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ESSAYS IN THE ECONOMICS OF BANKING

Antonios Michis

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I would like to express my deep gratitude to my advisor Dr. Panayiotis Theodossiou for his guidance and inspiring transfer of knowledge during the writing of this thesis. Our long discussions on financial phenomena have tremendously helped me acquire a new way of thinking about financial management. Needless to say, this thesis is also dedicated to my parents. Through their encouragement and support, during all these years, they have helped me fulfill my personal dreams and aspirations. My gratitude to them is a lifetime commitment.

ΠΕΡΙΛΗΨΗ

Στην παρούσα διδακτορική διατριβή εξετάζονται διάφορες πτυχές που αφορούν τους προσδιοριστικούς παράγοντες των επιτοκίων στον τραπεζικό τομέα, χρησιμοποιώντας τεχνικές ανάλυσης που σχετίζονται με τα ερευνητικά πεδία της Βιομηχανικής Οργάνωσης και των Χρηματοοικονομικών. Η ανάλυση που ακολουθεί βασίζεται στην εκτίμηση ότι τα επιτόκια ισορροπίας, όπως αυτά διαμορφώνονται στο τραπεζικό τομέα, είναι το αποτέλεσμα μιας σειράς πολύπλοκων οικονομικών φαινομένων και αλληλεπιδράσεων στο οικονομικό σύστημα που δεν μπορούν να επεξηγηθούν επαρκώς από μόνο ένα ερευνητικό πεδίο των οικονομικών. Συγκεκριμένα, τρία θέματα εξετάζονται διεξοδικά: (1) η μέτρηση του βαθμού μονοπωλιακής δύναμης στην τραπεζική βιομηχανία και οι επιπτώσεις αυτής στα δανειστικά επιτόκια, (2) η χρήση μη γραμμικής τιμολόγησης από τις τράπεζες και η συσχέτιση της με αλλαγές στη δομή της αγοράς και (3) εκτίμηση του αποτελέσματος ρευστότητας, όπως αυτό εκδηλώνεται σε διαφορετικές χρονικές κλίμακες, ως αποτέλεσμα αλλαγών στη νομισματική πολιτική.

Στο κεφάλαιο 1 αναπτύσσεται ένα υπόδειγμα με βάση την προσέγγιση της "εικαστικής απόκλισης" (conjectural variations) προκειμένου να μετρηθεί ο βαθμός μονοπωλιακής δύναμης σε ένα τραπεζικό σύστημα. Σε αντίθεση με προηγούμενες μελέτες που έχουν γίνει για το θέμα αυτό, και οι οποίες υιοθετούν εκτενείς συναρτήσεις κόστους παραγωγής για τα τραπεζικά ιδρύματα, η παρούσα διατριβή ορίζει το οριακό κόστος ως κόστος ευκαιρίας. Αυτό αντιπροσωπεύεται από το επιτόκιο που προσφέρεται για τα ελάχιστα αποθεματικά (minimum reserves) που υποχρεωτικά διατηρούν οι τράπεζες ως καταθέσεις στις Νομισματικές Αρχές μιας οικονομίας. Όταν οι καταθέσεις αυτές υπερβαίνουν το ελάχιστο όριο αποθεματικών, θεωρούνται στην παρούσα μελέτη ως μια εναλλακτική χρήση των διαθέσιμων κεφαλαίων των τραπεζών, τα οποία συνήθως δίδονται ως δάνεια με σκοπό την κερδοφορία. Στην εμπειρική ανάλυση του κεφαλαίου, η προτεινόμενη μεθοδολογία χρησιμοποιείται προκειμένου να μετρηθεί ο βαθμός μονοπωλιακής δύναμης που υπάρχει στο τραπεζικό σύστημα της Κύπρου. Με βάση τα οικονομετρικά αποτελέσματα η υπόθεση ύπαρξης μονοπωλίου απορρίπτεται.

Στο κεφάλαιο 2 παρουσιάζεται ένα οικονομετρικό υπόδειγμα για στατιστικό έλεγχο της υπόθεσης ύπαρξης μονοπωλιακής, μη γραμμικής τιμολόγησης (nonlinear pricing) στην Ευρωπαϊκή αγορά δανείων. Ο στατιστικός έλεγχος αναπτύσσεται με βάση το υπόδειγμα Baumol για την ζήτηση ρευστότητας και χρησιμοποιεί ένα δείκτη βιομηχανικής

συγκέντρωσης προκειμένου να εξετάσει, πως αλλαγές στην δομή της αγοράς, σχετίζονται με την παροχή εκπτώσεων στο επιτόκιο καθώς το ύψος του δανείου μεγαλώνει (quantity discounts). Η εμπειρική οικονομετρική ανάλυση και η διεξαγωγή του ελέγχου γίνονται χρησιμοποιώντας ένα δείγμα από χρονικά επαναλαμβανόμενα, διαστρωματικά στοιχεία (panel data) και ένα οικονομετρικό υπόδειγμα σταθερών χρονικών και διαστρωματικών επιδράσεων (fixed effects model). Τα αποτελέσματα της ανάλυσης συνηγορούν στο συμπέρασμα ότι η εφαρμογή μη γραμμικής τιμολόγησης από τα νομισματικά και χρηματοπιστωτικά ιδρύματα σχετίζεται με την ύπαρξη μονοπωλιακής δύναμης στην Ευρωπαϊκή αγορά δανείων.

Το κεφάλαιο 3 εστιάζεται στην μελέτη του αποτελέσματος ρευστότητας (liquidity effect) και συγκεκριμένα στον τρόπο με τον οποίο αυτό εκδηλώνεται και μεταβάλλεται σε διαφορετικές χρονικές κλίμακες στην οικονομία του Ηνωμένου Βασιλείου. Η ανάλυση βασίζεται στην θεωρία προτιμήσεως ρευστότητας (liquidity preference theory) για την συμπεριφορά των επιτοκίων και χρησιμοποιεί την μεθοδολογία των κυματιδίων (wavelets) στα πλαίσια ενός σταθμισμένου υποδείγματος παλινδρόμησης. Από τα αποτελέσματα της ανάλυσης προκύπτει ότι, βραχυπρόθεσμα τα επιτόκια επηρεάζονται κυρίως από μεταβολές στην προσφορά χρήματος στην οικονομία (αποτέλεσμα ρευστότητας). Μεσοπρόθεσμα και μακροπρόθεσμα, το αποτέλεσμα ρευστότητας εξασθενεί και τα επιτόκια παρουσιάζουν μεγαλύτερη ευαισθησία στις μεταβολές που υπεισέρχονται στο εισόδημα και στο επίπεδο τιμών της οικονομίας. Η μέση εκτιμώμενη διάρκεια του αποτελέσματος ρευστότητας δεν ξεπερνά τους 8 μήνες.

ABSTRACT

This thesis examines different aspects of the determinants of interest rates in the banking industry by using techniques associated with the fields of industrial organization and financial economics. It proposes that equilibrium interest rates, as observed in the banking industry, are the complex outcome of many different market forces in the economic system that cannot be adequately explained by only one branch of financial economics. Three topics are examined in detail: (1) measurement of market power and its impact on lending interest rates, (2) the use of nonlinear pricing tactics by banks in order to improve profitability and the influence of market structure and (3) time-scale estimation of the liquidity effect in an economy, following monetary policy changes.

