Index-futures arbitragers only enter the market when deviations from the no-arbitrage bound is sufficiently large to offset transaction costs and associated interest rates and dividends. They showed that the impact of equilibrium error increases with the magnitude of the error and that the information effect of lagged futures returns on index returns is significantly large when the mispricing error is negative.

2.9 Hansen test

The aim of this section is to test the significance of threshold behaviour identified using Tsay's (1989) test procedure as explained in subsection 2.8.1.1. Further analysis and TAR models are fitted conditional on whether the threshold behaviours are statistically significant.

Hansen (1997) developed a distribution theory for least square estimates of the threshold in non-linear TAR models for the following reasons. Firstly, conventional tests of the null hypothesis of the linear standard AR process against the TAR alternative have non-standard distributions, as the threshold parameter is not identified under the null hypothesis of linearity (Andrews and Ploberger, 1994; and Andrews, 1994). Secondly, sampling distributions of threshold estimates may not be free of nuisance parameters. Hansen (1997) suggests that if the threshold effects become smaller as the sample size increases, the sampling distribution of the threshold estimates can be approximated by an asymptotic distribution that is free of nuisance parameters. Hansen (1997) showed that a bootstrap method can be used to replicate asymptotic distributions.

Bootstrapped samples with mean zero and variance of one are used as dependent variable in a chow-type Monte Carlo simulation test. The observed test statistics are obtained from the grid search with observed sample data. From the simulated sample of test statistics, the asymptotic p-value is calculated by taking the percentage of the test statistics for which the test taken from the simulated sample exceeds the observed test statistics.

2.10 Multivariate threshold co-integration

The aim is to distinguish between symmetric and asymmetric threshold autoregression and examine the need to fit threshold vector error correction models.

In a univariate threshold co-integration, the adjustment process may exhibit symmetry. The neutral band relates to symmetric threshold band or regimes of no arbitrage, where the error correction term behaves like a random walk. This implies symmetric adjustment of coefficients and threshold values since arbitrage is induced in the same way no matter where the prices are higher. Some studies show that asymmetry can occur (Goodwin and Harper, 2000; Serra and Goodwin, 2002; Trenkler and Wolf, 2003; Abdulai, 2000, and Agüero, 2004). According to studies by Boorenste, Cameron and Gilbert, (1997), Eckert, (2002) and Palakas, (1995) as quoted in Aguero (2004), prices rise faster than they fall in various commodity markets. Peltzman (2000) found that U.S. output prices respond faster to input increases than to decreases. But why does asymmetric price transmission occur?

According to Goodwin and Harper (2000) and Goodwin and Piggott (2001), asymmetric shocks are assumed to be unidirectional because it is assumed that movement of goods between two spatial markets may occur in one direction, implying that some markets are importing and others exporting the goods. This suggests that market infrastructure is better suited for transactions in one direction. Hence, the transaction costs of exporting markets will tend to differ from that of importing markets. In such a situation threshold values may be asymmetric since movement of commodities in one direction will be more costly than movement in the opposite direction.

2.10.1 Threshold vector error correction models

Analysts have recently proposed the use of threshold vector error correction models (TVECM) to account for non-linear and threshold-type adjustments in error correction models. It can be used to investigate short run price dynamics. It allows for asymmetric adjustments to deviations in response to positive and negative price shocks (Goodwin and Piggott, 2001; Serra and Goodwin, 2002, and Bamba and Reed, 2004). Hansen and Seo (2002) suggest that Balke and Fomby's (1987) threshold co-integration autoregression approach discussed in previous sections involves univariate estimation and testing methods. It is not suitable for investigating asymmetric price adjustments in multivariate TAR. According to them, the approach is valid only for estimating a known co-integration vector even though no theory was provided by Balke and Fomby (1987) as to how the co-integration vector should be estimated.

Different approaches have been proposed for the estimation of threshold co-integration using error correction models. Hansen and Seo (2002) proposed a multivariate threshold vector error correction model. It is based on the assumption that the co-integrating vector and the threshold parameters are unknown and should be estimated. A joint two-dimensional grid search algorithm for each of the estimated parameters is first conducted over the threshold parameter. The trace or the determinants of the variance-covariance matrixes are computed. The values of the parameter estimates that minimize the trace or the log of determinant of the variance-covariance matrix is chosen. However, Hansen and Seo (2002) have been criticized for lack of proof of consistency of parameter estimates and the absence of a distribution theory for the maximum likelihood estimation (MLE). Hansen and Seo (2002) tested for the significance of the differences in parameters across regimes using the Sup-LM test. The hypothesis of linear co-integration against the alternative of no threshold co-integration was tested. Another approach was proposed by Enders and Granger (1998). This approach is based on the use of known co-integrating vectors and threshold parameters. Equality restrictions were imposed on the TAR parameters and a zero threshold parameter was assumed. OLS residual estimates were tested for stationarity, followed by the test of asymmetric adjustment with a null hypothesis of symmetry against the alternative of asymmetry threshold co-integration. If asymmetric threshold co-integration is confirmed, then a TVECM is estimated.

Granger and Teräsvirta (1993) proposed the use of a smooth transmission autoregressive (STAR) model to measure asymmetric price adjustment. In this case adjustment takes place in every period, but the speed of adjustment varies with the extent of the deviation from equilibrium. Granger and Teräsvirta (1993) suggest that even though economic agents react to bring the economic system back to equilibrium, the adjustment process is not instantaneous. Therefore, in contrast with the conventional TAR model, regime changes in a smooth transmission model occur gradually but not abruptly. The adjustment process in STAR is smooth rather than discrete. They argue that because of the discrete assumption in TAR models there is discontinuity at each threshold. This complicates tests for non-linearity and creates doubts on the inferential test of threshold effect (Michael, Nobay and peel, 1997). Michael *et al.* (1997) used an exponential form of the STAR model (ESTAR) to find evidence of mean-reverting behaviour for purchasing power parity (PPP) deviations. Baum, Barkoulas and Caglayan (2001) used ESTAR models to model the dynamics of adjustment to long-run PPP. Evidence of mean reverting dynamic processes was found for sizable deviations from PPP, with equilibrium tendency varying non-linearity with the magnitude of disequilibrium.

2.10.2 Threshold vector autoregression estimation procedures

Threshold vector error correction models (TVECM) can be used to model non-linear and threshold-type adjustments to long-run equilibrium. This happens when a linear TAR model becomes inadequate to measure co-integration because asymmetric price adjustments exist. Enders and Granger (1998) and Balke and Fomby (1997) have shown that standard co-integration tests will lack power in the presence of asymmetric adjustment. Asymmetric price transmission occurs mainly in one direction (Goodwin and Harper; 2000, Agüero, 2004, and Abdulai, 2000). Enders and Granger (1998) call this momentum threshold autoregression (M-TAR), that is, the series exhibit more momentum in one direction than the other. Threshold

VECM can be estimated using sequentially conditional iterative seemingly unrelated regression estimation (SURE). Firstly, a grid search is carried out to estimate the threshold parameters[7]. The search for the threshold is within the range (as specified in linear TAR estimation in subsection 2.8.1.2) of the positive and negative of the lagged error correction terms. Secondly, the co-integrating vectors are estimated (conditional on located threshold values) through OLS regression of the threshold model. The data are divided according to whether equilibrium error is above or below the threshold. From the estimation the log of the determinant of the variance-covariance matrix is obtained from the minimum residual sum of squares for error or from the sum of squares that maximize the likelihood function. Test for the significance of threshold effect is conducted using the Hansen (1997) multivariate simulation method.

2.10.3 Empirical Applications of TVECM

Tsay (1998) developed a generalized non-parametric test statistic based on a predictive residual and an arranged autoregression to detect threshold non-linearity in a multivariate TAR. The Akaike information criterion (AIC) procedure is used for estimating threshold and model selection. The model procedures are applied on security markets to study index futures arbitrage in finance. The data are intraday transaction data for the S&P 500 stock index in May 1993 and June futures contract traded at the Chicago Mercantile Exchange. Based on one minute intervals, an observation of 7060, adjusted for outliers were used. Threshold co-integration and model adequacy was reported, which is consistent with similar studies by other researchers.

Tsay (1998) extended the study to U.S monthly interest rates from 1959.1 to 1993.2 and to two daily river flow series of Iceland, from 1972 to 1974, with 1095 observations. For the U.S. bond market 3-month treasury bills and 3-year treasury notes were used, each having 109

[7] Many analyst estimate the threshold parameters using a grid search that minimizes the sum of squares error (SSE) criterion as proposed by Balke and Fomby (1997). In a multivariate model involving system equations like demand systems and VECM the SSE selection criteria is not appropriate (Serra and Goodwin 2002). Researchers use alternative methods of grid search that minimize the system sum of squared errors, that is, the trace of the variance-covariance matrix for the system's residual errors. In the maximum likelihood estimation, the log of the determinant of residual variance-covariance matrix is chosen. Serra and Goodwin (2002) suggest that the two criteria differ because the former ignores cross equation correlation while the latter accounts for it. As indicated, the latter is a better choice and in this study we choose it as our selection criterion.

observations. In the Iceland series, the dependent variables are the daily river flows (measured in meters cube per second.) In both cases threshold co-integration was found and the results are in agreement with common expectations.

Sera and Goodwin (2002) used an asymmetric threshold vector error correction model to analyze threshold co-integration among farm and retail market prices in a variety of Spanish dairy products. The analysis is conducted using both weekly and monthly price data from June 1994 to December 2000. A high asymmetric price relationship was intuitively expected, given the high perishable nature of the products. Their findings suggest that, although formal tests confirm asymmetry, their effects are modest. Results suggest that the transmission of shocks is unidirectional, running from farm to retail levels. The retail prices adjust to farm level shocks to the raw milk price, but the milk price modestly responds to retail market shocks. The weak response of farm price to retail price shocks are explained by a lack of an organized contracting system and scarce farm co-operative outlets.

Hansen and Seo (2002) tested for two-regime threshold co-integration in a vector error correction model with a single co-integrating vector and a threshold effect in the error correction term, using a monthly interest rate series from 1952 to 1991. The interest rate is estimated from U.S. treasury securities and corresponds to zero-coupon on bonds. Citing low power on non-parametric and non-linearity test of Tsay (1989, 1998), Hansen and Seo (2002) proposed a supLM test statistic for threshold co-integration. Applying various test procedures they found strong evidence of a threshold effect.

2.11 Regime switching

In an asymmetric threshold error correction model there is a possibility of regime switch. Recall that the influence of transaction costs induces threshold. The presence of thresholds represents non-linear breaks characterized as regimes. The asymmetric threshold models may consist of, say, three threshold-regimes. The inner threshold regime (regime I) corresponds to neutral band where the price difference is less than transaction costs and reduces incentives for arbitrage. In the outer regimes (regimes I and III), price differences are persistently more than the transaction costs. A particular regime characterizes the relationship among the prices at a point in time. The characteristics of each realization of the series depends upon whether the error correction term is less than or more than the threshold or whether it lies between the thresholds on either the negative and positive sides. Therefore, the switching model illustrates in which of the regimes each observation falls and the time of the switch among the regimes. Regime I corresponds to large negative errors that lie below the threshold. Regime III corresponds to the large positive errors that lie above the threshold. Regime II corresponds to the errors that are between the thresholds that define regimes I and II. However, the strength of the regime switch depends on the distance between the markets. Goodwin and Piggott (2001) observed that the greatest degree of switching among regimes and most frequent occurrence of price differences exceeding the neutral bands is realized by those markets that are widely separated. Regime switching is used to investigate the extent of integration between spatial markets. Observations that persistently fall in regime II indicate market price differentials that are less than or equal to transaction costs, an indication of integration.

2.12 Generalized impulse response function

The impulse response function is a concept that gives additional information about the dynamic interrelationships among prices. Impulse response analysis is employed to investigate the mechanism of shocks. This concept has been used to analyze the impact of price shocks and the way in which shocks are transmitted among market prices. According to Potter (1998), the impulse response function has been based on the foundation that the economy's dynamic behaviour can be well explained by random impulse generated over time by a constant linear structure. Gallant, Rossi and Tauchen (1993) extended the notion of impulse response to a non-linear time series. Potter (1998) improved the standard linear technique of impulse response function analysis to the non-linear case by defining a generalized impulse response function as a random variable on the underlying space of the time series. Impulse response functions can be applied to both univariate and multivariate time series. Kooper, Pesaran and Potter (1996) extend the use of the impulse response function to study multivariate time series.

In the case of univariate or symmetric adjustment models, the response to a price shock is independent of the history of the time series and the sign and magnitude of the postulated shock (Potter, 1995; Abdulai, 2000; Serra and Goodwin 2002; Goodwin and Piggott 2001; and Goodwin and Harper, 2000). However, asymmetric adjustment models produce impulse response functions that are functions of the history of the price series and the sign and magnitude of the shock (Potter, 1995). The possibility of asymmetric response implies that the size and sign of the shock will influence the nature of response (Goodwin and Piggott, 2001). Baum, Barkoulas and Mustafa (2001) observe that shocks of different magnitude have disproportionate effects and that positive and negative shocks (in absolute terms) of the same magnitude have different dynamic effects due to asymmetry based on the sign of the shock, as observed by Goodwin and Piggott (2001).

Bamba and Reed (2004) used impulse response functions to trace the effect of monetary shocks on current and future values of coffee and cocoa prices in the United States. Their impulse response function indicates that a shock in money supply has a negative and immediate impact on coffee prices. Baum *et al.* (2001) used an impulse response function to support evidence of non-linear dynamic structure to purchasing power parity (PPP). Convergence to log-run PPP in the post-Bretton Wood era was found to be low. Goodwin and Piggott (2001) used an impulse response function to study asymmetric price response to shock in U.S. corn and wheat markets. Strong evidence to support market integration was found. In their findings, responses to shocks are complete after fifteen days. Abdulai (2000) applied an impulse response function to study price transmission in Ghanaian maize markets. Wholesale maize prices in Accra and Bolgatanga were found to respond more quickly to increases than to decreases in central market prices.

2.13 Summary

The chapter illustrated price transmission mechanisms and its impact in characterizing market performance in a perfectly competitive economy. Several theories have found the transaction costs and arbitrage concepts very useful in the analysis of market integration. Faminow and Benson (1996) have observed that transaction costs play a key role in spatial market analysis and should not be ignored. The concept of arbitrage has been found to be consistent with the law of one price (LOOP). Varied conflicting opinions exist about the LOOP as a long-run equilibrium concept. Several analysts found no evidence to support LOOP as having a long-run relationship (Baffles, 1991; and Ardeni, 1989). Michael *et al.*(1997) and Baum *et al.*(2001) found evidence of mean-reversion to long-run deviation from equilibrium purchasing power parity (PPP). Generally, a long-run price relationship has been shown to posit market integration. Several reasons were provided in the text why markets should integrate. If markets are integrated the effects of policy intervention in one market would be rapidly transmitted to other markets. There will be regional balance between surplus and deficit producing areas. Different approaches have been applied to measure market integration. Earlier, analysts measured market integration using the bivariate price correlation approach. This was criticized for inferential bias and methodological flaws. The variance decomposition approach was considered. The approach was criticized because it does not allow for dynamic relationships between prices in different markets. Ravallion (1986) proposed a radial inter-market approach to measure market integration. Ravallion's (1986) approach was criticized for assumption of radial market structure and constant inter-market transfer. Market integration has also been related to structural factors (Goletti *et al.,* 1995).

The most commonly applied method of measuring market integration is co-integration analysis. Even though it has been regarded as a pretest to market integration, it has been shown that finding co-integration is a necessary condition for market integration (Asche, Bremnes and Wesells, 1999). Two methods have been identified as the most appropriate methods of measuring market integration.

These are the parity bound model (PBM) and the threshold co-integration method. The parity bound model posits that a co-integration test that ignores the influence of transaction costs is inconclusive and should not be used as a measure of market integration. Baulch (1997) proposed the estimation of transaction costs and use of regime switching models. Parity bound models are criticized for directly observing transaction costs. Fackler (1996) argues that there is no link between economic theory and the distribution assumptions of PBM. Tong (1978) originally introduced non-linear threshold time series models based on transaction costs. Tsay (1989) developed techniques for testing TAR models. Balke and Fomby (1997) introduced threshold co-integration to the study of market integration. Threshold co-integration has been used both in univariate and multivariate time series to study market integration. Threshold models are easier to estimate and are based on a simple dynamic principle compared to parity bound models. Both univariate and multivariate threshold models will be adopted in this study to investigate price transmission and adjustments to equilibrium in response to changes in the economic system.

However, a few weaknesses have been identified with threshold autoregression. Balke and Fomby (1997) relate the adjustment process prevalent in economic phenomenon to threshold co-integration and posit that presence of a fixed transaction costs may prevent economic agents from adjusting continuously. In other words, threshold co-integration assumes that transaction costsis fixed and that adjustment to deviation from equilibrium is a continuous process. If transaction costs are stochastic, assumptions of TAR become inappropriate. In TAR it is assumed that adjustment is instantaneous. Granger and Teräsvirta (1993) criticize TAR and suggest that adjustment in economic phenomena is not instantaneous but smooth. They argue that because of the discrete assumption in TAR models, there is discontinuity at each threshold. This complicates the test for non-linearity and creates doubts on the inferential test of threshold effect. Baulch (1997) and Sexton *et al.* (1991) criticize TAR for assuming continuous trade flow. According to Sexton *et al*. (1991), failure to observe short-run integration of prices in TAR amounts to normative inefficiency. They argue that even though long-run market integration is regarded as a more realistic concept, short run integration is a benchmark for market integration. The TAR model is criticized for the assumption of equal residual variance and for being overparametized (Enders, 2004). According to Enders (2004), model selection based on Akaike information criteria does not favour TAR because it is overparametized.

CHAPTER 3

INDUSTRY OVERVIEW

3.1 Introduction

This chapter provides an overview of production, distribution and domestic market sales for apples in South Africa. In addition, the impact of deregulation on real monthly prices, the price spread amongst selected markets and price volatility are measured.

3.2 Historical background

Fruit in the late 19th century was produced largely at a subsistence level in South Africa. Lack of a well established marketing system, and institutional and market infrastructure kept the development of the industry at a suboptimal level. The industry excelled to new levels in the early 20th century when the development of the rail system boosted trade and encouraged producers to engage in more production. This period witnessed large-scale planting of fruit trees. Despite the increase in production a lot of factors inhibited the growth of the industry, e.g. essential facilities necessary for good agricultural practices were lacking (Deciduous Fruit Producers Trust, 2003). Optimal growth in the sector was further constrained by the 1933 depression (Deciduous Fruit Board, 2003).

Under the 1937 Agricultural Marketing Act the Deciduous Fruit Board was established in 1939. In 1986 the powers of the Deciduous Fruit Board were delegated to the Universal Fruit Trade Cooperative (Unifruco) to carry out sole distribution, marketing and export of fresh produce on its behalf. During the 1990s Unifruco developed as the sole marketing organization handling all deciduous fruit in South Africa for exports. In 1996 all agricultural control boards were abolished following the deregulation of the industry. Unifruco merged with Outspan (citrus marketing organization) to form Capespan, which is currently one of the largest exporters of fruit in South Africa. Moreover, after deregulation the industry experienced mushrooming of several companies involved in fruit marketing and distribution (Louw and Fourie, 2004); in many cases coordination in the fruit value chain was poor which resulted in high transaction costs and financial losses to role players. Much emphasis was also put on the export market resulting in a general neglect of the role of the domestic market. This is evidenced by, amongst others, the general perception that poorer quality fruit are normally directed to the domestic market (largely the FPMs after deregulation), the generally poorer state of infrastructure available for fruit distribution and marketing on domestic markets and the lack of information on domestic markets (Fruit industry plan (FIP), 2004). Recent years saw a gradual increase in the level of sophistication in terms of the value chain for fruit marketed domestically with the introduction of new value chain methodologies, more sophisticated consumers and more advanced information technology (FIP, 2004). The fact, however, remains that very little is known on how prices between different markets are interlinked.

3.3 Apple production regions

Table 3.1 shows the major apple producing regions in South Africa. The total area planted to apples in South Africa in 2003 was estimated at 22 379 hectares (Optimal agricultural business system (OABS), 2004).This represents 29% of the total area planted to deciduous fruits. Groenland constitutes 34% of the total area under production; Ceres 22%, Langkloof-East 17% and Villiersdorp 16%. Of total apple production 89% comes from these four areas (OABS, 2004).

Table 3.1: SA apple producing regions

District	Area(ha)	% of Total
Groenland	7595	33.94
Ceres	5020	22.43
Langkloof East	3895	17.4
Villiersdorp/Vyeboom	3495	15.62
Langkloof West.	515	2.3
Little Karoo	490	2.19
Piketberg	384	1.72
Southern Cape	204	0.91
Hex Valley	148	0.66
Somerset West	143	0.64
Free state	142	0.64
Mpumalanga	88	0.39
Wolseley/ Tulbagh	69	0.31
Bergriver	55	0.25
Eastern Cape	47	0.21
Stellenbosch	34	0.15
Lower Orange River	19	0.09
Franschhoek	16	0.07

District	Area(ha)	% of Total
Cape Town	8	0.04
North-East Free State	6	0.03
Limpopo	3	0.02
Gauteng	2	0.01
Total	**22379**	**100**

Source: Optimal agricultural business system (OABS), 2004.

3.4 Apple cultivars grown in South Africa

The major apple cultivars grown in South Africa are Granny Smith, Golden Delicious, Royal Gala, Starking, Topred, Pink Lady, Fuji, Braeburn and Oregon Spur (OABS, 2004). Figure 3.1 shows that Granny Smith takes up 28% of the total area planted, followed by Golden Delicious (23%) and Royal Gala (11%). The other cultivars all take up less that 10% of total area planted with apples.

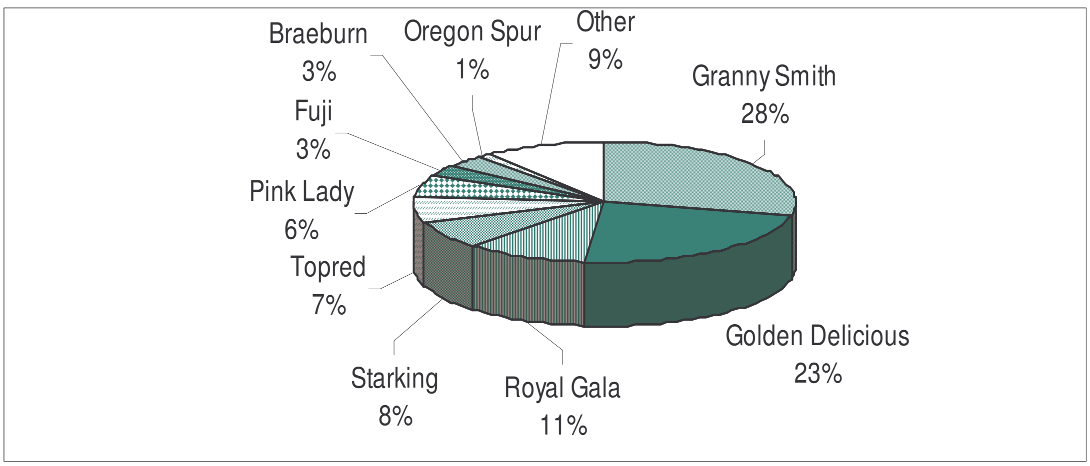

Figure 3.1: Different apple cultivars cultivated in South Africa
Source: Optimal Agricultural Business System (OABS), 2004.

3.5 Apple production and use

Figure 3.2 shows the total production and use of apples from 1991/92 to 2002/03. Production averaged 574 850 tons over this period with a standard deviation of 43 922 tons. Apple production shows an increasing trend over the depicted period. The apple production season extends from the beginning of December till the end of June. Apples are available on all markets throughout the year, mainly as a result of new technology developments in cold storage.

