THE EFFECT OF EXCHANGE RATE ON INDONESIAN AGRICULTURAL EXPORTS

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CHAPTER I

INTRODUCTION

1.1 Problem Statement

Since the breakdown of the Bretton Wood system in 1973, exchange rates have been allowed to float freely or almost freely. This has led to a substantial increase in the variation of currency values with respect to each other. Researchers have attempted to determine the effects of floating exchange rates on volume of trade. The hypothesis is that uncertainty of exchange rates could have effects on trade flows. Uncertainty of exchange rates is hypothesized to cause firms to increase prices, which will reduce export volumes. The other view is that firms view the opportunity to trade internationally as an option whose value rises with increases in volatility. In this view, international trade is increased when exchange rates are volatile.

Many analysts of international economics concur that the generalized floating of exchange rates in operation now for at least 15 years has increased volatility in both the nominal and real exchange rates for developed and developing countries. What remains controversial, however, is whether increased variability of the exchange rate, being indicative of the greater risk and uncertainty in international transactions, has had a negative effect on the volume and value of international trade flows. Has exchange rate variability contributed to the slowdown in the growth of international trade observed during the 1980s.

In recent years a significant volume of research has taken place in order to empirically evaluate the determinants of export demand in developing countries. The literature can generally be divided into papers that use conventional estimation procedures and those that recognize the non-stationary nature of real exports and its determinants. Studies which can be grouped into the former category include Kenen and Rodrik (1986), Pozo (1992), and Grobar (1993), while those included in the latter include Koray and Lastrapes (1990), Khumar and Dhawan (1991),Chowdury (1993), Arize (1995), Sukar (1998), Hassan and Tufte (1998), and Fountas and Bredin (1998) which use cointegration and error correction models (ECM) to estimate a long-run and short-run export demand function.

There is no consensus, theoretical or empirical, on the impact of exchange rate volatility on international trade (Hassan and Tufte, 1998). Hooper and Kholhagen (1978) derive and estimate a reduced form equation for trade between two countries starting from the assumption that utility is positively or negatively related to profits, and negatively related to variance of the profits. Similar equations have been estimated since then for different time periods, for bilateral and aggregate trade flows for different countries, for different measures of exchange rate on and for real and nominal variables.

On the other hand, a good deal of attention has been directed at the effect of currency depreciation on the agricultural sector of the economy. If currency depreciation strongly affects the agricultural sector, then the majority of the associated impact must come through agricultural exports. Schuh (1974) has demonstrated that overvaluation;

followed by devaluation could have important effects on the foreign component of the agricultural sector. This latter point is particularly relevant because many of the empirical studies of the effect of devaluation on U.S. agriculture have concentrated on price and export effects.

The question of the impact of exchange rate on the volume of international trade has been studied extensively since the late 1970s when the move to the floating exchange rates occurred. Whether exchange rate has an effect on real economic activity, such as trade, exchange rate management continues to be an important question in many countries since it has implications for the choice of exchange rate system and the conduct of exchange rate. policies. Economic theory has little to say to help understand the change of exchange rate. It seems excessive in the sense that existing models are not capable of generating the observed the change on exchange rate policy (Hassan and Tufte, 1998).

Nainggolan (1979) said that performance of the agriculture sector is affected by macroeconomic policies through its effect on inflation, real exchange rates and incentives to export and import. He has argued that informal sectors in general, and agriculture in particular, have been held back in many developing countries by policies that have contributed to capital market fragmentation, by inflation, administered interest rates, and exchange rate overvaluation. Timmer and Falcon (1983), argue in a similar fashion for greater focus on "macroprices", i.e., the inflation rate, interest rates, wage rate, the exchange rate, and the intersectoral terms of trade.

Studies of developing countries are of potential importance, in that much of their real exchange rate uncertainty stems from macroeconomics policies. In particular, many developing countries experience high variable rates of inflation as they expand domestic credit to finance fiscal deficits or to increase lending to the private sector.

This study improves on previous studies that have estimated Indonesian export demand functions in several ways. First, it recognizes that exports and their determinants are potentially non-stationary variables. Second, in contrast to all other studies, it includes a measure of exchange rate change to investigate the affect of such movement on exports.

Although several studies examine the effects of exchange rate on the levels of trade, there is no consensus of opinion on the consistency of the relationship. Indeed, some studies find that exchange rate change has positive effects, some find no effect and some find negative effects on volume of trade. Fewer studies examine agricultural commodities and the effects of exchange rate. And fewer yet look at disaggregated trade flows in agricultural in order to determine the effects of exchange rate on individual commodity trade flows. In fact, no published studies exist in the current agricultural economics literature that attempt to apply the same technique to several different commodities to estimate the extent that exchange rate affects an individual country's exports of individual commodities over an extended period of time.

For Indonesia, foreign exchange rates have been highly volatile since the abandonment of fixed exchange rates in November 1978. It is widely believed that increased uncertainty in exchange rates inhibits the growth of foreign trade. Ever since the floating exchange rate system has been adopted in Indonesia, there has been a substantial increase in nominal and real exchange rate, with little negative effect on international trade.

Indonesia has experienced both fixed and managed floating exchange rate systems. From July 1971 until November 1978, the value of the rupiah was pegged to the U.S dollar in a fixed exchange rate system. The system moved to a tightly managed float in November 1978, and to a more flexible but still managed float in March 1983.The current system, introduced in July 1997 is a floating exchange rate regime. The adoption of a floating regime in 1997 and the effects of exchange rate on the volume of Indonesia exports have been the subjects of both theoretical and empirical investigations. Much of the interest in this issue was stimulated by the increased uncertainty in exchange rates and by the potential for exchange rate to have lasting consequences on the volume of international trade. Nonetheless, no real consensus about the effects of exchange rate on Indonesia exports has emerged. Conventional theories by Hooper and Kohlhagen (1978) of the effect on trade volume imply that increases in uncertainty exchange rate will reduce the volume of trade. Indonesia has attempted to increase foreign earnings through increased exports, especially agricultural products.

One of the most dramatic events in the international arena over the last three years was the depreciation of Indonesia rupiah by almost 350 percent. The depreciation was largely a result of decreasing purchasing power parity for Indonesia. Dramatic as it was, however, the initial depreciation of the rupiah was followed by yet another depreciation. During this same period, Indonesia experienced deficits in the trade account balance. These facts, with the still weak position of Indonesian the rupiah in international currency markets, have led many trade theoreticians to question the overall effectiveness of exchange rate depreciation as a policy tool. In fact, there is a school of thought which suggests that depreciation can have only monetary effects, in which case depreciation likely causes a portfolio adjustment but is unlikely to affect seriously the trade balance (Laffer, 1976).

The motivation for estimating the Indonesian export demand function derives from the recent change in exchange rate. Most empirical studies of the determinants of Indonesian exports have used traditional estimation techniques ((e.g Timmer (1986) and Nainggolan (1987)), and have not considered the integration properties of the time series involved in the analysis. The general conclusion of the above papers is that the foreign exchange rate is the most important macro price affect. A lower real exchange rate decreases the costs to foreign consumers of the Indonesian products, improving agricultural exporters' competitive position.

Prior to 1987, oil and natural gas dominated Indonesia's export economy. Oil revenues have not been used to diversify the economy, and in fact, may have a detrimental effect on the economy by contributing to inflation, inadvisable government spending, conspicuous consumption and a general false sense of economic security (Vinick, 1991). Moreover, in 1987, the values of non-oil exports exceeded oil and gas exports for the first time. Growth in the non-oil sector has exceeded 15 percent annually since 1987, and in 1989 accounted for 62 percent of total exports, with oil and gas exports accounting for the remaining 38 percent. Table.1.1 and shows Indonesia's export value of oil and non-oil.

Agricultural exports remain the important, having averaged 9 percent of export value from 1994-1998. Low cost labor, soil, and climate have strongly influenced comparative advantage in agricultural exports.

Indonesia is a net exporter of agricultural commodities. During the 1994-1998 periods the growth of net agricultural exports in nominal values was 29 percent. The share of agricultural export earnings in 1998 was 8 percent. However in 1998 the agricultural component of total export earnings of Indonesia was 13 percent excluding export earnings from oil.

Table 1. 1 Export Value Oil and Non-Oil Indonesia 1994-1998 (US$ Millions)

Sectors		1994	1995	1996	1997	1998
Oil	Crude Oil	5,071.56	5,145.70	5,711.81	5,479.99	3,348.62
	Oil Product	932.92	1,296.74	1,516.09	1,302.45	708.07
	Natural Gas	3,689.12	4,021.97	4,493.91	4,840.10	3,815.46
	Sub total	9,693.61	10,464.41	11,721.81	11,622.55	7,872.16
Non Oil	Agricultural	2,818.33	2,887.32	2929.42	3,274.86	3,658.88
	Industrial	25,702.67	29,329.38	32,116.99	34,842.98	34,587.68
	Mining	1,837.11	2,735.30	3,054.21	3,170.54	2,724.44
	Others	1.71	1.55	1.30	532.69	4.45
	Sub total	30,359.82	34,953.56	38,092.93	41,821.05	40,975.47
Total Export		40,053.43	45,814.75	49,814.75	53,443.60	48,847.63

Source: Department of Trade and Industry, Nov. 1999

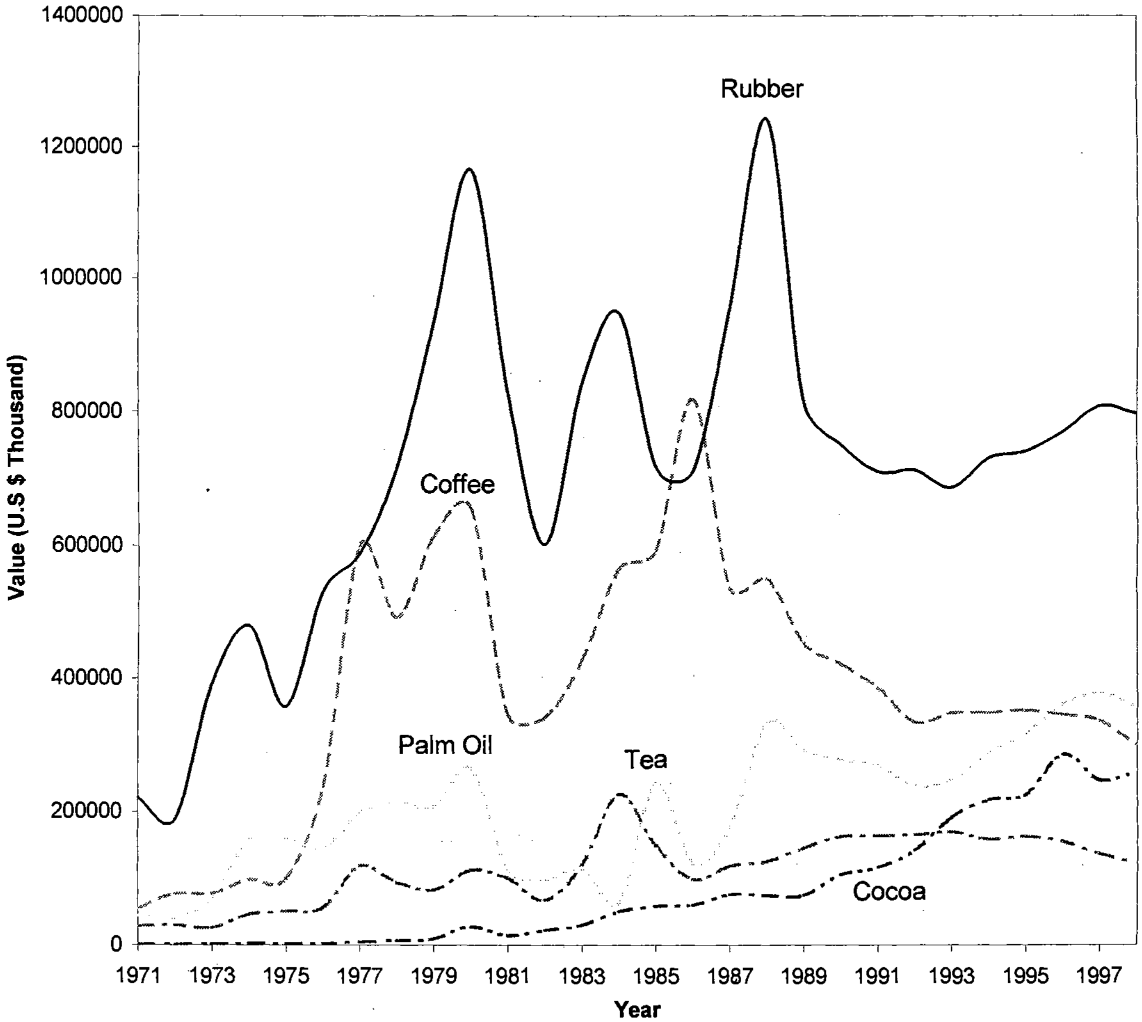
The most important cash crops for exports are natural rubber, palm oil, coffee, cocoa and tea. Traditionally, rubber has been the most valuable Indonesian agricultural export commodity. In 1997 rubber production was down 12 percent from 1996, and lower prices caused its export value to fall nearly 20 percent from $1,893 million to $1,505 million. Nevertheless, rubber retained its position as Indonesia's largest agricultural earner of foreign exchange, accounting for 26 percent of total agricultural export receipts. Table. 1.2 and Figure 1.1 show Indonesia's export value by major commodities.

Table 1. 2 Export Value by Major Commodities, 1994- 1998 (US$ Millions).

Commodities	1994	1995	1996	1997	1998
Plywood	3,650.25	3,451.51	3,544.12	3,476.80	2,327.25
Rubber	1,268.06	1,986.20	1,893.54	1,505.10	1,009.54
Garments	3,095.66	3,325.05	3,186.89	4,180.67	3,816.69
Coffee	750.40	621.75	597.76	582.58	606.79
Iron steel	453.66	521.78	608.27	659.69	990.38
Palm oil and Kernel	878.34	973.14	1,016.78	1,661.89	816.29
Copper	878.70	1,550.59	1,396.70	1,547.55	1,748.53
Pulp and Paper	782.48	1,503.65	1,369.42	1,952.99	2,469.19
Cocoa	218.18	225.35	286.45	247.34	259.34
Tobacco	62.13	76.46	81.93	123.88	139.32
Tea	106.71	94.16	106.22	150.20	169.28
Others	8,767.61	10,362.17	10,831.35	11,875.33	13,012.56
Total Non Oil Export	20,694.00	24,466.45	24,632.98	27,716.69	27,105.81

Source: Bank Indonesia, August 1999.

Figure 1. 1 Value of Indonesian Export by Commodities (1971-1998)



The unit root hypothesis has recently attracted a considerable amount of work in both the economics and statistics literature. Indeed, the view that most macroeconomic time series are stochastic rather than deterministic non-stationary has become prevalent. The seminal study of Nelson and Plosser (1982), that found most macroeconomic variables have a univariate time series structure with a unit root has catalyzed a burgeoning research program with both empirical and theoretical dimensions. It means if the series contained a unit root, the data are called non-stationary, which leads to spurious regression results. Their study applied a similar Dickey - Fuller (1979) statistical methodology to an economic time series. On the statistical front, they merged alternative approaches to test the unit root hypothesis, e.g they included a test proposed by Philips and Perron (1988) and methodology suggested by Campbell and Mankiw (1987,1988). Empirical applications of these methodologies generally reaffirmed that conclusion that most macroeconomic time series have a unit root.

Traditional cointegration tests have estimated a linear deterministic trend model without considering the possibility of structural changes in the data. Many macroeconomic and financial data appear to exhibit kinks in their trends due to a structural change in the potential growth rate. One major drawback of unit tests in that is all of them the implicit assumption is that the deterministic trend is correctly specified. Perron (1989) argued that if there is a break in the deterministic trend, then unit root tests would lead to the misleading conclusion that there is a unit root, when in fact there is not. He also tested the stationarity of some long-run U.S time series data that had been judged as non-stationary by previous studies. Since then, it has become apparent that the empirical results of many time series tests critically depend on the assumption of a deterministic trend.

Export demand functions have traditionally been estimated using standard regression models. However, several economists have pointed out the inappropriateness of applying regression models to non-stationary data because of the problem called "spurious relationship".

This study departs from the previously cited studies previously in at least five important respects. First, this study focuses upon the correct representation of the nature of non-stationarity evident in various time series across different agricultural commodities. Most specifications used in previous studies fail to recognize that real exports and some of its proposed determinants, such as real world trade or foreign real income, are potentially non-stationary integrated variables. Neglect of this point implies that inferences made concerning the long-run elasticity are potentially highly misleading (Granger and Newbold, 1974).

Second, special attention is given to the dynamic structure of the statistical model, which seems warranted in order to draw meaningful conclusions on the speed of adjustment. A common feature of most previous studies is the use of the log-level or log difference specifications. In this study, new econometric techniques that integrate the level and first-difference specifications are employed. To examine whether a long-run equilibrium relationship between real exports and exchange rate volatility exists, a cointegration technique is employed. The short-run dynamics by which real exports converge on the equilibrium long-term values are examined using error-correction procedures.

Third, the study contains a careful examination of the residuals. In particular, it tests for higher-order autocorrelation, functional form misspecification and non-normal residual. It is important to mention that previous studies failed to extensively examine the validity of their econometric model.

Fourth, this study uses quarterly data instead of annual data. It focuses on all exchange rates regimes in Indonesia from the fixed exchange rate period 1971:1 through the floating exchange period 1998:4.

Finally because of the policy change in the exchange rate regimes from fixed exchange rates to managed floating to floating exchange rates, alternative structural change regression estimations are conducted.

The purpose of this paper is to reexamine the assumptions of a cointegration model on which an error correction model is based, and to investigate whether a stable relationship exists between agricultural exports and exchange rate volatility. The thesis also emphasizes the importance of treating a deterministic trend properly, which has often been neglected in the past. It also focuses on the issue of structural change and cointegration on the deterministic model.

This study estimates an agricultural export demand model for Indonesia using recently developed cointegration and error correction techniques to examine the long-run and short-run relationship between exchange rate volatility and agricultural export movement. Unit root and cointegration tests are used appropriately with structural change. The advantages of this statistical approach are that it provides more efficient short-run and long-run coefficient estimates and avoids the problems of spurious regressions.

1.2 Objectives

The general objective of this analysis is to determine the influence of exchange rate on Indonesian agricultural export movement for individual products (cocoa, coffee, palm oil, rubber and tea) as well as aggregate products. The specific objectives are to evaluate the long-run relationship between exchange rate and Indonesian agricultural exports. Unit root and cointegration models with structural change are used to determine the effects of exchange rate on Indonesian short-run agricultural export growth using an error-correction model.

1.3 Organization of the thesis

The rest of the thesis consists of five chapters. In chapter II, a review of related empirical work is presented. This chapter contains three parts, the first covers the exchange rates and agricultural exports, the second deals with unit root and cointegration without structural change, and the third covers the unit root and cointegration without structural change. Chapter III presents relevant theories for the development of the model. Chapter IV describes the methodology for the analysis, details on data adjustments and methods for the model are presented. In chapter V, empirical results are examined and interpreted. Chapter VI, finally, concludes the study; the summary, conclusion, policy implications are presented based upon the results of the study.

CHAPTER II

LITERATURE REVIEW

2.1 Exchange rates and Agricultural Exports

From August 23, 1971 to November 15, 1978 Indonesia maintained a fixed exchange rate system rate of Rp. 415 for each U.S dollar. This rate was maintained until a major devaluation took place, which was Rp 625 for each U.S dollar until November 1986. Then a major change also took place in the system itself. A managed floating rate regime was implemented (Harinowo, 1985). From late 1986 until August 1997, the rupiah was on managed float, depreciating slowly against a basket of trading partner currencies. During this period, Bank Indonesia (the Central Bank) steadily widened the band between its buying and selling rate trade on the rupiah in an effort to encourage the development of an interbank foreign exchange market and discourage speculative short-term capital flows. However, with pressure on the rupiah and other currencies of the neighboring countries, Bank Indonesia decided on August 14, 1997 to eliminate its intervention band. Since then the rupiah has essentially floated, although Bank Indonesia continues to occasionally intervene in an effort to stabilize the exchange rate (Indonesia Economy Policy and Trade Practices Report, 1997).

The rationale behind the change from a fixed to a tightly managed floating system in November 1978 was to curb the rate of inflation. From August 1971 to November

1978, the rupiah pegged to the U.S dollar in fixed value was twice devalued. The high inflation rate in Indonesia before November 1978 caused a decline in the rupiah's purchasing power parity. In the context of free capital mobility, Indonesia was forced to abandon the fixed exchange rate system. Instead of moving to a flexible exchange rate system, Indonesia chose to move to a tightly managed floating system. Arnt (1978) has argued that the motivation behind this change was the prospect of the imposing Balance of Payments (BOP). It was projected that the BOP could be moved from a huge surplus to stationary or even declining in international reserves. Another argument (Dick, 1979) states that the change was a basis for developing export-oriented economy.