In Chapter 1 a conjectural variations model is developed to measure market power in the banking industry. Unlike previous studies, which use complete cost function specifications in the modelling framework, this study defines marginal cost based on an opportunity cost, which is represented by the interest rate on minimum reserves offered by the monetary authorities in a country. Deposits with the monetary authorities are considered to be an alternative use of available funds that are usually allocated to loans. The estimates of market power in the banking industry in Cyprus, using the proposed model, reject the monopoly hypothesis.

In Chapter 2 an econometric test is proposed to verify the existence of nonlinear pricing for loans in the European credit market. The test is based on Baumol’s model of cash demand and incorporates an industry concentration measure in order to examine the impact of market structure on the provision of quantity discounts (nonlinear pricing). The econometric results, using a panel dataset consisting of seven European countries and a fixed effects model specification, suggest that nonlinear pricing is associated with increasing monopoly power in the European banking industry.

Chapter 3 examines the existence of a liquidity effect in the UK economy over different time-scales. This analysis draws from the liquidity preference framework, an approach to interest rate determination, and uses wavelet multiscale analysis in the context of a standardised regression model. The results suggest that, in short-term cycles, interest rates are influenced primarily by changes in the money supply (i.e., the liquidity effect). In medium- and long-term cycles, the liquidity effect becomes less important and interest rates are found to be more sensitive to income and price effects. The average duration of the liquidity effect is estimated to be less than 8 months.

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ABBREVIATIONS

2SLS	2 Stage Least Squares
AC	Average Cost
ATMs	Automatic Teller Machines
BK	Baxter-King
CQ	Competitors’ Amount (Quantity) of Loans
CP	Competitors’ (Average) Price (Interest Rate)
COOPs	Cooperative Societies
DR	Deposit (Interest) Rate
DWT	Discrete Wavelet Transform
ECB	European Central Bank
EM	Efficient Markets
EOQ	Economic Order Quantity
EU	European Union
GDP	Gross Domestic Product
GMM	Generalized Method of Moments
HHI	Herfindahl-Hirschman Index
LB	Labour Wage
MC	Marginal Cost
MFIs	Monetary Financial Institutions
MODWT	Maximal Overlap Discrete Wavelet Transform
MR	Marginal Revenue
NEIO	New Empirical Industrial Organisation
OECD	Organization for Economic Cooperation and Development

OLS	Ordinary Least Squares
PC	Cost of Physical Capital
PMR	Perceived Marginal Revenue
S-C-P	Structure-Conduct-Performance
TR	Total Revenues
UK	United Kingdom
US	United States

INTRODUCTION

This study presents three essays that examine the determinants of interest rates in the banking industry and proposes that to properly examine such a complex subject it is necessary to use theorems and techniques associated with the fields of industrial organisation and financial economics. The banking industry, which is of fundamental global importance, presents some unique characteristics that distinguish it from other industries. Although traditionally associated with the field of financial economics, it is now commonly accepted that a proper analysis of the banking industry requires widespread knowledge regarding a combination of results from many subfields of economics and finance.

For example, the modelling and analysis of interest rate movements in the banking industry can be based on competitive models of pricing found in the industrial organisation literature, while, at the same time, the interest rates and the pricing of banking products are strongly influenced by the financial strategies of the banking institutions and the macroeconomic environment and monetary policies implemented within a country. Consequently, a proper examination and understanding of interest rate movements require theoretical knowledge as well as knowledge regarding the various techniques applied in the fields of industrial organisation, finance and macroeconomics.

The first essay (Chapter 1), using results from the new empirical industrial organisation literature, is concerned with measuring the degree of competition in a banking industry. The determinants of lending interest rates and econometric issues related to the estimation of own- and cross-interest rate elasticities are considered in detail. The measurement of competition in banking is a rapidly growing research field with many significant developments in recent years. Accordingly, these developments are reviewed in Subsection 1.2. To measure the degree of competition, an equilibrium interest rate model is developed from a representative bank's profit maximisation programme that incorporates conjectural derivatives, as such derivatives are commonly employed in conjectural variation models of competition.

An innovation associated with the model presented in Chapter 1 refers to the inclusion of an opportunity cost variable correlated with a bank's profit maximisation programme that represents an alternative use of funds that are usually allocated to loans. The opportunity cost variable is represented by the interest rate on minimum (or excess) reserves offered by the monetary authorities of a country. This is a risk-free, alternative investment for the excess reserves of banks. As a result of the inclusion of this variable in the profit maximisation programme, marginal cost is defined as the interest rate forgone on deposits that could be placed with the monetary authorities. Accordingly, the marginal cost is included in the first-order profit maximisation conditions. Consequently, marginal cost differs from most empirical studies in the conjectural variations literature, which define complete cost function specifications to measure the degree of competition in banking. This latter approach presents limitations in empirical applications as most available banking micro-data (e.g., from financial statements) include only aggregated cost variables that refer to the total cost of all operations associated with a bank (e.g., deposits, loans, credit card operations, underwriting services, foreign exchange services, etc.). As a result, the cost variables included in the associated econometric models that are commonly used to assess the level of competition complicate the analysis and cast doubt on the results.

While this approach of defining marginal cost based on an opportunity cost concept was first developed for a commodity market by Merel (2009), the existing study extends the modelling framework to incorporate competitive reactions (as in conjectural variations models) not included in Merel’s model. As a result, the explanatory power of the empirical econometric model is significantly enhanced. Empirical testing of the proposed method is performed in the context of a nonlinear, simultaneous equations system that consists of the following three equations (models): (1) a model for the equilibrium interest rate condition that includes the opportunity cost variable; (2) a model for the demand for loans; and (3) a model for competitive interest rates. An empirical study based on a panel dataset of 20 financial institutions operating in the Cyprus banking industry is presented and an estimation of the nonlinear, simultaneous equations system is performed using generalised method of moments. The results reject the monopoly hypothesis with respect to the Cyprus banking industry and instead suggest own- and cross-interest rate elasticities with a magnitude of approximately 2.

The second essay (Chapter 2) is also concerned with the application of industrial organisation methods in banking, particularly with respect to the pricing tactic of nonlinear pricing (a form of second-degree price discrimination). In fast moving consumer goods markets, this tactic is frequently used by manufacturers and concerns the provision of quantity discounts when higher quantities are purchased by the consumer. Regarding the theory of industrial organisation, this pricing tactic is associated with the exercise of monopoly power where a monopolist extracts additional consumer surplus by providing different package sizes of the same product. However, recent theoretical advances in the literature have shown that nonlinear pricing is also possible in oligopolistic markets where quantity discounts are used by manufacturers as a response to competition.

Even though there are a substantial number of empirical studies in the banking literature that provide evidence for the existence of different forms of price discrimination (reviewed in Subsection 2.2), thus far, no evidence regarding the existence of nonlinear pricing strategies has been found. Chapter 2 employs the Baumol model of cash demand to create a formal test for the existence of nonlinear pricing in European banking. An analysis of a panel dataset consisting of annual observations of seven European countries suggests that the interest rates on loans (cash demand) above the one million euro benchmark are consistently lower than the interest rates on loans below the one million euro benchmark.