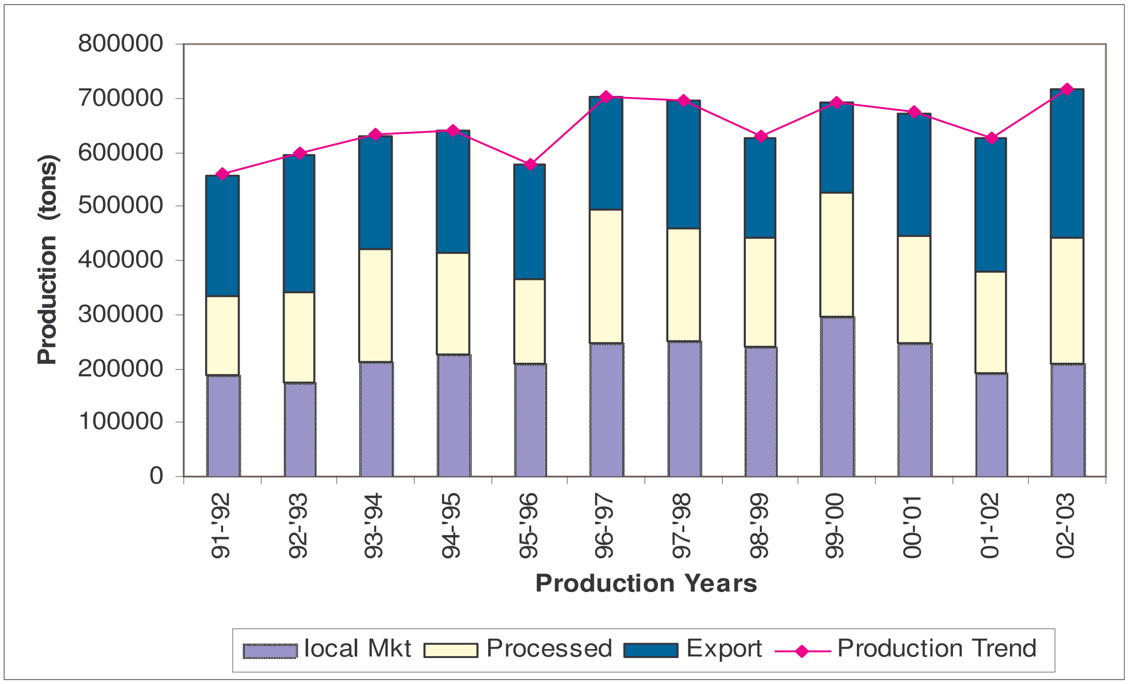

Figure 3.2: Distribution of South African apple production
Source: OABS, 2004.

Figure 3.2 shows that the largest proportion of production that is not exported is either consumed as fresh fruit or processed. It is expected that the local market for apples will become increasingly important in the short to medium-run due to the strengthening of the local currency that significantly reduced the profitability of exports (Fruit industry plan project team (IPPT), 2002).

The total average domestic consumption of apples over the period depicted in Figure 3.2 was 224 708 tons (OABS, 2004). This decrease could probably be attributed to the increase in exports over this period that resulted in lower availability of apples on the local market putting upward pressure on local prices (the increase in exports was fuelled by depreciation in the exchange rate that made exports more profitable).

3.6 Market price trends

Prices on the local market are largely influenced by seasonality in production, the perishability of produce and the amount of apples exported (availability of apples on the local market). The impact of seasonality is to some extent cushioned by cold storage facilities that ensure regular apple supplies in the local market. Demand factors such as consumer habits, substitution between products and per capita income also influence prices.

It can be seen from Figure 3.3 that prices are the lowest between March and June. This is the period when large volumes of apples enter the local market. The variability in prices in different markets increases as the distance from the surplus apple producing regions increases. For example, the Cape Town FPM which is located in a surplus apple production region has the lowest price movements, whereas the Johannesburg, Durban, and Port Elizabeth FPMs, which are all in deficit apple production regions and distant from the Western Cape, experience more price variability. Price spread among the markets also increases as the distance from the surplus region increases.

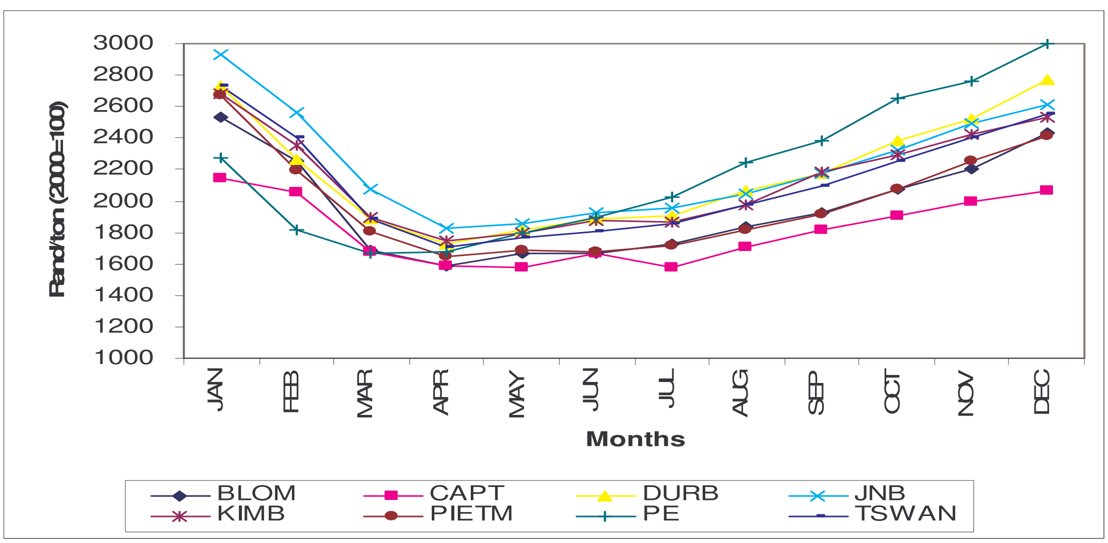

Figure 3.3: Real average market prices of apples in selected FPMs

The relationship between average volume sold in the FPMs and average real market prices is further illustrated in Figures C1 to C8 in Appendix C for the selected FPMs.

3.7 Local market sales

Figure 3.4 shows historical apple sales and prices in the domestic market from 1991/92 to 2002/03. Volumes sold increased from 1991/92 to 1999/2000 after which they decreased to around 200 000 tons. Average prices moved more or less sideways from 1994/95 to 1999/2000 after which they showed an increasing trend.

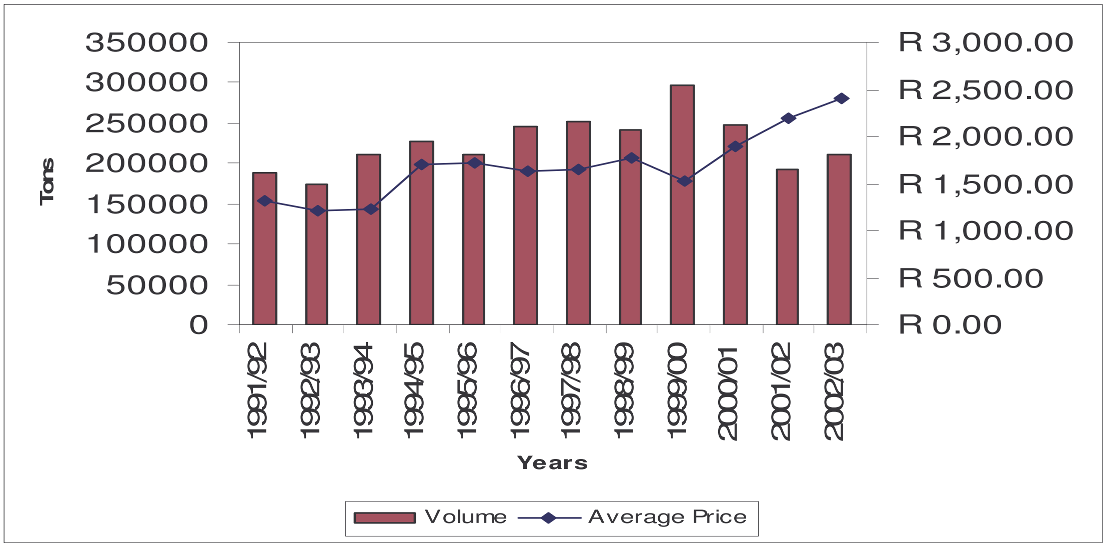

Figure 3.4: Local market historical price trends
Source: OABS, 2004.

3.8 Domestic market analysis

A number of changes have taken place within the domestic market. Internal competition has increased due to an increase in the number of role players that competes for apples. In addition, apple growers face a number of challenges on the demand side. There has been an increase in quality standards, consumer trends are changing, and there is a renewed demand for improved varieties of apple and shifts towards non-economic factors like technical and environmental issues, e.g. safety consciousness of consumers. In addition, various factors impact directly on farm level, e.g. the introduction of labour legislation, higher cost of water, the introduction of standards like EUREPGAP, etc. Within this milieu, as mentioned earlier, the domestic market will become an increasingly more important market segment.

According to Coyle, Bardi and Langley (1996), quoted in Bayley (2000), the fruit supply chain involves the management of a sequence of activities from the producer to the final consumers (quoted in Goedhals, 2003). Figure 3.5 shows the apply supply chain in South Africa.

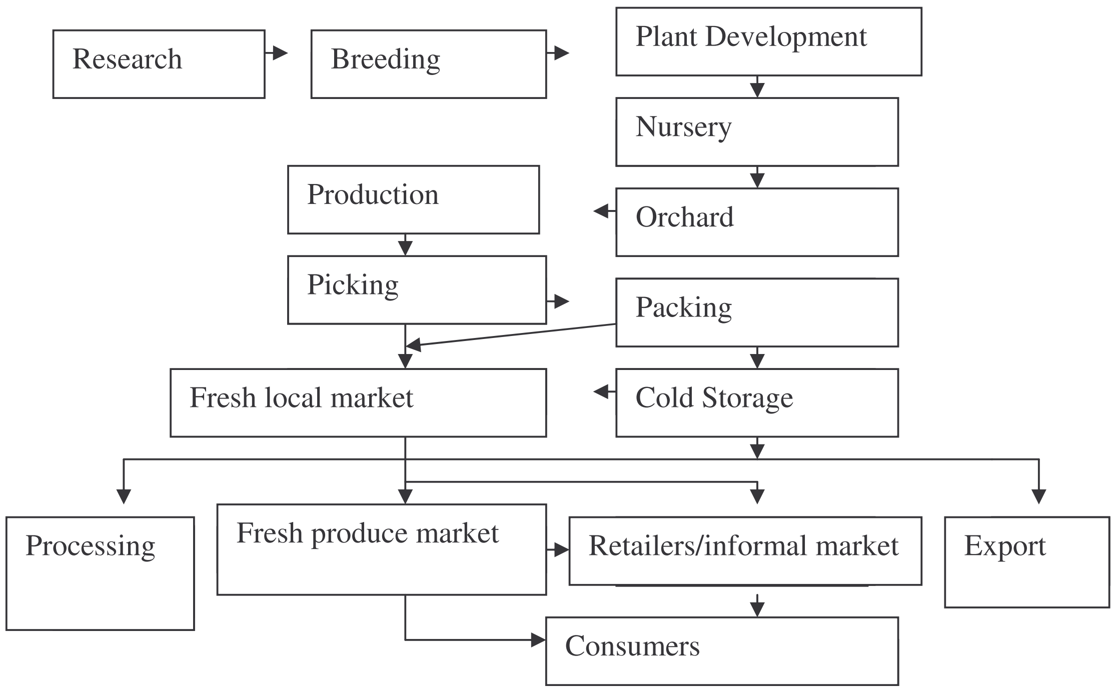

Figure 3.5 Fresh produce supply chain
Source: FIPPT, 2004.

3.8.1 The fresh produce markets (FPMs)

The 16[8] FPMs remain an important market outlet for fresh produce in South Africa. They serve both as a distribution channel and also as a platform for price determination. The FPMs are served by market agents affiliated to it. Producers send their produce to the market agent of their choice who transacts on their behalf at the produce market. The produce is consigned to the agent who operates under a lease of accommodation by the FPM to sell produce on behalf of producers. In most FPMs the producer pays charges as commission for using the market as a channel for marketing their produce. The market agents in turn levy a negotiable commission on the entire sales price from the producers for their services. All transactions are via a computerized system to avoid delays and to improve efficiency. Buyers at the FPMs consist of wholesalers, retailers, contract buyers, specialist and informal buyers, e.g. small private units or hawkers. Official documentation is issued to buyers in respect of every

[8] Johannesburg, Cape Town, Bloemfontein, Tshwane, Port Elizabeth, Durban, Kimberley, East London, Pietermaritzburg, Welkom, Klerksdorp, Vereeniging, Springs, Withbank, Nelspruit, Uitenhage.

transaction. Before any consignment of produce can be sold on FPMs it must be found to be fit for sale in all respects and must be properly sorted, stacked and displayed. This is to ensure a high level of quality and that most hygiene specifications are adhered to. Any condemned produce is disposed off after it has been ascertained unfit by health inspectors (Jooste et al., 2001).

Moreover, the activities of FPMs are governed by an umbrella organization, the Fresh Produce Marketing Association (FPMA) whose main objectives are:

- To promote the consumption of fresh fruit and vegetables.
- To provide a communication forum for all participants in the production and marketing of fresh fruit and vegetables.
- To provide an information network on all aspects of fresh produce handling, training and nutrition.
- To work on matters of interest to its members.
- To provide a forum for communication between the marketers of fresh fruit and vegetables and allied associations, government, consumer group, etc.

The markets selected for this study are described below.

(i) Johannesburg Fresh Produce Market (JFPM)

The Johannesburg fresh produce market is South Africa's major centre for the marketing of fresh produce. It is a private company wholly owned by the City of Johannesburg and was incorporated as a limited liability company in 2000 as part of the IGoli 2002 plan. It is the largest fresh produce market in South Africa and in Africa at large (Johannesburg Fresh Produce Market, 2004). It is a commission market where producers deliver their produce to any of the 14 market agents who sell the produce on their behalf to the buyers. It handles about 32% of all fresh produce marketed through formal channels in the Gauteng province. The total capacity of the JFPM is estimated at 65 000m^2. As a value-adding service, there are fruit ripening chambers and about 40 cold rooms which can accommodate 4 100 pallets of fresh produce. About 10 000 farmers send their produce to the JFPM. The buyer base is estimated at 6 000 at any given time. Accordingly, on an average trading day, the JFPM handles over 16 000 transactions (JFPM, 2004). The market realizes an estimated average daily turnover of R10.8 million and more than R1.8 billion a year. The JFPM makes a 5% non-negotiable commission on all fresh produce sold in the market, while the market agents receive a negotiable 7.5% commission on the selling price of all the fresh produce (JFPM, 2004).

(ii) The Cape Town Fresh Produce Market (CTFPM)

The Epping market is the only FPM in the Western Cape. It serves Cape Town, the whole Cape Peninsula and surrounding areas, as well as the most remote part of the North Western Cape and Namibia. It has the necessary infrastructure to handle large quantities of incoming and outgoing deliveries (Epping Fresh Produce Market, 2004). The market offers all the control measures vital to ensuring that all produce meets required quality standards and hygiene specifications. It offers office accommodation, banking facilities and ample storage accommodation for all its markets agents, plus a ripening complex to allow for all kinds of produce to be ripened, as well as a fully automatic cold storage facility. There was a marked growth in turnover from R8.4 million in 1962 to an estimated R226 million in 1997 (Epping Fresh Produce Market, 2004).

(iii) The Bloemfontein Fresh Produce Market (BFPM)

Producers from all over South Africa market their produce through the four independent market commission agents. The market has 12 cold rooms with a total capacity of 6 330 m^3 and one freezer of 200 m^3. It also has ripening rooms. Buyers, ranging from wholesalers, retailers, informal traders to individuals, have free access to the complete variety of fresh produce. It is estimated that more than 50% of all produce sold on the market is removed from the city to surrounding towns and cities in the Free State, as well as the neighbouring provinces and states. The annual turnover for 2001/02 was R130.7 million, while 91 330 tons of produce were handled (NDA, 2004).

(iv) Port Elizabeth Fresh Produce Market (PEFPM)

In Port Elizabeth buyers and sellers of fresh produce have traded in this market for more than 150 years. The market site of 28ha, buildings and equipment belong to the Nelson Mandela

Metropolitan Municipality. Four market agencies are operating in the market. The market facilities include ripening chambers, cold rooms (equivalent to 39 200 apple cartons and a market hall of 20 000m^3). The market achieved an annual turnover of R150 million during 2002/03 and 80 000 tons of produce were sold in this period (NDA, 2004).

(v) The Tshwane Fresh Produce Market (TFPM)

The market premises comprise 32.8 hectares. The trade area, which consists of two market halls, covers an area of 58 958m^3. The ripening centre with 49 ripening rooms and a capacity of 60380 boxes per week covers a total floor space of 6 377m^2. The cold rooms centre with a total floor space of 3 202m^2 consist of 25 cold rooms with a capacity of 2 130 pallets (NDA, 2004). The City of Tshwane Metropolitan municipality employs 206 permanent staff on the market. Eight market agencies, three fresh produce wholesalers, and two packaging wholesalers operate from the market (NDA 2004). The TFPM handles 16% of all fresh produce in Gauteng province (NDA, 2004). In 2002/2003 the market handled 411 219 tons of fresh produce with a turnover of R834 million (NDA, 2004).

(vi) The Pietermaritzburg Fresh Produce Market (PFPM)

The PFPM has a capacity of about 11 895 m^2 (NDA, 2004). However, due to the tremendous growth in the market activities, the capacity has been increased to include twenty ripening rooms and thirteen cold rooms. Currently there are six market agents operating at the PFPM (NDA, 2004). In 2003 the market achieved a record turnover in sales of about R181million and a total of 94 474 tons of produce were sold (NDA, 2004). This represents a 22% increase in turnover compared to the previous year.

(vii) Durban Fresh Produce Market (DFPM)

The DFPM's turnover came to R425.2 million, a growth of 16.7% compared to the previous year.

(viii) Kimberley Fresh Produce Market (KFPM)

The total floor space where three agents operate is approximately 4 800 m^2. Each agent has its own cold room on the floor with a capacity for 48 pallets. In addition to these cold rooms there are another five cold rooms which are also used by the agents and there are a further six ripening rooms. The larger presence of buyers is from the informal sector and hawkers. About 40% of the buyers distribute to wholesalers and retailers in the region. The 2001 to 2002 turnover was R40 million (Kimberley Fresh Produce Market 2004).

3.8.2 Market infrastructure

(i) Transport

South Africa has one of the most modern and extensive transport infrastructures in Africa. This infrastructure plays a crucial role in the country's economy (South African business guide (SABG), 2003).

Road transport is mostly preferred to rail transport by role players in the apple market since the rail transport system lacks sufficient infrastructure to meet specifics transport needs, i.e. the lack of special refrigerated trucks to transport fruits. In addition, road transport is much more flexible than rail transport because the road network has more internal linkages than rail (SABG, 2003; Goedhals, 2003).

A concerning factor is the general deterioration of the rail and road network in South Africa. According to the Rail Road Association of South Africa (2004), the regional rail and road infrastructures have deteriorated at a time when efficient and cost-effective transportation will play an important role in maintaining competitiveness in the marketing system. A recent Automobile Association report stated that 70% of South Africa's roads were in need of urgent repairs that would cost R65 billion. About 75% of South Africa's roads were in good condition in 1988 but by 1999 this had fallen to 30% (RRA, 2004). Roads in a poor to very poor condition increased from 5% in 1988 to 33% in 1999. The SA Bitumen Association has stated, "Motorists now face twice as many operating costs" (Rail Road Association of South Africa (RRA), 2004). This applies to all forms of transport, including commercial, freight and industrial, as well as public transport and emergency vehicles. This state of affairs could have a significantly negative impact on transaction costs to move produce between different market nodes.

The selected FPMs for this study are spread across wide distances. The local market characteristics and the road distance between these markets and the central market (Johannesburg) are shown in Table 3.2. Surplus producing regions are separated by large distances from the central market. Cape Town and Port Elizabeth are more than a thousand kilometers from Johannesburg. The deficit regions, such as Durban, Bloemfontein, Pietermaritzburg and Tshwane are separated by lower distances from the JFPM.

Table 3.2: Spatial separation of selected FPMs

Markets	Total capacity (Tons/year for the whole period)	Road distance between the JFPM and other markets (km)
Johannesburg	34129	
Cape Town	7525	1402
Port Elizabeth	1506	1075
Tshwane	25498	58
Bloemfontein	5151	398
Kimberley	1417	472
Pietermaritzburg	7278	499
Durban	13369	578

Source: Standard bank executive dairy 2004 and authors computation

3.9 Agricultural transformation and market institutions

During the last two decades there has been increased emphasis on the transformation of agriculture and the rural economy to stimulate economic growth. Eicher and Staatz (1990) recognize transformation of markets, institutions and sustainability as key elements for agricultural growth. Today it is recognized that some of the basic differences between industrialized and third world countries are the differences in economic organization (Eicher and Staatz, 1990). This difference is manifested in how factors of production interact and how the institutions mediate these interactions. Stigliz (1989), in Eicher and Staatz (1990), states that among the most important of these institutions are markets. The next section investigates the before and after effects of agricultural deregulation (transformation) on the apple industry with specific reference to price levels, price spread and price volatility (risk).

3.9.1 The effects of deregulation on price levels, price spread and price volatility (risk)

3.9.1.1 Price levels

Monthly time series price data obtained for the eight selected FPMs over 14 years (from 1991 to 2004) were used for the analysis. The data were deflated using the national price index for fruits and nuts collected in metropolitan areas by Statistics South Africa.

The results reveal that average real prices of apples for all the markets declined after deregulation (See Table 3.3). The changes in the average real prices were calculated by taking the difference between the before and after deregulation mean prices. A hypothesis test was conducted using conditional t-test statistics to determine whether there is a statistically significant difference in the real market price before and after deregulation. Firstly, F-Test statistics for two sample variance was conducted to determine whether the two series have equal variance. The null hypothesis of equal variance was rejected in four markets while in the remaining four markets it showed equal variance. Table 3.3 shows that there is a statistically significant difference in real market prices in six of the eight FPMs at different levels of significance. In the Cape Town and Kimberley FPM's there were no significant changes in real market price after deregulation. The largest declines in real prices were experienced in Port Elizabeth (4.6%), Pietermaritzburg (3.9%), Johannesburg (3.3%), Tshwane (2.5%) and Durban (2%). In Bloemfontein and Cape Town the decline was a little more than a percent and in Kimberley real prices declined by only 0.3%[9].

Table 3.3: Monthly real prices before and after market reform

Market	Before reform			After reform			Change		
	mean	**SD**	**CV (%)**	**mean**	**SD**	**CV (%)**	**mean**	**SD**	**CV (%)**
Johannesburg	3.149	0.196	6.233	3.046	0.180	5.918	-0.103*	-0.016	-0.315
Bloemfontein	3.03	0.224	7.391	2.896	0.175	6.039	-0.134*	-0.049	-1.351
Cape Town	2.912	0.186	6.387	2.881	0.190	6.594	-0.031	0.004	0.207
Durban	3.106	0.196	6.297	3.044	0.200	6.554	-0.061***	0.004	0.257
Kimberley	3.061	0.212	6.931	3.053	0.199	6.526	-0.009	-0.013	-0.405
Pietermaritzburg	3.039	0.195	6.407	2.922	0.181	6.206	-0.117*	-0.013	-0.2
Port Elizabeth	3.132	0.239	7.618	2.988	0.259	8.677	-0.145*	0.021	1.059
Tswane	3.086	0.204	6.602	3.008	0.184	6.128	-0.078**	-0.019	-0.474

Single, double and triple asterisks stand for statistical significance at α =0.01, 0.05, 0.10 levels respectively.

[9] These are % changes in the mean of the real market prices before and after reform.

3.9.1.2 Price spread

In this section the price spread (i.e., difference between prices) between the central market (Johannesburg) and the other markets[10] was investigated. The results in Table 3.4 show that there is a significant difference in the price spread in the majority of the FPMs. The price spread between Johannesburg and five of the other markets was statistically significant at the 1%, level while price spreads between Johannesburg and Pietermaritzburg and Johannesburg and Port Elizabeth were not significant, even at the 10% level. The results further show that the price spread between Johannesburg and the statistical significant FPMs declined following deregulation of the apple market.

Table 3.4: Monthly market spread before and after market reform

Markets	Before reform			After reform			Change		
	mean	SD	CV (%)	mean	SD	CV (%)	mean	SD	CV (%)
J-BLOEM	0.119	0.095	80.24	0.149	0.069	46.135	0.030*	-0.026	-34.11
J-CPT	0.237	0.09	38.066	0.166	0.134	80.672	-0.071*	0.044	42.606
J-DURB	0.043	0.081	186.325	0.002	0.069	3148.826	-0.041*	-0.012	2962.5
J-KIMB	0.088	0.113	129.333	-0.006	0.100	-1627.18	-0.094*	-0.014	-1757
J-PIETM	0.110	0.073	66.179	0.124	0.062	50.178	0.014	-0.011	-16
J-PE	0.017	0.154	927.786	0.059	0.249	426.267	0.042	0.095	-501.5
J-TSW	0.063	0.045	71.66	0.038	0.042	110.475	-0.025*	-0.003	38.815

Single asterisks stand for statistical significance at α =0.01 level.
Source: Author's computation

3.9.1.3 Price variability (risk)

The results obtained in the previous two sections show that price variability for the majority of FPMs declined after deregulation for both real market prices and price spreads. According to Table 3.3, the coefficient of variation of monthly real prices declined for five out of the eight FPMs. There was no decline in variability in the Cape Town, Durban and Tswane market prices. The coefficient of variation for price spreads declined for four out of seven FPMs (see Table 3.4). Moreover, the analysis reveals that for the majority of the FPMs investigated, price variability has reduced after deregulation.