The movement to the flexible managed floating exchange rate system in March 1983 was motivated by experience from using the previous system. The tightly managed floating system was considered a failure in reducing inflationary pressure from the second oil shock in fiscal year 1979/1980 (Arnt, 1983). Arnt (1979) argued reasons for this change as follows. First, the stability of the Indonesian exchange rate against major currencies that have moved freely since the collapse of the Bretton Wood system in 1973 is substantially determined by the major countries and has little to do with the external condition of Indonesia. Second, Indonesia was attempting to maintain an appropriate level of international reserves to provide assistance to industries producing tradable goods and to stabilize domestic prices. In pursuing these objectives, the fixed and the tightly managed floating system appeared to be ineffective.

In August 1997 the government of Indonesia decided to adopt a freely floating exchange rate. The flexible managed floating system did not maintain the desired level of international reserves and experience showed that the levels of foreign reserve changed with high variability, and the level of competitiveness for tradable goods fluctuated substantially.

Several previous studies of the relationship between agricultural exports and macroeconomic conditions have been undertaken to help explain domestic price and export expansion. One of the first articles examining the relationship between devaluation and agriculture was by Schuh (1974). He argued that the exchange rate affects the valuation of resources within a country, the distribution of benefits between consumers and producers, and the way that the benefits of technical change are shared between the domestic population and the world at large. He connected the devaluation of the early 1970's to agricultural price increases. The lower exchanges rates of the U.S. dollar cut the price of the agricultural exports.

Schuh argued that over-valuation and under-valuation of the dollar had been important in explaining the path of domestic agricultural prices. He showed that devaluation of the dollar during the fixed exchange rate regime to a floating one made the U.S more vulnerable to international economic policies and events. Nainggolan (1979) produced support for Schuh's reasoning for Indonesian agricultural exports, while Barnett, Bressler, and Thompson (1981), and Glecker (1988) did it for U.S. agricultural export, and Lin (1981) produced support for Canadian agricultural exports.

Clark (1974) studied the effect of the United States dollar on both manufactured goods and agricultural commodities exports. Grenshields (1974) studied the effect of the Japanese yen on United States exports of wheat, corn and soybean to Japan. Their results showed that the response to exchange rates changes was analogous to the response of price changes and the dollar devaluation had little effect on the exports.

Velliantis - Fidas (1976) reported econometric studies that measured the effect of exchange rates on U.S. agricultural exports. The results indicated that the exchange rate was not a significant explanatory variable for U.S. farm exports. Kost (1976) presented a theoretical framework to assess the trade impact of change in the exchange rate on commodity production, consumption, trade level, and prices for two trade partners. The theoretical model was used to analyze the possible effect of devaluation on the agricultural sector. As Henneberry, Drabenstott and Henneberry (1987) explained, the dollar's exchange value does not fully affect farm trade patterns because farm-trading countries are more likely to have fixed exchange rates, and these regimes are less responsive to exchange market forces. Gotur's (1985) estimation showed that the increased value of the U.S. dollar in "1981" and "1982" reduced the value of U.S. agricultural export. Also, Thursby and Thursby (1987) estimated that a dollar devaluation of 10 percent would increase on domestic prices of wheat in the U.S by 6.9 percent. These results compare closely with those obtained by Chambers and Just (1981).

Hooper and Kohlhagen (1978) were the first to analyze systematically the effects of exchange rate uncertainty on trade flows. The methodology proposed by Hooper and Kohlhagen (1978), in original or modified forms, is often used as the basis of empirical investigation. They modeled the behavior of exporters and importers operating under exchange rate risk and concluded that if traders are generally risk averse, an increase of exchange rate risk will unambiguously reduce the volume of trade.

Bredahl, Collins, and Myers (1980) explained why it is inappropriate to use simple exchange rate measures to infer exchange rate impacts for individual commodities. Their result indicated that much of the variability in commodity prices is caused by factors other than the exchange rate. They argued that the size of exchange rate impacts on trade and prices depends on the crop, the year, the country considered, and government influence, underlying elasticity and whether real prices or nominal prices are being measured.

Cushman (1983) modified the Hooper and Kohlhagen (1978) study and extended it to analyze 1965 to 1977 trade's flows. He also modified the study by using real exchange rates in his calculation of exchange rate risk. Using real exchange rates instead of nominal risk, he showed that as exchange rate uncertainty increases, trade quantities decrease. He also indicated that the risk effects usually occur with a lagged effect.

Cushman (1986) extended the Hooper and Kohlhagen (1978) framework to include what has been termed the "third country effects". He studied export flows from the U.S. to the U.K., the Netherlands, France, Germany and Japan. Using joint significance tests that account for both bilateral trade flows and third country effects, he found that third country risk effects are negative. He also concluded that the total cost of risk had grown steadily over the period of floating exchange rates.

Using quarterly data, Warner and Kreinin (1983) assess the effects of variations in the current and expected exchange rates on real trade flows by estimating import and export demand functions for 19 developed and 18 developing countries. They observe that generalized floating has had an impact on the volume of trade and that estimated coefficients change significantly from one period to another. Akhtar and Hilton (1984) however, reach more definite conclusions. Considering only bilateral trade between the United States and West Germany, the model is extended to include additional explanatory variables such as capacity utilization. Akhtar and Hilton conclude that exchange rate uncertainty, as measured by the standard deviation of effective exchange rates, has had a significant negative impact on the imports and exports of the two countries.

Kincaid (1984) estimated import demand and export supply equations, and analyzed the effectiveness of the exchange rate adjustment in promoting non-oil exports in Indonesia. Timmer (1986) emphasized the linkages between macroeconomics policy and food sector in Indonesia and found that the foreign exchange rate is the most important macro price effecting agricultural exports.

Batten and Belongia in 1984 as a part of a paper about exchange rate behavior and agricultural exports presented an extremely simple, single equation aggregate export model. Using quarterly data, their empirical model regressed the volume of U.S. agricultural exports on the trade-weight index of foreign real GNP, a deflated price index of U.S. agricultural exports and the trade-weight index of the dollar. Batten and Belongia conclude that importer affluence is the main factor that affects agricultural exports, not the exchange rate.

Gotur (1985), however, questioned the robustness of Akhtar and Hilton (1984) results. Gotur expanded the number of countries to include France, Japan and the United Kingdom, and varied the sample period, and changed the measures of exchange rate risk. Gotur failed to find conclusive evidence that exchange rate uncertainty had any significant impact on bilateral trade flows. She also developed the model later used by Asseery and Peel (1991) and Chowdhury (1993) and showed that it is a maximization solution of behavioral demand and supply functions for exports.

Bressler and Babula (1987) explored the relationships among the Federal Reserve Board's real trade-weighted exchange rate and cash prices, export sales, and shipment of wheat from a forecasting perspective explicitly. Bressler and Babula (1987) reported mixed results when comparing forecasts from the four variables vector autoregression (VAR) to those of univariate autoregression. They conclude that forecasts of wheat sales are not improved by including the exchange rate as an explanatory variable but that "exchange rates seem to have an impact on real wheat prices".

In their 1988 empirical study, Childs and Hammig used a model with simultaneous equations for five commodities to test the hypothesis that the exchange rate is a key explanatory variable affecting the level of farm exports. Childs and Hammig had conclusions similar to those of Batten and Belongia (1984), finding that exchange rates matter much less than do variables representing importing country income.

Khumar and Dhawan (1991) attempted an empirical examination of the impact of exchange rate uncertainty on Pakistan's exports to its major partners in the developed world for 1974-1985. Using monthly and quarterly data the results showed that the volume of Pakistan's exports to the developed world might have been adversely affected by increased variability of its bilateral exchange rates.

A study by Grobar (1993) used pooled time-series and cross-sectional data to conduct an empirical investigation of the relationship between real exchange rate uncertainty and manufactured exports of developing countries. Evidence is found that some categories of LDC manufactured exports are negatively affected by real exchange rate uncertainty.

2.2 Unit Root and Cointegration Without Structural Change

Studies that have measured the effects of exchange rate volatility on real exports have used export demand models that are very restrictive. The problem is particularly acute in the case of the exchange rate volatility and relative price estimates, because the effects of this variable are widely believed to build slowly with statistically significant lags. The specifications used in previous studies have not recognized that real exports and some of its proposed determinants such as real world income are, a priori, potentially non-stationarity integrated variables. Failure to consider the non-stationarity of the variables may, in part, explain the mixed conclusions on the effects of exchange-rate volatility. In this study, the properties of the individual time series will be established prior to testing for cointegration. Series that are integrated of a different order cannot be cointegrated. In the second step, the maximum likelihood framework for estimating cointegrating vectors between integrated series suggested by Dickey- Fuller (1981) and Johansen (1988) will be used.

The sample standard deviation of the exchange rate has been used as a proxy for exchange rate uncertainty in empirical studies that examine the effects of exchange rate uncertainty on the volume of foreign trade (Akhtar and Hilton (1984) and Gotur (1985)). These studies have provided conflicting evidence on this issue. It has been argued that using the sample standard deviation is inappropriate because the empirical distribution of exchange rates is not normal (Arize, 1997).

In this regard, subsequent empirical research by Cushman (1983), Kenen and Rodrik (1986), and Chowdhury (1993) employed a moving standard deviation of the rate of change of the exchange rate as a proxy for exchange-rate uncertainty. Chowdury obtained a clear pattern of results that supported the hypotheses that the higher volatility of exchange rates led to a reduction in international trade transactions after the breakdown of the Bretton Woods system in March 1973. Other studies have reported ambiguous results. The inconclusive empirical evidence in previous studies may also be due in part to a number of estimation problems, besides the use of different proxies for exchange-rate uncertainty. Researchers, with the exception of Chowdhury and Arize, have estimated the trade equation in the log-level form. They also implicitly assumed the data were stationary. However, it is highly unlikely that the utilized data have this desirable characteristic. As shown by Nelson and Plosser (1982) several macroeconomic variables generate spurious inferences in the absence of cointegration.

Dutt and Ghosh (1994) investigated the export and economic growth cointegration structure for a large sample of twenty-six low, middle, and high-income countries, including four newly industrialized countries over the period 1953-1991. The Dickey-Fuller and Phillips-Perron tests are conducted for stationarity. Then they performed cointegration tests based on the Phillips-Hansen fully modified OLS method, and ran the Phillips-Ourialis test of non-stationarity on the residuals. The results showed that for most countries in the post World War II period, export growth and economic growth have moved together.

Arize (1995) said that traditional export demand studies for other economies that do not include a variable representing the influence of exchange risk are potentially misspecified. This evidence further suggests that exchange rate volatility may have significant affects on the allocation of resources as market participants attempt to minimize exposure to the effects of exchange risks.

Arize (1995) said that Granger and Newbold (1974) have questioned the assumption of data stationarity because most time-series variables such as those included in the model contain one or more unit roots which make them non-stationary. In such circumstances, the use of standard t- ratios to judge the significance of a variable can be misleading. By using cointegration and error-correction techniques, conditional and unconditional measures of exchange rates, and by testing for structural stability in his study on the effects of exchange rate volatility on U.S exports, the results were favorable to the hypothesis that exchange rate volatility impedes trade.

Arize (1997) investigated the impact of exchange rate volatility on real exports by employing a multivariate cointegration and error-correction modeling. Arize (1997) used the quarterly export data of seven countries over the floating exchange rate period. In the specific function considered, real exports depended upon foreign economic activity, relative price and exchange rate volatility. Each estimated model satisfied several recently developed econometric tests in the analysis of time-series data for issues such as cointegration, stationarity, specification errors, residual autocorrelation, and heteroscedasticity. The resulting evidence strongly indicated the presence of a single unit root in virtually all variables at normal significance levels, a result consistent with the macroeconomic literature. It also suggested that there was a unique, statistically significant long-run relationship between real exports and exchange rate volatility in each country. In addition, in the majority of cases, exchange rate volatility had a short-run effect on export volume. The finding was consistent with the result report by Chowdhury (1993) and Arize (1995).

Arize (1997) examined the impact of real exchange rate volatility on the trade flows of the G-7 countries, in the context of a multivariate error-correction model. The results showed that the increase in the volatility of the real exchange rate exerted a significant negative effect upon export demand in both the short-run and the long-run in each of the G-7 countries. These effects may result in significant misallocation of resources by market participants.

Hassan and Tufte (1998) did not recognize that the trade flows and the variables explaining them were likely to be non-stationary and potentially integrated variables. Neglect of this point implies that inferences made concerning the long-run elasticities were potentially misleading as noted by Granger and Newbold.

Sukar (1998) investigated how U.S exports were dynamically associated with foreign income and the real effective exchange rate of the U.S. using cointegration and error correction models. Cointegration results indicated a direct relationship between exports and foreign income and an inverse relationship between U.S. exports and real exchange rates. The error correction model indicated a significant short run relationship between changes in exports and changes in foreign income.

There has long been concern over the volatility of exchange rates and their impact on the volume of foreign trade. However, very little attention has been paid to the choice of an appropriate volatility variable as well as the proper specification of the trade equation.

In this study techniques that integrate the level and first difference specifications are employed. To examine whether a long-run equilibrium relationship between real exports and exchange rate volatility exists, a cointegration technique is employed. The short run dynamics by which real exports converge on their equilibrium long-term values are examined using an error-correction model.

Most of the earlier studies specify trade models in levels or in log level model. These models have been criticized because the levels and log levels of many economic variables in trade models are non-stationary. The regression equation relating such variables could lead to spurious regressions, phenomena first described in Granger and Newbold (1986). This phenomenan refers to the possibility that inferences based on the ordinary least square parameter estimates in such models are invalid because t and F ratio test statistics do not converge to their limiting distribution as the sample size increases. In this case the null hypothesis of no relation would be rejected wrongly as discussed by Engle and Granger (1987).

2.3 Unit Root and Cointegration With Structural Change

In this section, unit root tests that are applicable in models with structural change are discussed. These tests differ from the usual unit root test in their treatment of the alternative hypothesis (Bacillar, 1996). The alternative hypothesis considered here is more general and allows for shifts in the level or the growth rate of the series.

Nelson and Plosser (1982) originally studied the trend stationary model. But they only explained characterization of the deterministic components of economic time series and the stochastic trend on the unit root. Rappoport and Reichlin (1987) and Perron (1989) argued that the heterogeneous behavior in the deterministic component of economic time series was mainly due to unusual events like the oil price shock of 1973.

They argued that most economic time series are not characterized by the presence of unit roots and that fluctuations are transitory.

Studies by Engle and Granger (1987), Johansen (1988), and Philips and Ourialis (1990) assumed no deterministic trend or a linear deterministic trend. Hansen (1992) and Johansen (1994) extended the earlier studies by introducing a higher order time trend. Perron (1989) first considered a structural change in deterministic trend in a unit root test. He presented evidence that most economic time series are trending stationary if one allows a single change in the intercept. He showed theoretically that if the data generating process has a kink or a jump in the deterministic trend, a unit test that ignored such a possibility tended to have a bias for accepting the null hypothesis of a unit root. In his research he found that many of the variables that had previously been judged as non-stationary were actually stationary.

Hogan (1990) argued that unit root rests for real exchange rates must span a period long enough to allow for the possibility that reversion takes a considerable amount of time. Using longer data series presents the problem of traversing obvious structural breaks. Even when accounting for the possibility of a structural break, at which exchange rate regimes changed during the early 1970s, he cannot reject the presence of unit roots in real exchange rate data. He also argued that the presence of roots does not give valuable insights into choosing one type of model over another. This would depend on the relative importance of the non- stationary components of the series. It is shown that a significant long-run relationship exists between real exchange rates for the period in which nominal exchange rates were fixed. However, for the floating period, there is no evidence that real exchange rates are related in the long-run.

Perron (1989) showed that the standard test of the unit root hypothesis against trend stationary alternatives could not reject the unit root hypothesis. The true data generating mechanism is that of stationary fluctuations around a trend function, which contains a one-time, break. He derived a test statistic that allowed distinguishing the two hypotheses when a break is present. He applied these tests to the Nelson-Plosser data set and to the postwar quarterly real GNP series. In the former, the break was due to the 1929 crash and takes the form of a sudden change in the level of the series. For 11 out of the 14 series analyzed by Nelson and Plosser, the unit root hypothesis cannot be rejected at high confidence level. In the case of the postwar quarterly real GNP series, the break in the trend function occurred at the time of the oil price shock (1973) and takes the form of a change in the slope. Here again he rejected the null hypothesis of a unit root. He concluded that the fluctuation is indeed stationary around a deterministic trend function. The only "shocks' which have had persistent effects were the 1929 crash and the 1973 oil price shock.

Perron (1990), with correction in Perron and Vogelsang (1992), considered testing for a unit root in a time series characterized by a structural change in the mean level (rather than in the trend). Again the analysis was for a known break point, but it was shown that allowing for a break reverses previous conclusions that a unit root characterizes the real interest rate for the U.S.

Hakkio and Rush (1991), Trehan and Walsh (1988,1991), Haug (1991) developed an alternative framework to test borrowing constraints. Imposing breaks they obtained cointegration between revenues and expenditures in the earliest years but no cointegration in the years starting from the mid 1970s. They interpreted this result to mean that the deficit had become a problem only in recent years and was not sustainable.

Kunitomo (1995), introducing a structural change to the cointegration test based on a maximum likelihood-ranking test, proved theoretically that the traditional test produced a bias toward reducing the rank if the data-generating process had a structural change. He emphasized the risk of "spurious cointegration." He also proposed a cointegration test for the variables with kinked linear deterministic trends, and presented some of the applied examples.

In addition to Ogaki and Park (1992), who first distinguished between the two kinds of cointegration, Johansen (1994) and Hansen (1992) also used a cointegration test that made the distinction between the two kinds of cointegration. Johansen (1994) proposed a testing method that introduced a deterministic trend to the error correction term. The model explained the following two cases: the case in which the cointegrating vector is linearly independent from the exogenous variables (consisting of a constant and a linear trend) and the case in which the cointegrating vector is linearly dependent on the exogenous variables. Although he did not use the terminology "stochastic cointegration", his case of linear dependence corresponds to stochastic cointegration. Extending the estimation method of Philips and Hansen (1990). Hansen (1992) proposed a stability test for the contegration vector based on the langrangian multiplier method. His method has the advantage that it can test the stability of the relationship between deterministic trends in addition to a cointegrating vector.

Quintos (1993) tested for structural breaks to determine whether a model with shifts is appropriate in U.S fiscal policy and whether there had been a structural change in deficit policy. She also applied a test for change in cointegration in the parameters of the cointegrating vectors. She found that there was a shift in deficit policy in the 1980's. In other words, no cointegration between revenue and spending in the early 1980's and cointegrating between revenue and expenditures.

Quintos (1993) tested for structural breaks to determine whether a model with shifts was appropriate. Quintos treated the break points as known since the test conducted was known to have higher power than the test used by Haug (1992) that treated break points as unknown. Haug's results showed no evidence of parameter instability over the sample periods 1960-1990 when the break points were treated as unknown, but use of the test with known breaks showed significant breaks in the early 1980's. Quintos justified the choice of the break date by statistically testing for its significance.

Kunitomo (1995) proposed a cointegration test for a multivariate time series model with structural changes. He found that if a structural change was assumed in the Japanese growth trend in the early 1970s on a postwar times series of data for real GDP and real private final consumption expenditure turned out to be stationary. If when it is a linear deterministic trend with a structural break, the long-run relationship between the two variables depends on maintaining the stable relationship before and after the structural change in the deterministic trend, rather than the cointegration between stochastic trends. This suggests the risk is high of a "spurious unit root" and a "spurious cointegration" arising from a misspecification of a deterministic trend when the traditional time series model is applied without appropriate caution.

Soejima (1995) found that real GDP might be stationary under the assumption of structural changes in the linear deterministic trend, but also that such an assumption is inappropriate for nominal variables such as money supply and price level.

Dropsy (1996) performed several structural stability tests for five foreign exchange rates relative to the dollar and five foreign exchange rate relative to the Deutsche Mark using quarterly data over 20 years (starting in the first quarter 1974). He missed the important structural break of the switch to flexible exchange rates in March 1973. He identified several series with structural breaks need to be analyzed in further study, whether they can be identified with major events or policy changes.

Soejima (1998) presented time series model with deterministic trend consisting of multiple linear and nonlinear parts as the appropriate model for Japans postwar real GDP, money supply and GDP deflator. This indicated that the cointegration between the three variables, which is supported by previous studies, arose from a misspecification of the time series model.