The Baumol model of cash demand is based on the economic order quantity model that was originally developed for inventory management problems. This specific model has been extended in the management science literature to allow for the possibility of quantity discounts, and it is also used in this study to model higher levels of cash demand that involve quantity discounts. To construct a test for nonlinear pricing in the banking industry, two versions of the economic order quantity model are considered: (1) the standard version without quantity discounts for loans below the one million euro benchmark and (2) a version that incorporates quantity discounts for loans above the one million euro benchmark. The two models are solved independently to derive the optimum (cost minimising) quantities of cash demand. By replacing the optimum quantities in the total cost functions, it is possible to solve for the difference in interest rates when the representative agents from the two groups (above and below the one million euro benchmark) minimise the cost of cash demand.

Following the empirical literature on nonlinear pricing, the difference in interest rates (spreads) depends on market concentration and other market and country specific characteristics in econometric models. Econometric estimates (which also account for endogeneity problems in the model specification) suggest that higher levels of market concentration (increasing monopoly power) in European banking tend to be associated with higher levels of interest rate spreads (quantity discounts). Therefore, nonlinear pricing provides a way for banks with monopoly power in Europe to increase their sales of loans and extract additional consumer surplus, as suggested in the industrial organisation literature.

The third essay (Chapter 3) is concerned with the determinants of short-term interest rates in an economy (specifically, the 91-day interbank equilibrium rate) and the measurement of the liquidity effect, as suggested in the financial economics literature. The main innovation associated with this chapter is the development of an econometric methodology to measure the existence of a liquidity effect over different time-scales. To achieve this, the wavelet methodology is used to decompose the observed sampling rate of the model's variables into different time-scale components (frequency bands) that cover short-term, medium-term and long-term periodicities in the data.

According to the liquidity preference theory for the determination of interest rates, increases in the money supply by the monetary authorities of a country result in a reduction in interest rates (i.e., the liquidity effect). On the other hand, changes in demand result from income and price effects that are positively related to interest rates. The liquidity effect generally has an immediate influence on interest rates, while income and price effects can involve substantial time lags. This is because it takes time before the increases in the money supply impact and increase prices and income in the economy. Consequently, to study the overall impact of a money supply increase on interest rates, it is important to estimate the impact of each effect over different time-scales.

In Chapter 3, a standardised, time-scale regression model is proposed that permits (1) the identification of the impact of each effect on short-term interest rates by time-scale; (2) comparisons of the magnitudes of the three effects; and (3) the measurement of the duration (in months) of the liquidity effect before interest rates begin to rise due to the income and price effects. Such an analysis is useful for investors as it facilitates an understanding of movements in interest rates as well as for predicting the impact of monetary policy changes on interest rates. The dependent variable in the model consists of the 91-day interbank equilibrium rate, and three variables, available with monthly regularity, are used as independent variables: the money supply (associated with the liquidity effect), inflation (associated with the price effect) and the industrial production index (associated with the income effect). By standardising the time-scale regression models so that all of the variables in a given time-scale are set on a common scale of measurement, it is possible to compare the three effects within each time-scale.

The proposed standardised time-scale regression model is empirically tested using data from the United Kingdom economy. The results suggest the existence of a liquidity effect with an average duration of approximately 8 months following a money supply increase. When moving to longer cycles, interest rates begin to rise in response to the income and price effects. An estimation of the proposed model is performed with the two-stage least squares method, using two lags for each one of the independent variables as instruments, as it is expected that the three independent variables are jointly determined (endogenous) and are therefore correlated with the error term of the model.

In Chapter 3, some additional econometric issues related to the proposed model are considered. First, an augmented model with additional financial variables is estimated. The additional variables refer to the exchange rate of the sterling against the dollar and the short-term interest rate in the US economy. Due to the size and importance of the US economy in the overall global economy, it is possible that developments in the US economy likely influence the interest rates in other countries. Therefore, it is of interest to examine whether the above-mentioned estimates of the liquidity effect are robust to an alternative model specification. Estimation results with the augmented model confirm the validity and robustness of the initial analysis.

Second, the Baxter-King filter is used to isolate the different time-scale components of the model's variables and to estimate the coefficients. This filter constitutes an interesting alternative to the proposed wavelet decomposition method and provides a useful basis for comparisons to better understand the advantages and limitations of wavelet analysis when applied to financial data. An estimation of the standardised model is again performed with the two-stage least squares method, but the results in this case are inconclusive. The liquidity effect is estimated with shorter duration (4 months), and in most time-scales, the results are not conclusive because some of the coefficients do not have the expected sign. As explained in Subsection 3.6, wavelets possess several key advantages as bases functions that enable a more accurate estimation of the liquidity effect by time-scale.

1 Measuring market power in the banking industry in the presence of opportunity cost

1.1 Introduction

developments in the industrial organisation literature. The measurement of market power and the development of competitive models for the banking industry using advances in the New Empirical Industrial Organisation literature (NEIO) have particularly generated important insights and improved understanding of several complicated issues that are of great importance to policy makers and regulators.

Initial efforts to model competition in the banking industry have relied mainly on the structure-conduct-performance (S-C-P) model and the relationship between interest rates and market concentration. The limitations associated with this approach led to the development of more sophisticated methods of modelling competition, most notably the Panzar-Rosse and conjectural variations approaches. Research using these approaches is still active and also addresses several econometric issues associated with the empirical measurement of market power. More recent research has adopted alternative approaches that can provide useful insights for competitive banking markets, such as discrete choice models and structural models of entry, which are reviewed along with other methods in Section 1.2 of this chapter.

In this study, a conjectural variations model for loans is developed, incorporating opportunity cost in the form of the interest rate on minimum reserves offered by the monetary authorities of a country. The model is derived from a profit maximisation programme, and the equilibrium price condition is incorporated into the perceived marginal revenue function proposed by Bresnahan (1982). This function depends on the opportunity cost, as in Merel (2009), and not on a complete cost function specification.

The proposed model has two important advantages relative to the existing literature: (i) Conjectural price reaction elasticities are incorporated in the modeling framework. As a result, the explanatory power of the simultaneous equations system, derived from the theoretical model, is significantly enhanced. (ii) Estimates of market power are based on an opportunity cost concept and not on a complete cost function specification. This is desirable in banking since most empirical studies estimate complete cost functions using micro-data from financial statements. These data are frequently characterized by measurement errors. As a result, many of the estimated coefficients are not consistent with economic theory and are sensitive to model specification changes.

The proposed conjectural variations model is used to estimate the degree of market power in the banking industry in Cyprus, and the results are compared to two alternative methods proposed in the literature. The first of these is based on the conjectural variations model proposed by Coccorese (2005), and the second is based on the Panzar-Rosse revenue test proposed by Delis et al. (2008). All estimates reject the monopoly hypothesis for the banking industry in Cyprus. The proposed conjectural variations model and the model of Coccorese suggests a competitive equilibrium while the revenue test suggests a market structure close to monopolistic competition.

The remainder of this chapter is organised as follows. Section 1.2 provides a detailed literature review of banking market power models and Section 1.3 introduces the conjectural variations model. Sections 1.4 and 1.5 develop two conjectural variations models using an opportunity cost concept. Section 1.6 addresses the empirical applications of the study, presents the data and estimates the models using data from the banking industry in Cyprus. Section 1.7 concludes the chapter with a discussion of the results and the associated policy implications.