[10] The null hypothesis in the F-Test for sample variance was rejected in three market spreads while four markets spreads showed equal variance.

Nevertheless, there is an opposing debate in the literature as to whether price deregulation causes an increase in price volatility. Akiyama, Baffles, Larson and Varangis (2003) suggest that reform might as well be followed by periods of high prices and low volatility or vice versa. Using Food and Agriculture Organization (FAO) data on domestic producer prices they calculated price variability for 1986-90 and 1991-95 for 35 African countries for cocoa, coffee, cotton and sugar. No evidence to show that price volatility increased between the periods were found. In another study on coffee in Kenya, Karanja (2001), as quoted in (Akiyama, Baffles, Larson, and Varangis 2003), found that price volatility increased in the domestic coffee market after reform. Therefore, the level of price volatility could be dependent on the commodity and the country under study. This is because deregulation policies are implemented differently in different countries. In some countries deregulation slows and commodity specific, market infrastructural developments are poor and are not fully liberalized.

3.10 Summary

Apple production is more or less equally distributed between exports and the local market. Average per capita consumption of apples declined from 2000/01 after showing an increasing trend in the decade before. Relatively little attention was paid to the local market for apples subsequent to deregulation and liberalization.

The impact of deregulation on real market prices, price spread and price volatility was also investigated. The investigation revealed the following:

i) A statistically significant decline in real prices in six of the eight markets investigated.

ii) A statistically significant relation in prices (price spread) between Johannesburg FPM and five other FPMs. The analysis further showed that the price spreads between these markets declined since deregulation; and

iii) Variation in real apple prices declined for five of the eight markets after deregulation.

CHAPTER 4

METHODOLOGY

4.1 Introduction

In the previous chapter, apart from providing an overview of the apple industry, the before and after affects of deregulation on real monthly prices, price spread and price risk were investigated. In this chapter the methodology to measure price transmission and market integration in the local apple market is discussed. The various statistical tests to be performed are also discussed.

Extensive work has been carried out in literature to analyze market integration using the concept of co-integration. According to Balke and Fomby (1997), co-integration captures the essence of the non-linear adjustment process that occurs in most economic time series. These adjustments can be characterized in terms of threshold co-integration in a threshold autoregression process (Balke and Fomby, 1997). Threshold co-integration defines regimes within which co-integration does or does not occur. Identification of the threshold and the regimes delineated by the thresholds form a major part of the study. This is because the error correction model and adjustment processes will depend on threshold values and threshold regimes. Threshold autoregression and co-integration models are adopted and econometric methods used in the analysis are discussed in this chapter.

The rest of the chapter is organized as follows. The method of analysis of time series characteristics on levels is discussed in section 4.2. The procedure for the least squares estimate of co-integration relationships is discussed in section 4.3. In section 4.4 threshold estimation is discussed. Threshold vector error correction models will be discussed in section 4.5. Section 4.6 provides a summary.

4.2 Statistical properties of the data in levels

Before time series price data for apples can be analyzed it is necessary to first check the statistical properties of the time series to determine the nature of data transformation necessary to perform further statistical analysis.

4.2.1 Unit root test on levels

The standard augmented Dickey-Fuller test is used to check the statistical properties of the individual price series on levels as follows:

$$P_t = \rho P_{t-1} + \mu_{t,} \tag{4.1}$$

Where P_t and P_{t-1} are the apple prices at time t and $t-1$ respectively, ρ is the coefficient of lagged apple prices and μ_t is a *iid* ~N $(0, \sigma^2)$ error term. If the null hypothesis of unit root[11] is not rejected $\rho = 1$, i.e. the time series is non-stationary.

A linear combination of the prices can also provide insight into the time series data properties. For a linear combination of the two prices knowledge of the order of integration is necessary. This is determined by using equation 4.1. A non-stationary series is integrated of order I(1), while a stationary series is integrated of the order I(0). A linear combination of two I(1) series in the short-run is non-stationary, but in the long-run they may be stationary I(0) (Balke and Fomby, 1997). Engel and Granger (1987) defines the dynamic long-run relationship between the two prices as

$$P_t^1 = \alpha_0 + \alpha_1 P_t^2 + \mu_t, \tag{4.2}$$

$$\mu_t = \rho \mu_{t-1} + v_t \tag{4.3}$$

Where v_t is a random error term, P_t^1 is a given local market price and P_t^2 is the price in a central market. μ_t is the error correction term. To confirm any stable equilibrium relationship between two prices a stationarity test is conducted on the residuals obtained from equation (4.2). If the residuals errors are stationary then the linear combination of the two prices is stationary (co-integrated).

[11] The term unit root refers to the root of the polynomial in the lag operator. If $\rho = 1$, one can write $p_t - p_{t-1} = \mu_t$. Using the lag operator L, $Lp_t = p_{t-1}$, $L^2 p_t = p_{t-2}$, therefore one can write equation 4.1 as (1-L) $p_t = \mu_t$, if (1-L) = 0 one obtains, L=1, hence the name Unit Root (Gujarati, 2003).

4.3 Test for long-term relationships between variables

In this section, the techniques used to test for long-run relationships between apple prices in two markets are shown. Two approaches are followed, namely the standard augmented Dickey-Fuller (ADF) test and Johansen multivariate test of co-integrating vectors. The standard augmented Dickey-Fuller test is used to test for a co-integration relationship in the residual of two co-integrating apple prices. The null hypothesis of the test is that the residuals μ_t in equation 4.3 are I(1) against the alternative that the residuals are stationary, i.e. I(0).

The Johansen test takes into account the number of co-integrating relationships amongst co-integrating variables. The test is based on the notion that economic variables are much more likely to be endogenously interdependent. Determining the number of co-integrating vectors will give an insight into the number of estimation equations to be fitted. Even though more than one co-integrating relationship might exist, at least the presence of one co-integration relationship is necessary for the analysis of long-run relationships of the prices to be plausible.

Johansen test utilizes two test statistics formulations, namely eigenvalues and the trace statistics. It is a maximum likelihood ratio test involving a reduced rank regression between two variables, say I(1) and I(0), providing n eigenvalues $\hat{\lambda}_1 > \hat{\lambda}_2 > \ldots > \hat{\lambda}_n$ and corresponding eigenvectors $\hat{V} = (\hat{v}_1, ..., \hat{v}_n)$, where the r elements of $\hat{V}$ are the co-integration vectors. The magnitude of λ_i is a measure of the strength of correlation between the co-integrating relations for $i = 1....r$. The test of the null hypothesis that there are r co-integrating vectors present can be stated as:

$$H_0 : \lambda_i = 0 \quad i = r + 1 \ldots\ldots, n$$

The maximal-eigenvalue ($\lambda - \max$) statistics is given by:

$$\lambda_{\max} = -T \log\left(1 - \hat{\lambda}_{r+1}\right) \qquad r = 0, 1, 2, ..., n-1 \tag{4.4}$$

Where T is the sample size, and $\left(1 - \hat{\lambda}_{r+1}\right)$ is the max-eigenvalue estimate.

The trace statistic is computed as:

$$\lambda_{trace} = -T\sum_{i=r+1}^{n}\log\left(1-\hat{\lambda}_i\right) \qquad r = 0, 1, 2, ..., n-1. \tag{4.5}$$

Testing the null hypothesis of r co-integrating vectors against the alternative of $r+1$.

4.4 Threshold models

In this section, the methodology used to fit univariate threshold error correction models is discussed.

The lagged residuals from equation (4.3) are used to define the error correction model. The co-integration of the P_t^1 and P_t^2 variables depends on the nature of the AR process for the error term $\mu_{t,}$, where $\mu_{t,}$ is the deviation from equilibrium. As ρ approaches one, deviations from equilibrium becomes non-stationary and thus the series are not co-integrated. Assuming that μ_t follows a threshold autoregression (TAR), a univariate TAR model of the following form, on which grid search is conducted, is specified to determine threshold values.

$$\mu_t = \alpha_0^{(j)} + \sum_{i=1}^{p}\alpha_i^{(j)}\mu_{t-i} + v_i^{(j)} \qquad r_{j-1} < \mu_{t-d} \leq r_j \tag{4.6}$$

Where $j = 1,..,k$ and the threshold lag d is a positive integer. Thresholds are $-\infty = r_0 < r_1 < \ldots < r_k = \infty$. $\{v_i^{(j)}\}$, is $i.i.d.\left(0, \sigma^{(j)2}\right)$.

Assuming $\rho = 1$ and two threshold values, μ_t can be divided into three regimes to reflect arbitrage transactions when there is large negative error correction terms in the previous period, that is, $\mu_{t-1} \leq c_2$, or to a large positive error correction term, $\mu_{t-1} > c_2$, we have,

$$\mu_t = \alpha_0^{(1)} + \alpha_1^{(1)}\mu_{t-1} + v_t^{(1)} \qquad \mu_{t-1} < -c_1 \tag{4.7}$$

$$\mu_t = \alpha_0^{(2)} + \alpha_1^{(2)}\mu_{t-1} + v_t^{(2)} \qquad -c_1 < \mu_{t-1} \leq c_2 \tag{4.8}$$

$$\mu_t = \alpha_0^{(3)} + \alpha_1^{(3)}\mu_{t-1} + v_t^{(3)} \qquad \mu_{t-1} > c_2 \tag{4.9}$$

where c is a potential threshold which represents transaction cost above which equilibrating adjustments will be triggered in the outer regimes. If $|\mu_{t-1}| < c_1$ no mean reversion to equilibrium occurs, hence μ_t will follow an AR (1) process. If $|\mu_{t-1}| > c_2$ a substantial degree of mean reversion will occur and μ_t becomes I(0), i.e. stationary with zero mean. $|\mu_{t-1}|$ will not change much when the deviations from zero are small. Hence, $\alpha_1^{(2)} = 1$ when deviations are small. The coefficients in (4.7) and (4.9) will be much smaller when $|\mu_{t-1}|$ is large to reflect mean reversion towards zero. The TAR model (4.7) to (4.9) is a band TAR equilibrium model. Equilibrium for μ_t is anywhere in a band or interval [-c, +c] and not at a point. Within the band error correction is not specified because the equilibrium error is zero.

4.4.1 Threshold estimation

Tsay's (1989) four main steps are used to select the threshold lags (p, k, d) to test for the presence of threshold and to locate threshold values.

4.4.1.1 Selecting an AR order p

The AR order is selected using the partial autoregression function (PACF) of μ_t by estimating equation (4.10)

$$\mu_t = \alpha_0 + \sum_{i=1}^{q} \alpha_i \mu_{t-1} + v_t \tag{4.10}$$

for increasing order of q, and choosing the order of q for which the coefficient α_q is significant. This is determined from the t-ratio. This is the q$^{\text{th}}$ partial autocorrelation coefficient, $\alpha_{qq.}$. The AR (p) is chosen such that

$$[\alpha_{qq} \neq 0 \quad for \quad q = p, \text{ and } \alpha_{qq} = 0 \quad for \quad q > p] \tag{4.11}$$

The α_{qq} are approximately normally distributed with mean zero and variance of 1/N for $q > p$, where N is the sample size. It is appropriate for the significant test of α_{qq}.

4.4.1.2 Test for threshold non-linearity[12]

The test for non-linearity of μ_t is based on arranged autoregression and predictive residuals. According to Tsay (1989), for the TAR model in equation 4.6, the arranged residual is useful if it is arranged according to the threshold variable μ_{t-d}. For the given TAR model, TAR (k, p, d) is estimated where

- k is the number of regimes separated by $k-1$ non-trivial thresholds (r_j) (k is assumed to be three);
- p is the AR order selected through the process described in section 4.4.1.1. In most cases the value of p is one; and
- the delay parameter (d) is chosen to be one.

Suppose we have the AR (p) model for z_t in equation (4.6). We use (z_t,1, z_{t-1},.........z_{t-p}) as a case of data for the AR (p) model. For a given TAR (3, p, d) model with n observations, the threshold variable z_{t-d} may assume the values $\{z_h,...,..z_{n-d}\}$, where $h = \max\{1, p+1-d\}$. Let π_i be the index of the i^{th} smallest observation of $\{z_h,..., z_{n-d}\}$. We can rewrite equation 4.6 as,

$$z_{\pi_i+d} = \alpha_0^{(1)} + \sum_{v=1}^{p} \alpha_0^{(1)} z_{\pi_i+d-v} + \mu_{\pi_i+d}^{(1)} \qquad \text{if } i \leq s \qquad (4.12)$$

$$z_{\pi_i+d} = \alpha_0^{(2)} + \sum_{v=1}^{p} \alpha_0^{(2)} z_{\pi_i+d-v} + \mu_{\pi_i+d}^{(2)} \qquad \text{if } s < i \leq g \qquad (4.13)$$

$$z_{\pi_i+d} = \alpha_0^{(3)} + \sum_{v=1}^{p} \alpha_0^{(3)} z_{\pi_i+d-v} + \mu_{\pi_i+d}^{(3)} \qquad \text{if } i > g \qquad (4.14)$$

Where s satisfies $z_{\pi_s} < r_1 \leq z_{\pi_{s+1}}$ and g satisfies $z_g < r_2 < z_{\pi_{g+1}}$.This is an arranged autoregression with the first s cases in the first regime, the second [g-s] cases in the second, the remainder in the third regime. The arranged autroregression orders the data in a way that all the observations in a group follows the same linear AR model. Tsay (1989) suggests that the parameter estimates are consistent OLS estimates if the number of observations in the first

[12] The test procedures are adopted from Tsay's (1989) publication. For details on this test refer to Tsay (1989).

regime is large. For example $\hat{\alpha}_i^{(1)}$ is consistent for $\alpha_1^{(1)}$, for many $i \leq s$. Therefore, the standardized residuals are $iid \sim \mathrm{N}\left(0, \sigma^2\right)$ asymptotically and orthogonal to the regressors z_{π_i+d-v} conditional on $\left[v = 1, ..., p\right]$. According to Tsay (1989), If $i \geq s$ the standardized residual for the observation with time index $\pi_{s+1} + d$ will be biased because of model change at $\pi_{s+1} + d$. At this point Tsay points out that the standardized residual will become a function of the regressors z_{π_i+d-v} conditional on $\left[v = 1, ..., p\right]$. Consequently, the orthogonality between the standardized residuals and the regressors will be destroyed once the autoregression goes on to the observation whose threshold value exceeds r_1.

For fixed p and d, the effective number of observations in arranged autoregression is $n - d - h + 1$, were $h = \max\{1, p + 1 - d\}$. Assuming that the autoregressions starts with b observations so that there are $\{ n - d - b - h + 1 \}$ standardized residuals available. Tsay suggests regressing the standardized residual on the arranged autoregression using OLS estimates for the first m cases as follows

$$\hat{e}_{\pi_i+d} = \omega_0 + \sum_{v=1}^{p} \omega_v z_{\pi_i+d-v} + \mu_{\pi_i+d}, \tag{4.15}$$

For $i = b + 1, ..., n - d - h + 1,$ the significance of the test is determined from F- statistics

$$\hat{F}\left(p, d\right) = \frac{\left(\sum \hat{e}_t^2 - \sum \hat{e}_t^2\right) / \left(p + 1\right)}{\sum \hat{e}_t^2 / \left(n - d - b - p - h\right)}, \tag{4.16}$$

This is summed over all observations in equation (4.15) and $\hat{e}_t$ is the least square residual of (4.15).However, the test depends on the value of $\left(p, d\right)$.Tsay (1989) suggests selecting d that maximizes the F-statistic. If the F-statistic exceeds the critical value of the F distribution with $\{ p + 1 \}$ and $\{ N - d - m - p - h \}$ degrees of freedom we reject the null of linearity and estimate the TAR model.

4.4.1.3 Locating the threshold value

The number and location of the threshold values are identified through a grid search algorithm. The search is conducted over an arranged threshold variable (in ascending order).

Ten percent of the data from the two extremes of the data points (i.e. the first 10 and the last 10 observations) are excluded. The search is restricted over at least a minimum of 20 observations. The approach is similar to Balke and Fomby's approach of using a grid search that minimizes the sum of squared error criterion.

4.4.1.4 Testing for the significance of threshold

After locating the threshold the data points are divided according to whether the equilibrium error is above or below the threshold and then ordinary least square regression is estimated. Firstly, dummy variables are incorporated and a TAR error correction model for each market combination is specified. The coefficient of the inner band (the neutral band) is not expected to be statistically significant. This is because there is no mean reversion in this regime and adjustment is expected to be low or zero. The outer band is expected to adjust faster than the inner band; therefore its coefficient is expected *a priori* to be larger and statistically significant.

4.4.1.5 Testing for threshold and Hansen Test[13]

The aim is to test the statistical significance of the threshold effect. First, the F distribution is obtained from a least square estimate of the TAR model. This is followed by obtaining a residual variance using sequential conditional least square with both TAR and standard AR models using 100 bootstrap simulations as dependent variables. The p-values are calculated by taking the percentage of the test statistics for which the test taken from the simulated sample exceeds the observed test statistics. The procedure outlined in Hansen (1997) is as follows:

Assuming a two-regime TAR model takes the form

$$y_t = (\alpha_0 + \alpha_1 y_{t-1} + \ldots + \alpha_p y_{t-1})\,1\,(q_{t-1} \leq \gamma) + (\beta_0 + \beta_1 y_{t-1} + \ldots + \beta_p y_{t-p})\,1\,(q_{t-1} > \gamma) + e_t \quad (4.17)$$

Where 1(.) is an indicator function, $q_{t-1} = q(y_{t-1}, \ldots., y_{t-p})$ depends on the time series data used. The AR order is $p \geq 1$ and the threshold parameter is γ. The parameters α_j are the AR

[13] The test procedure is adopted from Hansen (1997).

slopes when $q_{t-1} \leq \gamma$ and β_j are the slopes when $q_{t-1} > \gamma$. The error e_t is assumed *i.i.d* $N(0,\sigma^2)$. The test is based on whether the TAR model in equation (4.19) is statistically significant relative to a linear AR (p) process. Because the threshold is not identified under the null hypothesis the hypothesis of Ho: $\alpha = \beta$ can not be tested. Hansen (1997) recommends the F distribution test. It is a test with near optimal power against the alternative null hypothesis.

$$F_n = n\left(\frac{\hat{\sigma}_n^2 - \hat{\sigma}_n^2}{\hat{\sigma}_n^2}\right), \tag{4.18}$$

Where

$$\hat{\sigma}_n^2 = \frac{1}{n}\sum_{i=1}^{n}\left(y_t - x_t^{'}\tilde{\alpha}\right)^2, \tag{4.19}$$

and

$$\tilde{\alpha} = \left(\sum_{t=1}^{n} x_t x_t\right)^{-1}\left(\sum_{t=1}^{n} x_t y_t\right) \tag{4.20}$$

is the OLS estimate of α under the assumption that $\alpha = \beta$. Since F_n is a monotonic function of $\hat{\sigma}_n^2$,

$$F_n = \sup_{\gamma \in \Gamma} F_n(\gamma)$$

where

$$F_n(\gamma) = n\left(\frac{\hat{\sigma}_n^2 - \hat{\sigma}_n^2(\gamma)}{\hat{\sigma}_n^2(\gamma)}\right)$$

is the pointwise F-statistics against the alternative of $H_1: \alpha \neq \beta$ when γ is known. Since γ is not identified under the null hypothesis, F_n is not a χ^2. The asymptotic distribution may be approximated by following the bootstrap procedure. Let $u_t^*, t = 1,...,n$ be *i.i.d* $N(0,1)$ random draws, and let $y_t^* = u_t^*$. Using observation $x_t, t = 1,....,n$, regress y_t^* on x_t to obtain the residual variance $\tilde{\sigma}_n^{*2}$ on $x_t(\gamma)$ to obtain the residual variance $\hat{\sigma}_n^{*2}(\gamma)$ and form $F_n^*(\gamma) = n\left(\tilde{\sigma}_n^{*2} - \hat{\sigma}_n^{*2}(\gamma)\right) / \hat{\sigma}_n^{*2}(\gamma)$ and $F_n^* = \sup_{\gamma\in\Gamma} F_n^*(\gamma)$. Hansen shows that the distribution of F_n^* converges weakly in probability to the null distribution of F_n under local alternative for β. Therefore, repeated bootstrap draws from F_n^* may be used to approximate the asymptotic null distribution of F_n. The p-value is calculated by counting the percentage of bootstrap samples for which F_n^* exceeds the observed F_n.

4.5 Threshold vector error correction models (TVECM)

Unlike the univariate TAR, the multivariate TVECM models are used to model non-linear and asymmetric threshold adjustments to long run equilibrium. In a TVECM different regimes can be identified based on the size of the lagged error correction term. The variables in the model exhibit different types of behaviour in each regime based on whether the error correction term is less than or more than the threshold parameter. The TVECM model is estimated based on equations (4.7) to (4.9) as:

$$\Delta P_t = \alpha_0^{(j)} + \sum_{k=1}^{k} \alpha_k^{(j)} \Delta P_{t-1} + \varphi^{(j)} \mu_{t-1} + v_t^{(j)} \qquad r_j < \mu_{t-d} \le r_j \tag{4.21}$$

Where $P_t = (\ln P_{1t}, P_{2t})'$, Δ is the difference operator ($\Delta P_t = P_t + P_{t-1}$), $\alpha_0^{(j)}$ are 2 x 1 column vector of constants, $\alpha_k^{(j)}\ (k = 1,2...,k)$ are 2 x 2 matrixes of the parameters, $\varphi^{(j)}$ are a 2 x 1 column vector of parameters, $v_t^{(j)}$ are 2 x 1column vectors of residual terms, and k denotes the number of lags in the TVECM. The TVECM representation of the three regime models in equation (4.7)-(4.9) is as follows:

$$\Delta \ln P_t^1 = \alpha_1^{(1)} + \sum_{k=1}^{k} \alpha_{11,k}^{(1)} \Delta \ln P_{t-k}^1 + \sum_{k=1}^{k} \alpha_{12,k}^{(1)} \Delta \ln P_{t-k}^2 + \varphi_1^{(1)} \mu_{t-1} + v_{1,t}^{(1)}$$

$$\mu_{t-1} \le c_1 \tag{4.22}$$

$$\Delta \ln P_t^2 = \alpha_2^{(1)} + \sum_{k=1}^{k} \alpha_{21,k}^{(1)} \Delta \ln P_{t-k}^1 + \sum_{k=1}^{k} \alpha_{22,k}^{(1)} \Delta \ln P_{t-k}^2 + \varphi_2^{(1)} \mu_{t-1} + v_{2,t}^{(1)}$$

$$\Delta \ln P_t^1 = \alpha_1^{(2)} + \sum_{k=1}^{k} \alpha_{11,k}^{(2)} \Delta \ln P_{t-k}^1 + \sum_{k=1}^{k} \alpha_{12,k}^{(2)} \Delta \ln P_{t-k}^2 + \varphi_1^{(2)} \mu_{t-1} + v_{1,t}^{(2)}$$

$$c_1 < \mu_{t-1} \le c_2 \quad (4.23)$$

$$\Delta \ln P_t^2 = \alpha_2^{(2)} + \sum_{k=1}^{k} \alpha_{21,k}^{(2)} \Delta \ln P_{t-k}^1 + \sum_{k=1}^{k} \alpha_{22,k}^{(2)} \Delta \ln P_{t-k}^2 + \varphi_2^{(2)} \mu_{t-1} + v_{2,t}^{(2)}$$

$$\Delta \ln P_t^1 = \alpha_1^{(3)} + \sum_{k=1}^{k} \alpha_{11,k}^{(3)} \Delta \ln P_{t-k}^1 + \sum_{k=1}^{k} \alpha_{12,k}^{(3)} \Delta \ln P_{t-k}^2 + \varphi_1^{(3)} \mu_{t-1} + v_{1,t}^{(3)}$$

$$\mu_{t-1} > c_2 \quad (4.24)$$

$$\Delta \ln P_t^2 = \alpha_2^{(3)} + \sum_{k=1}^{k} \alpha_{21,k}^{(3)} \Delta \ln P_{t-k}^1 + \sum_{k=1}^{k} \alpha_{22,k}^{(3)} \Delta \ln P_{t-k}^2 + \varphi_2^{(3)} \mu_{t-1} + v_{2,t}^{(3)}$$

Where $\alpha_n^{(j)}$ denotes the nth element of $\alpha_0^{(j)}$ and $\alpha_{nm,k}^{(j)}$ the element of $\alpha_k^{(j)}$ in row n and m. The value of the error correction term is expected *a priori* to be larger in the outer regime than in the inner regime. This is because the impact of these mean reversions will depend on the size of the deviation from the no-arbitrage condition (Martens and Vorst, 1998). In the outer regime the error correction terms are greater than the threshold and thus show larger response. Therefore, it is expected that, $|\varphi_1^{(1)}| > |\varphi_1^{(2)}|$ and $|\varphi_1^{(3)}| > |\varphi_1^{(2)}|$ for the lag market P_t^1 and $|\varphi_2^{(1)}| > |\varphi_2^{(2)}|$ and $|\varphi_2^{(3)}| > |\varphi_2^{(2)}|$ for the lead market P_t^2.