The study examines the influences of exchange rate on the movement in the volume of Indonesia agricultural exports for individual commodities (coffee, tea, rubber, palm oil, and cocoa) as well as commodity aggregates. This study will provide additional empirical evidence on the effects of exchange rate on real exports. No previous study on the effect of exchange rate on Indonesian agricultural exports has been published using time series techniques of cointegration, error correction model and causality using the structural change method.

CHAPTER III

CONCEPTUAL FRAMEWORK

3.1 Exchange Rate Determination

Exchange rate determination is based upon two assumptions: (1) demand for money is a stable function of a limited numbers of aggregate economic variables, and (2) in the absence of tariffs, transportation costs and restrictions upon trade, the law of one price will hold in international markets. In the monetary approach, the law of one price appears in the form of purchasing power parity condition (PPP), in which the exchange rate equates the price of traded goods in alternative currencies. The absolute PPP hypothesis states that the exchange rate between currencies of two countries should equal the ratio of the price levels of the two countries. Specifically,

$$E = P / P^* \qquad (3.1)$$

Where P and P^* represent the domestic and foreign currency prices of traded goods, and E is the domestic currency price of foreign exchange. This definition implies that the exchange rate appreciation and depreciation refer to fall and rise in E. The relative PPP hypothesis states that the exchange rate should bear a constant proportionate relationship to the ratio of national price levels; in particular,

$$E = k\,P / P^{*} \tag{3.2}$$

where k is a constant parameter.

The real exchange rate (RER) can be defined (Dutton and Grennes) as:

$$RER = E(P_j / P_d) \tag{3.3}$$

where E is the nominal exchange rate. Since RER is determined by nominal exchange rates and the CPI foreign (P_j) and domestic (P_d) countries, the monetary model of exchange rate determination can be applied for the RER determination with the inclusion of the relative price ratio.

In calculating the export demand model, the trade weight for each commodity is calculated. These weights explain the index of total amount exported to one country compared to total amount export to all importing countries

$$V_{jt} = \sum_{t=1}^{p} W_{jt} \frac{E_{jt} P_{jt}}{PD_{jt}} \tag{3.4}$$

where:

V_{jt}= Real effective exchange rate

W_{jt} = Trade weight share corresponding to partner j.

E_{jt} = Nominal exchange rate in country j to domestic country

P_{jt}= CPI countri j.

PD_{jt} = CPI in the importing countries.

But the W_j needs to be calculated first in the Equation 3.5 as follow,

$$W_{jt} = \sum_{t=1}^{p} XPTS_{jt} / XPTS_{jt} \tag{3. 5}$$

where:

$XPTS_j$ = Total amount of the commodity exported to country j from Indonesia during each year.

$XPTS_r$ = Total amount of the commodity exported from Indonesia to all importing countries during each year.

3.2 Export Demand Model

Export demand functions have traditionally been derived from utility function. The demand for Indonesia goods by a trading partner j can be expressed in additive utility form as follows (Houthaker 1960)

$$u(X_{ij},\ldots\ldots,X_{nj}) = \sum_{i=1}^{n} U(X_{ij}) \tag{3. 6}$$

where X_i^j = quantity export of *ith* good shipped from Indonesia to country j.

Assuming a CES utility function, the above function can be maximized subject to a budget constraint and a system of equation of the following form can be obtained (Sato, 1976). However, the trade weighted income for importing countries needs to be established first as,

$$Y_{Jt} = \sum_{t=1}^{p} W_{Jt} GDP_{Jt} / CPI_{jt} \tag{3.7}$$

where

Y_{jt} = Income trade weight of country j.

W_{jt} = Trade weight share corresponding to partner j.

GDP_{jt}.= GDP country j as income measured.

CPI_{jt}= CPI country j as income measured.

$$ln X_{ij} = \phi_1 \, ln Y_j + \phi_{2i} \, ln P_{xi} / P_j + \varepsilon_t \tag{3.8}$$

where:

Y_j = Income trade weight of country j

P_j = Price index of country j

P_{xi} = The average price of Indonesia exports

Exchange rate volatility creates uncertainties about the size of profits that importers can realize from trade (Lanyi and Suss 1982). Abrupt changes the price of traded goods can cause the actual level of profit to deviate from the expected level. Thus, the volume of exports is expected to fall. So Equation 3.8 has been modified to include a variable of trade weight of foreign countries exchange rate.

$$ln X_{ij} = \beta_1 \, ln Y_j + \beta_{2i} \, ln P_{xi} / P_j + \beta_3 V_j + \varepsilon_t \tag{3.9}$$

where V_j represents real effective exchange rate with country j (trade weight index).

The export demand model is basically like any other demand model. Price and quantity are inversely related, ceteris paribus, with equilibrium price and quantity determined the interaction of supply and demand. In most empirical studies own price is assumed exogenous i.e. supply is perfectly elastic. Thus, the export supply equation is not explicitly considered in trade models (Murray and Ginnman, 1976; Houthaker and Magee, 1969; Warner and Kreinen, 1983; Krugmena, 1989; Arize, 1995; Chowdhury, 1993). Aggregating the export demand function over all the goods exported to aggregating over all the trading partners, the export demand function, from the Equation 3.8 which including the prices of commodity trading the export demand can be rewritten as:

$$X_t = \alpha + \beta_1 Y_t + \beta_2 P_t + \beta_3 V_t + \varepsilon_t \tag{3. 10}$$

Where :

X_t = Total exports to all trading partners (U.S $)

Y_t = Exports -weighted income of trading partners GDP (1995=100)

P_t = A Relative price variable (U.S$)

V_t = Real effective exchange rate between Indonesia and its trading partners (trading partner exchange rates /U.S$)

e_t = Error term

3.3 Non-stationarity and Unit Root

The econometric literature on unit roots took off after the publications of the paper by Nelson and Plosser (1982) that argued that most macroeconomic series have unit roots and that this is important for the analysis of macroeconomic policies.

Yule (1926) suggested that regressions based on trending time series data could be spurious. Granger and Newbold (1974) further pursued this problem and this also led to the development of concept of cointegration. The development of unit roots and cointegration has changed the way time series analysis is conducted.

There are many substantial differences between stationary and non-stationary data. Whether the time series data is stationary or not has an important implication. Under non-stationarity there are serious problems with interpreting standard regressions that attempt to explain the behavior of the time series data. In this case, the existence of unit roots implies an infinite variance and the standard errors of the estimated parameters are meaningless. The consequences for the statistical properties of estimators and tests are profound as evidenced by the substantial literature on " spurious regressions". To overcome the problem of non-stationarity, some researchers have suggested differencing the data to remove random walk and trends. However, by analyzing only differences of economic time series, all information about long run relationships between the levels of economic variables is lost. This is a solution to possible spurious regression. The Figure 3.2 shows the flowchart of diagnostics for time series regressions.

Engle and Granger (1987) suggest a two-step method of integrating the cointegration techniques with the error-correction mechanism. This has several advantages over the standard regression model in dealing with non-stationary data. Time series analysts usually advocate differencing of non-stationary series to estimate multivariate time series models. However, cointegration and error-correction modeling enable the researchers to study simultaneously the dynamics of short-run changes and the long-run equilibrium relationships. Since first differencing is not required to achieve stationarity, this procedure does not involve any loss of long-run information contained in the data.

Granger (1986) recommends cointegration tests and Engle and Granger (1987) as a technique for examining the long-run relationship and capturing the short-run dynamics. Lao (1993) suggested applying unit root tests to check the stationarity of data before performing the cointegration test. The most commonly used test of the null hypothesis of a unit root in an observed time series is a derivative of the Dickey-Fuller and Augmented Dickey-Fuller test. Dickey and Fuller, as well as Augmented Dickey-Fuller use Monte Carlo experiments to tabulate the sampling distribution of the regression "t statistic".

3.3.1 Stationarity

A time series sequence (x_t) is covariance stationary if the mean of the series is finite and independent of time. All periods of the variable have the same finite mean,

$$E(x_t) = E(x_{t-s}) = \mu_x \qquad (3.11)$$

The variance of the series is finite and time independent. That is:

$$Var(x_t) = Var(x_{t-s}) = \sigma_x^2, \text{ or} \tag{3. 12}$$

$$E(x_t^2) = E(x_{t-s}^2) = \sigma_x^2 \tag{3. 13}$$

All autocovariances are finite and time independent,

$$Cov(x_t, x_{t-s}) = Cov(x_{t-j}, x_{t-j-s}) = \gamma_s, \text{ or} \tag{3. 14}$$

$$E(x_t, x_{t-s}) = E(x_{t-j}, x_{t-j-s}) = \gamma_s \tag{3. 15}$$

where μ_x, σ_y^2, and γ_s all are constant and stationary.

If the three conditions above hold, this series sequence shows weak stationarity. If the probability distribution P ($x_1, x_2, \ldots x_t$) is also stationary, then the time series process is strictly stationary. A process whose joint probability distribution does not change through time is stationary. If the series x_t is stationary, then, for any $t,j,$ and s.

$$P(X_t,\ldots,X_{t-s}) = P(X_{t-j},\ldots,X_{t-j-s}) \tag{3. 16}$$

3.3.2 White Noise

A sequence $\{\varepsilon_t\}$ is a white noise process if each value in the sequence has a mean zero, a constant variance, and is serially uncorrelated. Formally, if $E\ (y_t)$ denotes the theoretical mean value of y_t, the sequence $\{\varepsilon_t\}$ is a white noise process if for each time period t,

$$E(\varepsilon_t) = \quad E(\varepsilon_{t-1}) = \ \dots \ = 0, \tag{3.17}$$

$$E(\varepsilon_t)^2 = \quad E(\varepsilon_{t-1})^2 = \ \dots \ = \sigma^2 \tag{3. 17}$$

$$E(\varepsilon_t, \varepsilon_{t-s}) = \quad E(\varepsilon_{t-j}\varepsilon_{t-j-s}) = 0 \tag{3. 18}$$

Hence, the autocorrelation function for a white noise random variable is zero for all non-zero lags. A white noise process is a particular form of a stationary process.

3.3.3 Non-stationarity

If the probability distribution of a time series process changes over time, then, it is a non-stationary time series. Most time series of economic variables exhibit non-stationary in level (variable before differencing). Such time series variables are subjected to detrending procedures to make them stationary before proceeding with further analysis.

The detrending procedure can take two forms. The first is regressing the time series as simple linear (or higher order) functions of time and then using the residuals as the detrended series. If stationarity is achieved after fitting a time trend, the variable is said to be trending stationary. The trend stationary process arises because of the effect of a deterministic trend. The second approach is to take the first difference of the series of interest and use the first difference as the detrended series. If stationarity is achieved after differencing, the variable is said to be difference stationary. A difference stationary process is a random walk, or it has a stochastic trend. An advantage of the second method is that if the series are in log levels, then the first difference series are approximately the percentage change over the previous period. Figure 3.1 shows the cocoa price movement in the level and first difference.

3.3.3.1 Removing time trend

The series, x_t, is generated according to the equation:

$$x_t = \alpha + \beta t + \varepsilon_t \qquad (3.19)$$

where β_t is deterministic trend and ε_t is white noise.

An appropriate way to transform this model is to estimate the regression equation. The regression of x_t on a constant and time results in residuals, which have a mean and are orthogonal to t.

3.3.3.2 Differencing

Consider the pure random walk model

$$x_t = x_{t-1} + \varepsilon_t \qquad (3.20)$$

Taking the first difference

$$\Delta x_t = x_{t-1} - x_{t-2} + \varepsilon_t \qquad (3.21)$$

Clearly, the $\{x_t\}$ sequence is stationary since the mean and variance are constant and the covariance between Δx_t and Δx_{t-s} depend solely on:

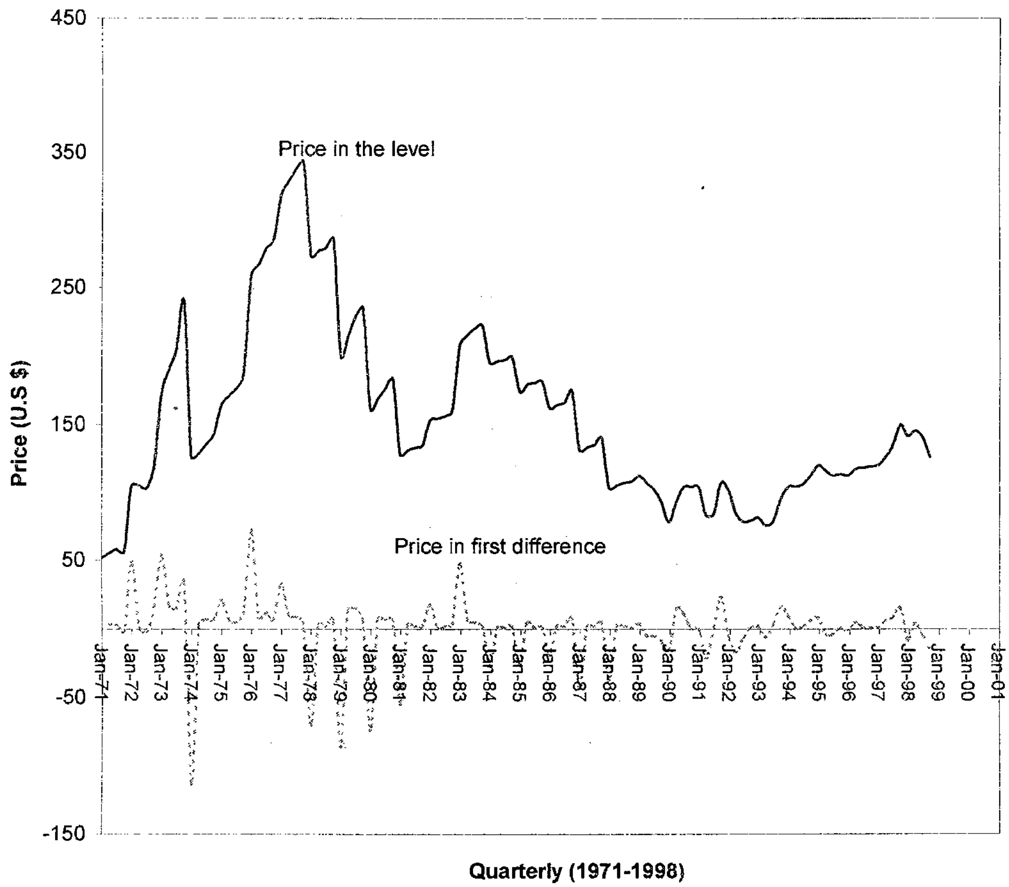
$$E(\Delta x_t) = E(x_t) = 0 \tag{3.22}$$

$$Var(\Delta x_t) = E(\Delta x_t)^2 = E(\varepsilon_t^{\,2}) = \sigma^2 \tag{3.23}$$

$$Cov(\Delta x_t, \Delta x_{t-s}) = E[(\Delta x_t, \Delta x_{t-s})] = E(\varepsilon_t \varepsilon_{t-s}) = 0 \tag{3.24}$$

Figure 3.1 shows the price of cocoa in the level and first difference.

Figure 3.1 Price of Indonesian Cocoa Export Quarterly (1971-1998).



3.4 Cointegration

Cointegration is a relatively new statistical concept, pioneered by Granger (1983), Engle and Granger (1987) as a technique for examining the long-run relationship and capturing the short-run dynamics. Cointegration is a property possessed by non-stationary time series data. In general terms, two variables are said to be cointegrated when linear combinations of the two are stationary, even though each variable is non-stationary. They also have examined the causal relationship between two variables when a common trend exists between them.

The coefficients of Equation 3.10 are usually estimated using traditional statistical procedures. The underlying assumptions of such tests are that the series in the equation is stationary in order to avoid the spurious regression coefficient result. Specifically, a time series is said to be covariance stationary if its mean, variance, and covariance are all invariant with respect to time and therefore it is integrated of order zero, I(0). If the time series requires first differencing to achieve stationarity, it is integrated of order one, I(1). Any linear combination of two I(1) time series will also be an I(1) time series. However, if there exists some linear combination of the two series, which is I(0), the series are said to be cointegrated. If the variables are non-stationary, standard test statistics such test t and F do not have their desirable limiting distributions and therefore, traditional tests of significance are not valid. Engle and Granger (1987) suggest that if the variables in their level are non-stationary, but linear combination of them are found to be stationary, a regression model among the non-stationary variables generate consistent estimates for the coefficients. Standard test statistics are then valid in making inferences without running into the "spurious regression" phenomenon (Ahmed, Haque and Talukder, 1993). The variables are said to be cointegrated which implies that there is a long run equilibrium relationship among the variables.

The estimation procedure suggested by Engle and Granger involves the following two stages. First, Equation 3.10 is used to establish the presence of cointegration among the variables. Dickey-Fuller (DF) and augmented Dickey-Fuller (ADF) test (Dickey an Fuller 1979; 1981) are done on each variable to check for non-stationarity. If the variables in their levels are integrated of order d, i.e., I(d) and the residual is found to be I(d') where d'< d and the variables are said to be cointegrated. Second an error - correction specification is estimated.

3.5 Error Correction Model

The Granger representation theorem (1981) proves that, if a cointegrating relationship exists among a set of I (1) series or stationarity of the data after first differencing, then a dynamic error-correction model of the data also exists.

According to Granger (1981) if the variables in Equation 3.10 are found to be cointegrated, a more general error correction mechanism (ECM) model should be used to model dynamic relationships. The residual from the estimated cointegrating equation in stage one is then included as an error correction term in estimating the ECM model. The error correction model specification for the model base on Equation 3.10 represented by the following equation:

$$X_t = \alpha_0 + \sum_{i=1}^{m} \beta_1 Y_{t-i} + \sum_{i=1}^{n} \beta_2 P_{t-i} + \sum_{i=1}^{q} \beta_3 V_{t-i} + \eta U_{t-1} + \varepsilon_t \qquad (3.25)$$

where:

$$U_t = X_t - \alpha - \alpha_1 Y_t - \alpha_2 P_t - \alpha_3 V_t \qquad (3.26)$$

Where U_{t-1} is the lagged error correction term of the residual from the cointegration regression equation as follows:

$$X_t = \alpha_0 + \sum_{i=1}^{m} \alpha_1 Y_{t-i} + \sum_{i=1}^{n} \alpha_2 P_{t-i} + \sum_{i=1}^{q} \alpha_3 V_{t-i} + \varepsilon_t \qquad (3.27)$$

Since the short-run and the long-run parts of this equation should provide same estimate of long-run elasticity, Mehra (1991) imposes the restrictions that the estimates of the long-run elasticity form the long-run part of the Equation 3.26.

If the variables have a cointegration vector than $U_t \sim I(0)$ represents the deviation from equilibrium in period *t*. The Error Correction Model shows how the system converges to the long-run equilibrium implied by the cointegrating regressions. The coefficient η in Equation 3.25 represents the response of the dependent variable in each period to departures from equilibrium. If the coefficient of the error-correction term η is found to be statistically significant, it implies that there is equilibrium in the long-run relationship. This approach has, so far, been the standard practice in the cointegration literature to distinguish between short-run and long- run relationships. Banarjee et. al (1986) note that the inclusion of an error correction term in the ECM model imposes restrictions on the coefficients, find serious problems of bias in the estimated coefficients, and suggest estimation of an unrestricted ECM model with lagged level variables from the cointegrating equations as regressors.

In addition to indicating the direction of causality among the variables, the ECM approach allows one to distinguish between "short-term" and "long-term" Granger causality. When the variables are cointegrating then in the short term, deviations from this long-term equilibrium will feed back on the change in the dependent variables in order to force the movement towards the long-term equilibrium. If the dependent variable is driven directly by this long-term equilibrium error, then it is responding to this feedback. If not, it is responding only to short-term shocks to the stochastic environment. The F-tests of the 'differences' in explanatory variables give us an indication of the 'short-term' causal effects, whereas the 'long term' causal relationship is implied through the significance of the t-test(s) of the lagged error correction term(s). This contain(s) the long-term information that is derived from the lagged error-correction term. However, the error correction term (μ_{t-1}) is a short-term adjustment coefficient and represents the proportion by which the long-term disequilibrium *ith* dependent variable in each short period.

3.6 Structural Break and Unit Root

It is well recognized that many economic time series have undergone structural breaks, due to economic crises, changes in institutional arrangements, wars, etc. These breaks have expressed themselves as alterations either in the level or in the trend of the series. Hence, in light of these breaks, and given the restriction of a constant trend function implied by a trend stationary model, there is a need to consider a breaking trend stationary process as a more realistic specification to the model. Deterministically, the model needs to include the nonfixed structure of the trend function of an economic time series in the model of structural break.

Perron (1989) first considered a structural change in a deterministic trend in a unit root test. He showed theoretically that if the data generating process has a kink or a jump in the deterministic trend, the unit root test might have a bias for accepting the null hypothesis of a unit root. Also it conduct a unit root test on the model that assumes a change in deterministic trends, using long-run data involving kinks in the data such as the great depression and the oil crisis.