1.2 Empirical measurement of market power in the banking industry

This section reviews several methods for measuring market power in the banking industry. Cabral (2000, p. 6) defines market power as the ability to set the price of a product or service above the marginal cost, suggesting that the lower the level of competition faced by a bank in a country, the higher its market power. The section begins with the S-C-P and efficient markets (EM) models, which represent the first attempts at measuring competition in the banking literature. It then proceeds to describe some of the NEIO methods that have been used for the banking industry, such as the Panzar-Rosse and conjectural variations approaches to measuring market power.

Some more recent methods based on alternative modelling frameworks are also presented. Each method is presented in a separate sub-section, which starts with a description of the methodology and an explanation of its main drawbacks. This is then followed by an examination of the main empirical studies that have been carried out with the method and their findings.

The structure-conduct-performance model

Description: The systematic study of market power and its determinants in the banking industry began in parallel with the development of the S-C-P model in the industrial organisation literature. In a banking context, this model suggests that changes in the market structure of a geographical area (e.g., country or state) affect the way banks structure their economic policies and consequently their performance. Market structure is usually represented by a concentration measure, and the leading hypothesis is that higher concentration (structure) is associated with more market power for banks (Cabral, 2000, p. 157). As a result, banks are able to set higher interest rates for loans and lower interest rates for deposits (conduct), thereby increasing their profits (performance).

The existence of market power ensures that firms can be inefficient without being forced out of the market, as would be the case in a competitive setting. An underlying assumption of the S-C-P model is that there are barriers to entry in the market (which generally characterise the banking industry), and few banks are able to raise lending interest rates as in a Cournot oligopoly setting (see, Heffernan, 2005, p. 495).

Drawbacks: There are four main shortcomings associated with this model:

1. Market power is measured indirectly through market concentration. It is assumed that higher concentration will almost surely lead to collusion. In many banking markets this is not true.
2. The model assumes that all firms in the market have constant and equal expectations concerning their rivals' reactions to changes in output. It is assumed that higher concentration will generate a common reaction by all firms, thus leading to collusive behaviour (see, Bikker and Bos, 2008). This is also an unrealistic assumption since individual bank strategies can differ considerably.
3. In many banking markets, increasing market power can have an opposite effect on interest rates. An alternative model for measuring the influence of market power on interest rates is the EM model proposed by Demsetz (1973). This model assumes that some banks are more efficient than others (e.g., due to superior management or technology) and, as a

result, are able to offer lower lending interest rates and higher deposit rates than competitors. These banks are able to progressively increase their market shares (thereby increasing concentration) and profits. In contrast to the S-C-P model, this model predicts that lending interest rates will decrease with increasing market concentration (Heffernan, 2005, p. 495).

4. A common problem encountered in studies examining price-concentration relationships, is the possible endogeneity problem associated with measures of market concentration (see, Evans et al., 1993). Endogeneity bias can result, either from the inverse effect of prices on concentration, or from errors in the measurement of both price and concentration, which are difficult variables to quantify in the banking industry.

Empirical studies: Empirical tests of the two models (S-C-P and EM) and their associated hypotheses are usually conducted by considering econometric models in which the dependent variables are either profits or interest rates. The independent variables usually include a measure of market concentration, such as the Herfindahl-Hirschman index, along with other control variables. The following is a simple regression representation of these tests:

$$r_{it} = \beta_0 + \beta_1 CONC_{it} + \sum_{k=2}^{K} \beta_k \chi_{itk} + \varepsilon_{it} . \qquad (1.1)$$

The dependent variable (r_{it}) is the interest rate on loans or the interest rate offered on deposits. The first independent variable represents a measure of market concentration ($CONC_{it}$) and the model includes additional control variables (χ_{itk}) that can vary across different banks (i) and periods (t).

If the dependent variable represents lending interest rates, then, according to the S-C-P model, the coefficient of the concentration variable (β_1) will be positive, implying that increasing market concentration is associated with higher interest rates. If the dependent variable represents deposit interest rates, then the coefficient will be negative. Different coefficients should be expected if the EM model holds. There will be a negative coefficient between concentration and lending interest rates and a positive coefficient between concentration and deposit interest rates. These cases are summarised in Table 1.1.

Table 1.1: Expected concentration coefficients: S-C-P and EM models

Model	Concentration - Lending interest rates	Concentration - Deposit interest rates
S-C-P	Positive	Negative
EM	Negative	Positive

Several empirical studies have tested these two models in banking markets, primarily using data from the United States (US). Berger and Hannan (1989) used the quarterly deposit data of 470 banks, from 1983 to 1985, to estimate the relationship between deposit interest rates and market concentration in the US. The Herfindahl-Hirschman index and the three firm concentration ratio were used in the analysis. The authors were able to find a negative and statistically significant relationship between market concentration and deposit interest rates in the majority of the metropolital areas examined; this relationship is consistent with the S-C-P model.

Most of the studies in this field adopted similar econometric methods in order to test the validity of the S-C-P and EM hypotheses. However, apart from interest rates some authors used different but equally informative dependent variables in econometric models such as: profits (Molyneux and Forbes, 1995), interest rate margins - the difference between lending and money market rates - (Covoisier and Gropp, 2002) and the Lerner Index (Angelini and Cetorelli, 2003).

Most empirical studies tend to confirm the validity of the S-C-P model. Notable studies in this direction include Berger and Hannan (1992), Molyneux and Forbes (1995), Goddard, Molyneux and Wilson (2001), Covoisier and Gropp (2002) and Angelini and Cetorelli (2003). There is considerably less empirical support for the EM hypothesis. The most important studies in support of the EM hypothesis were provided by Jackson (1992), Berger (1995), Goldberg and Rai (1996) and Angelini and Cetorelli (2003).

The Panzar-Rosse method

Description: The Panzar-Rosse (1987) method measures market power by examining how changes in input prices affect the equilibrium revenues of firms in an industry. The sensitivity of revenues to changes in input prices depends on the level of competition in the market and is calculated as the elasticity of revenues (R) with respect to a vector of input prices (see, Dick and Hannan, 2010, p. 410):

$$H_R = \sum_{i=1}^{k} \frac{\partial R}{\partial W_i} \frac{W_i}{R} . \tag{1.2}$$

The revenue function is defined as $R = f(W_1,...,W_k, Z, Y)$ and includes an exogenous vector Z of cost determinants, an exogenous vector Y of demand determinants and k input prices W. In the context of a monopoly, an increase in input prices (and therefore in marginal costs) reduces the revenues of the monopolist; it should therefore be expected that $H_R < 0$. This is because increases in marginal cost, reduce the level of output that the monopolist is willing to produce. In contrast, in a perfectly competitive market, revenues will increase proportionally with marginal cost and it should be expected that the Panzar-Rosse statistic will be equal to unity: $H_R = 1$.

Drawbacks: The Panzar-Rosse approach, although useful, is also characterised by some important limitations:

1. Values that are close to unity ($H_R = 1$) should be expected to hold only in the long-run (see, Dick and Hannan, 2010, p. 410). In contrast, most studies assume that there is no time lag in the adjustment of interest rates (they are assumed to change in parallel with changes in revenues) and that entry and exit by other firms in the market occurs within the same period.
2. Negative values of the Panzar-Rosse statistic can also arise in cases that are not associated with monopolistic market structures. Heffernan (2005, p. 508) describes one such case in the context of a perfectly contestable market.
3. Additional limitations include: possible endogeneity issues in the estimation procedure and correctly accounting for all the input factors in the estimation of the elasticity of revenues.