4.5.1 Estimating vectors and threshold parameters

Threshold vectors and threshold parameters are estimated using a similar approach to that followed by Serra and Goodwin (2002). A sequential conditional iterative Seemingly Unrelated Regression (SUR) is used in two steps. Firstly, a two-dimensional grid search is carried out to estimate threshold parameters $\gamma(c_1, c_2)$. The search is restricted to ensure adequate number of observations for estimating the parameters in each regime. A minimum of 20 observations is allowed in each regime. Threshold selection is based on a grid search that minimizes the log determinant of the variance-covariance matrix of the residual of the TVECM. The result is compared by considering a grid search based on the minimum of the trace of the covariance matrix. Secondly, the vectors of parameters φ^1, φ^2 and φ^3 are estimated conditional on the estimated threshold parameters γ using SURs. From the estimation the log of the determinant is obtained from:

$$\mathrm{S}(c_1, c_2, d) = \ln|\hat{\Sigma}(c_1, c_2, d)| \tag{4.25}$$

where $\hat{\Sigma}(c_1, c_2, d)$ is a multivariate iterative SUR estimate of $\Sigma = \mathrm{var}(e_1)$ conditional on (c_1, c_2, d), where $d = 1$. Trace of covariance matrix is obtained from:

$$S(c_1, c_2, d) = trace\left(\hat{\Sigma}(c_1, c_2, d)\right)$$

In the second step, c_1 and c_2 are obtained as:

$$(\hat{c}_1, \hat{c}_2, 1) = \arg \underbrace{\min}_{C_1, C_2} \quad S(c_1, c_2, 1) \tag{4.26}$$

The final estimate of the parameter are given by $\hat{\alpha}^{(i)'} = \alpha^{(i)'}(\hat{c}_1, \hat{c}_2, 1)$ and the estimation of the residual covariance matrix by $\hat{\Sigma} = \hat{\Sigma}(c_1, c_2, 1)$.

To test for the significance of the differences across relative regimes, Sup-LR statistics is used to test for a linear VECM against the alternative of TVECM. This test is a multivariate TVECM version of the Hansen test conducted for the univariate TAR in section 4.3. The null hypothesis of linear VECM is given as:

$$\Delta P_t = \beta' x_{t-1} + e_t \tag{4.27}$$

The alternative hypothesis of TVECM is thus given as:

$$\Delta P_t = \beta^{(1)} x_{t-1} d_{1t}(c_1, c_2, d) + \beta^{(2)} x_{t-1} d_{2t}(c_1, c_2, d) + \beta^{(3)} x_{t-1} d_{3t}(c_1, c_2, d) + e_{t.} \tag{4.28}$$

The sup-LR statistics is calculated from the formula:

$$LR = T\left(\ln|\hat{\Sigma}| - \ln|\hat{\Sigma}(c_1, c_2, 1)|\right), \tag{4.29}$$

where $\hat{\Sigma}$ and $\hat{\Sigma}(c_1, c_2, 1)$ represents the covariance matrix of the residual of the VECM and TVECM, respectively, and T is the number of observations. To calculate the p-values a similar procedure is adopted as in section 4.4.1.5. Bootstrapped 200 simulations are used instead of the dependent variable ΔP_t with *iid* $N(0,1)$ random draws. The proportion of the simulations under the null hypothesis for which the simulated LR statistics exceeds the observed LR statistics gives the asymptotic p-value of the sup-LR test.

4.5.2 Regime switching and impulse response function

In this section the frequency of observations falling in each regime from the error correction model representation is calculated conditional on whether the error correction term is less than or equal to the threshold in the first regime or greater than the threshold in the third regime. The second regime corresponds to the errors that are between the thresholds that define regimes 1 and 3. Observations falling in each regime are calculated for the entire time path and for each year from 1991 to 2004. The frequency of observations falling in each regime is discussed in chapter 5.

An impulse response function is also considered to evaluate the response in the Johannesburg market due to shocks in other local markets. The non-linear impulse response approach of Potter (1995) was adapted by Koop, Pesaran and Potter (1996). The procedure can be described as follows. A single observation corresponding to the last observation in the two alternate markets is chosen. The response to one-half standard deviation positive and negative shock on the price at the Johannesburg market is then evaluated. The response is related in a time plot involving all the markets. The Impulse response function can be expressed as the difference between two conditional first moment profiles (Baum *et al,*. 1996). It is defined in a similar way to standard impulse response functions, except one replaces the linear predictor with a conditional expectation (Potter 1995). It is expressed as:

$$I_{t+k}(v; y_t, y_{t-1}, \ldots) = E[Y_{t+n} | Y_t = y_t + v, Y_{t-1}, \ldots.] - E[Y_{t+n} | Y_t = y_t, Y_{t-1} = y_{t-1}, \ldots] \ldots\ldots \tag{4.30}$$

where I_{t+k} is the impulse response, $(y_t, y_{t-1}....)$ are observed data, v is the shock and $E[.]$ is the expectation operator. The impulse is produced by estimating $E[.]$.The aim is to investigate the effects of shocks in the Johannesburg market. Response to Johannesburg shocks are expected to be temporary. Due to the non-stationary nature of Johannesburg prices and some other market prices, shocks emanating from Johannesburg may produce either temporary or permanent responses in the alternate local market.

4.6 Summary

The chapter discussed the analytical procedures used in the study. The methods for checking the statistical properties of the time series were described. Long-term relations between price variables were investigated using ADF and the Johansen test. Threshold models and error correction representation of the models were fitted to evaluate the dynamic relationship among apple market prices. The statistical significance of the estimated thresholds was evaluated using Tsay's (1989) and Hansen's (1997) test statistics. To better explain the interrelationship among prices, switches between regimes and the response to price shocks among the markets were evaluated using regime switching and impulse response functions. The empirical result of the analysis is given in chapter 5.

CHAPTER 5

EMPERICAL RESULTS

5.1 Introduction

For markets to be integrated certain conditions must be fulfilled. There must be free flow of goods between markets and the markets must be linked by efficient arbitrage. The efficient arbitrage condition is fulfilled when price spreads between markets are less than or equal to transaction costs. Efficiency in spatial market arbitrage will cause the price risk to be spread amongst markets. Higher prices in deficit regions will be transmitted to surplus markets and lower prices from surplus regions will be transmitted to deficit markets, and by this process the prices in the markets will equalize through arbitrage.

As already discussed, if markets are co-integrated there is evidence that the markets have a long run relationship and that their prices in the long run will not diverge from each other. Studies based on the co-integration test alone have been criticized since they did not account for transaction costs. For this reason both the standard AR error correction and TAR error correction models are used in the analysis.

In terms of TAR models, both univariate and multivariate TVECMs were applied. Symmetric adjustments were studied using TAR error correction models while asymmetry was modeled using TVECM. TVECM incorporates short-run dynamics of price transmission and takes account of asymmetric interrelationships among prices. As previously noted in section 2.9 of chapter 2, symmetric price adjustments assume that there is free flow of goods among spatial markets in all directions, and that due to the volume of movement two alternate markets may face similar transaction costs implying that arbitrage is induced in the same way no matter where the prices are higher. The situation is different with asymmetric price adjustments according to which movement of goods between two spatial markets may occur in one direction, implying some markets are importing and others exporting.

Transaction costs of the exporting market will tend to differ from that of importing markets. In such a situation movement of commodities in one direction will be more costly than movement in the opposite direction. The threshold identified in the case of symmetric adjustment implies similar transaction costs between spatial markets, while asymmetric thresholds imply different transaction costs. The thresholds indicate the amount by which the prices in the alternate markets should increase to trigger symmetric or asymmetric adjustments.

Arbitrage efficiency and time of switching between regimes are also investigated. Finally, a non-linear impulse response function was applied to investigate how price shocks are transmitted and how long it takes for shocks to be eliminated in alternate markets.

The rest of the chapter is structured as follows. In section 5.2, data sources and statistical properties of the price data on levels are analyzed using the Johansen co-integration test and ADF unit root test. In section 5.3, results of Standard autoregression (AR) and Threshold autoregression (TAR) estimations are provided. This includes estimations of univariate threshold lags and threshold values. The significance of the threshold effect is illustrated using Hansen (1997) simulation test statistics. In section 5.4 multivariate TAR is estimated. Regime switching estimates are used to explain asymmetric price movements in the different regimes.

5.2 Data and statistical properties of the variables

The analysis described is based on monthly time series nominal price data obtained from the NDA for eight FPMs over 14 years (from 1991M1 to 2003M4). As stated, the markets included are Johannesburg (JNB), Cape Town (CAPT), Tshwane (TSW), Bloemfontein (BLOM), Port Elizabeth (PE), Durban (DURB), Pietermaritzburg (PIETM) and Kimberley (KIMB).

In the case of the symmetric TAR, pair wise autoregressive market integration is measured, i.e. the relation of two markets (lead and lag) is considered. For the symmetric TAR Johannesburg is taken as the central (lead) market. This is because it is the largest in terms of its capacity and actual volumes of trade. In the case of asymmetric adjustment, price movements can be in either direction and hence no market is regarded as being the lead or lag market.

In order to determine the data generating mechanism of the series, it was not decomposed into its component parts. The study was based on the assumption that the series are generated by non-stationary stochastic trend characterized by random walk and have unit root. Therefore, unit root test was embarked on to determine the time series characteristics of the data. Enders (2003) general to specific methods were considered in choosing the ADF lag length.

The ADF test confirmed non-stationarity in two of the eight price series, namely CAPT and TSW. The null hypothesis of non-stationarity was rejected for JNB, KIMB, BLOM, DURB, PIETM and PE.

Co-integrating relations in the price levels was analysed using the standard ADF and Johansen co-integration tests procedures. The standard ADF test was used to test for unit root in the residuals of the co-integrating variables. The test is mainly to detect the presence of only one co-integration vector in the residuals of the co-integrating variables. The null hypothesis of the test is that the two co-integrating series have unit root, i.e. I(1). The results in Table 5.1 show that the null hypothesis is rejected for all the markets at the 5% level of significance. This confirms that the linear combination of the price series in levels is stationary around a stochastic trend. The existence of a single or multiple co-integrating relations was also investigated using the Johansen multivariate co-integration test. The test is a maximum likelihood ratio test based on a maximal eigen value and Trace of the stochastic matrix in the vector autoregression. Using eigen value, the hypothesis that $r = 0$ was tested against the alternative where $r = 1$. In the Trace, the hypothesis of $r = 0$ against the alternative of $(r+1)$ co-integrating vectors was tested. A combined test of the null hypothesis of $r = 1$ against the alternative of $r > 1$ using eigen value and Trace statistics was also considered. Table 5.1 shows that co-integration was found using the Johansen co-integration test in all the markets.

Goodwin and Schroeder (1991) found that markets separated by longer distances have lower degrees of co-integration than do markets in close proximity. The combined test of the null hypothesis of $r = 1$ against alternative of $r > 1$ using eigen value and trace statistics showed that PE-JNB and DURB-JNB co-integrating relations might have more than one co-integrating vectors. This is because the null of $r = 1$ was rejected while for other pairs the null was not rejected. Given the results in Tables 5.1, it can be concluded that there is evidence of a long-run co-integration relationship in the apple market prices.

Table 5.1: Results of co-integration test

Markets	Test	Test statistics
JNB-BLOM	Max eigen value test r = 0	48.083**
	Trace test r = 0	49.36**
	max eigen value and trace test: r = 1	5.53
	ADF test of residual	-5.6769
JNB-CAPT	Max eigen value test r = 0	68.6206**
	Trace test r = 0	73.16**
	max eigen value and trace test: r = 1	4.54
	ADF test of residual	-6.8967
JNB-TSW	Max eigen value test r = 0	89.78**
	Trace test r = 0	97.84**
	max eigen value and trace test: r = 1	8.06
	ADF test of residual	-6.6515
JNB-PE	Max eigen value test r = 0	140.4**
	Trace test r = 0	155.52**
	max eigen value and trace test: r = 1	15.13**
	ADF test of residual	-7.6976
JNB-DURB	Max eigen value test r = 0	155.93**
	Trace test r = 0	170.5**
	max eigen value and trace test: r = 1	14.57**
	ADF test of residual	-7.1738
JNB-KIMB	Max eigen value test r = 0	33.46**
	Trace test r = 0	37.73**
	max eigen value and trace test: r = 1	4.27
	ADF test of residual	-4.4351
JNB-PIETM	Max eigen value test r = 0	61.84**
	Trace test r = 0	66.9**
	max eigen value and trace test: r = 1	5.06
	ADF test of residual	-6.3329

Single, double and triple asterisks indicate statistical significance at the $\alpha =$ 0.01, 0.05 and 0.10 levels, respectively. ADF critical value = -3.3771 @ 5% level .

5.4 Standard AR and TAR error correction models

As mentioned, both the standard AR and TAR error correction models are used in this study. The aim is to compare results from the two models to determine whether transaction cost has a significant effect on measuring market integration.

The standard AR error correction model was estimated from $\Delta\mu_t = \lambda\mu_{t-1}$, where $\Delta\mu_t$ is the change in the error term, λ is the adjustment coefficient and μ_{t-1} is the lag of the error correction term. The results are shown in Table 5.2, column 2.

The following steps were followed to fit a TAR error correction model. First the AR order p was selected using the partial autocorrelation function (PACF) described in section 4.4.1.1 of chapter 4[14]. The value of $p=1$ was selected for all the market pairs except JNB-KIMB for which $p=2$.

Secondly, the arranged residual (threshold variable) was arranged in ascending order and it was then refined to the order of the delay parameter before Tsay's (1989) non-parametric non-linearity test was conducted. The delay parameter d is chosen to be one[15]. According to the results found, Tsay's test rejects the null hypothesis of linearity (no threshold) in all the markets (see Table 5.2, column 8). This suggests that threshold behaviour exists in all the markets considered in the analysis. This means that a non-linear rather than a linear error correction model is appropriate.

Table 5.2: Results from the standard AR and TAR error correction models

Markets	AR λ	Half-life (Months)	TAR λ_{in}	TAR λ_{out}	Half-live (Months)	TAR C	Tsay's Test	Hansen's Test
Column 1	Column 2	Column 3	Column 4	Column 5	Column 6	Column 7	Column 8	Column 9
JNB-KIMB	-0.3644 (0.0632)	1.53	-0.2663 (0.2067)	-0.3743 (0.0666)	1.48	0.0875	15.94*	163.66*
JNB-PIETM	-0.1298 (0.0885)	4.98	-0.6989 (0.2586)	-0.6150 (0.0774)	0.73	0.0446	16.78*	153.41*
JNB-DURB	-0.3292 (0.0753)	1.74	-0.4085 (0.3094)	-0.5615 (0.0738)	0.84	0.0391	8.82*	95.68*
JNB-PE	-0.4006 (0.0681)	1.35	-0.3591 (0.2886)	-0.4374 (0.0687)	1.21	0.1129	82.84*	118.39*
JNB-CAPT	-0.1573 (0.0938)	4.05	-0.6014 (0.3439)	-0.7014 (0.07844)	0.57	0.0606	3.07**	98.07*
JNB-BLOM	-0.2323 (0.0731)	2.62	-0.6804 (0.2614)	-0.4274 (0.0688)	1.24	0.0512	106.77*	130.76*
JNB-TSHW	-0.2076 (0.0842)	2.98	-0.8458 (0.3059)	-0.6034 (0.07421)	0.75	0.0271	39.77*	159.37*

λ_{in} and λ_{out} are autoregressive parameters within and outside the neutral band, respectively.
C is the estimated threshold.
Half-live are the number of months required for one-half of a deviation from equilibrium to be eliminated.
Numbers in parentheses are standard errors.
Single, double and triple asterisks indicate statistical significance at the $\alpha =$ 0.01, 0.05 and 0.10 levels, respectively.

Thirdly, a grid search algorithm was conducted based on arranged residuals. Thresholds were identified using a grid search that minimizes the sum of squared error of the residuals (see Table 5.2, column 7). Following identification of an estimate of threshold values from the grid

[14] The autoregressive order p is the order of lag in an autoregression. That is, the value of p determines the number of lagged relations between a variable and its lag.

[15] The delay parameter d defines the number of lags appropriate to the error correction term in the threshold autoregression.

algorithm, the significance of the threshold effect was conducted using Hansen's simulation test. The test is conducted by first obtaining the least square estimates and residual variance using sequential conditional least squares with both standard AR and TAR models using 100 bootstrap simulations as dependent variables. The p-values are calculated by taking the percentage difference between the test statistics derived from the simulated sample and the observed test statistics. Hansen's test results are given in Table 5.2, column 9.

According to the test results the identified thresholds values are statistically significant at the 1% level of significance. The threshold values in column 7 of Table 5.2 represent the amount by which the prices in the spatial markets have to differ in order to trigger an equilibrating adjustment in the outer regime. For example, apple prices in the Johannesburg and Cape Town FPMs must be at least 6.1% different to exceed the threshold and trigger adjustment. For Johannesburg and Tswhane the threshold is just 2.7%, for Johannesburg and Port Elizabeth it is about 11.3% and for Johannesburg and Durban it is 4.0%. The results are consistent with *a priori* expectations that the greater the distance between markets the larger the threshold.

Fourthly, after identification of the threshold values, error correction representation of market prices was considered. A TAR error correction model was fitted (conditional on the identified thresholds) for each market. The threshold autoregressive error correction model was estimated from $\Delta\mu_t = \lambda_{in}\mu_{t-1} + \lambda_{out}\mu_{t-1}$, where the λ_{in} and λ_{out} are coefficients in the model that represent observations falling within (i.e. $|\mu_{t-1}| \leq c$) or outside the symmetric band (i.e.$|\mu_{t-1}| > c$) defined by the threshold c, the μ_{t-1} represents the error correction term, while c is the threshold. The TAR error correction model is then estimated using ordinary least squares. The results are shown in Table 5.2, columns 4 and 5

At this point note should be taken that Table 5.2 shows larger adjustment coefficients for the TAR model than for the standard AR model. This is an indication that price adjustments are faster in TAR models than standard AR models. This suggests that it is better to use TAR models than AR models because TAR models give a more reliable result. Moreover, this comparison provides proof that TAR models are superior to standard AR models in analyzing time series price data. In addition, the results show that transaction costs have a significant influence on spatial market linkages. These findings are in line with the results found by Goodwin and Piggott (2001) and Obstfeld and Taylor (1997).

As a next step the deviation half-life were calculated using the formula $\ln(0.5)/\ln(1+\lambda)$ (Obstfield and Taylor, 1997; Goodwin and Piggott, 2001), where ln is the natural log and λ represents the adjustment in the outer regime. According to Table 5.2, columns 3 and 6, half-life ranging from less than a month to more than one month were detected. The findings are consistent with the expectation that distance is positively correlated with slower speed of adjustment and larger threshold behaviours (Goodwin and Piggott, 2001), i.e. on average it takes a longer time for one-half of a deviation from equilibrium to be eliminated for markets that are more distant from Johannesburg. Cognisance should however also be taken that distance is not necessarily the only factor that will determine the speed of adjustment. The speed of adjustment could be determined by a combination of the distance between two points, the quality and level of maintenance of transportation infrastructure, communication links and the existing supply and demand balances among surplus and deficit producing regions. A good example in this case is Cape Town (a more distant region from Johannesburg than all other regions) where the calculated half-life is less than the values obtained for regions in closer proximity of Johannesburg (see Table 5.2). The Western Cape is a surplus apple production area, but is at the same time very well linked with the Johannesburg FPM in terms of road infrastructure and communication. The same infrastructure serves, for example, the Bloemfontein FPM but this market is a deficit market. Moreover, the results indicate that prices in surplus regions will respond quicker to price changes in the reference market provided that infrastructure and communication linkages are more or less homogeneous. In other words, if prices in Cape Town FPM are generally under pressure due to surpluses and the price on the Johannesburg FPM increases marketers in the former region will react faster to remove surpluses at a specified threshold value to ensure better returns.

A comparison of the deviation in half-life for the AR and TAR models shows that threshold models adjust much faster in response to deviations from the equilibrium when threshold behaviour is taken into account (Table 5.2, compare columns 3 and 6).

5.4 Multivariate TVECM

In the previous section univariate TAR was used to investigate a linear threshold-type adjustment. Goodwin and Holt (1999), however, suggest that the presence of asymmetry may result in dynamic responses that may be non-linear. In such a case, various adjustments from alternative threshold regimes may be involved. Multivariate TVECM are used to capture these dynamic adjustments in an error correction representation.

5.4.1 Estimation of asymmetric threshold parameters

In section 5.3 symmetric threshold values were identified using a grid procedure. This is a residual based criterion and is inappropriate in multivariate vector error correction models involving systems of equations. Therefore, criteria that contain information about the variance-covariance matrix of the residual of TVECM are preferred by many researchers to identify thresholds (Serra and Goodwin, 2002). Therefore asymmetric thresholds were selected using a two-dimensional grid search method[16]. Two approaches were used to select threshold values. Firstly, a similar approach to that by Serra and Goodwin (2002) was used, and thresholds were selected by minimizing the logarithm of the determinant of the variance-covariance matrix of the residuals $(\sum)$. The second approach involves minimization of the trace of $(\sum)$ matrix, i.e. the thresholds are searched over 10% and 90% of negative and positive lagged error correction terms. The search was restricted to a minimum of 20 observations in the each of the outer regimes. The asymmetric thresholds values are given in Table 5.3.

According to the results, the threshold values identified using both the minimum of log determinant and the Trace are approximately the same for all the markets, i.e. it appears that the one modeling framework is not significantly superior over the other.

Calculated threshold values are consistent with *a priori* expectations, i.e. they are found to be positively correlated with distances between markets. For example, the threshold values for Johannesburg-Cape Town and Johannesburg-Port Elizabeth are the largest. Also,

[16] Two-dimensional search allows for an iterative search on both positive and negative of the arranged residual of the price series to select positive and negative thresholds

Johannesburg-Tswane are the least separated markets and the threshold values found for these two markets are expectedly the lowest.

Table 5.3: Threshold values and sup-LR Test

Markets	Tsay's Test	Minimum of log determinant of covariance matrix			Minimum of trace of covariance matrix		
		Negative threshold	Positive threshold	Sup-LR Test (p-value)	Negative threshold	Positive threshold	Sup-LR Test (p-value)
JNB-CAPT	3.07**	-0.0575	0.08498	215.23 (0.0000)	-0.1170	0.02554	204.65 (0.0000)
JNB-BLOM	106.77*	-0.0514	0.0502	241.75 (0.0000)	-0.0335	0.0453	235.82 (0.0000)
JNB-TSW	39.77*	-0.0245	0.0171	284.95 (0.0000)	-0.0245	0.0315	283.38 (0.0000)
JNB-PE	82.84*	-0.0871	0.11005	222.06 (0.0000)	-0.0871	0.1064	220.55 (0.0000)
JNB-DURB	8.82*	-0.0636	0.0212	236.22 (0.0000)	-0.0581	0.0339	225.06 (0.0000)
JNB-KIMB	15.94*	-0.0664	0.0852	244.23 (0.0000)	-0.0547	0.10896	220.67 (0.0000)
JNB-PIETM	16.78*	-0.0446	0.0432	365.63 (0.0000)	-0.0356	0.0432	330.94 (0.0000)

Single, double and triple asterisks indicate statistical significance at the $\alpha =$ 0.01, 0.05 and 0.10 levels, respectively.