One major drawback of unit root tests is an implicit assumption that the deterministic trend is correctly specified. Perron (1989) argued that if there were a break in the deterministic trend, then unit root tests could lead to a misleading conclusion that there is not unit root. He developed a methodology to test for a structural break in apparent non-stationary series that enables one to use the complete sample period rather than splitting the sample into two parts. In the Perron methodology, the null hypothesis is the presence of a unit root against the alternative that the series is trending stationary. In his test, he allowed for more than one time break in the level and or slope of the trend under both the null and alternative hypothesis.

If the deterministic component is misspecified, inference from unit root tests will be misleading because the detrended series will not be purely stochastic and will depend on some nuisance parameters. A misspecified trend function heavily distorts the test results. It is well known that many economic time series display heterogeneous behavior in their deterministic component. This heterogeneous behavior is, in large part, the result of unusual events such as the Great Depression and the first oil price shock. Radical policy change also produces heterogeneous behavior in the deterministic component of many economic time series. It has been observed that this heterogeneous behavior displays itself in the form of level shifts, trend shifts, or both. A level shift corresponds to a change in the mean of the series and a trend shift corresponds to a change in the growth rate of the series. This type of behavior is described as a variable trend by Stock and Walson (1988), and breaking trends by Perron (1989).

If a series undergoes a shift in its deterministic component, traditional ways of detrending, in addition to the cyclical component, will produce residuals that display non-stationary behavior. In the absence of shifts in the trend function, traditional detrending would produce residuals that are purely cyclical. Therefore, the shifting trend stationary (STS) models with finitely many shifts in the trend function are in sharp contrast with the trend stationary (TS) models, and approximate better the behavior of non-stationary series. Now, consider a unit root process. Such a process is potentially capable of producing a finite member of shifts in the level and growth rate of the series since the innovations have permanent impacts. Each shock is potentially capable of shifting the level and growth rate of the series indefinitely, with complicated patterns if we also allow innovations to be weakly dependent and heterogeneously distributed.

The argument put forward by Perron (1989) is that is that the level and trend shifts or a combination for the two, exogenously occurring at a given date cause, the Dickey-Fuller unit root test statistics to incorrectly fail to reject the null hypothesis of a unit root. This approach is very much along the lines of the intervention analysis of Box and Tiao (1975). The level and trend shifts can be modeled as exogenous shocks.

Intervention analysis can be used to detect the effects of policy changes. The effects of any kind of exogenous events occurring at known dates, such as the definitional change of a monetary aggregate, which is a level shift in nature, or events occurring at unknown dates that manifest their effects as outliers can be modeled as changes in the deterministic component of the series.

A weakly stationary process has a mean and variance, which do not change over time. If the mean of a series undergoes a drop due to a sudden economic crash then the above form of stationarity will cease to hold, since the level of the series will be different after the break takes place. A similar argument holds for a change in the growth rate of the series, reflected in a shifting trend. It will refer to the shifts in level and trend as shifts in the trend function of a series, where the deterministic function will be composed of an intercept and linear trend. Therefore, using the traditional method of unit a root time series will produce a residual displaying, apart from a cyclical component, non-stationary behavior if the shifts in its trend function will be called a breaking trend stationary process.

Perron (1989) said for testing in the presence of a unit root in time series data against the hypothesis of stationary fluctuation around a deterministic trend function, the use of a long span of data has definite advantages. It allows tests with larger power compared to using a smaller span, in most cases even if the latter allow more observation. The data set with large span has more change to include a major event which one would rather consider as an outlier or as exogenous given its relative importance. Therefore, it is considered as a relevant alternative for a trend function with a change in the intercept and slope.

To assess the effects of the presence of a shift in the intercept or a shift in the slope (as single point of time) on tests for the presence of a unit root, Perron (1989) first performed a Monte Carlo experiment. The Monte Carlo results show that if the magnitudes of the shifts are significant, one could hardly reject the unit root hypothesis, even if the series is stationary with a broken trend and identically independent distribution (iid) disturbances. Perron extended the Dickey - Fuller testing strategy to ensure a consistent testing procedure against shifting trend functions. He expanded it to include detrending the series first, then analyze the behavior of the estimated residuals.

Perron applied the modified Dickey - Fuller test for the same US macroeconomic series used by Nelson and Plosser (1982) and found the quite strikingly different result that the unit root hypothesis can be rejected. Perron's procedure is a conditional test given a known break point. This assumption of a known break date (treated as an exogenous event) raised the problem of pre-testing and data–mining regarding the choice of the break date. After Perron (1989), several methods have been developed for endogenizing the choice of a break point into testing procedures. These procedures incorporate the estimation of a break point and use recursive method (using sub samples) or sequential methods (using full sample with dummies).

Perron also said the estimation model about the change in the trend function is an important avenue of future research. As the research by Hamilton (1987) and Lam (1988) explained where the slope of the trend function is allowed to take two different values and the changes are modeled as a binomial process. In fact, any test for the presence of a unit root against trend-stationary alternatives is subject to another type of observational equivalence, as recently argued by Cochrane (1987) and Blough (1988).

Perron (1989) concluded it might be more advantageous to adopt the trend-stationary view with breaks and detrend the series accordingly prior to analyzing the remaining noise.

Rappoport and Reichlin (1989) explained that economists are accustomed to attributing changes in trend rates of growth to events that occur infrequently. If the events are the source of permanent shocks in the date, then a segmented trend captures their effects better than a difference stationary model. In addition, macroeconomic time series are found to undergo in infrequent structural change, rather than follow difference stationary processes. Using a segmented trend model immediately raises the problem of selecting the dates at which the trend changes.

Hendry and Nealy (1991) showed that inference on the existence of a unit root is affected by structural change (the unit root tests tend to under reject the null of a unit root), in the same for cointegration. However, considering cointegrated relationships one has to distinguish between breaks in the relationships, and breaks in the individual variable. In the latter case, there is the problem that the dates of the breaks in the different variables may not coincide (Hendry, 1996).

Muro (1993) studied the effects of structural breaks on unit root test. He said the main point is the difficulty to distinguish between a random walk (difference stationary process) and a stationary model with structural breaks (break trend stationary process). This means that in general, it is possible to misspecify a break trend stationary model as an integrated process. Moreover, it has been proven that the difference stationary specification is the default model: it will appear to fit the data best if competing models are not adequately parameterized (Rappoport and Reichlin, 1987). This means that, when choosing between difference stationary and trend stationary models when there is a trend break in the stationery process, and then the ADF test will tend to favor the difference stationary model finding a spurious unit root.

Quintos (1993) also tested or structural breaks and a change in cointegration to determine whether a model with shifts is appropriate. He treated the breakpoints as known since tests conducted in this manner are known to have higher power than the mean tests for unknown break points used by Haug (1992).

Gregory, Nason, and Watt (1996) find the sensitivity of the ADF test for cointegration in the presence of a single break in studied Monte Carlo results show that the rejection frequency of the ADF test decreases substantially. That is, in the presence of a break, the ADF tends to under reject the null of no cointegration. The under rejection is similar to the under rejection of the null in the case of unit root tests. However, in this case, the under rejection of the null indicates correctly that the constant parameter cointegration relation is not appropriate.

3.7 Effect of Structural Break on Cointegration Tests.

As noted earlier, by Rappoport and Reichelin (1989), Hendry and Nealy (1991), and Perron (1989) show that inference on unit roots is affected by structural change (the unit root tests tend to under reject the null of a unit root). The same is the case with the test for cointegration. However, when considering cointegrated a relationship one has to distinguish between breaks in the relationships and breaks in the individual variable. In the latter case, there is the problem that the dates of the breaks in the different variables may not coincide (Hendry, 1996).

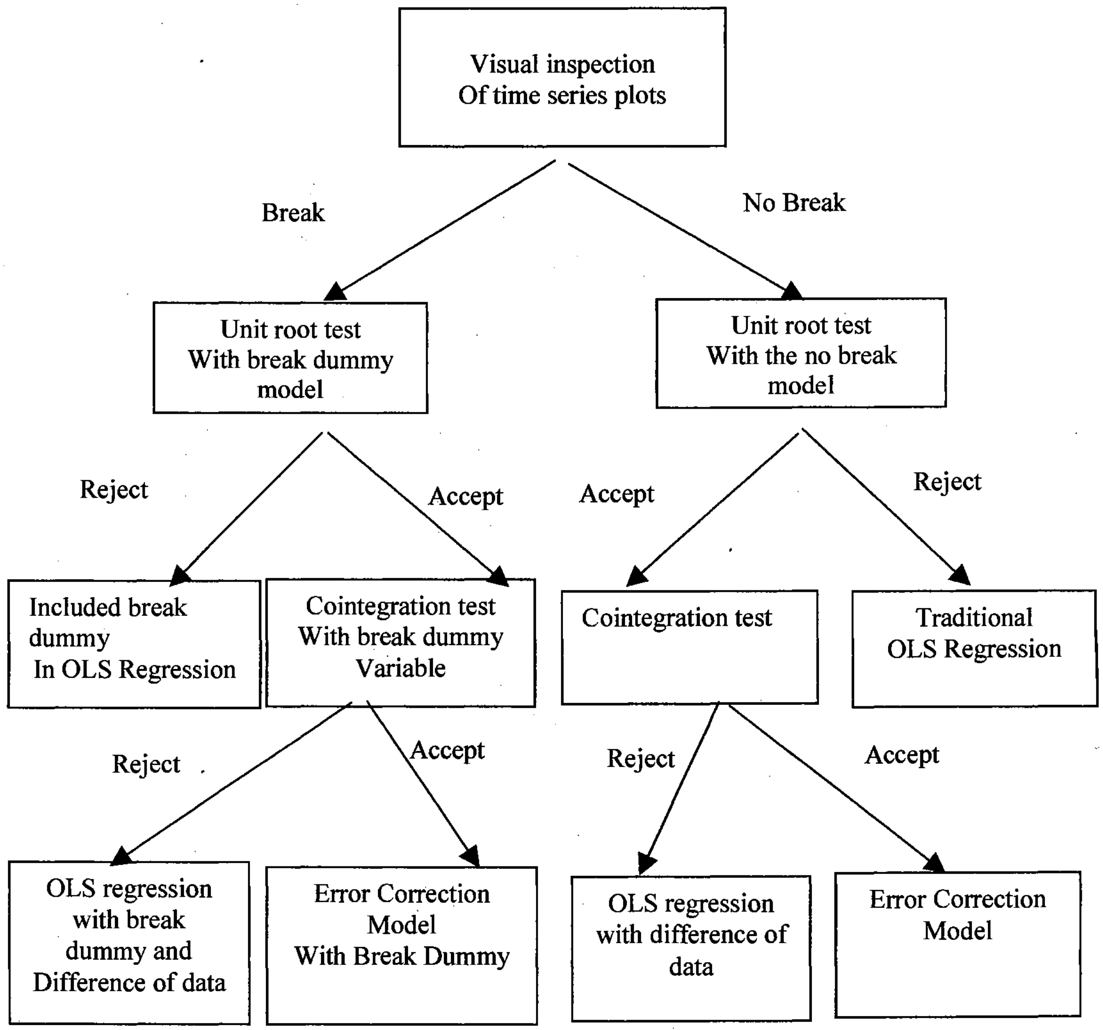
Gregory, Nason, and Watt (1996) studied the sensitivity of the ADF test for cointegration in the presence of a single break. Their Monte Carlo results show that the rejection frequency of the ADF test decreases substantially. That is, in the presence of a break, the ADF test tends to under reject the null of no cointegration. However, in this case the under rejections of the null hypothesis indicate correctly that the constant parameter cointegration relation is not appropriate.

Campos, Ericsson, and Hendry (1996) investigate the properties of several cointegration tests when the marginal process of one of the variables is stationary with a structural break. They find that the break has little effect of the test size. However, the test based on the ECM is more powerful than the Engle-Granger two-step procedure employing the DF unit root test.

Soejima (1996) applied Hansen's (1992) stability test for cointegration and examines the validity of a linear deterministic trend. In the case of variables with linear deterministic trends without structural change, their linear combination becomes stationary around a constant if deterministic cointegration holds. However, in the case of variables with a deterministic trend with structural change, even if cointegration exists between their stochastic trends, their linear combination based on the cointegrating vector does not necessarily have a constant or linear deterministic trend. If the pattern and timing of structural change do not coincide, the time series will exhibit a shift in the constant or trend term. Hansen's (1992) stability test for the cointegrating vector can check the stability of the constant and trend parameters, using a test for structural change in the deterministic trend.

Hansen's stability test involves three tests. First, Hansen tests the null hypothesis of no changes in the parameters (including the cointegrating vector) during the sample period against an alternative hypothesis of a shift in the parameter at an unknown date. This is appropriate for finding the incidence of a sudden structural change. The other two tests assume that each parameter follows a stochastic process and the test the null hypothesis of zero variance in the parameters (constant parameters).

Figure 3. 2 The Flowchart of Diagnostics for Time Series Regressions



CHAPTER IV

METHODS AND PROCEDURES

This chapter identifies data sources and variables constructed in this study. The hypotheses tested in this study are that the real exchange rate has a significant effect on agricultural exports. Equations for agricultural exports and the real exchange rate are specified. Non-stationary and unit root tests, non-stationary and unit root tests with structural change, and cointegration tests are conducted. Error correction models are estimated.

The export models are specified for five Indonesian commodities (coffee, cocoa, palm oil, rubber, and tea). As well as aggregate exports. For these commodities, exporters are assumed to be price takers on the world market because Indonesia has a small share of world trade. The export quantity reflects the equilibrium condition between the domestic and foreign markets. The foreign demand for Indonesia's exports is hypothesized to be a function of the current or lagged (1) real income of foreign (importing) countries, (2) price of Indonesia's agricultural exports, and (3) real exchange rate between Indonesia and its trading partners, and (4) price of each commodity from competing countries. Other things equal, the higher the level of foreign real income, the larger is foreign demand for Indonesia's agricultural exports. On the other hand, the higher the price of Indonesia's exports, other things equal, the smaller is the demand quantity for Indonesia's agricultural exports. The higher the real exchange rate, the lower the demand for Indonesia's agricultural exports.

In each commodity and aggregate trade model, export equations are developed separately to yield a more meaningful policy analysis of the effect of real exchange rates on the volume of exports. By assuming infinite supply elasticity, Indonesian agricultural export demand from the rest of the world is reduced to a single equation. As such, the export model for each commodity can be presented as,

$$X_t = f(Y_t, P_t, V_t, PR_t, \varepsilon_t) \tag{4. 1}$$

where

X_t = Volume of Indonesia agricultural export in metric tons.

Y_t= Real foreign income (countries imported agricultural product from Indonesia) as calculated in Equation 3.7.

P_t= Real Indonesian agricultural commodities price expressed in \$ metric ton. It measured by the unit value of Indonesia's agricultural exports price deflated by Indonesia CPI (1995=100)(Chamber and Just,1981).

V_t= Agricultural trade-weight exchange rate index of the Rupiah versus the currencies of agricultural importing countries (1995=100), as calculated in Equation 3.4.

PR_t= Price of competitive (relative) price express in \$ per metric ton deflated by its countries CPI (1995=100).

ε_t= Error term.

4.1 Data Sources.

Wholesale price indices, export quantity, and prices of exports and world prices for rubber, tea, coffee, cocoa and palm oil were obtained from the bulletin of Quarterly Statistics for Asia and Pacific, FAO trade year book, Quarterly Bulletin of Cocoa Statistics, and also from Indonesia's Central Bureau of Statistics. The exchange rate data is taken from the Internet through the Bank Indonesia exchange rate home page. GDP, GNP deflator, CPI, trade balance, money supply, budget deficit, export price index, export volume index, and population are available from the Main Economic Indicators and International Financial Statistic International Monetary Fund CD- ROM.

4.2 Model Specification

The long-run equilibrium relationship between Indonesia's real export volume to importing countries, the real activity of importing countries, the bilateral real exchange rate weight between importing countries and Indonesia, the real Indonesia agricultural export price, the real relative price is specified in first difference of natural log. It is written as:

$$\begin{aligned}\Delta \ln X_t = \alpha_0 &+ \sum_{i=1}^{m} \alpha_{1i} \Delta \ln Y_{t-i} + \sum_{j=1}^{n} \alpha_{2j} \Delta \ln P_{t-i} + \sum_{k=1}^{q} \alpha_{3k} \Delta \ln V_{t-i} \\ &+ \sum_{l=1}^{r} \alpha_{4l} \Delta \ln PR_{t-i} + U_t \end{aligned} \qquad (4.2)$$

Where

X_t = Volume of Indonesia agricultural export in metric tons.

Y_t = Real foreign income for countries imported agricultural product from Indonesia as calculated in Equation 3.7.

P_t = Real Indonesia agricultural commodities price expressed in $ metric ton. it measured by the unit value of Indonesia's agricultural exports price deflated by Indonesia CPI (1995=100) (Chamber and Just,1981).

V_t = Agricultural trade-weighted exchange rate index of the Rupiah versus the currencies of agricultural importing countries (1995=100), as calculated in Equation 3.4.

PR_t= Price of competitive (relative) price express in $ per metric ton deflated by its countries CPI (1995=100).

U_t = Error term.

Equation 4.2 can be derived as a long-run solution of the demand function for exports (Chowdhury (1993)). Since the higher real income in the importing countries lead to higher imports, it is expected $\alpha_{1i} > 0$.

The prices in the equation are deflated with each countries CPI. The relative prices are used to reduced multicollinearity and thereby decrease standard errors (Konandreas, Bushnell, and Green). Double logarithmic functional form is used where the coefficients are elasticities.

4.3 The Real Exchange Rate Weight Calculation

The short-run elasticity of exports is determined with respect to world real level of economic activity, real price, and the real exchange rate. The long-run elasticity can be calculated by combining the short-run elasticity with the lag coefficients for each independent variable.

The exchange rate variable used in the model is a real effective exchange rate where the weight is used constant constructed form the data from 1971 to 1998. The weight is equal to each country's total share from Indonesia agricultural exports. The seven most important trading partners (Japan, USA, Germany, Singapore, Britain, Netherlands, and Australia) are included. The real exchange rate weight is calculated as follows:

$$V_{jt} = \sum_{t=1}^{p} W_{jt} \frac{E_{jt} P_{jt}}{PD_{jt}} \tag{4. 3}$$

Where W_{jt} is the relevant weight which sums to unity, and is import shares of seven major trading partners constructed by Warr (1984). E_{jt} is the nominal exchange rate between Indonesia and each of its trading partners (foreign currency/rupiah). P_{jt} is refers to the *CPI* of each of Indonesia's major trading partners, and PD_{jt} is Indonesia's Consumer Price Index.

The inclusion of the real exchange rate weight as a separate regressor is based on Orcutt's argument that the market reacts more quickly to exchange rate changes than to price changes. Furthermore, exchange rate changes are usually larger than price fluctuations in the short run (Chambers and Just, 1979). This approach allows for estimation of changes in exports that arise directly from either exchange rate movement or from the real price movement in the exporting country.

From the Equation 4.2, the coefficient of V_t may be positive or negative depending on whether supply or demand response is greatest. As the real level of world economic activity improves, demand for agricultural commodities increases. This would increase the export quantity from an exporting country. The higher (appreciate) the real exchange rate (*RER*) the lower the export volume.

4.2 Non-Stationarity and Unit Root Test

Among the many tests for unit roots available, the most widely used is the one proposed by Dickey and Fuller (1979) and extended by Said and Dickey (1984).

The stationarity of the data need to be determined, because most of the time series data is non-stationary. The unit root test was applied to see if the data has a unit root (non-stationary) or does not have a unit root (stationary). The Dickey - Fuller (DF) and Augmented Dickey-Fuller (ADF) test with and without structural change is utilized to test for the stationarity of the data.

The most commonly used tests of the null hypothesis of unit root in an observed time series are derivatives of the Dickey- Fuller (DF). Engle and Granger (1987) suggest the following Dickey - Fuller test of stationarity:

$$X_t = \alpha + \rho X_t + \varepsilon_t \tag{4. 4}$$

Substract X_{t-1} from Equation 4.4 from both sides of the equation

$$\Delta X_t = \alpha + \beta_1 X_{t-1} + \varepsilon_t \tag{4. 5}$$

Where X_t is represent the Volume of agricultural commodities export, $\Delta X_{t-j} = X_{t-j} - X_{t-j-1}$ (first difference) and coefficient $\rho = 1$ in Equation 4.4 if there is a unit root. In principle, a test of hypothesis $\rho = 1$ in Equation .4 can be done by test $\beta_1 = 0$ in the Equation 4.5, since $\beta_1 = (\rho - 1)$ from ρ in Equation 4.4. With the formulation in Equation 4.5, the Dickey-Fuller test for a unit root is carried out by testing the hypothesis that $\beta_1 = 0$. The standard t statistic is refereed to the Dickey - Fuller table. One cannot, however use the usual t test to test $\beta_1 = 0$ in the Equation 4.5 because under the null hypothesis, X_t is I(1), and hence the t statistic does not have an asymptotic normal distribution.