Empirical evidence: Shaffer (1982) estimated the Panzar-Rosse statistic in a study of New York banks. He obtained a value of 0.318, which suggests that this specific market is neither a monopoly nor a perfectly competitive market but something between these two extremes.

Nathan and Neave (1989) examined a cross section of Canadian banks and estimated the Panzar-Rosse statistics for 1983 and 1984. They obtained positive values of the statistic that were different from unity, which led them to reject the monopoly power hypothesis.

Similar values were also obtained for European markets. Molyneux et al. (1994) analysed banking data from Germany, the UK, Italy and Spain for the period from 1985 to 1989. They concluded that the market structure in these countries is compatible with monopolistic competition. An exception to the previous findings is the study by Molyneux et al. (1996) on the Japanese banking market, which did not reject the monopoly hypothesis. More recently, Claessens and Laeven (2004) estimated the Panzar-Rosse statistic for fifty countries and then proceeded to examine the regulatory and country-specific characteristics that influence competitiveness. Their results suggest that countries with fewer restrictions on entry and a greater presence of foreign banks tend to be associated with more intense competition.

The conjectural variations method

Description: This method is based on the empirical conjectural variations model, proposed by Bresnahan (1982) and Lau (1982) that allows for strategic interactions among firms (see, Varian, 1992, p. 302-303). This model considers the assumptions (conjectures) that a firm might have concerning the reactions of other firms to its price or quantity decisions. For example, in the case of a quantity (q_A) change decision by bank A, attention is paid to the conjectural derivative $\partial q_B / \partial q_A$, which quantifies the reaction of competitive bank B.

The first order condition for profit maximisation is more complex in this framework relative to when there are no strategic interactions among firms ($\partial q_B / \partial q_A = 0$). It takes the following form (Dick and Hannan, 2010, p. 411):

$$P = MC(Q,Z) - D_1(Q,Y)\,Q\,\lambda . \tag{1.3}$$

In this equation, the price (P) of the product is equal to the marginal cost (MC) minus the derivative of the inverse demand (D_1) multiplied by the quantity (Q) and the conduct parameter (λ). The conduct parameter is closely related with the conjectural derivative (as will be explained in Section 1.3) and can be estimated using a system of equations that includes: the demand equation associated with the financial product, a cost equation and the first order condition mentioned above. The conduct parameter will have a value of zero ($\lambda = 0$) in a perfectly competitive market (because the marginal cost will be equal to the price of the product) and a value of one ($\lambda = 1$) in a monopoly. Varying degrees of imperfect competition will produce values between zero and one ($0 < \lambda < 1$).

Drawbacks: Although this is a promising approach for measuring market power in banking, it also presents some limitations.

1. Banks hold arbitrary conjectures about their rivals.
2. It is assumed that banks choose their output decisions simultaneously and only once and there are no strategic interactions among multiple periods.
3. It is sensitive to the correct specification of the demand function.
4. It is sensitive to the geographic definition of a market, as is frequently the case in banking (see, Dick and Hannan, 2010, p. 412).

A more detailed analysis and discussion of the conjectural variations model is included in the next section. It will provide the basis for developing the opportunity cost model in Section 1.3.

Empirical studies: This approach was used by Shaffer (1989, 1993) to study market power in the US and Canadian banking industries. Based on the results, the collusion hypothesis was rejected in both countries.

Other methods

This section concludes with some more recent methods that have been proposed in the literature. While interesting and promising, more empirical research is needed in order to evaluate their usefulness.

Discrete choice models: One family of models that has been used with promising results in banking is that of random coefficients discrete choice models, as proposed by Berry et al. (1995). These models were used by Dick (2008) in an effort to estimate a structural model for deposits and to identify how the deregulation of branching networks in the US during the nineties affected consumer welfare. The empirical results did not find any evidence of a reduction in consumer welfare as a result of the deregulation. Another application of these models was provided by Adams et al. (2007), who compared the deposit demand of banks and thrift institutions and found that they cannot be considered to be close substitutes.

Structural entry models: Another area of industrial organisation applications in banking involves structural entry models, such as those used in Bresnahan and Reiss (1987). In these models, the emphasis is on the relationship between price, the number of firms in the market and the size of the market necessary to accommodate the entry of new firms. The level of competition in the market is indirectly identified by estimating entry thresholds, which refer to the size of the market per bank when an additional bank enters the market (Dick and Hannan, 2010, p. 416).

The main idea characterising these models is that when a new bank enters the market, price competition will increase and banks will be expected to realise lower variable profits as a result. Consequently, the size of the market per bank (entry thresholds) must increase to provide the opportunity for variable profits to cover fixed cost. If estimates indicate that entry thresholds have increased, then it can be inferred that price competition has also increased. Structural entry models have been used in a banking context by Cetorelli (2002) and Cohen and Mazzeo (2007).

Boone indicator: In a promising article, Boone (2008) proposed a new way for measuring competition that combines elements of the structural entry and efficiency hypothesis models. Based on this measure, competition can be intensified in two ways: (i) through a fall in the barriers to entry that characterise an industry and (ii) through more aggressive market interactions between the participating firms. In the first case, lower entry barriers in an industry are associated with higher firm participation and therefore with more intense competition. In the second case, more aggressive interactions between firms force inefficient firms out of the market and market concentration increases as a result. By capturing both effects, this competitive measure is shown to be more robust than the price-cost margin.

In empirical applications, the profit elasticity provides a way of capturing both effects and can be estimated using log-linear models (see, Degryse et al. 2009, p. 36): $\log \pi_i = a + \beta \log c_i$. In this model, the profit elasticity (β) is estimated as the coefficient of the marginal cost variable (c_i). More efficient firms with lower marginal cost will be expected to realize higher profits (π_i) and therefore $\beta < 0$. Consequently, higher absolute values of this negative coefficient will be associated with stronger bank competition. In contrast, the existence of market power (as in monopolies) ensures that firms can be inefficient without being forced out of the market. In this case the absolute value of the profit elasticity will be low. Van Leuvensteijn et al. (2007) used the Boone indicator to measure banking competition in several countries. They found considerable variability among different markets. As suggested by Degryse et al. (2009), more research is needed in order to properly evaluate the usefulness of the Boone indicator.

1.3 The conjectural variations model in banking

The conjectural variations approach to oligopoly modelling was first introduced in Section 1.2. Here it is presented in more detail together with a discussion of its advantages and disadvantages. In sub-sections 1.4 and 1.5 it will be used in order to develop methodologies for measuring market power using an opportunity cost concept.