As can be seen from Tables 5.2 and 5.3, the estimated threshold values for symmetric bands (TAR model) lie exactly within those implied for asymmetric bands (TVECM) in two markets, namely Johannesburg-Bloemfontein and Johannesburg-Pietermaritzburg. The bands are also very close for other market pairs. The similarity in the threshold values for the symmetric TAR and asymmetric TVECM models implies that the market structure is such that apples flow in both directions and that no market can be totally considered as importer or exporter. The sup-LR statistics test in Table 5.3 provides an indication of whether a linear VECM or asymmetric TVECM is better to use. According to the results, there is a significant asymmetric threshold effect in the TVECM. The p-values of the sup-LR test indicate that thresholds identified by both log determination and trace criteria are statistically significant at 1% level.

5.4.2 Parameter estimates of TVECM

In this section parameter estimates of TVECM are analyzed. Price adjustment in the TVECM is also compared to VECM. The aim is to see if the TVECM adjust faster to deviations from equilibrium than the VECM[17] as is *a priori* expected. This test augments the Sup-LR test conducted in section 5.4.1 and the findings in section 5.3 where TAR was found to be more appropriate (Table 5.2, columns 2 and 4). Note that the TVECM models are fitted on the basis of the asymmetric threshold values that were calculated in section 5.4.1. Parameter estimates and their standard errors are presented in Tables 5.4 to 5.10.

The parameter estimates are obtained conditional on the thresholds which are not identified under the null hypothesis. The null hypothesis entails that co-integrating relations are linear in contrast to the alternative that it exhibits threshold nonlinearities[18] (Balke and Fomby, 1996). Therefore, Goodwin and Harper (2000) state that care should be taken with the direct use of standard error estimates to test for the significance of individual parameter estimates. Results in Tables 5.4 to 5.10 indicate that there are significant dynamic relations among the price series studied for the different market pairs. In the TVECM bi-directional causality was found in five market pairs, namely, JNB-DURB, JNB-CAPT, JNB-BLOM, JNB-PIETM and JNB-TSW. This means that lagged price changes in JNB have a significant influence on price changes in these markets and vice versa. For the remaining two market pairs, i.e. JNB-KIMB and JNB-PE, only unidirectional causality was found. In addition, as expected, the error correction terms are found to be significant in all market pairs. The results for the VECM indicate bi-directional causality exist only between Johannesburg and Cape Town. Lagged price changes in Johannesburg influences price changes in Kimberley and Cape Town. On the other hand, lagged price changes in Durban, Tshwane and Bloemfontein significantly influences Johannesburg prices. No significant price leadership was identified between Johannesburg and Pietermaritzburg and Johannesburg and Port Elizabeth.

Price adjustment estimated with the VECM and TVECM differs for different markets pairs. The estimated coefficient for adjustment to long-term equilibrium in the VECM show a

[17] The VECM was estimated with equation 4.16 in chapter 4, while TVECM was estimated with equations 4.17 to 4.19.
[18] Tsay's (1989) F-test is used in most similar studies to test for the presence of threshold-type nonlinearity.

significant reaction to apple prices in the Johannesburg market in five market pairs namely, Johannesburg-Durban, Johannesburg-Cape Town, Johannesburg-Kimberley, Johannesburg-Pietermaritzburg, and Johannesburg-Port Elizabeth. Significant reactions to apple prices also occurred in Bloemfontein, Pietermaritzburg and Tshwane (Tables 5.6, 5.8, & 5.9 columns 5)[19].

In the TVECM significant adjustment to apple prices in Johannesburg in the 1st and 3nd regimes was found in six markets pairs (JNB-DURB, JNB-CAPT, JNB-BLOM, JNB-KIMB, JNB-PIETM, and JNB-TSW)[20]. The result is consistent with expectations because the absolute deviation from long-term equilibrium is expected in the outer regimes where adjustment occurs. However, significant adjustments were also found in the second regime in three market pairs, namely, JNB-DURB, JNB-KIMB, and JNB-PE, which is surprising because significant adjustment is not expected in the second regime. The error correction coefficients in the 1st and 3rd regimes of the TVECM indicate a much faster adjustment compared to VECM. This result is slightly different in the JNB-PE pair. This could be attributed to the larger distance separating the two markets, i.e. the slow speed of adjustments has been found to be positively correlated with distance (Goodwin and Piggott, 2001).

Table 5.4: TVECM and VECM parameter estimates & summary statistics: Durban-Johannesburg FPMs

	TVECM			VECM
Variables	Regime I	Regime II	Regime III	
$\Delta Durb$				
α_t	-0.114701* (0.024546)	0.003602*** (0.014047)	0.10371* (0.012719)	0.003341 (0.00908)
$\Delta Durb_{t-1}$	0.69159* (0.274296)	0.740194* (0.186138)	-0.217432 (0.215803)	0.478129** (0.17177)
ΔJnb_{t-1}	-0.31985 (0.197548)	-0.309058** (0.15362)	0.414117*** (0.222987)	-0.172041 (0.14457)
μ_{t-1}	-0.147276 (0.328078)	0.143616 (0.404493)	-0.398056 (0.246054)	0.085332 (0.19458)
ΔJnb				
α_t	-0.000423 (0.021006)	0.025261** (0.012021)	0.038302* (0.010884)	0.002878 (0.00594)
$\Delta Durb_{t-1}$	0.477159** (0.234738)	0.686047* (0.159294)	-0.040911 (0.184681)	0.383209** (0.11241)
ΔJnb_{t-1}	-0.123238 (0.169058)	-0.230405*** (0.131465)	0.18111 (0.190828)	-0.040502 (0.09461)
μ_{t-1}	0.769571* (0.280764)	0.960478* (0.346159)	0.408884*** (0.210569)	0.586609** (0.12734)

Single, double and triple asterisks indicate statistical significance at the $\alpha =$ 0.01, 0.05 and 0.10 levels, respectively.
Figures in parenthesis are standard errors.

[19] Note the error correction term is considered here for the significance of price adjustment
[20] See the last row of Tables 5.4 to 5.10

Table 5.5: TVECM and VECM parameter estimates & summary statistics: Cape Town-Johannesburg FPMs

TVECM				VECM
Variables	Regime I	Regime II	Regime III	
$\Delta Capt$				
α_t	-0.080812* (0.018164)	0.022385 (0.014964)	0.138245* (0.021858)	0.006252 (0.01066)
$\Delta Capt_{t-1}$	-0.321534** (0.150605)	-0.388761* (0.152397)	-0.06252 (0.133909)	-0.342833** (0.09701)
ΔJnb_{t-1}	0.773361* (0.18582)	0.848634* (0.176196)	0.234531 (0.215362)	0.645945** (0.13074)
μ_{t-1}	-0.236306 (0.181629)	-0.668675** (0.322002)	-0.809955* (0.158754)	-0.134255 (0.10307)
ΔJnb				
α_t	0.03852* (0.015463)	0.001715 (0.012739)	-0.013804 (0.018608)	0.00299 (0.00697)
$\Delta Capt_{t-1}$	-0.306558* (0.128212)	-0.198715 (0.129738)	-0.015581 (0.113999)	-0.179219** (0.06345)
ΔJnb_{t-1}	0.76531* (0.158191)	0.64361* (0.149998)	0.425307** (0.183341)	0.688812** (0.08551)
μ_{t-1}	0.559239* (0.154623)	0.239 (0.274124)	0.241351*** (0.135149)	0.405169** (0.06741

Single, double and triple asterisks indicate statistical significance at the $\alpha = 0.01$, 0.05 and 0.10 levels, respectively.
Figures in parenthesis are standard errors.

Table 5.6: TVECM and VECM parameter estimates & summary statistics: Bloemfontein-Johannesburg FPMs

TVECM				VECM
Variables	Regime I	Regime II	Regime III	
$\Delta Blom$				
α_t	-0.090021* (0.018505)	0.005616 (0.010885)	0.125861* (0.018654)	0.003 (0.00928)
$\Delta Blom_{t-1}$	-0.224439 (0.193624)	0.469023* (0.153606)	1.082346* (0.236835)	0.248405 (0.13099)
ΔJnb_{t-1}	0.600538* (0.235348)	0.161978 (0.159934)	-0.17621 (0.266466)	0.251554 (0.14617)
μ_{t-1}	-0.429039** (0.17953)	-0.842144** (0.352575)	-1.502978* (0.202728)	0.248405** (0.13099)
ΔJnb				
α_t	0.000206 (0.018291)	0.009238 (0.010759)	0.0214 (0.018438)	0.003105 (0.00757)
$\Delta Blom_{t-1}$	-0.322855*** (0.191383)	0.384221* (0.151828)	0.989362* (0.234095)	0.291236** (0.10695)
ΔJnb_{t-1}	0.714465* (0.232625)	0.192794 (0.158083)	-0.201588 (0.263383)	0.192852 (0.11934)
μ_{t-1}	0.375928** (0.177452)	0.038705 (0.348495)	-0.5971* (0.200382)	-0.096291 (0.09988)

Single, double and triple asterisks indicate statistical significance at the $\alpha = 0.01$, 0.05 and 0.10 levels, respectively.
Figures in parenthesis are standard errors.

Table 5.7: TVECM and VECM parameter estimates & summary statistics: Kimberley-Johannesburg FPMs

	TVECM			VECM
Variables	Regime I	Regime II	Regime III	
$\Delta Kimb$				
α_t	-0.096178* (0.02264)	0.032024** (0.012708)	0.099032* (0.021929)	0.004797 (0.00946)
$\Delta Kimb_{t-1}$	-0.568269* (0.171362)	0.067968 (0.143608)	0.259525 (0.240866)	-0.26246** (0.10535)
ΔJnb_{t-1}	0.667688* (0.175503)	0.497195* (0.150521)	0.358872 (0.327496)	0.657538** (0.12591)
μ_{t-1}	-0.185136 (0.156761)	-0.629726* (0.283489)	-0.751765* (0.177966)	0.099429 (0.06273)
ΔJnb				
α_t	-0.096178** (0.02264)	0.02427** (0.01135)	-0.03656*** (0.019585)	0.002666 (0.00707)
$\Delta Kimb_{t-1}$	-0.182704 (0.153043)	0.099794 (0.128255)	0.208563 (0.215117)	0.034925 (0.07871)
ΔJnb_{t-1}	0.472139* (0.156741)	0.321353** (0.134429)	0.294316 (0.292486)	0.48682** (0.09407)
μ_{t-1}	0.542944* (0.140002)	0.434381*** (0.253184)	0.099707 (0.158941)	0.238091** (0.04687)

Single, double and triple asterisks indicate statistical significance at the $\alpha = 0.01$, 0.05 and 0.10 levels, respectively.
Figures in parenthesis are standard errors.

Table 5.8: TVECM and VECM parameter estimates & summary statistics: Pietermaritzburg-Johannesburg FPMs

	TVECM			VECM
Variables	Regime I	Regime II	Regime III	
$\Delta Pietm$				
α_t	-0.054998* (0.018197)	0.012704 (0.011094)	0.12236* (0.016252)	0.003739 (0.00906)
$\Delta Pietm_{t-1}$	0.866767* (0.194652)	0.074579 (0.174485)	0.166931 (0.236804)	0.226029 (0.14491)
ΔJnb_{t-1}	-0.344536*** (0.197713)	0.55548* (0.190191)	-0.113568 (0.235176)	0.162239 (0.15513)
μ_{t-1}	-0.7059* (0.22934)	-0.767202*** (0.400748)	-1.347636* (0.249297)	-0.313847** (0.15487)
ΔJnb				
α_t	0.03218*** (0.017879)	0.011887 (0.0109)	0.03752** (0.015968)	0.00282 (0.00713)
$\Delta Pietm_{t-1}$	0.915601* (0.191249)	0.141739 (0.171435)	0.308748 (0.232664)	0.397309 (0.11407)
ΔJnb_{t-1}	-0.37993*** (0.194257)	0.486921* (0.186866)	-0.077971 (0.231065)	0.108973 (0.12211)
μ_{t-1}	0.133959 (0.225331)	0.170263 (0.393743)	-0.342747 (0.244939)	0.169119** (0.12191)

Single, double and triple asterisks indicate statistical significance at the $\alpha = 0.01$, 0.05 and 0.10 levels, respectively.
Figures in parenthesis are standard errors.

Table 5.9: TVECM and VECM parameter estimates and summary statistics: Tshwane-Johannesburg FPMs

	TVECM			VECM
Variables	Regime I	Regime II	Regime III	
		$\Delta Tshwane$		
α_t	-0.074036* (0.018546)	0.026643* (0.013795)	0.071993* (0.015314)	0.002608 (0.00841)
ΔTsw_{t-1}	0.448985 (0.35473)	0.797675* (0.396668)	0.549634*** (0.320187)	0.458085 (0.23546)
ΔJnb_{t-1}	-0.044514 (0.344901)	-0.244586 (0.357642)	-0.073153 (0.340757)	-0.026188 (0.22778)
μ_{t-1}	-0.061624 (0.384187)	0.646965 (1.056646)	-1.215303* (0.432014)	0.079921** (0.25844)
		ΔJhb		
α_t	-0.02208 (0.017777)	0.026896* (0.013223)	0.027055*** (0.014679)	0.002543 (0.007)
ΔTsw_{t-1}	0.379948* (0.340017)	0.827881* (0.380216)	0.538137 (0.306907)	0.492228** (0.19605)
ΔJnb_{t-1}	0.024541*** (0.330596)	-0.299433 (0.342808)	-0.08712*** (0.326624)	-0.058869 (0.18965)
μ_{t-1}	0.745741** (0.368253)	1.251705 (1.012822)	-0.191538 (0.414097)	0.575542 (0.21518)

Single, double and triple asterisks indicate statistical significance at the $\alpha = 0.01$, 0.05 and 0.10 levels, respectively.
Figures in parenthesis are standard errors.

Table 5.10: TVECM and VECM parameter estimates and summary statistics: Port Elizabeth-Johannesburg FPMs

	TVECM			VECM
Variables	Regime I	Regime II	Regime III	
		$\Delta Port\ Elizabeth$		
α_t	-0.213034* (0.022447)	0.04046* (0.015248)	0.227929* (0.02)	0.003544 (0.01444)
ΔPe_{t-1}	0.263001* (0.087428)	0.210229 (0.144841)	-0.029234 (0.130592)	0.217237 (0.11788)
ΔJnb_{t-1}	0.145754 (0.103909)	-0.542003* (0.173031)	0.018772 (0.278726)	-0.119161 (0.14282)
μ_{t-1}	-0.699857* (0.095877)	-0.735174* (0.253584)	-0.743875* (0.129777)	-0.149221 (0.10858)
		ΔJnb		
α_t	-0.007496 (0.015792)	0.041199* (0.010727)	0.020183 (0.01407)	0.004016 (0.00551)
ΔPe_{t-1}	0.086809 (0.061506)	-0.053501 (0.101897)	0.06417 (0.091873)	0.045861 (0.04502)
ΔJnb_{t-1}	0.29989* (0.073101)	0.148634 (0.121729)	0.133331 (0.196087)	0.232924** (0.05455)
μ_{t-1}	0.31649* (0.06745)	0.378318** (0.178399)	0.164735*** (0.0913)	0.318825** (0.04147)

Single, double asterisks indicate statistical significance at the $\alpha = 0.01$ and 0.05 levels, respectively. Figures in parenthesis are standard errors.

5.4.3 Arbitrage efficiency and time of switching between regimes

In this section, arbitrage efficiency and time of switching between regimes are considered to determine whether major apple markets in South Africa are integrated with the central apple market located in Johannesburg FPM. Firstly, the aggregate frequency of observations occurring in one of the three regimes is determined by using asymmetric thresholds discussed in previous sections. This is useful to determine whether the width of the neutral band is consistent with the theory, i.e. that the most distant markets have the widest neutral band compared to those markets located closer to one another (Goodwin and Piggott, 2001). Secondly, the persistence of price differences occurring in each regime and the time it takes for price differences to switch between regimes will be considered. If price differences persistently occur in the neutral band it is considered as a case where markets are strongly integrated (Goodwin and Piggott, 2001). Furthermore, according to Goodwin and Piggott (pg.311, 2001) "T*he greatest degree of switching among regimes ... is realized generally by those markets that are the most distant*".

As explained in section 5.1, market integration implies that price differences between trading regions should be less than or equal to transaction costs (Goodwin and Piggott, 2001). This is the case when price differences occur in the neutral threshold regime band. According to the results shown in Table 5.11, the neutral band (i.e., regime II) has registered the highest frequency in terms of the percentage of observations that occur in it for any given threshold in all the market combinations except in the case of JNB-TSW.

The results found in respect of the relationship between the width of neutral bands and distances between markets were inconclusive. For example, the neutral band for the most distant FPM markets, namely JNB-CAPT, accounted for 48 percent of the price differences, while the closest FPM markets, namely JNB-TSW, accounted for 38 percent of the price differences. This finding is in line with *a priori* expectations. However, further investigation reveals that the neutral band for the JNB-BLOM FPMs accounted for about 58 percent of the price differences which is more than that observed for the JNB-CAPT FPMs. In terms of kilometers separating the JNB-BLOM and JNB-CAPT FPMs, the latter markets are the most distant and hence its neutral band should be larger than the one for the JNB-BLOM FPMs.

This implies that there is not sufficient evidence to support the hypothesized relationship between distances and width of neutral bands.

TABLE 5.11: PERCENTAGE OF OBSERVATIONS FALLING IN REGIME I, II AND III

Markets	Log Determinant			Trace		
	Regime I	**Regime II**	**Regime III**	**Regime I**	**Regime II**	**Regime III**
JNB-CAPT	29.76	47.62	22.62	14.29	43.45	42.26
JNB-BLOM	22.62	58.33	19.05	33.33	45.83	20.83
JNB-TSW	24.40	37.50	38.10	24.40	47.62	27.98
JNB-PE	32.74	35.71	31.55	32.74	35.12	32.14
JNB-DURB	14.29	44.64	41.07	16.67	46.43	36.90
JNB-KIMB	25.60	51.19	23.21	30.36	50.00	19.64
JNB-PIETM	22.02	51.19	26.79	26.79	46.43	26.79

In order to measure the degree of market integration between the different FPMs for apples the persistency of price differences occurring in each regime and the time it takes for price differences to switch between regimes was investigated. If monthly price differences persistently occur in the neutral band it is a sufficient condition for market integration. On the other hand, if monthly price differences persistently occur in either Regime I or Regime III it is indicative that monthly price differences are higher than transaction costs and hence signify a disequilibrium situation. Moreover, a situation of persistent disequilibrium indicates that markets are not integrated.

In addition, the timing of switching of monthly price differences between Regimes was investigated to determine whether Regime switching could be attributed to the structural breaks, e.g. deregulation of the market.

Figure 5.1 shows the results pertaining to persistency of price differences occurring in each regime and the timing of switching of monthly price differences between Regimes for the Cape Town and Johannesburg market pairs . The rest of the results are shown in Figures D1 to D6 of Appendix D. The results indicate that monthly price differences persistently occur in Regime II for all markets, although it appears to be weaker in the case of the JNB-PE FPMs, especially since 2002. This is indicative that the FPMs are relatively well integrated. With regard to the timing of switching between Regimes, there is not sufficient evidence to indicate that structural change, such as deregulation, has impacted on the level of market integration.

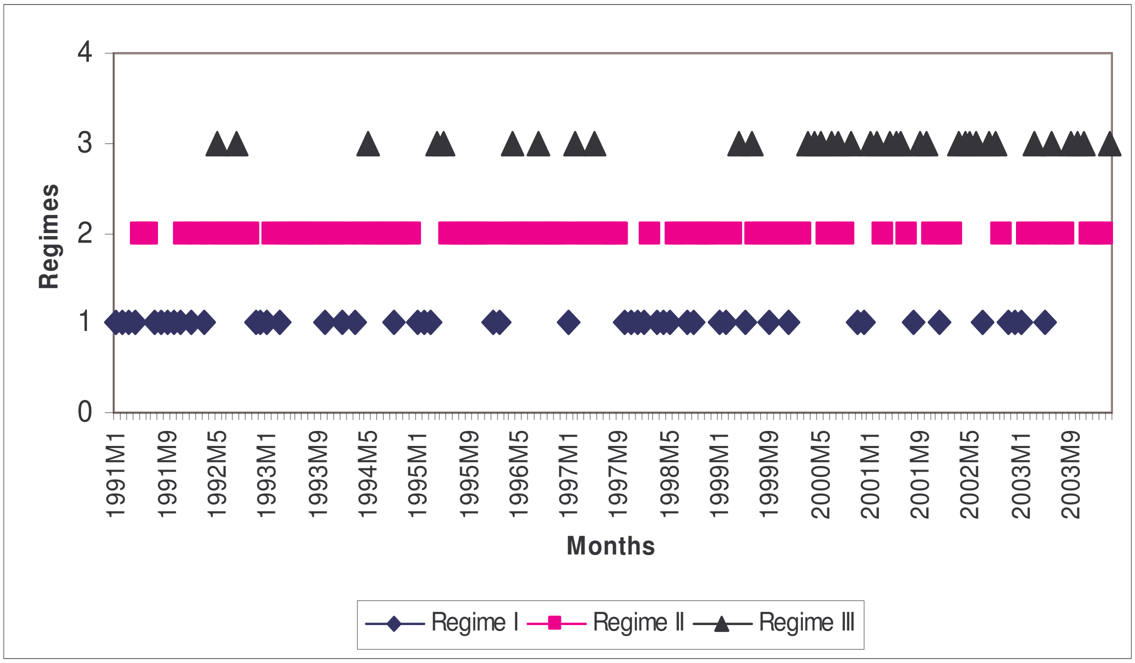

Figure 5.1: Persistency of price differences occurring in each Regime and Regime switching: Cape Town-Johannesburg FPMs

5.4.4 Generalized impulse response function

In this section, in an attempt to get a better understanding of the dynamic price interrelationships, an impulse response function was estimated. This was done by adjusting the last observation in the time series for the Johannesburg FPM by one-half standard deviation to represent positive and negative shocks.

Figure 5.2 depicts the response in the Cape Town FPM due to price shocks in the Johannesburg FPM (Note that the Figures for the other FPMs are in Appendix E and the discussion below refers to all FPMs). Considering all markets, the time plot shows that both positive and negative shocks lead to almost a similar pattern of adjustment. For example, it takes about five to seven months for the Cape Town and Kimberley FPMs to respond completely to both positive and negative price shocks in the Johannesburg FPM. Generally, it takes about six to twelve months for other FPMs to adjust completely in response to a price shock in the Johannesburg FPM. The positive and negative shocks converge to equilibrium and do not show any tendency to deviate from equilibrium in the long run. These responses are consistent with long-run market integration. It should, however, be noted that these responses are conditional on the size of the shock due to asymmetric price relationships among the markets.

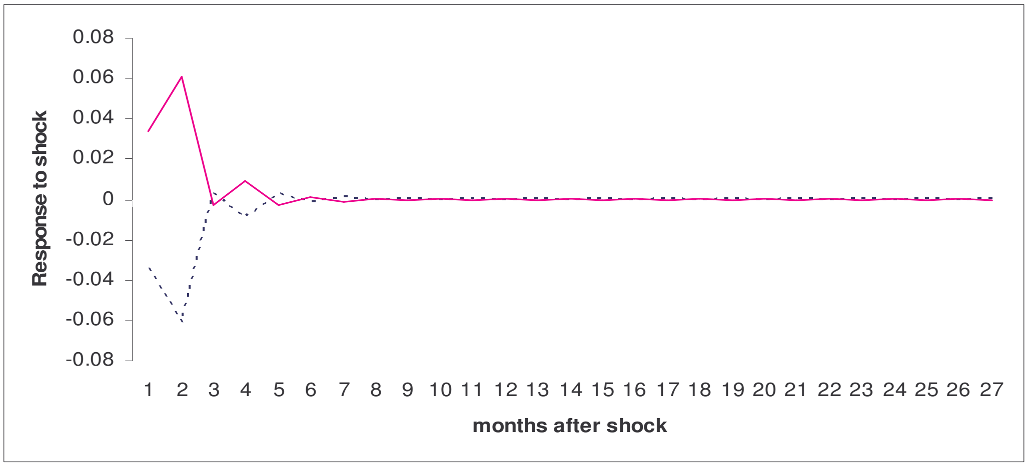

Figure 5.2: Response of the Cape Town FPM to positive and negative price shocks in the Johannesburg FPM

5.5 Summary

In this chapter the results of the procedures discussed in chapter 4 were presented. The aim was to determine the level of market integration that exists between selected FPMs that trade apples.