If the model using the Dickey - Fuller test by Dickey (1976), Fuller (1976) and Dickey and Fuller (1979) which include time trends with no autoregressive of the X_t,, the equation will be:

$$X_t = \alpha + \rho_1 X_{t-1} + \rho_2 t + \varepsilon_t \tag{4. 6}$$

The null hypothesis is

$$H_0: \rho_1 = 1, \qquad Ha: \rho_1 < 1 \tag{4. 7}$$

The test statistic for the unit root is given by:

$$\hat{\tau}_\tau = \frac{\hat{\rho}_1 - 1}{SE(\hat{\rho}_1)} \tag{4. 8}$$

The null hypothesis of a unit root is rejected if the value of the t-statistic of ρ_1 is negative and below the critical value presented by Dickey-Fuller. The t statistic for $\hat{\rho}_1 - 1 / SE(\hat{\rho}_1)$ is not asymptotically normal or symmetric. Tables of critical values tabulated by D. A. Dickey reported in Fuller (1976).

The OLS F test of the joint null that statistic, $\rho_1 = 1$ and $\rho_2 = 0$ can be estimated to check the if the Augmented Dickey - Fuller with time trend test is also consistent with the unit root specification, using the Dickey - Fuller critical values for OLS F statistic (Dickey and Fuller (1981)).

For the Augmented Dickey-Fuller of stationary unit roots tests developed by Dickey (1976), Fuller (1976), and Dickey and Fuller (1979) include time trends and the autoregressive of change in X_t. The equation is:

$$\Delta X_t = \alpha + \beta_1 X_{t-1} + \beta_2 t + \sum_{j=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t \tag{4. 9}$$

The number of autoregressive (AR) lag p in the Equation 4.9 is calculated using AIC (Akaike Information Criterion) in the SAS package program. The AIC is used to determine the autoregressive order, which essentially contains all the information relevant for prediction of future values of the time series.

The AIC is calculated as,

$$AIC = -2\ln(L) + 2k \tag{4. 10}$$

where L is the value of the likelihood function evaluated at the parameter estimates.

k is the number of estimated parameters. (Judge et. al,1985).

The null hypothesis

$$H_0: \beta_1 = 0, \qquad Ha: \beta_1 < 0. \tag{4. 11}$$

The test statistic for the unit root is given by:

$$\hat{\tau}_\tau = \frac{\hat{\beta}_1}{SE(\hat{\beta}_1)} \tag{4. 12}$$

where β_1 is the Dickey - Fuller level of variables, t is the time trends.

The ADF test of a unit root corresponds to the null hypothesis that $\beta_1 = 0$ in Equation 4.9. The ratio of the estimate of β_1 to its standard error is pseudo t-statistic. ADF test is based on testing the hypothesis $\beta_1 = 0$ under the assumption that ε_t is white noise error. The value of the t- test needs to be absolutely be greater than the Dickey - Fuller critical value of the OLS t- statistic (Fuller (1976)).

The OLS F- test of the joint null that statistic $\beta_1 = 0$ and $\beta_2 = 0$ can be estimated to check the if Augmented Dickey - Fuller F-test is also consistent with the unit root specification. If the null hypotheses that the commodities have a unit root are rejected, the cointegration on linear combinations of the variable series can be pursued.

4.3 Cointegration Tests

As mentioned earlier in the previous chapter if a linear combination of the two data series of I (1) is stationary or I(0) then the variables are said to be cointegrated.

The Engle and Granger (1987) test is a popular way to test whether variables are cointegrated. They suggest that as a starting point for a unit root test between exchange rate and export, one can start by modeling the static relationship between the two series, estimates of "cointegration regression as":

$$\Delta \ln X_t = \alpha_0 + \sum_{i=1}^{m} \alpha_{1i} \Delta \ln Y_{t-i} + \sum_{j=1}^{n} \alpha_{2j} \Delta \ln P_{t-i} + \sum_{k=1}^{q} \alpha_{3k} \Delta \ln V_{t-i} + \sum_{l=1}^{r} \alpha_{4l} \Delta \ln PR_{t-i} + U_t \tag{4. 13}$$

where the entire variables are as defined in Equation 4.2.

The first thing to do is to estimate the cointegration regression by ordinary least squares to obtain the residual U_t. Under the null hypothesis of no cointegration, U_t will be I(1) series. The series X_t and V_t , P_t , PR_t are cointegrated individually they are integrated of order one, denoted I(1) (the data is stationary after first differencing) and their linear combination which can be expressed as Equation 4.13 is integrated of order zero denoted I(0)(i.e. stationary).

If there exist β such that U_t in Equation 4.13 is stationary (does not contain unit root), then X_t and V_t are said to be cointegrated, and the long run relationship between export and exchange rate is

$$\Delta \ln X_t - \alpha_0 - \sum_{i=1}^{m} \alpha_{1i} \Delta \ln Y_{t-i} - \sum_{j=1}^{n} \alpha_{2j} \Delta \ln P_{t-i} - \sum_{k=1}^{q} \alpha_{3k} \Delta \ln V_{t-i} - \sum_{l=1}^{r} \alpha_{4l} \Delta \ln PR_{t-i} - U_t = 0 \tag{4. 14}$$

Therefore, the U_t term in Equation 4.13 measures the cointegrating linear relationship among the export volume and real effective exchange rates. If U_t is not stationary (contain a unit root), then X_t and V_t are not cointegrated. Each series of X_t and V_t is first checked for stationarity, by testing the null hypothesis of a unit root, using the Dickey - Fuller (DF) test. One uses the Dickey – Fuller test base on the regression

$$\Delta\hat{U}_t = \rho\hat{U}_{t-1} + \varepsilon_t \tag{4. 15}$$

If the null hypothesis of a unit root cannot be rejected, then the cointegration test on the residual of Equation 4.13 can be pursued.

The Durbin - Watson (D-W) statistic, Dickey - Fuller (DF), augmented Dickey-Fuller (ADF) is used to examine cointegration between exchange rate and export volume series. Based on the Monte Carlo studies, Engle and Yoo (1987) argued that, for a first-order system, both D-W, and DF test are appropriate approaches to test cointegration of the series of exchange rates and export volume. If the D-W statistic of the cointegration regression Equation 4.16 were significantly greater than zero, which would be its probability limit, U_t contains a unit root as required by the null hypothesis (Engle and Yoo, 1987). The ADF test on the residual with included the autoregressive variable of the error term of the cointegrating regression is:

$$\Delta\hat{U}_t = \alpha_0 + \alpha_1\hat{U}_{t-1} + \sum_{j=1}^{p}\beta_j\Delta\hat{U}_{t-j} + W_t \tag{4. 16}$$

where W_t is a white noise error term. The null hypothesis of no cointegration is rejected if absolute value of the calculated t-statistic on the coefficient for the lagged of the error term (α_1) in Equation 4.16 is greater than the absolute value of the critical value reported by Engle and Yoo (1987). The number of lags that are appropriate is determined using Akaike information criterion (AIC) criteria. If the variables are cointegrated, the error correction model will be estimated. Then the process of cointegration test without structural break is explained in Figure 4.3. The cointegration test is conducted using SAS software in SAS MACRO for % dftest.

4.4 Error Correction Model

Engle and Granger (1987) conduct the fourth test of cointegration using the ECM (error correction model) test. The test is designed to test whether the error correction terms from the cointegrating regression are significant in the error correction model.

According to Granger (1981), if the variables in Equation 4.2 are found to be cointegrated, a more general error correction mechanism (ECM) model should be used to model dynamic relationships. The residual from the estimated cointegrating equation in stage one is then included as an error correction term in estimating the ECM model. The error correction model specification for the model base on Equation 3.10 is represented by the following equation:

$$\begin{aligned}\Delta \ln X_t = \alpha_0 &+ \sum_{i=1}^{m} \beta_{1i} \Delta \ln Y_{t-i} + \sum_{j=1}^{n} \beta_{2j} \Delta \ln P_{t-i} + \sum_{k=1}^{q} \beta_{3k} \Delta \ln V_{t-i} \\ &+ \sum_{l=1}^{r} \beta_{4l} \Delta \ln PR_{t-i} + \eta U_{t-1} + \varepsilon_t \end{aligned} \tag{4. 17}$$

where

$$U_t = \Delta ln X_t - \alpha_0 - \sum_{i=1}^{m} \alpha_{1i} \Delta lnY_{t-i} - \sum_{j=1}^{n} \alpha_{2j} \Delta lnP_{t-i} - \sum_{k=1}^{q} \alpha_{3k} \Delta lnV_{t-i} - \sum_{l=1}^{r} \alpha_{4l} \Delta lnPR_{t-i} \tag{4. 18}$$

as explained in Equation 4.13.

If the variables have a cointegration vector then $U_t \sim I(0)$ represents the deviation of equilibrium in period t. The error correction model shows how the system converges to the long-run equilibrium implied by the cointegrating regressions. The coefficient η in Equation 4.17 represents the response of the dependent variable in each period to departures from equilibrium. If the coefficient of the error-correction term η is found to be statistically significant, it implies that there is equilibrium in the long-run relationship.

Equation 4.17 also represents as a demand function with error correction term that gives the short-run determinants of export demand and embodies both the short-run dynamic and the long-run relationships of the series. The presence of U_{t-1} in equation (4.17) reflects the presumption that actual exports do not adjust instantly to the long-run determinants. Therefore, in the short-run, an adjustment is made to correct for disequilibrium in the long-run export demand. The parameter η in the Equation 4.17 measures the response of the regression in each period to departures from equilibrium conditions. The ECM therefore reflects how the system converges to the long-run equilibrium implied in the Equation 4.17, with convergence being assured when η is between zero and minus one (Arize, 1996).

4.5 Test for the Unit root Under Structural Break

According to Maddala (1998) because of events like the great depression, oil price shocks, policy change, and so on, models with constant coefficients have been found to perform poorly, either for forecasting purposes or for the purpose of analyzing the effect of policy change or the exchange rate regime changes. The solution to this problem have been modeled as:

(i) Model with continuous parameter changes: these are estimated using some recursive algorithm like the Kalman filter. The problem with these models is that they do not capture sudden shifts.

(ii) Outlier models: these models argue that sudden shocks produce outliers (with temporary or permanent level shifts).

(iii) Switching regression models, with abrupt switches and gradual switches: one popular modal during recent years in this category has been the Markov switching regression (MSR) model.

There are an enormous number of statistical test to test the structural change .The tests can be conveniently classified under the categories:

(i) Known break points versus unknown break points.

(ii) Single breaks versus multiple breaks.

(iii) Univariate versus multivariate relationships.

(iv) Stationary versus non- stationary variables.

In this example, the break points are known.

4.5.1 Test of Unit Root under Structural Break with Single Known Break

The break point for a structural change is known, Perron (1989) has proposed a modified Dickey Fuller test for a unit root in three different types of deterministic trend functions. The null hypothesis considered is that a given series, Y_t (of which a sample of size T+1 is available) is a realization of a time series process characterized by the presence of a unit root and possibly a nonzero drift. The approach allows one time change in the structure occurring at ($1<T_c<T$). The time of a structural change is referred to as T_c, the period at which the change in the parameters of the trend function occurs. Three different models are considered under the null hypothesis. First, model (A), allows for a one-time change in the intercept of the trend function (drift term structural change), T_c, is the year 1987 (I) when the government of Indonesia changed the exchange rate regime from fixed to managed floating.

$$f(t) = \mu_0 + \mu_1 DU_t + \delta t \tag{4. 19}$$

The hypothesis for model (A) is parameterized

$$Y_t = \mu_0 + \mu_1 DU_t + Y_{t-1} + \delta t \tag{4. 20}$$

where

$$DU_t = \begin{cases} 1 & \text{if } t > T_c \\ 0 & \text{otherwise} \end{cases}$$

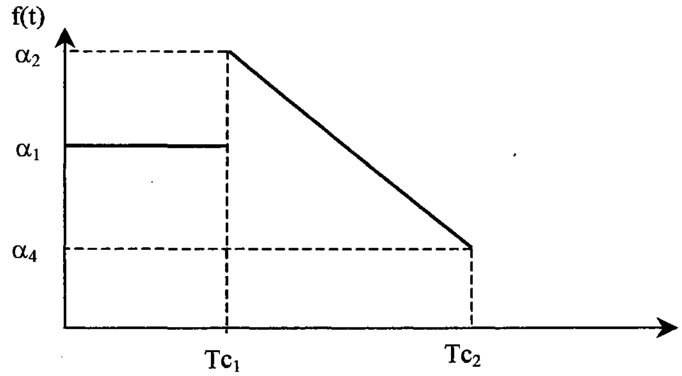


Figure 4. 1 Drift Term

The changing growth model, model (B), allowed for a change in the slope of the trend function, without any sudden change in the level at the time of the break (deterministic trend structural change).

$$f(t^*) = \mu + \delta_1 DT_t + \delta t \tag{4. 21}$$

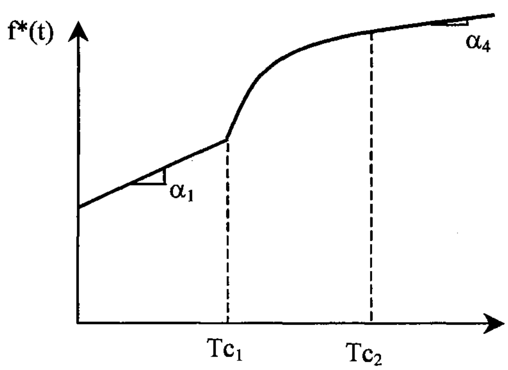
The hypothesis for model (B) is parameterized

$$Y_t = \mu_0 + \mu_1 DT_t + \mu_2 Y_{t-1} + \delta t \tag{4. 22}$$

where

$$DT_t = \begin{cases} t & if\ t > T_B \\ 0 & otherwise \end{cases}$$

Figure 4. 2 Deterministic Trend



Model (C) allows for both effects to take place simultaneously, i.e., a sudden change in the level followed by a different growth path.

$$Y_t = \mu_0 + \mu_1 DU_t + \mu_2 DT_t + \delta t \tag{4. 23}$$

Instead of considering the alternative hypotheses that Y_t is a stationary series around a deterministic linear trend with time invariant parameters, the following three alternative possible models are analyzed:

$$\text{Model(A) } Y_t = \mu_0 + \beta t + \mu_2 D(T_b)_t + \delta t \tag{4. 24}$$

$$\text{Model(B) } Y_t = \mu_0 + \beta t + \mu_2 DU_t + \delta t \tag{4. 25}$$

$$\text{Model(C) } Y_t = \mu_0 + \beta t + \mu_2 D(T_b)_t + \mu_3 DU_t + \delta t \tag{4. 26}$$

$$D(T_b)_t = \begin{cases} 1 & if\ t = T_c + 1 \\ 0 & otherwise \end{cases}$$

From the equations model A (change in intercept) and model B (change in the slope) De Jong (1996) starts with Perron's model. However, he considers a third-order autoregression can be expressed as:

$$Y_t = \mu_0 + \mu_1 DU_t + \mu_2 T + \mu_3 DT_t + \mu_4 D(T_b)_t + \beta Y_{t-1} + \sum_{i=1}^{p} \gamma_i \Delta Y_{t-1} + V_t \tag{4.27}$$

where DU_t and DT_t are dummies for the breaks in the intercept and slope coefficient. With this model, the unit root hypothesis and trend stationary hypothesis can be expressed in H_0 and H_1:

$$\begin{aligned} H_0 : &\quad \beta = 1, (\mu_2, \mu_3, \mu_4) = (0,0,0) \\ H_1 : &\quad -1 < \beta < 1, (\mu_2, \mu_3, \mu_4) \neq (0,0,0) \end{aligned} \tag{4.28}$$

The null hypothesis of a unit root is different, since the deterministic trend function included dummy variables (DU_t , $DT_{t,}$, $D(T_b)_t$.) The alternative hypothesis is a broken-trend stationary system, which also incorporates the same dummy variables. Under the null hypothesis there is a restriction of β=1 and $\mu_2 = \mu_3 = \mu_4 = 0$ whereas under the alternative hypothesis of a trend stationary process it is that expected that β is less than one, μ_1, μ_2 , μ_3, may be non zero and μ_4 is close to zero. To test the presence of the unit root using the *t*-statistic from the Equation 4.27 requires the critical values established by Perron (1989). Perron is able to show that the normalized bias $T(\hat{\beta} - 1)$ has a probability limit that series with $\lambda = (T_c/T)$. Given this result, procedures that allow the unit root hypothesis to be tested, in the presence of a structural break, are clearly desirable.

4.6 Test for Contegration with Structural Break

Wright (1993) extends the test to non-stationary trended variable and to integrated variables. Hoa and Inder (1996) extended the test for non-stationary regressor and since the test does not explicitly specify the nature of the alternative, they suggest its use as diagnostic test for structural change. Hoa and Inder derive the asymptotic distribution of the fully modified OLS test statistic, tabulate the critical values, and show that the test has non-trivial local power irrespective of the particular type or structural change. They tabulate the asymptotic critical values for the two models.

Model 1 (M_1) $$y_t = \alpha_t + \beta_t x_t + u_t \qquad (4.29)$$

$$x_t = x_{t-1} + v_t, \qquad t=1,2,\ldots,T \qquad (4.30)$$

Model 2 (M_2) $$y_t = \alpha_t + \beta_t x_t + u_t \qquad (4.31)$$

$$x_t = \mu + x_{t-1} + v_t, \qquad t=1,2,\ldots,T \qquad (4.32)$$

In M_1, x_t is I (1) without drift, M_2, x_t is I (1) with drift. The asymptotic critical values are shown in Table 4.1. The bootstrap-based small sample of critical values can be computed and compared with these asymptotic critical values.

Soejima (1996) introduces the cointegration test of Kunitomo (1995), which incorporates structural changes in the exogenous variable of a multivariate time series model. The model can be expressed by using dummy variables as exogenous variables. The model is expressed;

$$\begin{aligned} \Delta \ln X_t &= \alpha_0 + \sum_{i=1}^{m} \alpha_{1i} \Delta \ln Y_{t-i} + \sum_{j=1}^{n} \alpha_{2j} \Delta \ln P_{t-i} + \sum_{k=1}^{q} \alpha_{3k} \Delta \ln V_{t-i} \\ &\quad + \delta_1 DU_t + \delta_2 DT_t \sum_{l=1}^{r} \alpha_{4l} \Delta \ln PR_{t-i} + U_t \end{aligned} \tag{4. 33}$$

Similar to the unit root test of Dickey - Fuller, the error term is tested using the Augmented Dickey - Fuller test to see if there is a long-run relationship between variable X_t and all other exogenous variables. The test for cointegration with structural break is explained in Figure 4.4. The cointegration test is expressed as,

$$\Delta \hat{U}_t = \alpha_0 + \alpha_1 \hat{U}_{t-1} + \sum_{j=1}^{p} \beta_j \Delta \hat{U}_{t-j} + W_t \tag{4. 34}$$

The lag length number of p is estimated using the Akaike Information Criterion (AIC) from the Statespace procedure in SAS.