Conjectural variations models incorporate a firm’s strategic concerns into the modelling framework. This is achieved by considering the beliefs or expectations (conjectures) that a firm might have, concerning the reactions of other firms to its price and quantity decisions. These strategic reactions are captured by the partial derivative $\partial q_j / \partial q_i$ which quantifies the strategic reaction of bank j (the rival) to a quantity change by competitive bank i. To see this, consider a homegeneous product duopoly where banks compete over quantities and face identical costs. The profit maximization program for bank i is ($p = price$, $q = quantity$, $C = total\ cost$)

$$\max \Pi_i = p_i . q_i - C_i . \qquad (1.4)$$

The first-order condition for profit maximization after allowing for strategic interactions among firms is

$$\frac{\partial \Pi_i}{\partial q_i} = p_i + q_i \left[\frac{\partial p_i}{\partial q_i} + \frac{\partial p_i}{\partial q_j} \frac{\partial q_j}{\partial q_i} \right] - MC_i = 0 . \tag{1.5}$$

Letting $v_i = \partial q_j / \partial q_i$ (bank i's conjecture) and using straightforward algebra, the first-order condition can be written as follows (see, Chrurch and Ware, 2000, p. 272)

$$p_i + \frac{\partial p}{\partial q_i} [1 + v_i] q_i = MC_i . \tag{1.6}$$

Within this framework four different models are nested, depending on the value of the conjectural variations paparameter v_i:

1. <u>Cournot model</u>: $v_i = 0$ (bank i treats bank j as fixed in its decisions).
2. <u>Bertrand model</u>: $v_i = -1$ (bank i believes that an increase in its quantity is exactly offset by a decrease in the quantity of bank j, so price remains unchanged).
3. <u>Cartel model</u>: $v_i = 1$ (bank i believes that an increase in its quantity is matched by an equal increase in the quantity of bank j, as in collusion).
4. <u>Conjectural variations model</u>: $v_i \neq 0$ (bank i believes that an increase in its quantity will generate an unequal response by bank j).

Another related approach for estimating market power uses the conduct parameter in marginal revenue models (see, Belleflamme and Peitz 2010, p. 70):

$$MR_i = p_i + \lambda \frac{\partial p_i}{\partial q_i} q_i . \tag{1.7}$$

Depending on the value of the conduct (or market power) parameter λ, different market structures can be represented as follows: $\lambda = 0$ - competitive market, $\lambda = 1$ - monopoly market and $\lambda = 1/n$ - symmetric Cournot model with n firms. Since in equilibrium

$MR = MC$, by equating the two equations above and using the symmetry assumption ($q_i = q/n$), it is possible to establish a relationship between the conjectural variations parameter and the conduct parameter: $\lambda = (1+v_i)/n$. The equilibrium condition can also be written with reference to the Lerner index. Let $MR = MC$ and re-arrange the equilibrium condition (1.7) as follows

$$p_i - MC_i = -\lambda \frac{\partial p_i}{\partial q_i} q_i . \tag{1.8}$$

Dividing the left-hand side with p_i provides the following expression for the Lerner index.

$$L = \frac{p_i - MC_i}{p_i} = -\lambda \frac{\partial p_i}{\partial q_i} \frac{q_i}{p_i} = \frac{\lambda}{n} . \tag{1.9}$$

The Lerner index is equal to the ratio of the conduct parameter (λ) to the price elasticity of demand (n). Consequently, the conduct parameter provides an index for the degree of market power (see, Belleflamme and Peitz 2010, p. 71).

Assuming profit-maximization in production, the Lerner index measures a firm’s degree of monopoly power based on the difference between price and marginal cost. Therefore, its sole emphasis is on price competition, which suggests two important limitations. First, it completely ignores other forms of monopoly behaviour such as product differentiation strategies. Second, the difference between price and marginal cost might not be related with monopoly power. Instead, it can be related with increasing returns in production or the need to cover fixed costs (see, Elzinga and Mills, 2011).

With regards to the conjectural variations model, there are two basic disadvantages:

1. The beliefs (conjectures) that banks hold about other banks are arbitrary and contradict with game-theoretic models.

2. The analysis is static and excludes any considerations of strategic interactions among multiple periods. It is therefore assumed that banks choose their output decisions simultaneously and only once.

Carlton and Perloff (2005, appendix of Ch. 6) consider point 1 less problematic, simply because arbitrary assumptions are also used in game-theoretic models of oligopolies. The second point is certainly more problematic and has been the subject of serious criticism. Vives (2001, p. 185) considers conjectural variations models as an attempt to consider dynamics within the framework of a static model, which is conceptually ambiguous. On the contrary game theoretic models are based on credible strategies and equilibria.

Despite these disadvantages, the conjectural variations approach is characterised some important practical advantages. First, it provides a framework for empirically measuring market power through the estimation of the conduct parameter. Greater values of this parameter imply greater exercise of market power. Bresnahan (1982, 1989) and Lau (1982) clarified several econometric issues associated with the correct estimation of market power in structural models, such as the presence of endogenous variables and the fact that the conduct parameter is the ratio of two estimated parameter.

Second, this thesis proposes that conjectural derivatives and price reaction elasticities are useful quantities in econometric models. This is because they provide substantial explanatory power in simultaneous systems of demand and price equations. For this reason they are used in the context of Bresnahan's empirical framework in order to measure market power in banking. As will be explained in the next section, such measurements need not rely on complete cost function specifications for which micro-data from financial statements are subject to measurement errors. Instead, an opportunity cost concept will be used.

Most studies that estimate conjectural variations models, tend to use cost equations with several proxies[1] in the context of simultaneous equations systems that also include demand and equilibrium price equations. As a result, many of the estimated coefficients are not consistent with economic theory and are sensitive to model specification changes. For

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[1] For example in Coccorese (2005) the cost of physical capital is approximated by the value of operating cost (excluding deposit and labour expenses) divided by the total amount of funds under management. Canhoto (2004) admits that the "price of capital" is a controversial topic and conveniently defines it as a portion of the operating cost.

example, Canhoto (2004) obtained a negative wage rate coefficient in an equation for the marginal cost. She used balance sheet data from Portuguese banks and attributed this unexpected result to data quality problems.

This thesis develops a model with conjectures with respect to price. Therefore strategic reactions are captured by the partial derivative $\partial p_j / \partial p_i$, which quantifies the strategic price reaction of bank j to a price change by competitive bank i. The profit maximization program in this case is

$$\max \Pi_i = p_i q_i - C_i(q_i, w_i) \tag{1.10}$$

where ,

$C_i(q_i, w_i)$ = cost function of bank i.

q_i = quantity demanded from bank i.

p_i = price charged by bank i.

w =vector of input prices.

The first-order condition is

$$\frac{\partial \Pi_i}{\partial p_i} = q_i + \left(p_i - MC_i\right)\left(\frac{\partial q_i}{\partial p_i} + \frac{\partial q_i}{\partial p_j}\frac{\partial p_j}{\partial p_i}\right) = 0. \tag{1.11}$$

Rearranging this first order condition yields the following equality (see, Coccorese, 2005)

$$\frac{p_i - MC_i}{p_i} = -\frac{1}{\varepsilon_{ii} + \lambda \varepsilon_{ij}(p_i / p_j)} \tag{1.12}$$

where

$\varepsilon_{ii} = -(\partial q_i / \partial p_i)(p_i / q_i)$ is the own-price elasticity of demand, $\varepsilon_{ij} = +(\partial q_i / \partial p_j)(p_j / q_i)$ is the cross-price elasticity of demand and $\lambda = \partial p_j / \partial p_i$ is the conjectural derivative. The profit function for bank i in (1.10) can also be written as follows

$$\pi = q_i(.)p_i - C_i(q_i(.),\omega_i)\,. \tag{1.13}$$

In this model $q_i = f(p_i, p_j, \chi_i, a)$ is defined as the quantity demanded. It is a function of the price charged for the financial product (p_i), the price index for competition (p_j), an exogenous variable (χ_i) and a parameter of the demand system that will be estimated (a).