Results show that adjustments to deviations from equilibrium in threshold models are faster than in standard autoregressive models. This suggests that transaction cost significantly influences spatial price linkages. A comparison of the deviation in half-life for the AR and TAR models shows that threshold models adjust much faster in response to deviations from the equilibrium when threshold behaviour is taken into account

Regime switching estimates to see whether arbitrage efficiency improved over time, especially after deregulation in 1996, show that the inner band (regime II) has the highest frequency in terms of the percentage of observations that fall in it for any given threshold in all but one market combination. This implies that deviations from equilibrium are hardly large enough to exceed the neutral band implied by transaction costs. The results also show that the average percentage that fall in the inner band after deregulation increased in three out of seven markets considered.

The non-linear impulse response function to investigate how price shocks are transmitted and how long it takes for shocks to be eliminated in alternate markets revealed that it takes about six to twelve months for one-half standard deviation shocks to Johannesburg prices to be completely eliminated.

CHAPTER 6

Summary, Conclusions and Recommendations

6.1 Introduction

It is increasingly recognized that the formulation of market-enhancing policies to increase the performance of the local market requires a better understanding of how the market functions. Aggregate market performance is better understood by studying the level of market integration that exists, which in turn is affected by transaction costs in the value chain. In the context of this study market performance was investigated by studying the impact of deregulation on average market prices, price spreads and market volatility in the domestic apple industry. A review of the literature showed that no study has been specifically conducted to measure the extent of market integration and price transmission in this industry. The primary objective of this study was to measure market integration for apples on the South African FPMs to determine the existence of long-run price relationships and spatial market linkages.

6.2 Literature review

Various approaches that have been used to study market integration have been criticized because transaction costs were not considered. Recent improvements have been the introduction of the parity bound and threshold co-integration models that account for transaction costs. The price relationships between spatially separated markets are generally analyzed within the framework of spatial price equilibrium theory. The key assumption underpinning spatial price equilibrium theory is that price relationships between spatially separated competitive markets depend on the size of transaction costs. This implies that transaction cost plays a key role in the study of spatial price relationships and should not be ignored. When the price difference between different markets exceeds transaction costs, arbitrage opportunities will be created and profit seeking merchants will seek to exploit such opportunities, by purchasing commodities from a low-price surplus market and transferring them to a higher-priced deficit market. Several theories found the transaction cost and arbitrage concepts very useful in the analysis of market integration. Different approaches can be applied to measure market integration.

Two methods have been identified as the most appropriate methods of measuring market integration. These are the parity bound model (PBM) and the threshold co-integration method. The parity bound model posits that a co-integration test that ignores the influence of transaction cost is inconclusive and should not be used as a measure of market integration. Baulch (1997) proposed the estimation of transaction cost and use of regime switching models. Parity bound models are criticized for directly observing transaction costs. Tong (1978) originally introduced non-linear threshold time series models based on transaction costs. Tsay (1989) developed techniques for testing TAR models. Balke and Fomby (1997) introduced threshold co-integration to the study of market integration. Threshold co-integration has been used both in univariate and multivariate time series to study market integration.

Both univariate and multivariate threshold models were used in this study to investigate price transmission and adjustments to equilibrium in response to changes in economic system.

6.3 Industry overview

The total area planted to apples in South Africa in 2003 was estimated at 22 379 hectares (OABS, 2004).This represents 29% of the total area planted to deciduous fruits. Groenland constitutes 34% of the total area under production; Ceres 22%, Langkloof-East 17% and Villiersdorp 16%. Of the total apple production 89% comes from these four areas (OABS, 2004). From 1991/92 to 2002/03 apple production averaged 574 850 tons with a standard deviation of 43 922 tons. Apple production is more or less equally distributed between exports and the local market. The largest proportion of production that is not exported is either consumed as fresh fruit or processed.

Prices on the local market are largely influenced by seasonality in production, the perishability of produce and the amount of apples exported (availability of apples on the local market). The impact of seasonality is to some extent cushioned by cold storage facilities that ensure regular apple supplies in the local market. Demand factors such as consumer habits, substitution between products and per capita income also influence prices. Prices are the lowest between March and June since this is the period when large volumes of apples enter the local market. The variability in prices in different markets also increases as the distance from the surplus apple producing regions increase.

The 16 FPMs remain an important market outlet for fresh produce in South Africa. They serve both as a distribution channel and also as a platform for price determination.

6.4 Methodology and data used

Long-term relations between price variables were investigated using ADF and Johansen test. Threshold models and error correction representation of the models were fitted to evaluate the dynamic relationship among apple market prices. The statistical significance of the estimated thresholds was evaluated using Tsay's (1989) and Hansen's (1997) test statistics. To better explain the interrelationship among prices, switches between regimes and the response to price shocks among the markets were evaluated using regime switching and impulse response functions.

The estimations were based on time series price data spanning from 1st January 1991 to 30th March 2004 for selected FPMs. The markets included are Johannesburg (JNB), Cape Town (CAPT), Tshwane (TSW), Bloemfontein (BLOM), Port Elizabeth (PE), Durban (DURB), Pietermaritzburg (PIETM) and Kimberley (KIMB).

6.5 Summary of results

6.5.1 Effect of market reform on average price, price spread and risk

The impact of deregulation on real market prices, price spread and price volatility was investigated. Monthly time series price data obtained for the eight selected FPMs over 14 years (from 1991 to 2004) were deflated using the national price index for fruits and nuts collected in metropolitan areas by Statistics South Africa. The changes in the average real prices were calculated by taking the difference between the before and after deregulation mean prices. The investigation revealed that there was a statistically significant decline in real prices in six of the eight markets investigated. A statistically significant relation in prices (price spread) existed between the Johannesburg FPM and five other FPMs. The analysis further showed that the price spreads between these markets declined after deregulation; and that variation in real apple prices declined for five of the eight markets after deregulation.

6.5.2 Threshold co-integration

Standard AR and TAR models were estimated. Adjustments coefficients in the outer regimes of the TAR model are faster than in the inner band indicating outer regimes adjust faster to deviation from equilibrium than the inner regime. Standard autoregressive and threshold autoregressive error correction models were compared to determine whether transaction cost has significant effects in measuring market integration. Larger adjustment coefficients were found in the TAR model. This is an indication that price adjustments are faster in threshold autoregressive TAR models than in standard autoregressive AR models. Also half-life deviations in the TAR model are much smaller than the AR model. The TAR model requires less time for one-half of the deviation from equilibrium to be eliminated than the standard AR model. Therefore, it is better to use TAR models than AR models because TAR models give a more reliable result. Goodwin and Piggott (2001) and Obsfeld and Taylor (1997) found similar results.

The TAR models estimated are symmetric are in nature meaning that arbitrage is induced in the same way between the markets no matter where prices are higher. For example, a threshold value of 5.3% identified for JNB-BLOM indicates that Johannesburg and Bloemfontein prices must differ at least by 5.3% to exceed the threshold and trigger equilibrating price adjustments. Symmetric thresholds identified for other markets are JNB-PE (23%), JNB-DURB (5.8%), JNB-CAPT (13%), JNB-TSW (0.01%) and JNB-KIMB (16%).

If markets are linked by an asymmetric relationship the threshold for the two markets will differ. This is seen in a situation were one market is characterized as an importing market and the other as an exporting market. In an asymmetric relationship causality is in either direction, hence no market is regarded as the lead or lag market. However, the threshold values identified for symmetric and asymmetric models in this study are close approximates. This implies that no two selected apple markets in South Africa can be characterized as purely importing or exporting to each other.

With the threshold vector error correction model bidirectional and unidirectional causality was found among market prices. There are bidirectional relationships between Johannesburg and five markets (Durban, Cape Town, Bloemfontein, Pietermaritzburg, and Tshwane). Unidirectional relationships exit between Johannesburg and two markets (Port Elizabeth and Kimberley).

Arbitrage efficiency in the selected apple markets was also analyzed. The objective was to use the concept of arbitrage efficiency to determine whether the major apple markets in South Africa are integrated with the central market in Johannesburg. According to the result, the neutral band (regime II) has the highest percentage of observations that fall in it for any threshold value in all the combination of markets considered in the analysis.

Similar results were obtained with both log determinant and trace threshold selection criteria[21].

Following the investigation of arbitrage efficiency, the relationship between the neutral band and distance between markets was analyzed. The aim was to determine whether the width of the neutral band is consistent with the theory that the most distant markets have the widest neutral band compared to those at closer proximity (Goodwin and Piggott 2001). According to the results, the most distant market JNB-CAPT accounted for 48 percent of the number of observations falling into the neutral band, while the closest market JNB-TSW accounted for 37 percent. Even though this is in line with expectations the overall result could be regarded as being inconclusive. This is because JNB-BLOM accounted for about 58 percent of the price differences which is more than that for the most distant market JNB-CAPT. This contradicts *a priori* expectations. This implies that there is not enough evidence to support the hypothesized relationship between distance and width of the neutral band.

To further investigate market integration in the selected apple markets, regime switching estimates were analyzed. The objective was to investigate, firstly, the persistence of deviations from equilibrium and secondly, the timing of switching of observations among regimes. The results show that no persistent deviation from equilibrium existed for all the market pairs studied, except for JNB-PE, which showed some degree of deviation from equilibrium from 2002 M12.

For the second objective, the aim was to determine whether regime switching could be attributed to the structural break in the apple industry due to market deregulation in 1996. With regard to the timing of switching between markets no clear evidence could be found to support improved market integration after market deregulation in 1996. Therefore, it can be concluded that apple markets in South Africa are strongly and spatially integrated regardless of the reform program.

[21] Except for the Johannesburg and Tshwane pair in the log determinant selection criterion, where there is a slight difference in the result.

To better interpret the dynamic interrelationship among prices, an impulse response function was estimated. The objective was to evaluate the response to positive and negative price shocks in the Johannesburg market in alternate markets. According to the result, it takes about six to twelve months for positive and negative shocks to be completely eliminated in all the markets. Generally positive and negative shocks converge to equilibrium and do not show any tendency to diverge in the long-run. This confirms a high degree of market integration.

6.6 Conclusions and recommendations

6.5.1 Major conclusions

The study confirmed that market price linkages and the inter-relationship among spatial markets are important in economic analysis. Inter-market price linkages and the speed of adjustment to shocks show that transaction costs have a significant impact in determining the degree of market integration in the South African fresh produce markets for apples. Market integration was found to be negatively correlated to distance between markets due to transaction costs. Bidirectional and unidirectional causality was found among the selected apple markets. The result shows that no particular FPM market can be said to be purely importing or exporting from each other.

Sufficient arbitrage was found to exist in the South African FPMs for apples. Price differences between the selected FPMs were in the majority of cases found to be less than or equal to transaction costs. This is consistent with spatial market integration. The impact of deregulation on real market prices, price spread and price volatility revealed that (i) a statistically significant decline in real prices in six of the eight markets investigated, (ii) a statistically significant relation in prices (price spread) between the Johannesburg FPM and five other FPMs and that the price spreads between these markets declined since deregulation; and (iii) variation in real apple prices declined for five of the eight markets after deregulation. No evidence was found that there was improved market integration among the selected South African fresh produce apple markets after deregulation in 1996.

6.5.3 Recommendations

The following recommendations stem from this study:

- Similar analyses to the ones conducted for this study should be done for the remaining FPMs. This will give a broader picture of the interrelationships among all FPMs in South Africa and could show the overall efficiency of the FPM system in South Africa. The analysis should also include other fresh produce.
- The composite commodity theorem should be further explored and implemented. Other fruits should be included in the study to analyze their substitutability and to determine if the commodity (fruit) baskets belong to the same market.
- In this study continuous trade flow and fixed transaction costs were assumed. Assuming there is discontinuity in trade and non-constant transaction costs, there will be variability in the results of TAR. Therefore, it is recommended to improve the TAR model to account for variable transaction costs and trade flows. In such a case variable and multiple threshold values (transaction costs) should be calculated.

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APPENDIX A

Other unit root test procedures

Granger (1987) proposed seven test statistics for unit root test, namely (i) The Durbin Watson co-integration regression test (CRDW), (ii) Dickey Fuller test (DF), (iii) Augmented Dickey Fuller test (ADF), (vi) Restricted vector Autoregression test (RVAR), (v) Augmented Restricted Vector Autoregression (ARVAR), (vi) Unrestricted Vector Autoregression (UVAR) and (vii) Augmented Unrestricted Vector Autoregression (AUVAR). According to Granger, the powers of these test statistics are computed under various alternatives. Their use depends on their performance and relevance to the researcher. A brief description of the characteristics of the various test statistics are as follows.

Let y_t and x_t represent two economic variables. A linear combination of the two series using Ordinary least square (OLS), will give,

$$y_t - \alpha - \beta x_t = e_t \tag{A1}$$

Where,

e_t is the residual series.

Estimates of residual error $\hat{e}$, can be obtained from the estimates of parameters of the co-integrating regression;

$$\hat{e} = y_t - \hat{\alpha} - \hat{\beta} x_t \tag{A2}$$

$$\text{(i) CRDW} = \left(\sum_{t=2}^{T} (\hat{e}_t - \hat{e}_{t-1})^2 \right) / \left(\sum_{t=1}^{T} \hat{e}_t^{\,2} \right) \tag{A3}$$

Using this test statistics, the residual of co-integration regression of levels price series is tested for stationarity. If they are non-stationary, the CRDW statistics will approach zero. The null hypothesis of no co-integration is rejected for CRDW values significantly different from zero. The CRDW statistics is approximately equal to $2 - 2\hat{\rho}$, where $\hat{\rho}$ is the estimated autoregressive parameter for the residual errors from the co-integration regression (Judge et al 1985) Quoted in Goodwin and Schroeder (1991). According to Engel and Granger (1987) if DW exceeds 0.386, the null is rejected and co-integration is confirmed. CRDW is used where the time series is observed and the null and the alternative are of the first order models.

However, CRDW have limitations. Its critical value is usually sensitive to the particular parameter within the null and as such, choice of critical value to use could be practically problematic (Granger 1987). It is unstable and not recommended for use by Engel and Granger.

(ii) DF: In this case the residual is tested for stationarity by running an auxiliary regression. If there is a unit root, then the two series are not co-integrated. DF test depends on the estimate of;

$$\Delta\hat{e}_t = -\phi\hat{e}_{t-1} + \in_t \tag{A4}$$

Where $\hat{e}_t$, is the first estimate of the residual from equation (A.1) and Δ is the first difference. A test statistics DF is obtained from the ratio of the estimated ϕ to its standard error. The null hypothesis of no co-integration is rejected for values of ϕ significantly different from zero.

(iii) ADF: It allows for more dynamics in the DF regression. ADF test is over-parametized in the first order case but correctly specified in the higher order. It differs from DF test by containing p lagged value of the differenced residual errors.

$$\Delta\hat{e}_t = -\phi\hat{e}_{t-1} + \theta_1\Delta\hat{e}_{t-1} + ...+ \theta_p\Delta\hat{e}_{t-p} + \in_t \tag{A5}$$

The addition of the lagged differences ensures that $\in_t$ series are not serially correlated. It is recommended to regress the dependent variable on a sufficiently large number of lags (Bamba and Reed, 2004, Engel and Granger, 1987, Engel and Yoo, 1987). Test statistics (ADF) is obtained from the t-ratio for the estimates of ϕ. Engel and Granger recommends ADF test, because, it has the same critical value for both finite sample experiments, and has a good performance properties.

(iv) RVAR: This is a two-step estimator procedure. Co-integrating vector from the co-integrating regression is first estimated then; error correction coefficients are estimated based on the estimated vectors. A test of co-integration is based on the joint significance of the error correction coefficients β_1 and β_2. If they are jointly different from zero, the null hypothesis of no co-integration is rejected.

$$\Delta y_t = \beta_1 \hat{e}_{t-1} + \in_{1t}, \tag{A6}$$

and,

$$\Delta x_t = \beta_2 \hat{e}_{t-1} + \gamma \Delta y_t + \in_{2t} \tag{A7}$$

The test is based on the sum of the squared t statistics of the coefficients β_1 and β_2. By making the system triangular, the disturbances are uncorrelated, and under normality the t statistics are independent (Engle and Granger, 1987).

(v) The ARVAR: The ARVAR is similar to RVAR save for added lagged values of the differences of the variables Δy_t and Δx_t, to ensure stationarity in the error terms of the vector autoregressive system. The equations are specified as,

$$\Delta y_{t-p} = \beta_1 \hat{e}_{t-1} + \in_{1t} \tag{A8}$$

and,

$$\Delta x_{t-p} = \beta_2 e_{t-1} + \gamma \Delta y_{t-p} + \in_{2t} \tag{A9}$$

The test statistics is same as RVAR.

(vi) UVAR: This test is based on a vector autoregression in the levels which is not restricted to satisfy the co-integration constraints (Engle and Granger, 1987).

$$\Delta y_t = \theta_1 y_{t-1} + \theta_2 x_{t-1} + c_1 + \in_{1t} \tag{A10}$$

and,

$$\Delta x_t = \theta_3 y_{t-1} + \theta_4 x_{t-1} + \gamma \Delta y_t + c_2 + \in_{2t} \tag{A11}$$

The null hypothesis of no co-integration is rejected if the parameters θ_1, θ_2, θ_3, and θ_4 are jointly significant from zero. If the null is not rejected, no significant relationship exists between current and past values of the variables and thus, no co-integration. A test statistics from the two equations is constructed by taking the sum of the F-test multiplied by the degree of freedom, 2 (Goodwin and Schroeder, 1991; Engel and Granger, 1987).

(vii). AUVAR: This is a higher order version of UVAR with additional lags of Δy_t and Δx_t;

$$\Delta y_{t-p} = \theta_1 y_{t-1} + \theta_2 x_{t-1} + c_1 + \in_{1t} \tag{A12}$$

And $\Delta x_{t-p} = \theta_3 y_{t-1} + \theta_4 x_{t-1} + c_2 + \in_{2t}$ (A13)

APPENDIX B

Conceptual Framework of Parity Bound Model

If goods move freely between markets and are competitive, then price differential across spatially dispersed markets for a homogeneous good should not exceed transaction cost. This is the spatial arbitrage conditions. The parity bound model test for market integration to determine whether these assumptions are violated or if market integration is detected, the extent to which the integration occurs. This is carried out by considering two markets located in different region (i and j) that may engage in trade for a given homogeneous commodity (A). When there is market equilibrium (i.e. supply equals demand) no trade will occur and the prices in the two the two markets are represented as, autarky prices,

Autarky prices:

$$P_{it}^{A} = \alpha_{it} + \pi_{it} \tag{B1}$$

$$P_{ij}^{A} = \alpha_{ij} + \pi_{ij} \tag{B2}$$

Where,

α_{it}, α_{ij} are time varying mean autarky prices and π_{it}, π_{ij} are stochastic disturbances affecting the autarky prices. The transaction cost incurred in trade between the two markets at time t is specified as a random variable as follows.

$$TC_{jit} = \gamma_{jit} + \varepsilon_{jit} \tag{B3}$$

Where,

γ_{jit} is the time varying mean transaction cost, ε_{jit} is the random component and is normally distributed with mean zero and constant variance.

The trade regimes:

Given the formulations of autarky prices and transaction cost three mutually exclusive and exhaustive spatial arbitrage conditions or trade regimes could be identified based on the contemporaneous spatial price differential and transaction cost (Negassa *et al*., 2003).

Regime 1

The spatial price differential is equal to transaction cost, whether trade occurred or not.

$$P_{it} - P_{jt} = TC_{jit} \tag{B4}$$

Where P_{it}, P_{jt} are the contemporaneous prices in the i^{th} and j^{th} markets respectively. This is a condition for a spatial efficient market with or without trade occurring. With trade, the two prices P_{it}, P_{jt} may differ from the autarky prices, in which case price movements are related due to changes in supply and demand or stochastic disturbance terms.

Regime 2:

Spatial price differential is less than transaction cost.

$$P_{it} - P_{jt} < TC_{jit} \tag{B5}$$

There are no profitable arbitrage opportunities and no incentives to engage in trade since any trade will result in a loss. Autarky price and the actual price are equal, therefore the two prices are independent and no shocks will be transmitted across the market. In this case prices could be efficient but the two markets may not be integrated.

Regime 3: In this regime spatial price differential is greater than transaction cost. Trade may or may not occur.

$$P_{it}^{\;A} - P_{ji}^{\;A} > TC_{jit} \tag{B6}$$

The spatial arbitrage condition is violated. Market are not efficient but may be integrated to some extent if trade occur. The inefficiencies result in some profitable arbitrage opportunities not being exploited (Baulch, 1997; Negassa *et al.,* 2003, Penzhorn and Arndt, 2002). Tomek and Robinson (1990), Baulch (1997), classified these inefficiencies as. (i) Existence of transportation bottleneck. (ii). Non-competitive pricing practices. (iii)Government controls on product flows between regions. (iv) Government price support systems. (v) Licensing requirements. (vi) Government Quotas.

Empirical specification of Parity bound model

The extent of market integration can be assessed by distinguishing the three trade regimes (Baulch, 1997). Regime 1, occur at parity bound, (in which price differentials equal transaction cost.) Regime 2, occur inside the parity bound, (in which price differentials are less than transaction cost.) Regime 3, occur outside the parity bound, (in which price differentials exceed transaction cost.) The parity bound model seeks to determine the probability that an observation will fall into one of these three regimes specified above. There are no arbitrage opportunities in region 2, therefore; only two parity bounds exist for spatial arbitrage conditions between the markets i and j, (i.e. upper and lower parity bounds). Assuming the spatial prices and transaction cost are stochastic (as specified earlier above), and the transaction cost between the markets are not dependent on the direction of trade flows. The deviations of the inter-market price spread from the extrapolated transaction cost in any time period can be decomposed into three components according to the three regimes as follows;

$$|P_{it} - P_{jt}| - \gamma_{jit} = \varepsilon_{jit} \quad \text{(B7)}$$

$$|P_{it} - P_{jt}| - \gamma_{jit} = \varepsilon_{jit} - \mu_{jit} \quad \text{(B8)}$$

$$|P_{it} - P_{jt}| - \gamma_{jit} = \varepsilon_{jit} + \nu_{jit} \quad \text{(B9)}$$

Where the error terms ε_{jit}, μ_{jit}, and ν_{jit}, are assumed to be normal(ε_{jit}), half-normal(μ_{jit}) and half-normal(ν_{jit}), independently distributed random variables with means equal to zero and standard deviations equal to σ_{ε}, σ_{μ}, and σ_{ν}, respectively. The error term ε_{jit} allows transaction cost to vary between periods, in response to seasonality or transportation utilization (Penzhorn and Arndt, 2002; Baulch, 1997). The μ_{jit} and ν_{jit} are composite error terms of the disturbance terms in the demand and supply functions for the pairs of markets i and j under consideration. Their magnitude depends on the imbalances in demand and supply in each market. While μ_{jit} captures the extent to which price differences fall short of the parity bound in regime 2 (where there is no incentive to trade), ν_{jit} measures the degree price differential exceeds transaction when spatial conditions are violated (Baulch, 1997, Negass *et al.,* 2003; Penzhorn and Arndt, 2002).