Figure 4. 3 Process of Cointegration Test Without Structural Break

Determine if the data series
Are stationary
Using ADF

$$\Delta X_t = \alpha + \beta_1 X_{t-1} + \beta_2 t + \sum_{i=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t$$

↓

Do Regression to obtain
The Residual

$$X_t = \alpha_0 + \sum_{i=1}^{m} \alpha_1 Y_{t-i} + \sum_{i=1}^{n} \alpha_2 P_{t-i} + \sum_{i=1}^{q} \alpha_3 V_{t-i} + \varepsilon$$

$$U_t = X_t - \alpha - \alpha_1 Y_t - \alpha_2 P_t - \alpha_3 V_t$$

↓

Determine if residual
Does not have unit root I(0)
(Cointegration test)

$$\Delta \hat{U}_t = \alpha_0 + \alpha_1 \hat{U}_{t-1} + \sum_{j=1}^{p} \beta_j \Delta \hat{U}_{t-j} + W_t$$

↓

Do ECM to See LR Relationship

$$\Delta X_t = \alpha_0 + \sum_{i=1}^{m} \beta_i \Delta Y_{t-i} + \sum_{i=1}^{n} \chi_i \Delta P_{t-i} + \sum_{i=1}^{q} \delta_i \Delta V_{t-i} + \eta U_{t-1} + e_t$$

Figure 4. 4 Process of Cointegration Test With Structural Break

Determine if the data series with break
Are stationary
Using ADF

$$Y_t = \mu_0 + \mu_1 DU_t + \mu_2 t + \mu_3 DT_t + \mu_4 D(T_b)_t + \beta Y_{t-1} + \sum_{i=1}^{p} \Delta Y_{t-1} + V_t$$

↓

Do Regression with structural break
to obtain
The Residual

$$\Delta ln X_t = \alpha + \beta_1 \Delta ln Y_t + \beta_2 \Delta ln P_t + \beta_3 \Delta ln V_t + \delta_1 DU_t + \delta_2 DT_t + \mu_t$$

↓

Determine if residual
Does not have unit root I(0)
(Cointegration test)

$$\Delta \hat{U}_t = \alpha_0 + \alpha_1 \hat{U}_{t-1} + \sum_{j=1}^{p} \beta_j \Delta \hat{U}_{t-j} + W_t$$

↓

Do ECM With break dummy to See LR
Relationship

$$\Delta ln X_t = \alpha_0 + \sum_{i=1}^{m} \beta_1 \Delta lnY_{t-i} + \sum_{i=1}^{n} \beta_2 \Delta lnP_{t-i} + \sum_{i=1}^{q} \beta_3 \Delta lnV_{t-i} + \partial_1 DU_t + \partial_2 DT_t + \eta U_{t-1} + \varepsilon_t$$

CHAPTER V

EMPIRICAL RESULTS

This chapter reports the results of estimating the impact of exchange rate on Indonesian agricultural exports for each commodity (cocoa, coffee, palm oil, rubber, and tea) and aggregate agricultural exports. The sample period is from 1971:I to 1998:IV. The starting point of the sample period corresponds to the time of the new economic and sociopolitical system ("Orde Baru") under President Suharto's leadership. In 1971, Indonesia adopted the fixed exchange rate regime.

For each commodity an error-correction model for exports is developed. However, a prerequisite for developing a model is first to test the stationarity of the variables (unit root test) and then to determine whether a long run relationship among the variables exists (cointegration test). The unit root tests procedures developed by Dickey and Fuller (1981) with no structural break (Equation 4.9) and by Perron (1989) with a structural break (Equation 4.27) are employed. To determine whether a long run relationship among variable exists, the cointegration test constructed by Engle and Granger (1987) is used with and without structural breaks for changes in exchange rate policy.

Implementation of the Augmented Dickey-Fuller and also Philips-Perron procedure requires the determination of a lag length for the VAR (Vector Autoregression) for the unit root and cointegration models for each variable in each commodity. The order of integration for the variables entering each of the VAR models must also be determined. The order of integration of the individual time series are determined using the Augmented Dickey-Fuller test. Following Lutkepohl (1982), Akaike's Information Criterion (AIC) is used to determine the lag length for the VAR model. The optimum lag length needs to be determined in order to have the test statistics with higher power (Perron, 1989).

5.1 Unit Root Test.

Unit roots tests of stationarity are presented in Table 5.1 to 5.12. In all cases, the variables were tested using both the Augmented Dickey-Fuller test for data without a structural break and the Philips-Perron test for the data with structural breaks. These results are based on the critical t value for $\beta_1=0$ from both Equation 5.1 for the unit root test with out break with time trend and the critical t value of $\beta=1$ Equation 5.2 below for the unit root test with structural break as modified by Perron (1989).

$$\Delta X_t = \alpha + \beta_1 X_{t-1} + \beta_2 t + \sum_{j=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t \tag{5. 1}$$

$$Y_t = \mu_0 + \mu_1 DU_t + \mu_2 t + \mu_3 DT_t + \mu_4 D(T_b)_t + \beta Y_{t-1} + \sum_{i=1}^{p} \gamma \Delta Y_{t-i} + V_t \tag{5. 2}$$

Both of the models are to test for the null hypotheses if there is a unit root. To reject the null hypothesis at any given confidence interval, the test statistic observed must be greater in absolute value than the critical test statistic. The t value is called pseudo t

value which is compared to MacKinnon critical values from Fuller (1976) for unit root with out structural break and Philips-Perron critical values for a unit root with structural break. In Philips-Perron tests the division between the break time (T_c) and the whole observation (T), which is $\lambda = T_c / T$ must be employed. The observed test statistics for MacKinnon and Perron critical values are reported in this chapter. The calculation of *t*-value in Equation 5.2 is:

$$\hat{\tau}_{\tau} = \frac{\hat{\beta} - 1}{SE(\hat{\beta})} \tag{5.3}$$

Many measures of exchange rates are non-stationary variables. What is highly uncertain is whether trade flows of individual agricultural commodities are unit root processes as well (Parrish, 1999). If so, there might be a relationship that could be represented as a cointegrating regression and error-correction model. All five individual commodities and aggregate agricultural exports for Indonesia were tested for a unit root.

If the series of all variables are stationary after first differencing of either levels or log levels for both unit root tests with structural breaks or without structural breaks, cointegration tests are conducted. The cointegrating critical values corresponding to the dominant long-run relationship are reported. If the variables are cointegrated, the short-run dynamic interactions between those variables in the model are also examined using the error correction model.

5.1.1 Unit Root Results

All series are from form January 1971 to December 1998 and are tested for a unit root. Cocoa results are shown in Tables 5.1-5.2 are the Perron (1989) critical values for unit root with structural break and ADF critical values without a structural break. The λ value in Table 5.1 is the division of T_c number of observation before the break and T, number of total observations. The exchange rate policy changed in 1987:I. Lag length are selected using the Akaike information criterion and shown in the parentheses.

For cocoa most variables are stationary after first differencing both in level of logarithmic data. The variables are volume of cocoa export, for foreign income, for price of cocoa, for exchange rate weight of trading partners countries, and the prices of competitive cocoa exporters, and world cocoa production. The results show that most all the variable are stationary after the first difference, except for the log of the Ghana coca price with structural break. Variables that are I(0) are dropped from the cointegration model since it will not be possible to estimate the cointegration regression unless all the series included are integrated of the same order (Sukar, 1998).

The coffee variables results are reported in Tables 5.3-5.4. Sugar prices with linear and logarithmic transformation were found to be I(0) are dropped from the model.

Unit results for palm oil are reported in Tables 5.5-5.6. The null hypothesis of the root is rejected at the 1-10 percent level for first difference and log of first difference for tests with a structural break or without a structural break. Variable that are I(0) *log wpp* and *log dpc* with structural break, and , *dpc* and *logdpc* without structural break are deleted for the cointegration and error correction models.

Tables 5.7 and 5.8 are the tables for rubber with and without a structural break. The results for unit root with structural break shows that only *logwpd* is I(0). Results without a structural break indicate that six variables(*dpir, dpmr, dptr, log pir*, *log dpmr, log dptr*) are I(0).

Tables 5.9 and 5.10 show similar results for tea. The variables that are I(0) are, *wpp*, and log *wpp,* and they are not to be used in the cointegration test.

Tables 5.11 and 5.12 show similar results for aggregate agricultural exports. All variables in the structural break unit root are stationary only after first differencing and all are used for the cointegration test as well as error correction model.

For the error correction model, the logarithms of the variables with a structural break are used. The exchange rate policy changed in fourth quarter of year 1986 and the unit root result with structural break is appropriate for the error correction model. Natural log variables were chosen because their first differences then reflect the rate of change of each variable.

Table 5.1. Philips-Perron unit root test result with a structural breaks for cocoa variables, and the time trend of quarterly data (1971-1998)

Variable	Level	Δ Level	Log	Δ Log
(bc) Quantity of cocoa export	-3.27(9)	-4.87(7) ***	-3.92(2)	-8.01(2) ***
(fec) Real foreign income	-1.04(1)	-4.04(1) *	-1.35(3)	-4.08(1) *
(erc) Real exchange rate weight	-2.56(4)	-4.88(4) ***	-2.52(4)	-6.28(2) ***
(wpr) World cocoa production	-3.58(10)	-5.12(9) ***	-3.57(10)	-4.89(9) ***
(dpc) Indonesian cocoa price	-3.42(1)	-6.66(8) ***	-3.38(1)	-10.52(0) ***
(dpbz) Brazilian cocoa price	-2.92(2)	-5.20(3) ***	-3.27(2)	-5.74(3) ***
(dprg) Ghana cocoa price	-3.46(1)	-4.68(7) **	-7.27(4) ***	-7.31(10) ***

- The unit root equation is

$$Y_t = \mu_0 + \mu_1 DU_t + \mu_2 t + \mu_3 DT_t + \mu_4 D(T_b)_t + \beta Y_{t-1} + \sum_{i=1}^{p} \gamma \Delta Y_{t-i} + V_t$$

- Critical Value 1%= -4.88,5%=-4,24, 10%=-3.95. λ = 0.6 (Perron, 1989)
- Δ is the first differences of the data.
- Number of lags in parentheses
- Number of lags are determined using Akaike Information Criterion (AIC)(Perron, 1989)

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5.2 Augmented Dickey Fuller (ADF) unit root test results trend with no structural breaks and a time trend for cocoa variables, quarterly data (1971-1998)

Variable	Level	Δ Level	Log	Δ Log
(bc) Quantity of cocoa export	-0.48(9)	-4.80(7) ***	-2.92(2)	-8.01(2) ***
(fec) Real foreign income	-1.72(1)	-3.87(1) **	-2.37(3)	-3.96(1) **
(erc) Real exchange rate weight	-1.66(4)	-5.16(4) ***	-2.61(4)	-6.17(2) ***
(wpr) World cocoa production	-2.97(10)	-3.85(9) ***	-2.75(10)	-4.95(9) ***
(dpc) Indonesian cocoa price	-3.18(1) *	-6.89(8) ***	-3.68(1) *	-7.60(0) ***
(dpbz) Brazilian cocoa price	-2.73(2)	-5.12(3) ***	-1.52(2)	-7.42(3) ***
(dprg) Ghana cocoa price	-2.02(1)	-4.73(7) ***	-2.10(4)	-3.72(10) **

- The unit root equation is

$$\Delta X_t = \alpha + \beta_1 X_{t-1} + \beta_2 t + \sum_{j=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t$$

- Critical Value 1%= -4.04, 5%=-3.73, 10%=-3.15. (Fuller, 1976)
- Δ is the first differences of the data.
- Number of lags in parentheses
- Number of lags are determined using Akaike Information Criterion (AIC)(Perron, 1989)

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5. 3 Philips-Perron unit root test result with a structural breaks for coffee variables, and the time trend of quarterly data (1971-1998)

Variable	Level	Δ Level	Log	Δ Log
(bcf) Quantity of coffee exports	-3.46(9)	-4.70(10)**	-3.37(7)	-5.03(6)***
(ficf) Real foreign income	-0.66(10)	-4.45(5)**	-0.23(1)	-3.96(8)*
(ercf) Real exchange rate weight	-2.44(4)	-4.59(4)***	-2.26(4)	-6.52(1)***
(wpr) World coffee production	-3.16(8)	-10.77(7)***	-2.84(8)	-7.63(5)***
(dpcf) Indonesian coffee prices	-2.56(2)	-10.53(0)***	-2.83(1)	-10.1(0)***
(dprs) World sugar prices	-3.44(5)**	-4.95(4)***	-3.79(0)	-3.96(4)***
(dpbz) Brazilian coffee prices	-3.02(2)	-8.35(1)***	-3.61(1)	-17.71(0)***

- The unit root equation is

$$Y_t = \mu_0 + \mu_1 DU_t + \mu_2 t + \mu_3 DT_t + \mu_4 D(T_b)_t + \beta Y_{t-1} + \sum_{i=1}^{p} \gamma \Delta Y_{t-i} + V_t$$

- Critical Value 1%= -4.88,5%=-4,24, 10%=-3.95. $\lambda = 0.6$ (Perron, 1989)
- Δ is the first differences of the data.
- Number of lags in parentheses
- Number of lags are determined using Akaike Information Criterion (AIC)(Perron, 1989)

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5.4 Augmented Dickey Fuller (ADF) unit root test results trend with no structural breaks and a time trend for coffee variables, quarterly data (1971-1998)

Variable	Level	Δ Level	Log	Δ Log
(bcf) Quantity of coffee exports	-2.73(9)	-4.56(10)***	-3.19(7)	-5.08(6)***
(ficf) Real foreign income	-2.41(10)	-4.04(5)***	-2.39(1)	-4.70(8)***
(ercf) Real exchange rate weight	-1.71(4)	-4.97(4)***	-2.62(4)	-6.48(1)***
(wpr) World coffee production	-2.56(8)	-5.36(7)***	-3.01(8)	-11.15(5)***
(dpcf) Indonesian coffee prices	-2.11(2)	-4.97(0)***	-3.21(1)	-7.95(0)***
(dprs) World's sugar prices	-3.14(5)	-4.97(4)***	-3.22(0)	-4.53(4)***
(dpbz) Brazilian coffee prices	-2.83(2)	-7.63(1)***	-3.13(1)	-7.49(0)***

- The unit root equation is

$$\Delta X_t = \alpha + \beta_1 X_{t-1} + \beta_2 t + \sum_{j=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t$$

- Critical Value 1%= -4.04, .5%=-3.73, 10%=-3.15. (Fuller, 1976)
- Δ is the first differences of the data.
- Number of lags in parentheses
- Number of lags are determined using Akaike Information Criterion (AIC)(Perron, 1989)

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5.5 Philips-Perron unit root test result with structural breaks for palm oil variables, and the time trend of quarterly data (1971-1998)

Variable	Level	Δ Level	Log	Δ Log
(bp) Quantity of palm oil export	-3.52(10)	-5.59(9) ***	-2.95(5)	-6.54(3) ***
(fep) Real foreign income	-1.73(2)	-2.12(7)	2.31(1)	-3.96(10) *4
(erp) Real exchange rate weight	-2.66(4)	-4.88(4) ***	-2.78(4)	-6.13(2) ***
(wpp) World palm oil production	-2.79(2)	-4.73(8) **	-8.21(2) ***	-6.82(5) ***
(dpep) Indonesian palm oil price	-3.37(2)	-5.60(7) ***	-3.46(1)	-6.82(1) ***
(dpc) World coconut oil price	-4.68(5) **	-4.78(5) **	-5.14(7) ***	-5.51(6) ***
(dpmc) Malaysian palm oil price	-3.80(1)	-4.28(10) **	-2.70(1)	-5.73(1) ***

- The unit root equation is

$$Y_t = \mu_0 + \mu_1 DU_t + \mu_2 t + \mu_3 DT_t + \mu_4 D(T_b)_t + \beta Y_{t-1} + \sum_{i=1}^{p} \gamma \Delta Y_{t-i} + V_t$$

- Critical Value 1%=-4.88,5%=-4,24, 10%=-3.95. λ = 0.6 (Perron, 1989)
- Δ is the first differences of the data.
- Number of lags in parentheses
- Number of lags are determined using Akaike Information Criterion (AIC)(Perron, 1989)

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5.6 Augmented Dickey Fuller (ADF) unit root test results trend with no structural breaks and a time trend for palm oil variables, quarterly data (1971-1998)

Variable	Level	Δ Level	Log	Δ Log
(bp) Quantity of palm oil export	-1.78(10)	-6.43(9) ***	-3.09(5)	-11.93(3) ***
(fep) Real foreign income	-1.58(2)	-3.79(7) **	-1.74(1)	-3.83(10) **
(erp) Real exchange rate weight	-1.91(4)	-5.14(4) ***	-2.82(4)	-6.01(2) ***
(wpp) World palm oil production	-2.16(10)	-14.08(2) ***	-5.18(9) ***	-6.91(8) ***
(dpep) Indonesian palm oil price	-3.35(2) **	-5.27(7) ***	-2.26(1)	-6.47(1) ***
(dpc) World coconut oil price	-4.71(5) ***	-5.96(5) ***	-3.36(7) **	-5.32(6) ***
(dpmc) Malaysian palm oil price	-2.70(1)	-4.46(10) ***	-3.14(1)	-6.84(1) ***

- The unit root equation is

$$\Delta X_t = \alpha + \beta_1 X_{t-1} + \beta_2 t + \sum_{j=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t$$

- Critical Value 1%= -4.04, .5%=-3.73, 10%=-3.15. (Fuller, 1976)
- Δ is the first differences of the data.
- Number of lags in parentheses
- Number of lags are determined using Akaike Information Criterion (AIC)(Perron, 1989)

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5.7 Philips-Perron unit root test result with a structural breaks for rubber variables, and the time trend of quarterly data (1971-1998).

Variable	Level	Δ Level	Log	Δ Log
(brr) Quantity of rubber export	-3.62(3)	-5.11(3)***	-3.62(1)	-9.56(1)***
(fer) Real foreign income	-1.91(1)	-3.99(1)*	-2.07(1)	-3.93(10)***
(err) Real exchange rate weight	-2.14(4)	-4.34(4)**	-2.16(4)	-6.35(2)***
(wpd) World rubber production	-3.69(5)	-4.21(10)**	-4.16(1)*	-8.96(3)***
(dpir) Indonesian rubber price	-3.52(1)	-5.50(7)**	-3.20(1)	-4.45(4)**
(dpmr) Malaysian rubber price	-3.50(4)	-5.52(5)***	-3.69(4)	-5.39(6)***
(dptr) Thailand rubber price	-3.76(2)	-5.76(5)***	-3.95(6)	-5.56(4)***

- The unit root equation is

$$Y_t = \mu_0 + \mu_1 DU_t + \mu_2 t + \mu_3 DT_t + \mu_4 D(T_b)_t + \beta Y_{t-1} + \sum_{i=1}^{p} \gamma \Delta Y_{t-i} + V_t$$

- Critical Value 1%= -4.88,5%=-4,24, 10%=-3.95. λ = 0.6 (Perron, 1989)
- Δ is the first differences of the data.
- Number of lags in parentheses
- Number of lags are determined using Akaike Information Criterion (AIC)(Perron, 1989)

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5. 8 Augmented Dickey Fuller (ADF) unit root test results trend with no structural breaks and a time trend for rubber variables, quarterly data (1971-1998)

Variable	Level	Δ Level	Log	Δ Log
(brr) Quantity of rubber export	0.76(3)	-5.08(3)***	-1.87(1)	-9.37(1)***
(fer) Real foreign income	-1.75(1)	-3.73(1)**	-2.14(1)	-4.02(10)**
(err) Real exchange rate weight	-1.52(4)	-4.72(4)***	-2.53(4)	-6.24(2)***
(wpd) World rubber production	-2.06(5)	-6.29(10)***	-2.69(5)	-9.33(3)***
(dpir) Indonesian rubber price	-3.37(1)*	-5.37(7)***	-1.81(1)	-3.91(4)**
(dpmr) Malaysian rubber price	-3.70(4)*	-5.39(5)***	-3.08(4)	-5.08(6)***
(dptr) Thailand rubber price	-4.16(2)*	-5.67(5)***	-3.53(6)*	-5.03(4)***

- The unit root equation is

$$\Delta X_t = \alpha + \beta_1 X_{t-1} + \beta_2 t + \sum_{j=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t$$

- Critical Value 1%= -4.04, .5%=-3.73, 10%=-3.15. (Fuller, 1976)
- Δ is the first differences of the data.
- Number of lags in parentheses
- Number of lags are determined using Akaike Information Criterion (AIC)(Perron, 1989)

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5. 9 Philips-Perron unit root test result with a structural breaks for Tea variables, and the time trend of quarterly data (1971-1998)

Variable	Level	Δ Level	Log	Δ Log
(bt) Quantity of tea export	-2.73(4)	-6.28(3)***	-3.00(2)	-8.33(2)***
(fet) Real foreign income	-2.14(1)	-4.13(1)**	0.82(1)	-4.03(8)**
(ert) Real exchange rate weight	-2.13(4)	-4.32(4)**	-2.47(4)	-6.49(1)***
(wpp) World tea production	-4.58(9)**	-4.54(10)**	-5.06(4)***	-4.98(9)***
(dpt) Indonesian tea price	-3.73(2)	-5.67(6)***	-3.89(1)	-6.84(1)***
(dpst) World sugar price	-3.82(5)	-5.39(4)***	-3.56(1)	-3.97(8)***
(dptt) Sri Lankas tea price	-2.44(4)	-5.1(4)***	-2.65(1)	-7.27(1)***

- The unit root equation is

$$Y_t = \mu_0 + \mu_1 DU_t + \mu_2 t + \mu_3 DT_t + \mu_4 D(T_b)_t + \beta Y_{t-1} + \sum_{i=1}^{p} \gamma \Delta Y_{t-i} + V_t$$

- Critical Value 1%= -4.88,5%=-4,24, 10%=-3.95. λ = 0.6 (Perron, 1989)
- Δ is the first differences of the data.
- Number of lags in parentheses
- Number of lags are determined using Akaike Information Criterion (AIC)(Perron, 1989)