In a seminal paper, Bresnahan (1982) demonstrated how demand identification can be achieved in order to consistently estimate the market power parameter. To see this, note that when banks are not price takers the perceived marginal revenue function derived from the quantity demanded has the following form

$$p_i = c_i(q_i, w_i, \beta) - \lambda.h_i(q_i, \chi_i, a) + \eta_i\,. \tag{1.14}$$

The marginal revenue is $p_i + h_i(.)$. The marginal cost, c_i, includes the supply side exogenous variables w_i and the parameter β. The demand side exogenous variables and parameters are in $h(.)$ and λ is the market power parameter. If $\lambda = 1$ then a monopoly exists. Other values of this parameter correspond to different market structures as explained above.

The problem in applied practice is to estimate a simultaneous equations system with equations for q_i and p_i as defined above. Bresnahan (1982) showed that a linear demand equation and the market power parameter can be identified by rotating as well as shifting the demand equation. In such a setting, the estimated market power parameter corresponds to the conjectural variations elasticity. In a banking context, Shaffer (1989) proposed using the interaction term variable $p_i.p_j$ in order to rotate and the national income variable Y in order to shift the demand equation.

A similar strategy will be followed in this study. In addition to the interaction term ($p_i.p_j$), which will be used as an elasticity shifter in the estimation procedure, the level of income in the economy will be used as an exogenous demand shifter in the econometric analysis of Section 1.6. Bresnahan (1982) emphasises that elasticity and demand shifters are necessary to change (rotate and shift) the demand slope in such a way that distinguishes between the hypotheses of competition and monopoly. Joint movements in the two exogenous variables enable the estimation of the degree of market power. In a perfectly competitive market, the rotations will not change the equilibrium price, but the opposite is true if market power exists.

1.4 A conjectural variations model with opportunity cost

In this section, a conjectural variations model that incorporates opportunity cost in the form of interest income forgone by the provision of loans is developed. This is a reasonable assumption in banking markets because the total deposits received by banks in a period are usually allocated into two channels. The first is the provision of loans to economic agents in the market and the second is the placement of deposits with the monetary authorities of the country. At a minimum, these deposits should be equal to the minimum (or required) reserves requested by the monetary authorities. As emphasised by Mishkin and Eakins (1998, p.325), banks also hold additional amounts of reserves with the monetary authorities because these deposits (called excess reserves) are considered to be the most liquid assets in a bank's balance sheet. In this way, these deposits are readily available to meet unexpected obligations that may arise, especially in case of large withdrawals.

The idea of incorporating opportunity cost in models that examine competition in banking was also proposed by Heffernan (2002), who used the Libor rate in generalised pricing econometric models. In these models, the Libor rate is used as a proxy for the deposit rate forgone by a bank when it provides an additional monetary unit of loans instead of withholding it as part of its deposits in another institution. In this study an alternative opportunity cost concept is used: the interest rate that is offered by monetary authorities on minimum reserve deposits (see, Heffernan, 2005, p.290). In Europe, banks in countries that are members of the Eurosystem (as is the case for Cyprus) are required to keep a minimum reserves ratio as defined by the European Central Bank. This ratio is calculated based on their short-term liabilities and banks earn interest on these reserves that is equal to the average rate of the weekly tenders over the maintenance period (Matthews and Thompson, 2008, p. 54).

The interest rate on minimum reserves is used in a conjectural variations model that incorporates competitive responses. Mérel (2009) used a similar modelling framework for a commodities market by incorporating the opportunity cost of the revenue forgone by not producing a substitute product in the context of a profit maximisation programme. However, his study did not consider competitive responses, as is common in conjectural variations models. This section develops one such model for banking that also incorporates an opportunity cost measure in the form of the interest rate on minimum reserves.

An advantage of the proposed model is that it does not require a complete cost function specification to estimate the market power parameter in a system of equations. This is a frequently encountered problem in empirical studies because financial statements and other banking industry data sources do not separate expenses by product or service category (e.g., operating costs attributed to loans and deposits). As a result, it is difficult to accurately measure bank output (see, for example, Heffernan, 2005, p. 476), which introduces several problems for econometric estimation.

Consider the following profit maximisation programme for bank i:

$$\max_{p} \Pi_i = p_i\, q_i + k_i\,(D_i - q_i) - C_i\,(q_i, \omega_{1i}, \omega_{2i}, \omega_{3i}, z_i, D_i - q_i)\,. \tag{1.15}$$

The total loan revenues are equal to the quantity demanded $q_i = f(p_i, p_j, \chi_i)$ multiplied by the price charged for the financial product p_i. The quantity demanded is also a function of a price index for competition p_j (e.g., competitors' weighted average interest rate) and a vector of exogenous factors χ_i. The total loans provided by the bank are derived from the deposits it receives ($D_i = s_i + q_i$) after subtracting the amount placed by the bank as an interest-bearing deposit (equal to or in excess of the minimum reserves) with the monetary authorities (s_i). Similarly, the cost function $C_i(q_i, \omega_{1i}, \omega_{2i}, \omega_{3i}, z_i, D_i - q_i)$ depends on several factors: the loan output q_i; input factors such as interest paid on deposits (ω_{1i}), labour wages (ω_{2i}) and the cost of physical capital (ω_{3i}); and the deposits placed with other institutions ($D_i - q_i$), which also involve a managing cost. The deposits at other institutions are associated with an interest rate (k_i) that also contributes to profitability. This will constitute the opportunity cost concept in the analysis and will be measured by the interest rate on minimum reserves. This profit maximisation programme is similar to that used by

Coccorese (2005), who also estimated a conjectural variations model. However, this model includes the opportunity cost variable, which further complicates the first-order profit maximisation conditions.

The partial derivative of the profit function with respect to price is as follows:

$$\frac{\partial \Pi}{\partial p_i} = q_i + p_i \frac{\partial q_i}{\partial p_i} + p_i \frac{\partial q_i}{\partial p_j}\frac{\partial p_j}{\partial p_i} - k_i \frac{\partial q_i}{\partial p_i} - k_i \frac{\partial q_i}{\partial p_j}\frac{\partial p_j}{\partial p_i} - \frac{\partial C_i}{\partial q_i}\frac{\partial q_i}{\partial p_i} - \frac{\partial C_i}{\partial q_i}\frac{\partial q_i}{\partial p_j}\frac{\partial p_j}{\partial p_i}.$$

Setting this expression equal to zero and letting $\frac{\partial C_i}{\partial q_i} = MC_i$ provides the first order condition:

$$q_i + \left[p_i - k_i - MC_i\right]\left[\frac{\partial q_i}{\partial p_i} + \frac{\partial q_i}{\partial p_j}\frac{\partial p_j}{\partial p_i}\right] = 0. \tag{1.16}$$

This can be written in a price-cost margin format as:

$$p_i - k_i - MC_i = -\frac{q_i}{\frac{\partial q_i}{p_i} + \frac{\partial q_i}{\partial p_j}\frac{\partial p_j}{\partial p_i}}. \tag{1.17}$$