According to Baulch (1997) and Penzhorn and Arndt (2002), utilizing a result derived by Weistein (1964) for the density of a normal plus half normal distribution, and following Sexton, Kling and Carman (1991), the likelihood function for the parity bound model may be specified as;

$$L = \prod_{t=1}^{T}\left[\lambda_1 f_t^1 + \lambda_2 f_t^2 + (1-\lambda_1-\lambda_2) f_t^3\right] \qquad \text{(B10)}$$

Where regime 1(at the parity bound) is specified as;

$$f_t^1 = \frac{1}{\sigma_\varepsilon}\Theta\left[\frac{|P_{it}-P_{jt}|-\gamma_{jit}}{\sigma_\varepsilon}\right], \qquad \text{(B11)}$$

Regime 2 (inside the parity bound) is specified as;

$$f_t^2 = \left[\frac{2}{\left(\sigma_\varepsilon^2+\sigma_\mu^2\right)^{1/2}}\right]\Theta\left[\frac{|P_{it}-P_{jt}|-\gamma_{jit}}{\left(\sigma_\varepsilon^2+\sigma_\mu^2\right)^{1/2}}\right] * \left\{1-\Phi\left[\frac{\left(|P_{it}-P_{jt}|-\gamma_{jit}\right)^{\sigma_\mu}/\sigma_\varepsilon}{\left(\sigma_\varepsilon^2+\sigma_\mu^2\right)^{1/2}}\right]\right\}, \qquad \text{(B12)}$$

Regime 3 is specified as;

$$f_t^3 = \left[\frac{2}{\left(\sigma_\varepsilon^2+\sigma_\nu^2\right)^{1/2}}\right]\Theta\left[\frac{|P_{it}-P_{jt}|-\gamma_{jit}}{\left(\sigma_\varepsilon^2+\sigma_\nu^2\right)^{1/2}}\right] * \left\{1-\Phi\left[\frac{\left(|P_{it}-P_{jt}|-\gamma_{jit}\right)\sigma_\nu/\sigma_\varepsilon}{\left(\sigma_\varepsilon^2+\sigma_\nu^2\right)^{1/2}}\right]\right\}, \qquad \text{(B13)}$$

Where, from equation (B10), λ_1 and λ_2 denotes the probabilities for regimes 1 and 2. The absolute value of the natural logarithm of the price spread between the markets i and j in periods t (i.e., $In\{|P_{it}-P_{jt}|\}$) is utilized. σ_ε, σ_μ and σ_ν are the standard deviations of the three error terms ε_t, μ_t, and ν_t. γ_{jit}, represents the logarithm of mean transaction cost in time period t. While $\Theta(.)$ and $\Phi(.)$ denotes the standard normal density and distribution functions. The probability estimate of the three regimes and the standard errors of various error component of the parity bound model can be obtained by maximizing the likelihood function using numerical optimization with respect to λ_1, λ_2, σ_ε, σ_μ and σ_ν The relative sizes of the three variances terms and their standard errors will indicate the volatility of the price spread in the three regimes. Generally, inferential statistical hypothesis test can be conducted to test the null hypothesis that $\lambda_1 + \lambda_2 = 1$. Baulch, (1997) and Penzhorn and Arndt (2002) suggests it is more logical to focus on the frequency of violation of spatial arbitrage conditions which may be interpreted as an index of market efficiency.

APPENDIX C

Average volume of apple sold and real average market prices in Eight selected South African apple markets.

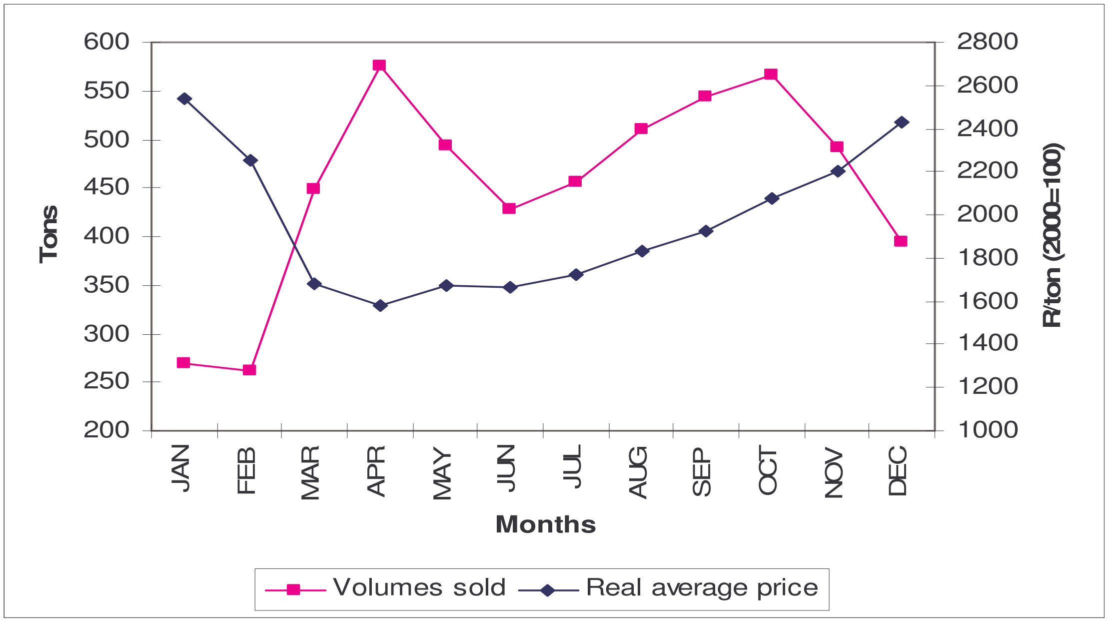

Figure C1: **Average volume sold and real average prices in Bloemfontein market.**
Source: NDA (2004) and authors computation

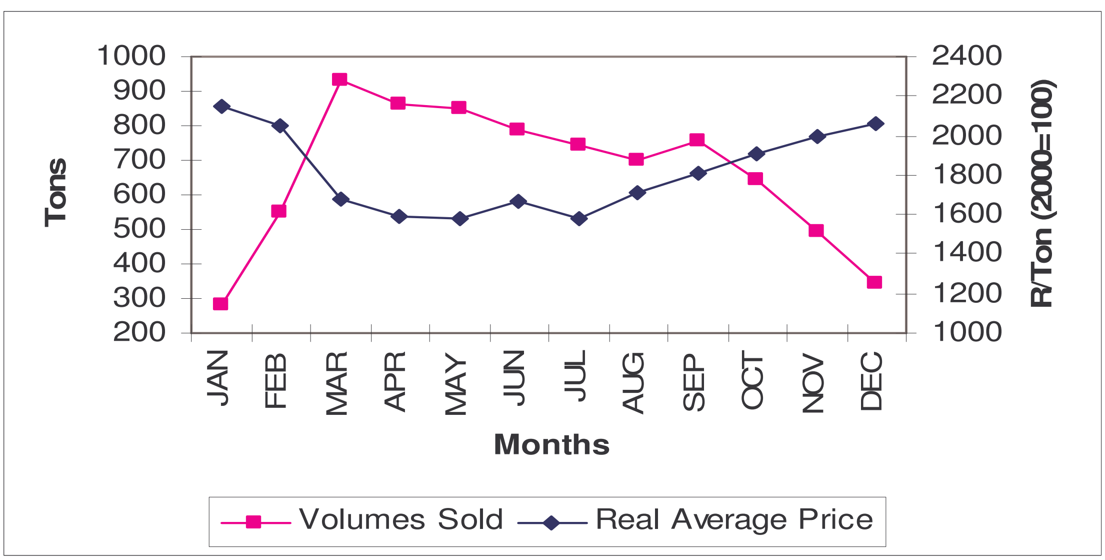

Figure C2: **Average volume sold and real average prices in Cape Town market.**
Source: NDA (2004) and authors computation

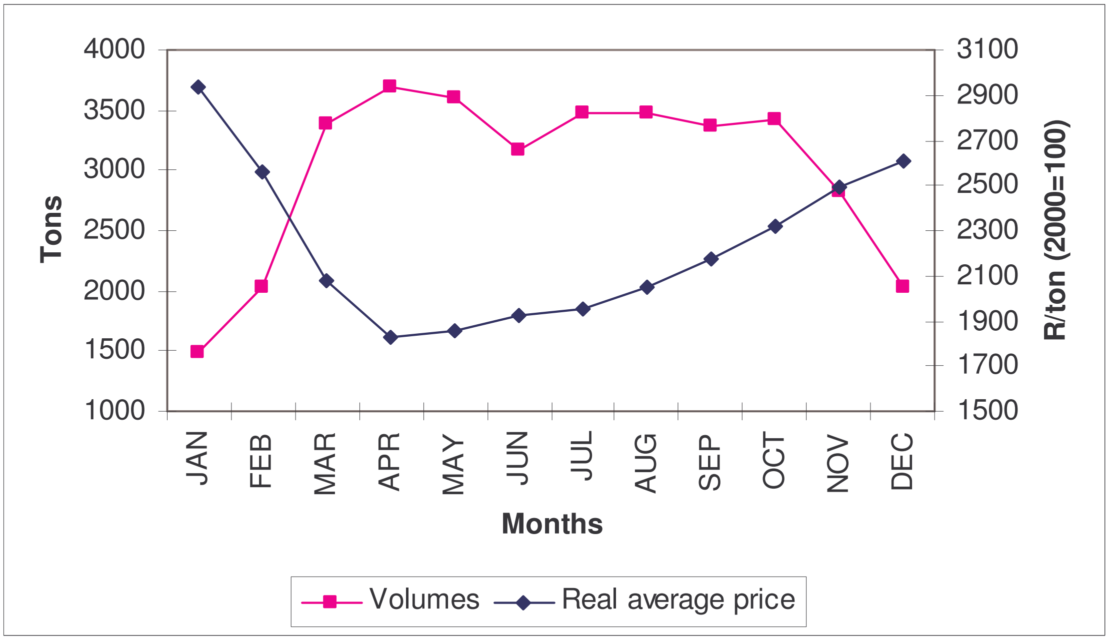

Figure C3: **Average volume sold and real average prices in Johannesburg market**
Source: NDA (2004) and authors computation

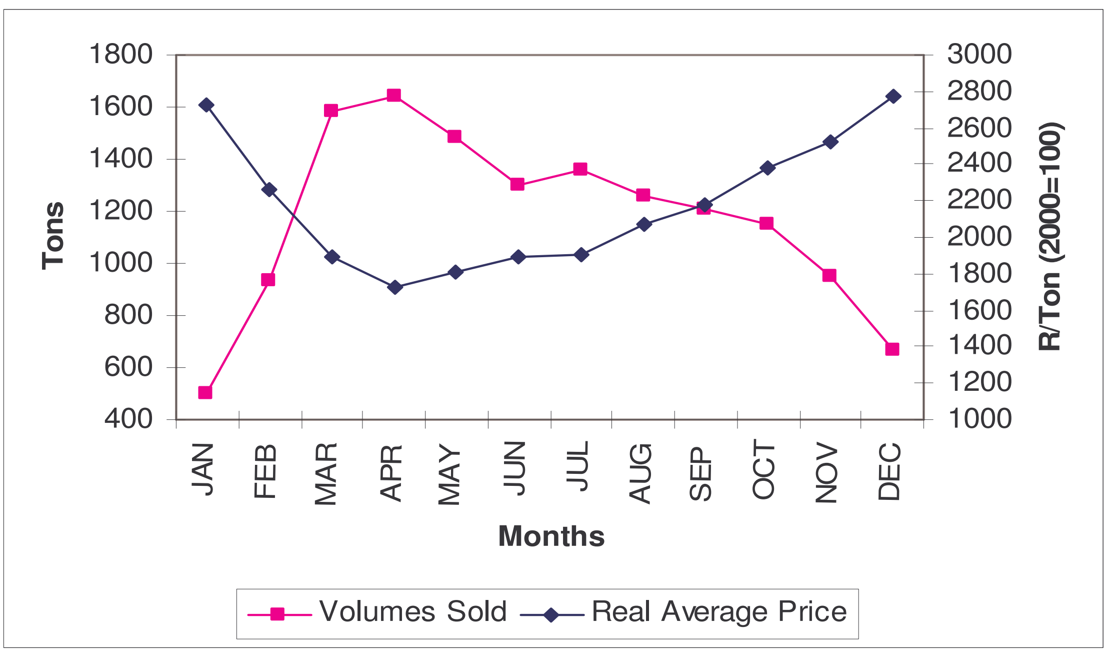

Figure C4: **Average volume sold and real average prices in Durban market**
Source: NDA (2004) and authors computation

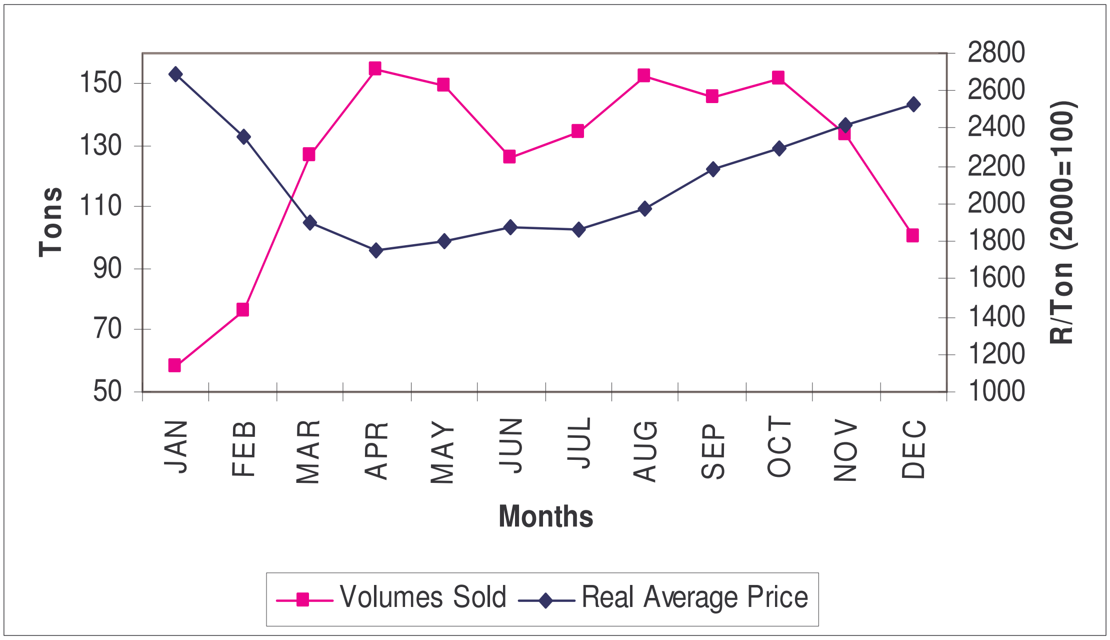

Figure C5: **Average volume sold and real average prices in Kimberley market**
Source: NDA (2004) and authors computation

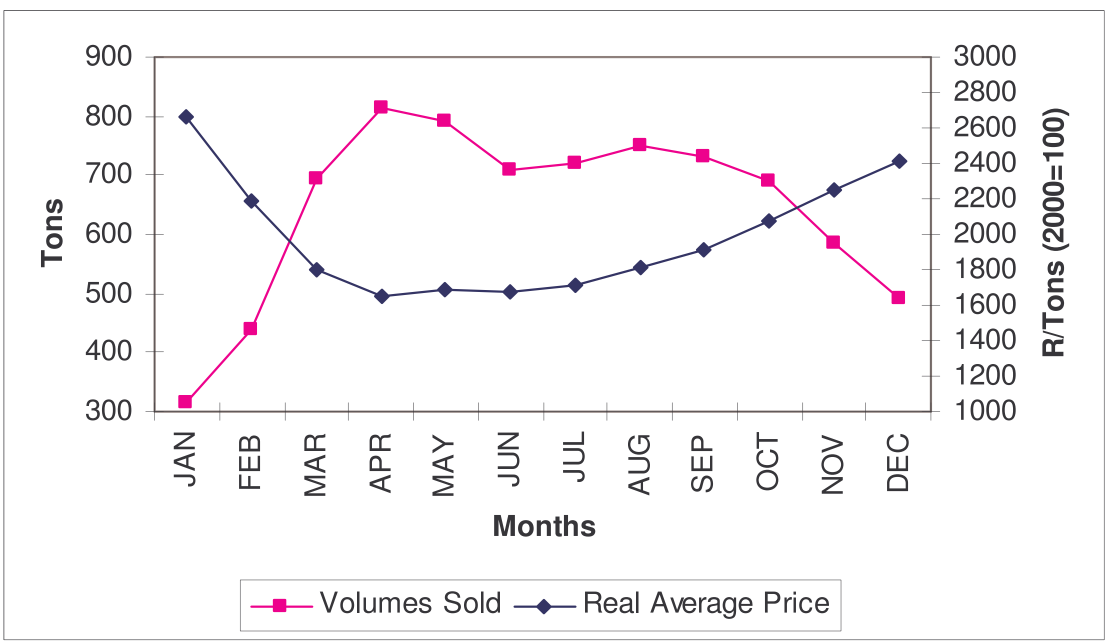

Figure C6: **Average volume sold and real average prices in Pietermaritzburg market**
Source: NDA (2004) and authors computation

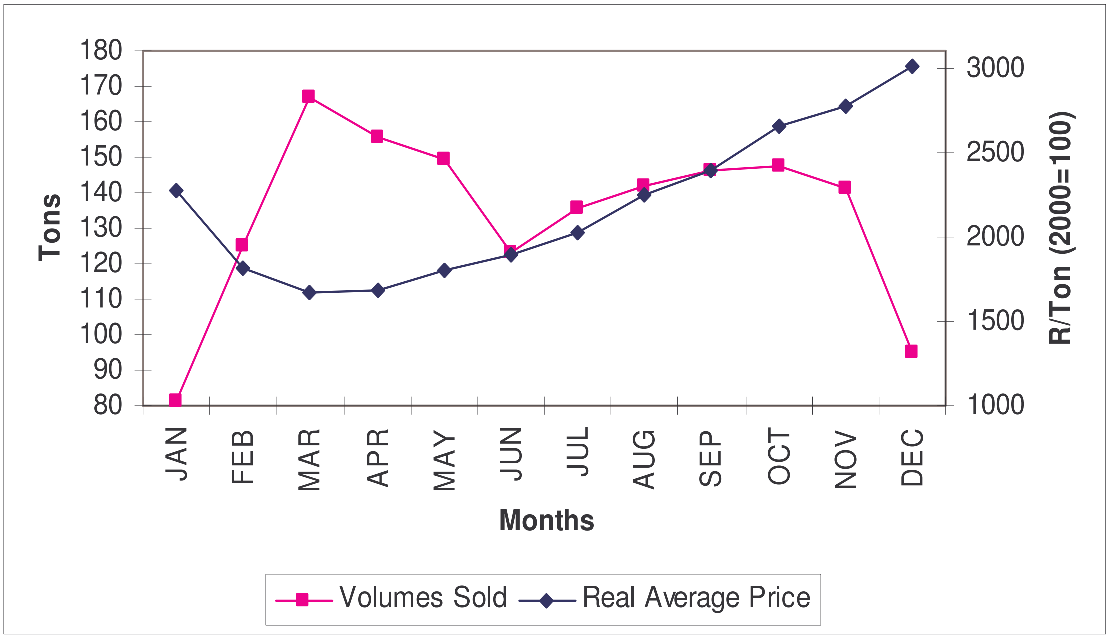

Figure C7: **Average volume sold and real average prices in Port Elizabeth market**
Source: NDA (2004) and authors computation

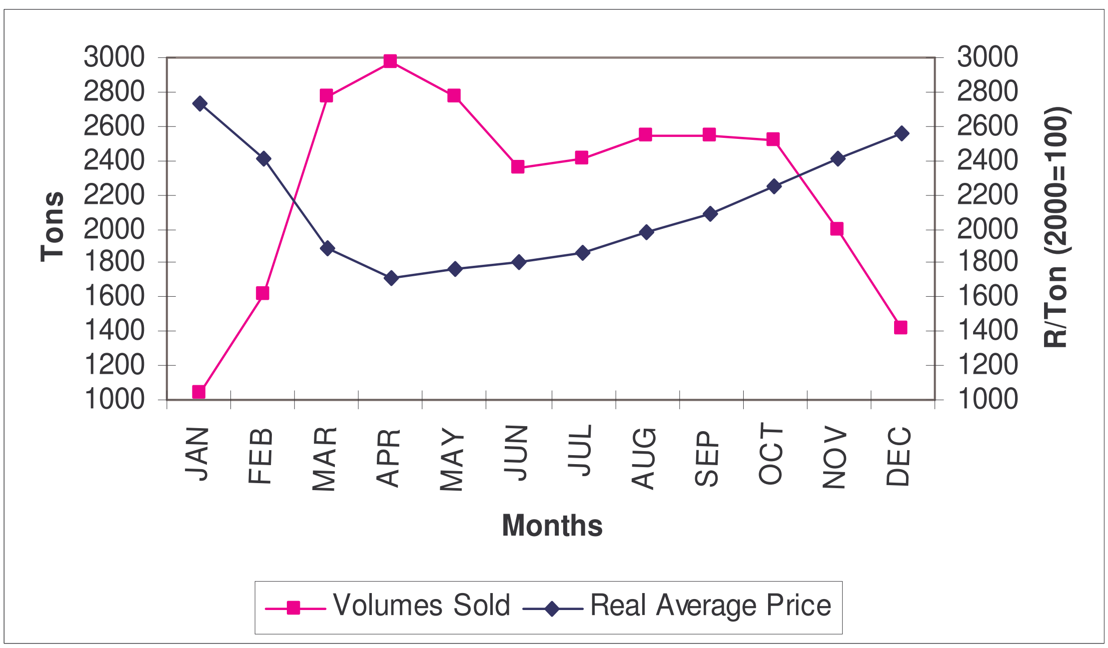

Figure C8: **Average volume sold and real average prices in Tshwane market**
Source: NDA (2004) and authors computation

APPENDIX D

Persistency of price differences occurring in each Regime and Regime switching between Johannesburg and other Fresh Produce Markets in South Africa

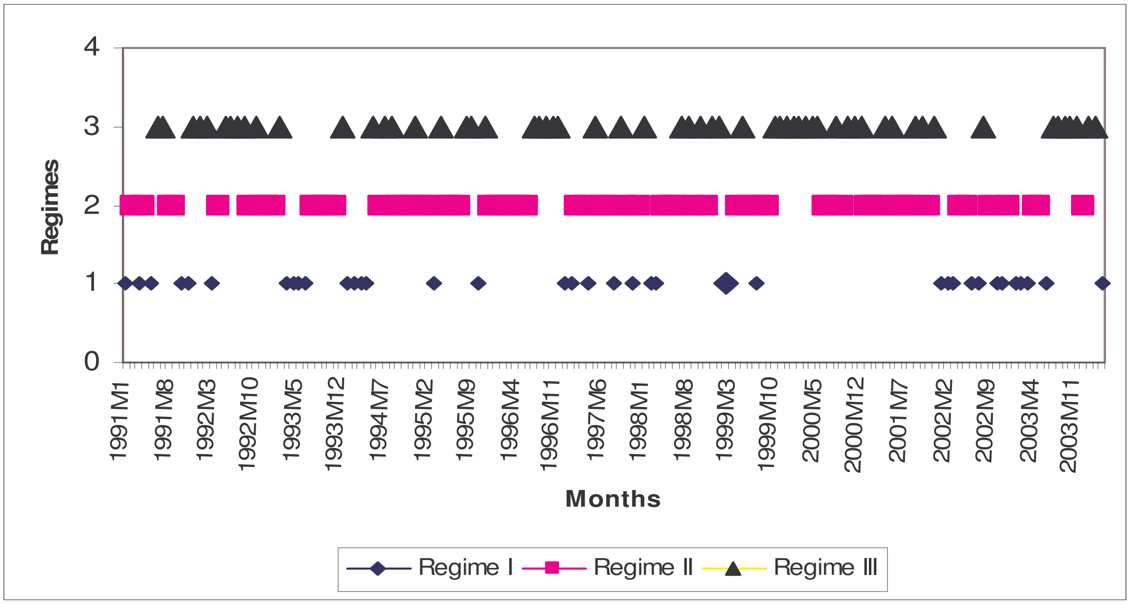

Figure D1: **Persistency of price differences occurring in each Regime and Regime switching: Tshwane and Johannesburg FPMs**

Source: NDA (2004) and authors computation

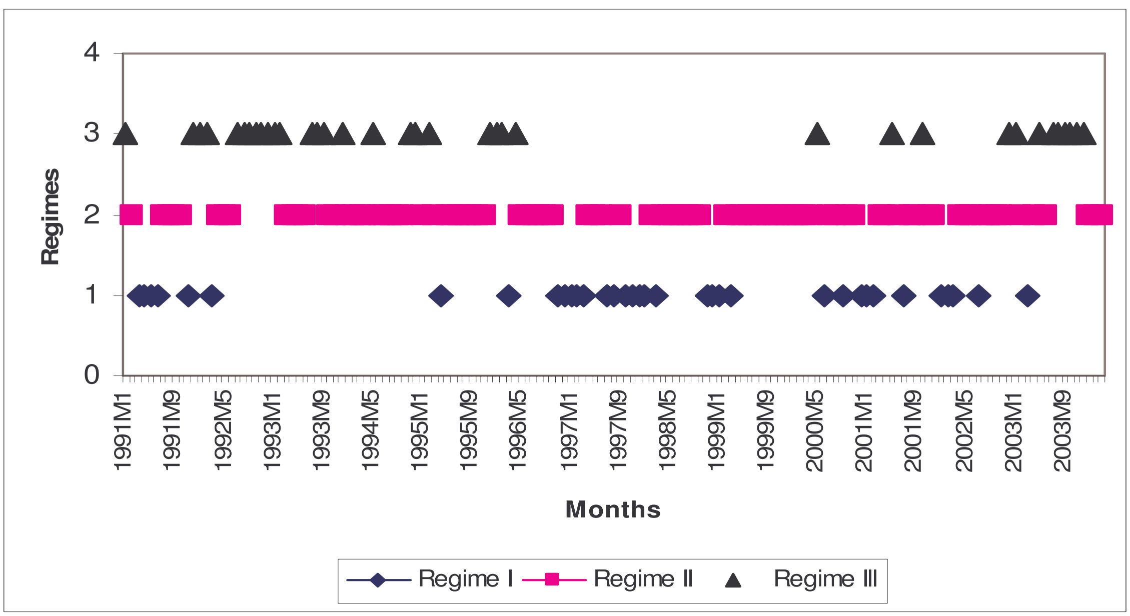

Figure D2: **Persistency of price differences occurring in each Regime and Regime switching: Bloemfontein and Johannesburg FPMs**