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5. 10 Augmented Dickey Fuller (ADF) unit root test results trend with no structural breaks and a time trend for tea variables, quarterly data (1971-1998)

Variable	Level	Δ Level	Log	Δ Log
(bt) Quantity of tea export	-2.32(4)	-7.13(3)***	-2.29(2)	-7.99(2)***
(fet) Real foreign income	-0.22(1)	-3.19(7)*	-0.34(1)	-3.76(8)***
(ert) Real exchange rate weight	-1.64(4)	-4.69(4)***	-2.63(4)	-6.47(1)***
(wpp) World tea production	-3.87(9)***	-4.44(10)***	-3.41(4)*	-5.05(9)***
(dpt) Indonesian tea price	-3.08(2)	-5.74(6)***	-4.19(1)***	-6.90(1)***
(dpst) World sugar price	-2.54(5)	-5.45(4)***	-2.99(1)	-3.55(8)***
(dptt) Sri Lankas tea price	-3.13(4)	-4.61(4)***	-2.61(1)	-7.26(1)***

- The unit root equation is

$$\Delta X_t = \alpha + \beta_1 X_{t-1} + \beta_2 t + \sum_{j=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t$$

- Critical Value 1%= -4.04, .5%=-3.73, 10%=-3.15. (Fuller, 1976)
- Δ is the first differences of the data.
- Number of lags in parentheses
- Number of lags are determined using Akaike Information Criterion (AIC)(Perron, 1989)

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5.11 Philips-Perron unit root test result with a structural breaks for aggregate export variables, and the time trend of quarterly data (1971-1998)

Variable	Level	Δ Level	Log	Δ Log
(be) Quantity aggregate export	-3.46(3)	-7.44(3) ***	-3.41(2)	-7.30(3) ***
(fe) Real foreign income	0.35(1)	-4.84(5) ***	-0.70(1)	-4.82(8) ***
(ere) Real exchange rate weight	-2.13(4)	-13.99(4) ***	-2.18(4)	-14.49(2) ***
(dpe) Real aggregate export price	-3.82(1)	-12.22(1) ***	-3.12(1)	-5.02(2) ***

- The unit root equation is

$$Y_t = \mu_0 + \mu_1 DU_t + \mu_2 t + \mu_3 DT_t + \mu_4 D(T_b)_t + \beta Y_{t-1} + \sum_{i=1}^{p} \gamma \Delta Y_{t-i} + V_t$$

-

Critical Value 1%= -4.88,5%=-4,24, 10%=-3.95. λ = 0.6 (Perron, 1989)
- Δ is the first differences of the data.
- Number of lags in parentheses
- Number of lags are determined using Akaike Information Criterion (AIC)(Perron, 1989)
*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5. 12 Augmented Dickey Fuller (ADF) unit root test results trend with no structural breaks and a time trend for aggregate export variables, quarterly data (1971-1998)

Variable	Level	Δ Level	Log	Δ Log
(be) Quantity aggregate export	-1.21(3)	-7.42(3) ***	-2.96(2)	-7.23(3) ***
(fe) Real foreign income	-1.10(1)	-4.12(1) **	-3.98(1)	-3.28(8) **
(ere) Real exchange rate weight	-1.52(4)	-4.76(0) ***	-2.52(4)	-6.27(2) ***
(dpe) Real aggregate export price	-1.66(1)	-7.63(5) ***	-2.25(1)	-7.80(2) ***

- The unit root equation is

$$\Delta X_t = \alpha + \beta_1 X_{t-1} + \beta_2 t + \sum_{j=1}^{p} \gamma_j \Delta X_{t-j} + \varepsilon_t$$

- Critical Value 1%= -4.04, 5%=-3.73, 10%=-3.15. (Fuller, 1976)
- Δ is the first differences of the data.
- Number of lags in parentheses
- Number of lags are determined using Akaike Information Criterion (AIC)(Perron, 1989)
*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

5.2 Cointegration Tests

From Table 5.1 to 5.12 the results indicate that for all five commodities and aggregate agricultural exports, first differences and first difference of logarithmic what of variables are I(1). The variables that are not stationary are not used in the cointegration regression. In this section the degree of integration of the residual from the cointegration equation (Equation 4.18) are reported. If the variables are to be cointegrated, the residuals from the cointegration equation must be stationary, I(0). The structural break regression is to be used for the cointegration using the dummy variable dum_u and dum_t dummy variables for structural break as shown in the table. The error term μ_t is tested for a unit root using the ADF test. The lag length for the ADF test are determined using Akaike Information Criterion (AIC) from Statespace Procedure in SAS.

In each of the five commodities and aggregate exports, volume of exports is the dependent variables. Variables that are also I(1) are sequentially included in the equation the equation as shown in table 5.13 to 5.18.

(1) Tables 5.13 to Table 5.18 the results of the ADF test applied to the residual cointegration equations, R^2 and cointegration regression Durbin Watson (DW) statistic are presented for five commodities and aggregate exports.

(2) If the critical value of the ADF statistics for the entire residuals on the cointegration regression equation is greater than the critical value reported by Fuller (1976), there is a long-run relationship between all stationary variables and the volume of exports.

(3) The critical value table by Fuller (1976) is presented in Table 5.24 in the end of the chapter.

(4) If the degree of integration of the residual is less than the degree of integration of all variables in the cointegration equation, all series are cointegrated, and there is a long-run relationships between export volume and all others variables in the cointegration equation.

(5) The DW test statistic from the cointegration regression indicates whether the residuals from the cointegration to indicate whether the residual from the cointegration regressions are stationary.

(6) If the DW is greater than the critical 0.51 the null hypothesis of no cointegration is rejected (Engle and Yoo, 1987).

For all of the commodities and aggregate exports evidence supports cointegration among the variables. The ADF statistics and DW statistics are all uniformly greater than the critical values. This suggests volume of exports is influenced by exchange rate, competitive prices, price of export and foreign income. There is a long-run relationship detected between volume of exports and exchange rate, price and foreign income long-run relationship in the cointegration regression.

Although it has been established that there is the long-run relation equation exists, the question remaining is which variables provides short-run dynamic adjustment toward the long run equilibrium. Estimating the error correction models described by Equation 5.4 provides the answers.

Table 5. 13 Cointegration test results for cocoa exports demand variables with structural break, quarterly (1971-1998)

Cointegration Equation	ADF Statistics	R^2	DW
bc=erc,dum_u,dum_t	-8.06(3)***	0.0024	2.98
bc=fec, erc,dum_u,dum_t	-7.99(3)***	0.0025	2.98
bc=fec, erc, dpc, dum_u,dum_t	-7.97(3)***	0.0025	2.98
bc=fec, erc, dpc, dpbz, dum_u,dum_t	-8.04(3)***	0.0139	2.96
bc=fec, erc, dpc, dpbz, dps,dum_u,dum_t	-9.66(3)***	0.1653	2.88
bc=fec, erc, dpc, dpbz, dps, wpr, dum_u,dum_t	-10.04(3)***	0.2090	2.88

The complete cointegration equation is :

$$\ln bc = \alpha_0 + \alpha_1 \ln fec + \alpha_2 \ln erc + \alpha_3 \ln dpc + \alpha_4 \ln dpbz + \alpha_5 \ln dps + \alpha_6 \ln wpr + \delta_1 dum\_u + \delta_2 dum\_t + U_t$$

bc= volume of cocoa exports
erc= exchange rate weight
fec= foreign income weight
dpc= price of Indonesian cocoa
dpbz= price of Brazilian cocoa
dps= price of sugar
wpr= world cocoa production
dum_u= intercept dummy
dum_t= slope dummy
a. Break Year is 1987 (I).
b. Number in parentheses is the number of lags.
c. Critical Values cointegration .
*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level (Engle and Granger, 1987).

Table 5. 14 Cointegration test results for coffee exports demand variables with structural break, quarterly (1971-1998)

Cointegration Equation	ADF Statistics	R^2	DW
bcf=ercf,dum_u,dum_t	-4.94(10)***	0.002	2.16
bcf=ficf, ercf, dum_u,dum_t	-4.73(6)***	0.030	2.13
bcf=ficf, ercf, dpcf, dum_u,dum_t	-5.17(4)***	0.113	2.17
bcf=ficf, ercf, dpcf, dpbz, dum_u,dum_t	-5.25(4)***	0.119	2.17
bcf=ficf, ercf, dpcf, dpbz, wpr, dum_u,dum_t	-5.43(3)***	0.154	2.09

The complete cointegration equation is :

$$\ln bcf = \alpha_0 + \alpha_1 \ln ficf + \alpha_2 \ln ercf + \alpha_3 \ln dpcf + \alpha_4 \ln dpbz + \alpha_5 \ln wpr + \delta_1 dum\_u + \delta_2 dum\_t + U_t$$

bcf= volume of coffee exports
erc= exchange rate weight
fec= foreign income weight
dpcf= price of Indonesian coffee
dpbz= price of Brazilian coffee
wpr= world coffee production
dum_u= intercept dummy
dum_t= slope dummy
a. Break Year is 1987 (I).
b. Number in parentheses is the number of lags.
c. Critical Values cointegration .
*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level (Engle and Granger, 1987).

Table 5. 15 Cointegration test results for palm oil exports demand variables with structural break, quarterly (1971-1998)

Cointegration Equation	ADF Statistics	R^2	DW
bp=erp,dum_u,dum_t	-6.67(3)***	0.002	2.40
bp=fep, erp,dum_u,dum_t	-6.70(3)***	0.003	2.41
bp=fep, erp, dpep, dum_u,dum_t	-6.76(3)***	0.043	2.42
bp=fep, erp, dpep, dpmc, dum_u,dum_t	-7.09(3)***	0.025	2.46
bp=fep, erp, dpep, dpmc, wpr, dum_u,dum_t	-7.43(3)***	0.068	2.46

The complete cointegration equation is :

$$\ln bp = \alpha_0 + \alpha_1 \ln fep + \alpha_2 \ln erp + \alpha_3 \ln dpep + \alpha_4 \ln dpmc + \alpha_5 \ln wpr + \delta_1 dum\_u + \delta_2 dum\_t + U_t$$

bp= volume of palm oil exports
erc= exchange rate weight
fec= foreign income weight
dpep= price of Indonesian palm oil
dpmc= price of Malaysian Palm Oil
wpr= world palm oil production
dum_u= intercept dummy
dum_t= slope dummy
a. Break Year is 1987 (I).
b. Number in parentheses is the number of lags.
c. Critical Values cointegration .
*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level (Engle and Granger, 1987).

Table 5.16 Cointegration test results for rubber exports demand variables with structural break, quarterly (1971-1998)

Cointegration Equation	ADF Statistics	R^2	DW
brr=err,dum_u,dum_t	-7.22(2)***	0.020	2.60
brr=fer, err,dum_u,dum_t	-7.18(2)***	0.024	2.63
brr=fer, err, dpir, dum_u,dum_t	-7.55(1)***	0.028	2.59
brr=fer, err, dpir, dpmr, dum_u,dum_t	-7.15(2)***	0.034	2.60
brr=fer, err, dpir, dpmr, dptr,dum_u,dum_t	-7.21(2)***	0.041	2.59
brr=fer, err, dpir, dpmr, dptr, wpd, dum_u,dum_t	-7.13(2)***	0.047	2.58

The complete cointegration equation is :

$$\ln brr = \alpha_0 + \alpha_1 \ln fer + \alpha_2 \ln err + \alpha_3 \ln dpir + \alpha_4 \ln dpmr + \alpha_5 \ln dptr + \alpha_6 \ln wpd + \delta_1 dum\_u + \delta_2 dum\_t + U_t$$

brr= volume of rubber exports
err= exchange rate weight
fer= foreign income weight
dpir= price of Indonesian rubber
dpmr= price of Malaysian rubber
dptr= price of Thailand rubber
wpdr= world rubber production
dum_u= intercept dummy
dum_t= slope dummy
a. Break Year is 1987 (I).
b. Number in parentheses is the number of lags.
c. Critical Values cointegration .
*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level (Engle and Granger, 1987).

Table 5. 17 Cointegration test results for tea exports demand variables with structural break, quarterly (1971-1998)

Cointegration Equation	ADF Statistics	R^2	DW
bt=ert,dum_u,dum_t	-6.95(3)***	0.013	2.63
bt=fet, ert,dum_u,dum_t	-6.97(3)***	0.023	2.62
bt=fet, ert, dpt, dum_u,dum_t	-8.02(2)***	0.068	2.56
bt=fet, ert, dpt, dpst, dum_u,dum_t	-8.00(2)***	0.107	2.53
bt=fet, ert, dpt, dpst, dptt,dum_u,dum_t	-8.19(2)***	0.111	2.53

The complete cointegration equation is :

$$\ln bt = \alpha_0 + \alpha_1 \ln fet + \alpha_2 \ln ert + \alpha_3 \ln dpt + \alpha_4 \ln dpst + \alpha_5 \ln dptt + \delta_1 dum\_u + \delta_2 dum\_t + U_t$$

bt= volume of tea exports
ert= exchange rate weight
fet= foreign income weight
dpt= price of Indonesian tea
dptt= price of Sri Lankan tea.
dpst= price of sugar
dum_u= intercept dummy
dum_t= slope dummy
a. Break Year is 1987 (I).
b. Number in parentheses is the number of lags.
c. Critical Values of cointegration .
*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level (Engle and Granger, 1987).

Table 5. 18 Cointegration test results for aggregate exports demand variables with structural break, quarterly (1971-1998)

Cointegration Equation	ADF Statistics	R^2	DW
be=ere,dum_u,dum_t	-7.42(3)***	0.023	2.77
be=fe, ere,dum_u,dum_t	-7.23(3)***	0.025	2.77
be=fe, ere, pe, dum_u,dum_t	-7.06(3)***	0.040	2.77

The complete cointegration equation is :

$$\ln be = \alpha_0 + \alpha_1 \ln fe + \alpha_2 \ln ere + \alpha_3 \ln pe + \delta_1 dum\_u + \delta_2 dum\_t + U_t$$

be= volume of Indonesian agricultural exports
ere= exchange rate weight
fe= foreign income weight
pe= index price of Indonesian agricultural exports
dum_u= intercept dummy
dum_t= slope dummy
a. Break Year is 1987 (I).
b. Number in parentheses is the number of lags.
c. Critical Values of cointegration .
*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level (Engle and Granger, 1987).

5.3 Error Correction Model

The error correction model estimates for the five commodities and aggregate agricultural exorts reported in Tables 5.19 to 5.23. All the prices are deflated with the country price index using 1995(I) as the base. The analysis of the model's residuals and model misspecification tests are also reported. The error correction parameters link the long-run relationships of the system to the short-run dynamics.

An insignificant error correction parameter would imply that movement in the variable does not resolve deviation from the long-run equilibrium. The fact that all of the parameters of the error correction model are significantly different from zero implies that they all participate in restoring long-run equilibrium. This means that permanent change in any one of the independent variables will lead to a temporary deviation in the long-run relationship between the export demand variables. The extent to which each independent variable will move in response to long-run disequilibrium is manifested in the relative size or the error correction parameter. The regression model of the error correction model is:

$$\begin{aligned} \Delta ln X_t &= \alpha_0 + \sum_{i=1}^{m} \beta_{1i} \Delta lnY_{t-i} + \sum_{j=1}^{n} \beta_{2j} \Delta lnP_{t-i} + \sum_{k=1}^{q} \beta_{3k} \Delta lnV_{t-i} \\ &\quad + \delta_1 dum\_u + \delta_2 dum\_t + \eta U_{t-1} + \varepsilon_t \end{aligned} \tag{5. 4}$$

where U_{t-1} is the error correction term that derive from the Equation 5.5 below, dum_u is the dummy variable that if t >T_c is one and zero otherwise, and dum_t is t-T_c if t > T_c or zero otherwise, where T_c is the time break which are the dummy variable for structural break.

$$\Delta \ln X_t = \alpha_0 + \sum_{i=1}^{m} \alpha_{1i} \Delta \ln Y_{t-i} + \sum_{j=1}^{n} \alpha_{2j} \Delta \ln P_{t-i} + \sum_{k=1}^{q} \alpha_{3k} \Delta \ln V_{t-i} + \vartheta_1 dum\_u + \vartheta_2 dum\_t + \varepsilon_t \qquad (5.5)$$

The error correction model estimated for this research is the restricted model from the Engle-Granger two-step procedure. The model allows various lag lengths for each variable to allow for lagged response of importers in other countries to a change in values of independent variables.

5.3.1 Cocoa

The results for the error correction model for the cocoa are exports are reported in Table 5.19. An experiment was conducted to select the length of lags on foreign income (*Δlnfec)* and weighted exchange rate (*Δlnerc)* using the Equation 5.4. Initially three lags were included in foreign income variable, three lags in the exchange rate variable and two lags in cocoa price variable. The foreign income variable is not significant in the equation. It indicates that effect of foreign income on cocoa export volume in the long-run model is small since cocoa is a basic commodity. Trading partner income does not influence quantity of cocoa imported from Indonesia. Weight exchange rate *Δlnerc* has a negative coefficient at the three lagged level, and is statistically significant at one percent. This seems to imply that the exchange rate has three quarter delay effect on the volume of exports, and as the exchange rate depreciates by ten percent, export volume will increase by 6.9 percent. Since variable *Δlndpc* has a negative coefficient as expected and significant at the one percent level for the current period. The parameter estimates is

(-0.79) which is the own price elasticity. The negative sign indicates that excess demand elasticity greater than excess supply. The significance of the parameter indicates that the price of the Indonesian export has a strong effect on volume of e cocoa exports.

The one lagged error correction term Ec_{t-1} is statistically significant and displays the appropriate negative sign. This finding supports the validity of an equilibrium relationship among the variables in the error correction model. The dum_u and dum_t have a negative sign and are not statistically significant. This mean there is no strong evidence of a structural break dummy variable in the model.

The overall F test is (9.09) statistically significant at one percent, and the R^2 is 0.47. In particular, the estimated model fulfills the conditions of serial noncorellation (DW=2.57), which indicates low probability of serial correlation and normality of residual (χ^2=2.81).

5.3.2 Coffee

The results for the coffee error correction model are is reported in Table 5.20. The coffee export model included foreign income ($\Delta lnficf$), weighted exchange rate ($\Delta lnercf$), price of coffee export ($\Delta lndpcf$), and price of Brazilian coffee as the competitive price ($\Delta lndpbz$). Initially,up to five lags have been included for the $\Delta lnficf$ and none of the coefficient are statistically significant. Foreign income does not influence the volume of coffee exports. The exchange rate variable ($\Delta lnercf$) coefficient is negative and statistically significant. The exchange rate does not affect the exports instantly, but is significant after the fifth quarter. The own price coefficient (current period) coefficient is negative significant at the one percent level. The current price of coffee influences the volume of export. The Brazilian coffee price (five lags) coefficient is positive sign statistically significant. As the price of Brazilian coffee increases, exports of Indonesian coffee increases. Both countries are major coffee exporters and importers will buy more Indonesian coffee if the Brazil coffee price increases.

The one lagged error correction term Ec_{t-1}, is statistically significant and displays the appropriate negative sign. This finding supports the validity of an equilibrium relationship among the variables in the error correction model. This implies that overlooking the cointegratedness of the variables would have introduced misspecification in the underling dynamic structure. A higher power of Ect-1 is statistically insignificant. The change in real exports per quarter is attributed to the disequilibrium between the actual and equilibrium level is measured by the absolute values of the error correction term equation. There is considerable inter-commodity variation in the adjustment speed to the past period's disequilibrium in the export demand model. The dum_u and dum_t have a negative coefficient and are not statistically significant. There is no strong evidence of a structural break in the model.

The F test is (8.5) that statistically significant at one percent and the R^2 is 0.49. In particular the estimated model fulfills the conditions of serial noncorellation (DW=2.75), normality of residual (χ^2=2.75).

5.3.3 Palm Oil

The result for the palm oil the error correction model is reported in Table 5.21. An experiment is used to guide to select the length of lags on foreign income *(dlnfep),* weighted exchange rate *(dlnerp)* using Equation 5.7. Initially up to of the foreign income, three lags in exchange rate and four lags in palm oil price ($\Delta lndpep$) variables. Four lags in the Malaysian palm oil ($\Delta lndpmc$) price is the competitive price. Foreign income (two lags) is statistically significant. It indicates that foreign income does an impact on palm oil exports. Trading partner countries will import more from Indonesia as they increase the income of their countries. Variable $\Delta lnerp$ has a negative sign and is statistically significant at one percent. The exchange rate has a one quarter delay effect on the volume of exports. As the exchange rate increases by ten percent, it will reduce the export volume by 6.2 percent. The variable $\Delta lndpep$ as a negative sign as expected, and is significant at one percent level. The parameter estimates is -0.90. The price of the Indonesian exports has a strong negative effect on the volume of the palm oil volume exports. Malaysian Palm oil price coefficient has a positive sign as expected and is statistically significant at the five lagged periods. As the price of Malaysian Palm oil increases, exportation of Indonesian Palm oil exports increase. Both countries are the major palm oil exporters, so the importers will buy more Indonesian Palm oil if Malaysian prices increase.