Source: NDA (2004) and authors computation

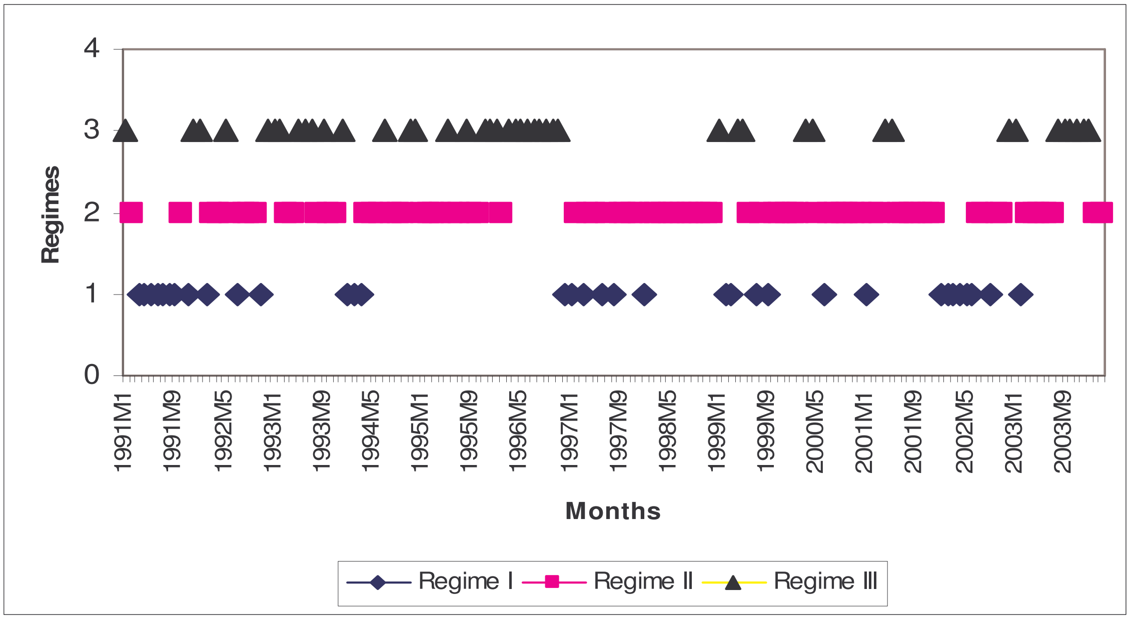

Figure D3: **Persistency of price differences occurring in each Regime and Regime switching: Pietermaritzburg and Johannesburg FPMs**

Source: NDA (2004) and authors computation

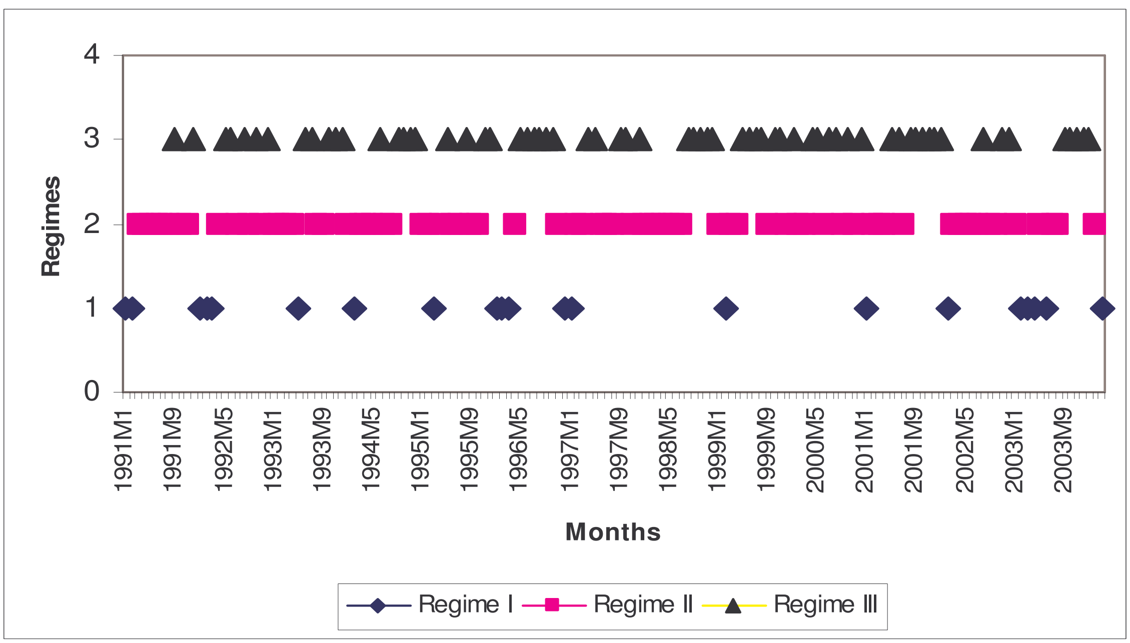

Figure D4: **Persistency of price differences occurring in each Regime and Regime switching: Durban and Johannesburg FPMs**

Source: NDA (2004) and authors computation

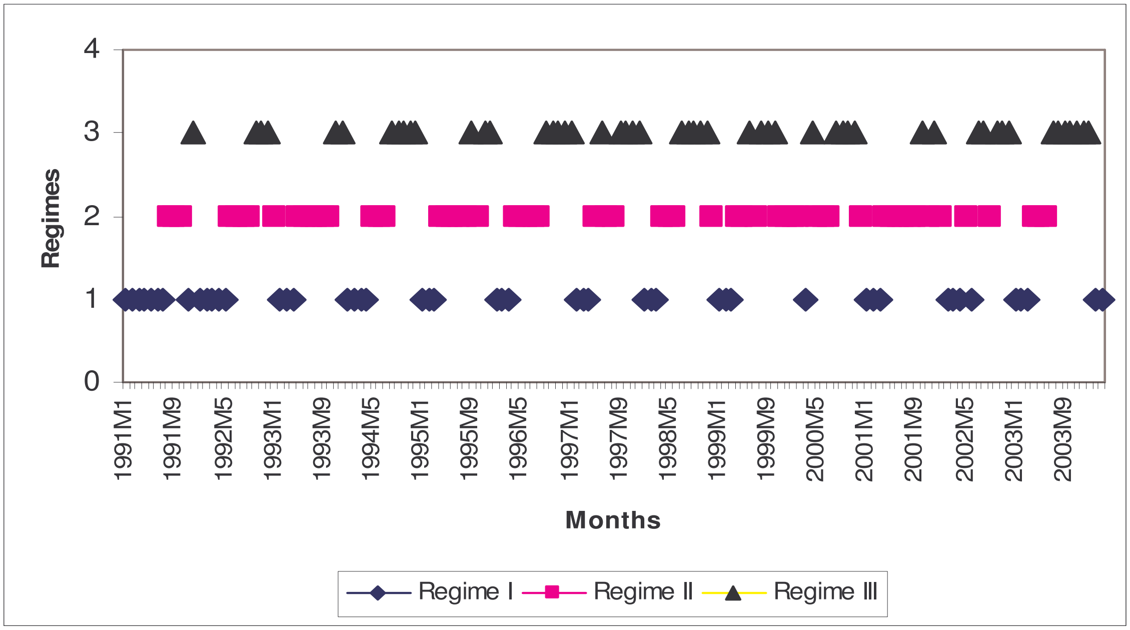

Figure D5: **Persistency of price differences occurring in each Regime and Regime switching: Port Elizabeth and Johannesburg FPMs**

Source: NDA (2004) and authors computation

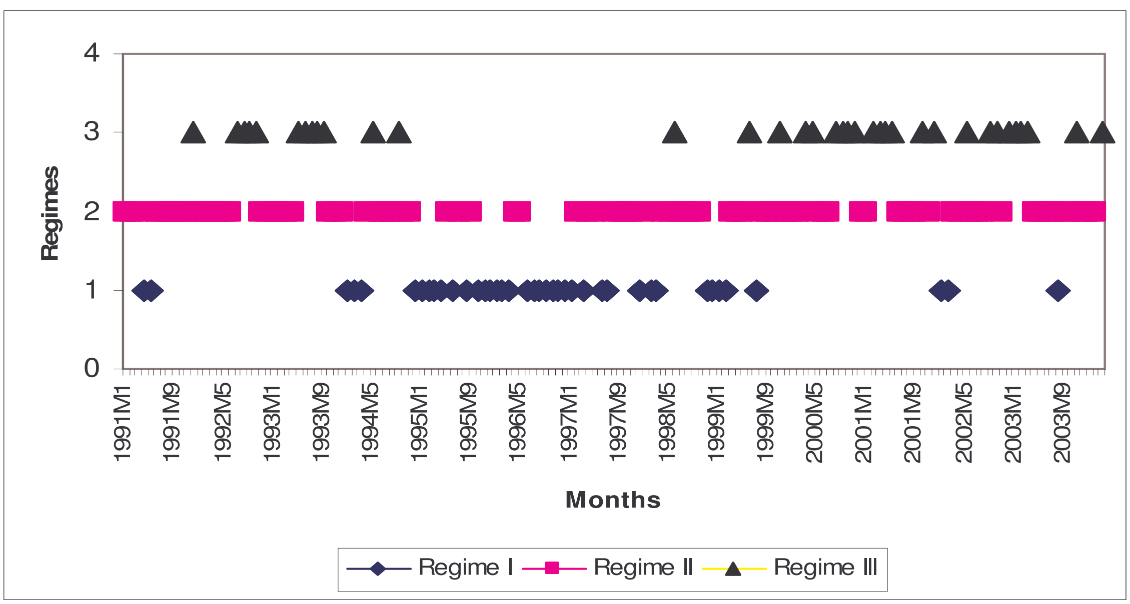

Figure D6: **Persistency of price differences occurring in each Regime and Regime switching: Kimberley and Johannesburg FPMs**

Source: NDA (2004) and authors computation

APPENDIX E

Impulse Response Function: Response of selected Fresh produce markets to one-half positive and negative price shock from Johannesburg fresh produce market.

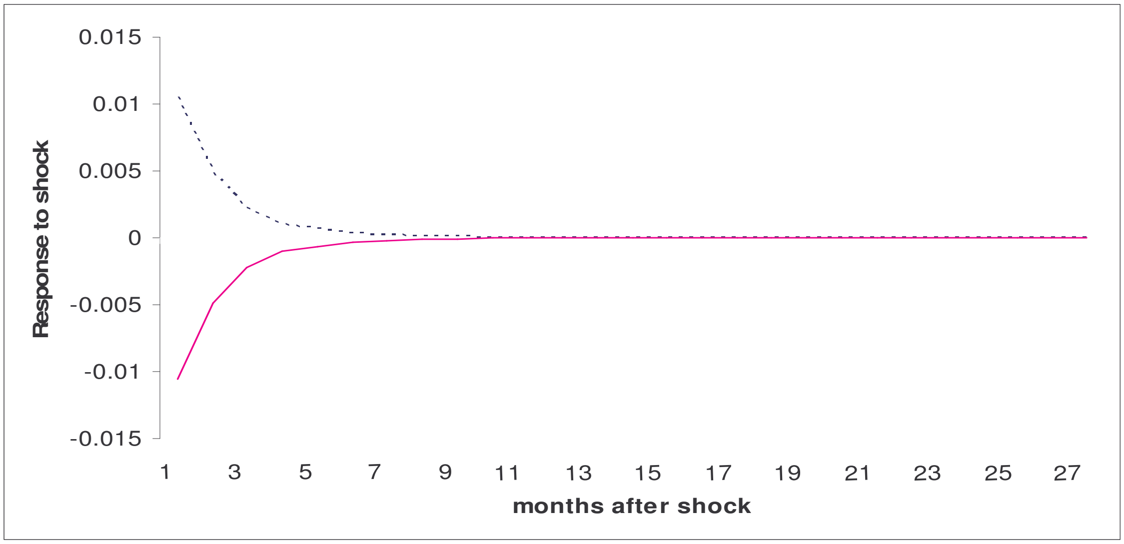

Figure E1: **Response of the Tshwane FPM to positive and negative price shocks in the Johannesburg FPM**

Source: NDA (2004) and authors computation

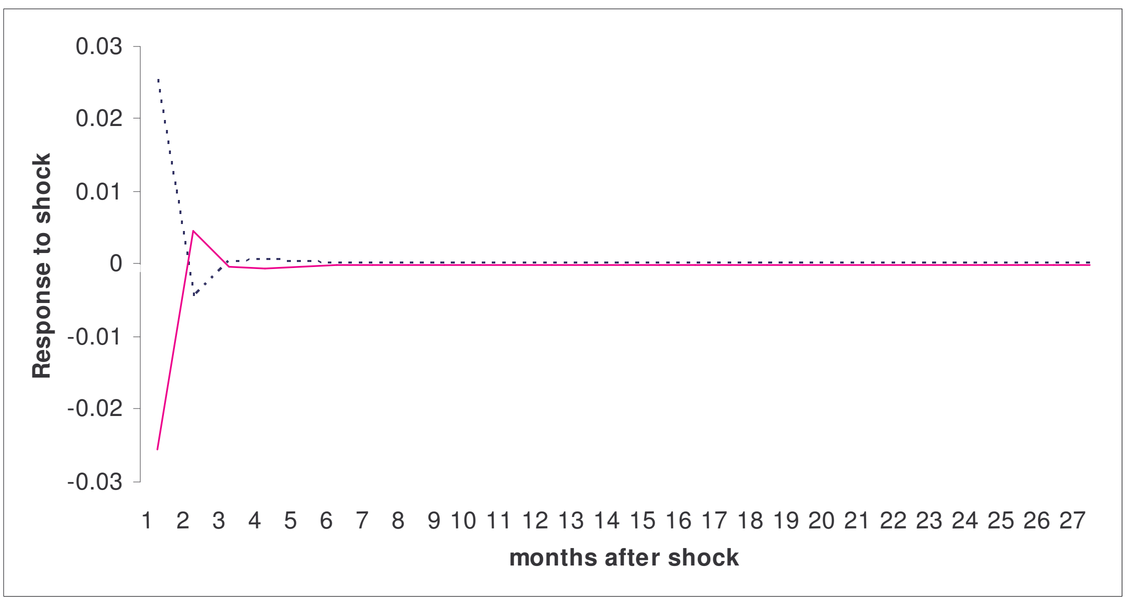

Figure E2: **Response of Bloemfontein FPM to positive and negative price shocks in the Johannesburg FPM**

Source: NDA (2004) and authors computation

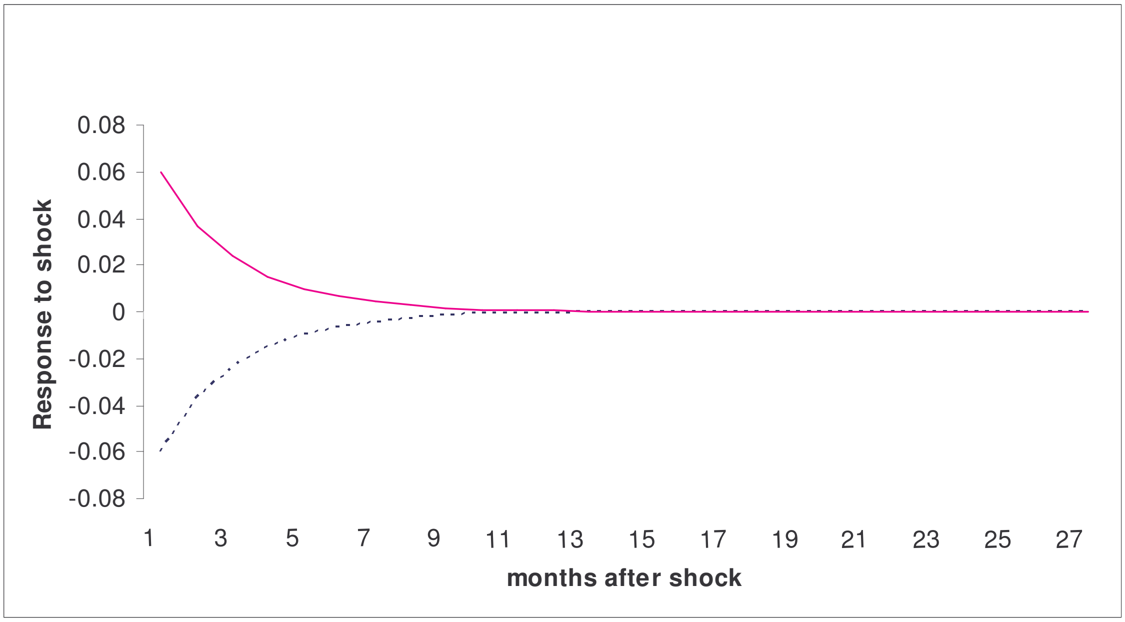

Figure E3: Response of Durban FPM to positive and negative price shocks in the Johannesburg FPM

Source: NDA (2004) and authors computation

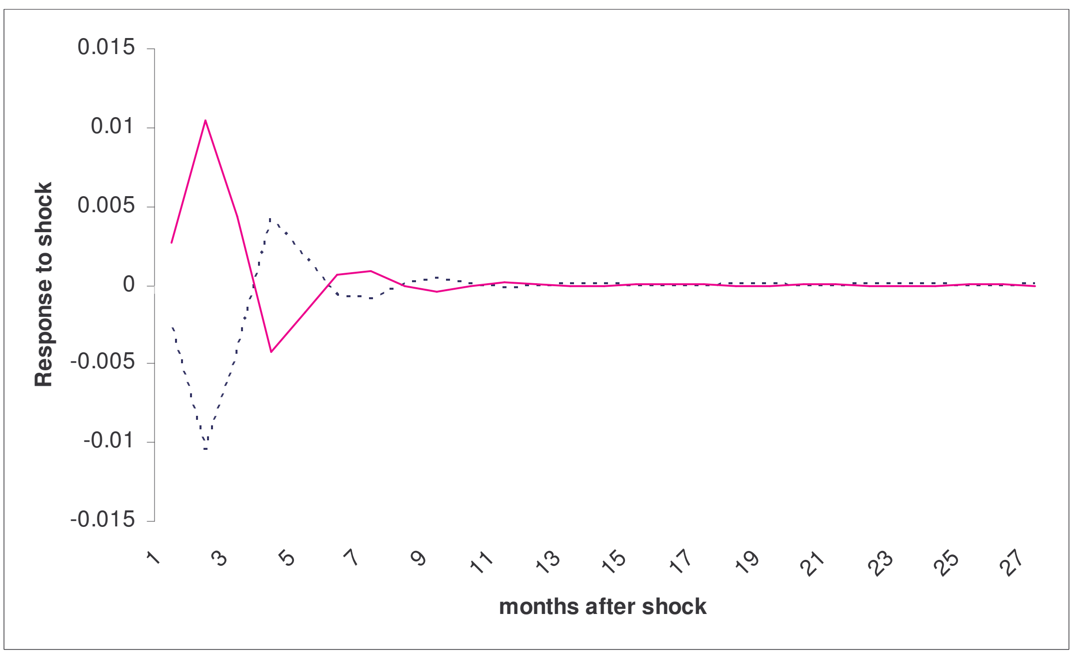

Figure E4: Response of Port Elizabeth FPM to positive and negative price shocks in the Johannesburg FPM

Source: NDA (2004) and authors computation

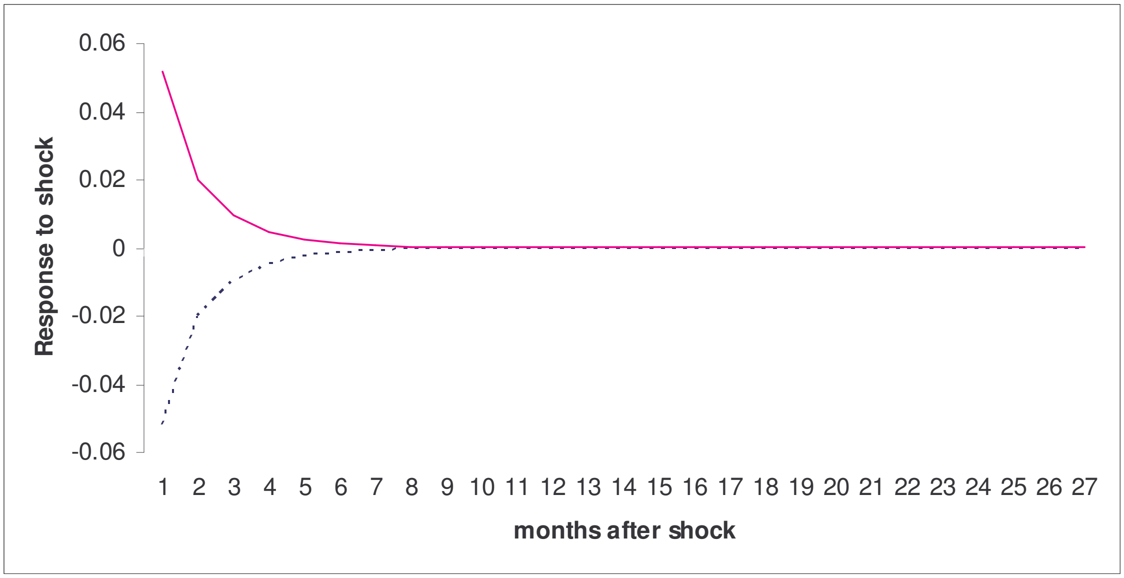

Figure E5: Response of Kimberley FPM to positive and negative price shocks in the Johannesburg FPM

Source: NDA (2004) and authors computation

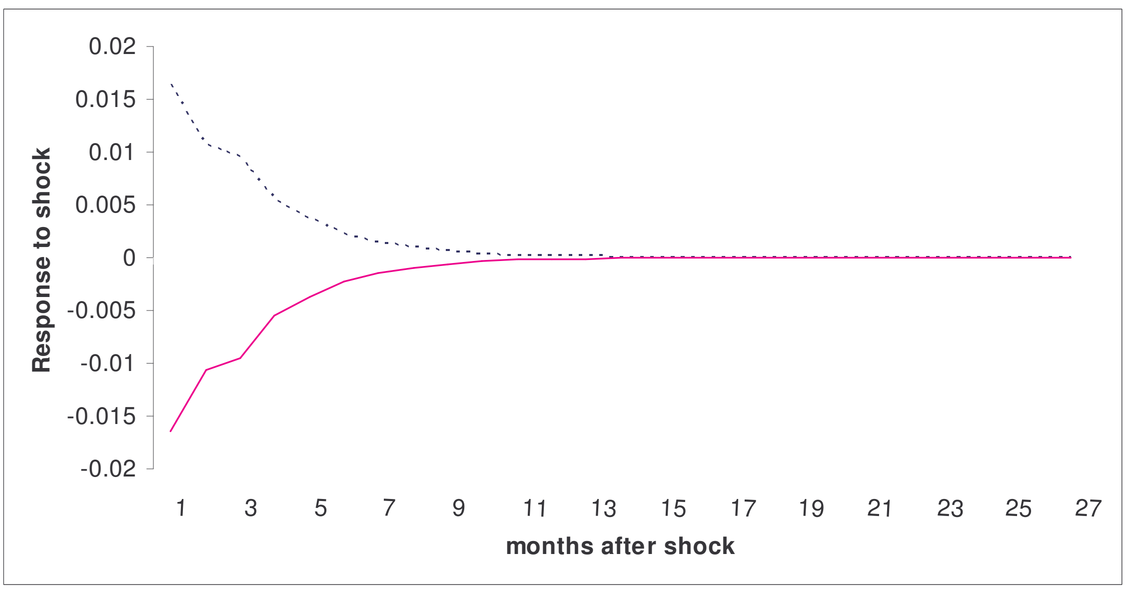

Figure E6: Response of Pietermaritzburg FPM to positive and negative price shocks in the Johannesburg FPM

Source: NDA (2004) and authors computation