The lagged error correction term $Ec_{t-1,}$ is statistically significant and the coefficient is appropriately negative. This finding supports the validity of an equilibrium relationship among the variables in the error correction model. The dum_u and dum_t have a negative sign and are not statistically significant. There is no strong evidence supporting a structural break dummy in the model.

The F test is (2.5) that statistically significant at one percent, the R^2 is 0.22 in particular the estimated model fulfills the conditions of serial noncorellation (DW=2.14), normality of residual (χ^2=3.67).

5.3.4 Rubber

The results for the rubber error correction model are reported in Table 5.22. Estimation was established to select the length of lags on foreign income ($\Delta lnfer$), exchange rate weight ($\Delta lnerr$) using Equation 5.4. Initially, four lags were included in the foreign income variable, along with one lag in exchange rate variables and two lags in cocoa price ($\Delta lndpir$) variables. The foreign exchange variable is not significant in the equation, even with three lags of the variable. Variable $\Delta lnerr$ has a negative sign only at current period levels and has a positive sign at one level of lags, it is found to be statistically significant at one percent. This implies that the exchange change immediately influences the volume of exports. The exchange rate increases by ten percent export volume decreases by 2.1 percent. Rubber price variable $\Delta lndpir$ is not significant even for the four lagged periods. Which suggest Indonesian rubber demand is also very inelastic.

The one lagged error correction term Ec_{t-1}, is statistically significant coefficient and has the appropriate negative sign. This finding supports the validity of an equilibrium relationship among the variables in the error correction model. The dum_u and dum_t have a negative sign but are not statistically significant. There is no strong evidence a structural break dummy variable in the model.

The misspecification test for the model suggests that it is statistically fit for the data. The F test is (3.54) statistically significant at one percent, and the R^2 is 0.47. In particular, the estimated model fulfills the conditions of serial noncorellation (DW=2.11). But the normality of residual (χ^2=14.67) give the result reject the null hypothesis its mean the residual distribution is not normal.

5.3.5 Tea

The results for the error correction model for tea regression are reported in Table 5.23. The variables estimated in the model tea export included foreign income ($\Delta lnfet$), weighted exchange rate ($\Delta lnert$), price of tea exports ($\Delta lndpt$), and the price of Srilanka Tea ($\Delta lndptt$). Up to four lags for the $\Delta lnfet$ variable are allowed and the coefficient are not statistically significant. The exchange rate (five lags)($\Delta lnert$) coefficient is negative and significant, suggesting that the exchange rate affects exports. The own price coefficient has a negative sign and is significant at one percent. Price of tea influences the volume of tea exports. Sri Lanka tea price coefficient has a positive sign, as expected, and is significant. It seems that as the price of Sri Lanka tea increases, then Indonesian tea exports will also increase. As both countries are major tea exporters.

The one lagged error correction term Ec_{t-1}, is statistically significant coefficient and displays the appropriate negative sign. This finding supports the validity of an equilibrium relationship among the variable in the error correction model.
This implies that overlooking the cointegratedness of the variables would introduce misspecification in the underling dynamic structure. Indeed, higher on lags of Ec_{t-1} were included in the regression equation, but proved statistically insignificant. Secondly, the change in real exports per quarter is attributed to the disequilibrium between the actual and equilibrium level is measured by the absolute value of the error correction term equation. There is considerable the inter-commodity variation in adjustment speed to the past period's disequilibrium. The dum_u and dum_t coefficients are negative and not statistically significant. This mean there is no strong evidence of a structural break in the model.

The F test is (4.4) statistically significant at one percent and the R^2 is 0.29. In particular, the estimated model fulfills the conditions of serial noncorellation (DW=2.05), normality of residual (χ^2=3.75).

5.3.6 Aggregate Agricultural Exports

The results for the error correction model for aggregate agricultural exports are reported in Table 5.23. The equation includes foreign income ($\Delta lnfe$), weighted exchange rate ($\Delta lnere$), and the index of agricultural exported ($\Delta lndpe$). Initially, five lags have been conducted for $\Delta lnfe$ variable and the coefficient is positive and significant. The exchange rate variable (five lags)($\Delta lnere$) coefficient gave the negative sign and is statistically significant at one percent. It indicates that the exchange rate affect the exports fifth quarter after the payment made. The own price coefficient is negative and significant at one percent at current period. The demand is very inelastic (-0.30) and only a little change in exports occurs if the price increases.

The one lagged error correction term Ec_{t-1} coefficient is negative and significant. This finding supports the validity of an equilibrium relationship among the variables in each error correction equation. This implies that overlooking the cointegratedness of the variables would have introduced misspecification in the underling dynamic structure. Higher powers of Ec_{t-1} were included in the regression equation, but prove statistically insignificant. The change in real exports per quarter is attributed to the disequilibrium between the actual and equilibrium level is measured by the absolute value of the error correction term. The dum_u and dum_t have a negative sign, and are not statistically significant. There is no strong evidence of a structural break in the model.

The F test is (4.4) is statistically significant (one percent), and the R^2 is 0.8. The d model fulfills the conditions of serial noncorellation (DW=2.35), normality of residual (χ^2=0.49).

5.4 Result Summary

In this study, it have been taken explicitly into account of stationarity by employing new techniques of unit root, cointegration and error-correction modeling with a structural break.

As evident from the unit root tests for stationarity using the Augmented Dickey - Fuller and Philip-Perron test with a structural break, most variables for the level and log of the level are non-stationary. Most became stationary after taking the first difference of the level and logarithm variables. The variables that are stationary after first differencing are integrated of order one, and can be used for the cointegration technique.

The cointegration test, which applied the ADF test to the residual of cointegration equation, indicates that all are statistically significant at one percent. This suggests there is a direct relationship between volumes of exports, and foreign income, weighted exchange rate, own price, and price of competitive suppliers. Having established the evidence of the long-run relation between the variables in the equation, the error-correction model to provide short-run dynamic adjustment need is estimated.

The empirical results for the error correction model indicate that the weighted exchange rate of importer countries significantly affects agricultural exports. All the commodities and the aggregate exports have negative statistically significant coefficient for exchange rate. This suggests that the depreciation of the Indonesia rupiah relative to the importer country currencies increases of agricultural exports. The effect of the number of the lag on exchange rate varies from none to five quarters. Foreign income has no strong impact for four of the five commodities. For palm oil and aggregate exports, foreign income has a positive sign, as is significant. The real price of the commodity variable reveals the strong effect on the agricultural export with a negative sign, except for rubber, with no evidence of significance. Competitive prices, which are included in the coffee, rubber, and tea models, have a positive sign and coefficients are significant.

The significance of the of the one lagged of the error correction term indicates the change in every variables in the error correction models per quarter attributed to the disequilibrium between the actual and equilibrium levels is measured by the absolute value of the error correction term of each equation. There is considerable inter-commodity variation in the adjustment speed to the last period disequilibrium. The bigger the absolute value of the error correction term coefficients the faster is the adjustment of export volume to changes in the regressors. This indicates the existence of market forces in the export market that operates to restore long-run equilibrium after short-run disturbances.

Table 5. 19. Regression Results from Error Correction Model for Cocoa Export Demand Model with Structural Break, Quarterly (1971-1998). Dependent variable is Δlnbc (Quantity of cocoa export)

Lag	Variable					
	Ec	Δlnfec	Δlnerc	Δlndpc	dum_u	dum_t
0				-0.79***	-0.01	-0.08
				(5.62)	(0.20)	(0.57)
1	-0.65***					
	(-7.87)					
2		-2.27	0.76	-0.07		
		(-0.50)	(1.06)	(-0.56)		
3		-4.56	-0.69*			
		(-1.04)	(-1.73)			

The equation for the model is:

$$\Delta \ln bc_t = \alpha_t + \sum_{i=1}^{p} \beta_i \Delta \ln fec_{t-i} + \sum_{j=1}^{q} \delta_j \Delta \ln erc_{t-i} + \sum_{k=1}^{r} \mu_k \Delta \ln dpc_{t-i} + \vartheta_1 dum\_u + \vartheta_2 dum\_t + \eta Ec_{t-1}$$

bc= volume of cocoa exports
erc= exchange rate weight
fec= foreign exchange rate weight
dpc= price of Indonesian cocoa
Ec=error correction term
dum_u= intercept dummy
dum_t= slope dummy
- Δ is first differencing of the variables
Summary statistics:
R^2= 0.47, Adjusted R^2 = 0.42
DW= 2.57
F = 9.09
χ^2(Jarque-Bera)=2.81
. *** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level (Engle and Granger, 1987).

Table 5. 20 Regression Results from Error Correction Model for Coffee Export Demand Model with Structural break, quarterly (1971-1998). Dependent variable is Δlnbcf (Quantity of coffee export)

Lag	Variable						
	Ec	Δlnficf	Δlnercf	Δlndpcf	Δlndpbz	dum_u	dum_t
0				-0.81		0.021	-0.001
				(5.33)***		(0.61)	(0.87)
1	-0.57***			-0.05			
	(-6.92)			(-0.04)			
2			0.71				
			(1.69)				
3			-0.35				
			(-1.14)				
4			0.05				
			(0.15)				
5		-3.21	-0.83***		0.05***		
		(-0.83)	(-1.93)		(2.80)		

The equation for the model is:

$$\Delta \ln bcf_t = \alpha_t + \sum_{i=1}^{p} \beta_i \Delta \ln ficf_{t-i} + \sum_{j=1}^{q} \delta_j \Delta \ln ercf_{t-i} + \sum_{k=1}^{r} \mu_k \Delta \ln dpcf_{t-i} + \sum_{l=1}^{s} \mu_k \Delta \ln dpbz_{t-i}$$
$$+ \vartheta_1 dum\_u + \vartheta_2 dum\_t + \eta Ec_{t-1}$$

bcf= volume of coffee exports
erc= exchange rate weight
fec= foreign exchange rate weight
dpcf= price of Indonesian coffee
dpbz= price of Brazilian coffee
Ec= error correction variable
dum_u= intercept dummy
dum_t= slope dummy
- Δ is first differencing of the variables
-.Ec is error correction variable
Note: All variables converted to logs
Summary statistics:
R^2= 0.49, Adjusted R^2=0.43
DW= 2.61
F=8.5
χ^2(Jarque-Bera)=2.75
*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level

Table 5. 21 Regression Results from Error Correction Model for Palm Oil Export Demand Model with Structural break, quarterly (1971-1998). Dependent variable is Δlnbp (Quantity Palm Oil export)

Lag	Variable						
	Ec	Δlnfep	Δlnerp	Δlndpep	Δlndpmc	dum_u	dum_t
0						-0.014 (-0.20)	-0.003 (0.002)
1	-0.25*** (2.69)		-0.62*** (-1.70)	-0.49 (-1.14)			
2		2.53*** (2.84)		-0.90 (2.02)***			
3			0.23 (0.55)	0.49 (1.01)			
4		2.72*** (-2.78)			1.04*** (-2.50)		

The equation for the model is:

$$\Delta \ln bp_t = \alpha_t + \sum_{i=1}^{p} \beta_i \Delta \ln fep_{t-i} + \sum_{j=1}^{q} \delta_j \Delta \ln erp_{t-i} + \sum_{k=1}^{r} \mu_k \Delta \ln dpep_{t-i} + \sum_{l=0}^{s} \varphi_l \Delta \ln dpep_{t-i}$$
$$+ \vartheta_1 dum\_u + \vartheta_2 dum\_t + \eta Ec_{t-1}$$

bp= volume of palm oil exports
erc= exchange rate weight
fec= foreign exchange rate weight
dpep= price of Indonesian palm oil
dpmc= price of Malaysian palm oil
Ec= error correction variable
dum_u= intercept dummy
dum_t= slope dummy
- Δ is first differencing of the variables

Summary statistics:
R^2= 0.22 , Adjusted R^2 = 0.13
DW= 2.14
F= 2.5
χ^2(Jarque-Bera)=3.67
. *** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5. 22 Regression Results from Error Correction Model for Rubber Export Demand Model with Structural break, quarterly (1971-1998). Dependent variable is Δlnbrr (Quantity of Rubber export)

Lag	Variable					
	Ec	Δlnfer	Δlnerr	Δlndpir	dum_u	dum_t
0			-0.21		0.02	-0.005
			(-1.68)***		(0.99)	(-0.83)
1	-0.35		0.30			
	(-3.70)***		(1.20)			
2						
3						
4		-1.50		0.02		
		(-0.98)		(0.15)		

The equation for the model is:

$$\Delta \ln brr_t = \alpha_t + \sum_{i=1}^{p} \beta_i \Delta \ln fer_{t-i} + \sum_{j=1}^{q} \delta_j \Delta \ln err_{t-i} + \sum_{k=1}^{r} \mu_k \Delta \ln dpir_{t-i} + \vartheta_1 dum\_u + \vartheta_2 dum\_t + \eta Ec_{t-1}$$

brr= volume of rubber exports
err= exchange rate weight
fer= foreign exchange rate weight
dpir= price of Indonesian rubber
Ec= error correction variable
dum_u= intercept dummy
dum_t= slope dummy
- Δ is first differencing of the variables
Summary statistics:
R^2= 0.20 , Adjusted R^2 = 0.15
DW= 2.11
F= 3.54
χ^2(Jarque-Bera)=14.67
*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5. 23 Regression Results from Error Correction Model for Tea Export Demand Model with Structural break, quarterly (1971-1998). Dependent variable is Δlnbt (Quantity of Tea exports)

Lag	Variable						
	Ec	Δlnfet	Δlnert	Δlndpt	Δlndptt	dum_u	dum_t
0				-0.88 (-3.80)***	1.01 (3.18)***	-0.82 (0.41)	-0.01 (-1.66)*
1	-0.30 (-3.17)***						
4		-1.92 (-0.40)					
5			(-0.63)*** -2.03				

The equation for the model is:

$$\Delta \ln bt_t = \alpha_t + \sum_{i=1}^{p} \beta_i \Delta \ln fet_{t-i} + \sum_{j=1}^{q} \delta_j \Delta \ln ert_{t-i} + \sum_{k=1}^{r} \mu_k \Delta \ln dpt_{t-i} + \sum_{l=0}^{s} \varphi_l \Delta \ln dptt_{t-i} + \vartheta_1 dum\_u + \vartheta_2 dum\_t + \eta Ec_{t-1}$$

bt= volume of tea exports
ert= exchange rate weight
fet= foreign exchange rate weight
dpt= price of Indonesian tea
dptt= price of Srilanka tea.
Ec= error correction variable
dum_u= intercept dummy
dum_t= slope dummy
- Δ is first differencing of the variables

Summary statistics:
R^2= 0.29 , Adjusted R^2 = 0.23
DW= 2.05
F= 4.44
χ^2(Jarque-Bera)=3.87

*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level

Table 5. 24 Regression Results from Error Correction Model for Aggregate Export Demand Model with Structural break, quarterly (1971-1998). Dependant variable is Δlnbe (Quantity of Aggregate export demand)

Lag	Variable					
	Ec	Δlnfe	Δlnere	Δlnped	dum_u	dum_t
0				0.032 (0.28)	-0.01 (-0.71)	0.007 (1.10)
1	-0.41 (-4.71)***					
2		0.81 (0.44)	-0.44 (-2.87)***			
3		3.81 (1.35)				
5		2.92** (1.91)		-0.30 (-2.37)***		

The equation for the model is:

$$\Delta \ln be_t = \alpha_t + \sum_{i=1}^{p} \beta_i \Delta \ln fe_{t-i} + \sum_{j=1}^{q} \delta_j \Delta \ln ere_{t-i} + \sum_{k=1}^{r} \mu_k \Delta \ln dped_{t-i} + \vartheta_1 dum\_u + \vartheta_2 dum\_t + \eta Ec_{t-1}$$

be= volume of Indonesian agricultural exports
ere= exchange rate weight
fe= foreign exchange rate weight
pe= index price of Indonesian agricultural exports
Ec= error correction variable
dum_u= intercept dummy
dum_t= slope dummy

- Δ is first differencing of the variables
-.Ec is error correction variable
Summary statistics:
R^2= 0.28 , Adjusted R^2 = 0.23
DW= 2.35
F= 4.17
χ^2(Jarque-Bera)=0.49
*** Significant at the 1% level.
** Significant at the 5% level.
* Significant at the 10% level.

Table 5. 25. Critical Values for Cointegration Test

Number of var N	Sample size T	Significance level 1%	5%	10%
1[a]	50	2.62	1.95	1.61
	100	2.60	1.95	1.61
	250	2.58	1.95	1.62
	500	2.58	1.95	1.62
	∞	2.58	1.95	1.62
1[b]	50	3.58	2.93	2.60
	100	3.51	2.89	2.58
	250	3.46	2.88	2.57
	500	3.44	2.87	2.57
	∞	3.43	2.86	2.57
2	50	4.32	3.67	3.28
	100	4.07	3.37	3.03
	200	4.00	3.37	3.02
3	50	4.84	4.35	3.73
	100	4.45	4.22	3.59
	200	4.35	4.18	3.47
4	50	4.94	4.35	4.02
	100	4.75	4.22	3.89
	200	4.70	4.18	3.89
5	50	5.41	4.76	4.42
	100	5.18	4.58	4.26
	200	5.02	4.48	4.18

[a] critical values of τ
[b] critical values of τ_μ
Both cited from Fuller (1976, p.373)

CHAPTER VI

CONCLUSIONS

Since the advent of current floating exchange rates, many authors analyzed the effects of exchange rates on the agricultural exports. Developing countries have received little attention. Therefore, any study that deals with developing countries will be an addition to the literature.

Previous approaches to study of the relationship between export volume and exchange rate have been based on inappropriate econometric procedures, in that relationships between non-stationary variables have been estimated. In this study, I have taken explicit account of this non-stationarity by employing unit root tests, multivariate cointegration, and error correction modeling with structural break.

Export demand functions, estimated on quarterly export data for five commodities (cocoa, coffee, palm oil, rubber, tea), and aggregate export are as for the 1971-1998 periods. Quantity of exports is hypothesized to depend upon foreign economic activity, own price, competitive price, and exchange rate. Each estimated model follows recently developed time-series tests with or with out structural breaks. The empirical results suggest the following conclusions.

First, the evidence strongly indicates the existence of a single unit root in virtually all variables at normal significance levels for all commodities and aggregate agricultural exports. This is true if they are tested with and without structural break for change in exchange rate policy. Results are consistent with the macroeconomic literature that most time-series data are non-stationary (Nelson and Plosser, 1982). Previous equations relating to non-stationary variables could be subject to the spurious regression phenomenon first described in Granger and Newbold (1974).

Second, the results suggest that there is a statistically significant long-run relationship between export volume and exchange rate, as well as relative price and foreign income, which is explained in a cointegration equation with structural break. In addition for each commodity exchange rate change has a negative short-run effect. These findings are similar to the results of other researchers (Nainggolan, 1987). The foreign income result indicates no short-run effect, since the commodities in the model are basic commodities, and the income of importers has little impact on the export and represents a small percentage of their import volume. The price of each commodity has a negative short-run effect in every case. The coefficient of the competitor price was positive and has a short-run effect on export volume. Importers of Indonesian agricultural commodities are able and willing to import from alternative sources if the price is more favorable than the Indonesian price.

Third, the results attributed the difference to the more appropriate way in which I have implemented the model for each commodity, which included the structural break for an exchange rate policy change. The results, therefore, provide different insights into the relationship between exchange rate and real exports. Moreover, the results suggest that a statically robust demand for exports can be estimated using error-correction dynamic specifications. This approach was found to reduce misspecification error.

Fourth, the major finding is that exchange rate changes have a statistically significant negative impact on the real export of all five commodities and aggregate exports. This finding is consistent with the theoretical considerations discussed in Chapter 4. Importing countries reduce their activities and switch to other sources of agricultural import supply to reduce the effect of the exchange rate risk.

6.1 Policy Implications

Because of the strong effect of real exchange rate on the agricultural exports, a policy that dampens the real exchange rate can enhance agricultural exports. This implies that if the domestic inflation rate is higher in domestic countries than the trading partners, the depreciation in the nominal exchange rate will be adjust instantly to maintain the competitiveness, encourage exports, helps correct the trade deficit. The expansionary monetary policy to increase the money supply seems less effective.

An expansionary monetary policy to influence the exchange rate alone is not sufficient to increase agricultural exports.

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