To obtain an equilibrium condition in terms of own and cross-price elasticities, as is common in conjectural variations models, both sides of the above equation are multiplied by q_i / p_i. In addition, the second term in the denominator of the right side is multiplied by $p_j / p_j = 1$. The equation then becomes:

$$\frac{p_i - k_i - MC_i}{q_i}\,\frac{q_i}{p_i} = -\frac{1}{\frac{\partial q_i}{p_i}\frac{p_i}{q_i} + \frac{\partial q_i}{\partial p_j}\frac{\partial p_j}{\partial p_i}\frac{p_i}{q_i}\frac{p_j}{p_j}}. \tag{1.18}$$

When written in terms of elasticities, the equation is equal to:

$$p_i - k_i - MC_i = -\frac{p_i}{\varepsilon_{ii} + \varepsilon_{ij}\, n_{ji}}. \tag{1.19}$$

The own-price elasticity is $\varepsilon_{ii} = -\frac{\partial q_i}{p_i}\frac{p_i}{q_i}$, the cross-price elasticity is $\varepsilon_{ij} = +\frac{\partial q_i}{p_j}\frac{p_j}{q_i}$ and the competitive price-reaction elasticity is $\eta_{ji} = +\frac{\partial p_j}{\partial p_i}\frac{p_i}{p_j}$. This elasticity shows how price changes by bank i generate a competitive price reaction by competitor j (which also influences the market price).

In Coccorese (2005), the equilibrium condition (1.19) does not include the interest rate on minimum reserves on the left side and only two elasticities are included on the right side. Instead of the elasticity η_{ji}, a conjectural variations parameter $\lambda = \partial p_j / \partial p_i$ is defined and is multiplied by p_i / p_j. However, estimation of this parameter in the context of equation (1.19) requires a system of equations with a cost function specification, an approach not adopted by this study due to the problems associated with cost measurement in banking. Leaving only marginal cost on the left side and rearranging provides the following equilibrium condition:

$$MC_i = p_i\left(1 + \frac{1}{\varepsilon_{ii} + \varepsilon_{ij}\,\eta_{ji}}\right) - k_i. \tag{1.20}$$

To introduce a market power parameter in the modelling framework, a perceived marginal revenue function is defined, as proposed by Bresnahan (1982) and used by Shaffer

(1989) and Mérel (2009) in the banking and commodities markets, respectively. The perceived marginal revenue function is

$$PMR_i = (1-h)\, p_i + h\, MR_i\,. \tag{1.21}$$

Equation (1.21) includes the market power parameter h (also referred to as the conduct parameter in the literature). Values of this parameter that are close to zero ($h = 0$) indicate a competitive industry while values that are close to one indicate monopoly power ($h = 1$). Accordingly, intermediate values represent equilibria between these two extremes. Because profit maximisation in equilibrium (see, Pepall et al., 2008, p. 247) implies equality between marginal revenue and marginal cost ($MR = MC$), expression (1.20) can be substituted into (1.21):

$$PMR_i = p_i - h\, p_i + h\, p_i \left(1 + \frac{1}{\varepsilon_{ii} + \varepsilon_{ij}\, \eta_{ji}}\right) - h\, k_i\,. \tag{1.22}$$

By cancelling out equal terms, the equation becomes:

$$PMR_i = p_i \left(1 + \frac{h}{\varepsilon_{ii} + \varepsilon_{ij}\, \eta_{ji}}\right) - h\, k_i\,. \tag{1.23}$$

Another expression for marginal revenues can be obtained by taking the partial derivative of the profit function with respect to the quantity demanded:

$$\frac{\partial \Pi}{\partial q_i} = p_i + \frac{\partial p_i}{\partial q_i} q_i - k_i - \frac{\partial C_i(q_i,.)}{\partial q_i} + \frac{\partial C_i(.,D_i - q_i)}{\partial q_i}\,.$$

The first partial derivative of the cost function $\frac{\partial C_i(q_i,.)}{\partial q_i}$ refers to the first quantity expression q_i in the cost function $C_i(q_i, \omega_{1i}, \omega_{2i}, \omega_{3i}, z_i, D_i - q_i)$ and, as is the case of the second partial derivative, which refers to the deposits with the monetary authorities $D_i - q_i$ that are also included in the cost function. The first two terms in the above first order condition are equal to the marginal revenue $MR_i = p_i + \frac{\partial p_i}{\partial q_i} q_i$ (see, for example Cabral, 2000, p. 25). Equating the first order condition to zero and substituting for the marginal revenues expression gives:

$$MR_i = k_i + \mu . \tag{1.24}$$

The term $\mu = \frac{\partial C_i(q_i,.)}{\partial q_i} - \frac{\partial C_i(., D_i - q_i)}{\partial q_i}$ in equation (1.24) is the difference between the two partial derivatives of the cost function, and according to Mérel (2009), represents the difference in marginal cost associated with the two alternative uses of inputs. In the context of this study, the deposits collected by bank i are directed to two alternative uses: (i) for the production of loans and (ii) as deposits held with the monetary authorities. Consequently, according to equation (1.24), in equilibrium, the marginal revenue from producing an additional unit of loans is equal to the marginal revenue forgone if the unit was deposited with the monetary authorities (the opportunity cost k_i) plus the difference in marginal cost (μ_i) between producing an additional unit of loans instead of managing an additional unit of deposits. By equating the PMR in (1.23) with the MR in (1.24), the following price equation is obtained, which is to be used in a nonlinear system of equations to estimate the market power parameter:

$$p_i \left(1 + \frac{h}{\varepsilon_{ii} + \varepsilon_{ij} \eta_{ji}} \right) = h k_i + k_i + \mu . \tag{1.25}$$

In addition to equation (1.25), the nonlinear system of equations will also include a demand function for bank i based on the previously defined specification for the quantity demanded $q_i = f(p_i, p_j, \chi_i)$. Following Coccorese (2005), a log linear model will be used as follows:

$$\log q_i = a_0 + a_1 \log p_i + a_2 \log p_j + a_3 \log Y + a_4 \log p_i \log p_j + \varepsilon_q . \tag{1.26}$$

However, in addition to the interaction term ($\ln p_i \ln p_j$), which is to be used as an elasticity shifter in the estimation procedure, equation (1.26) also includes the exogenous variable Y that represents the level of income in the economy. Bresnahan (1982) emphasises that elasticity and demand shifters are necessary to change (rotate and shift) the demand slope in such a way that distinguishes between the hypotheses of competition and monopoly. Rotation is achieved through the new exogenous variable p_j (competitors' price index) in the demand equation, which enters interactively with the price variable p_i, whereas shifts in the demand function are achieved through the income variable Y. Joint movements in the two exogenous variables enable the estimation of the degree of market power. In a perfectly competitive market, the rotations will not change the equilibrium price, but the opposite is true if market power exists.

Equations (1.25) and (1.26) can be estimated as a nonlinear system of equations. However, unlike the studies by Coccorese (2005) and Mérel (2009), the equilibrium price condition (1.25) additionally includes the competitive price elasticity $\eta_{ji} = \frac{\partial p_j}{\partial p_i} \frac{p_i}{p_j}$ that captures the price reaction of competition when bank i changes the price of its product. To estimate this elasticity, an additional equation will be included in the system of equations, in log-linear format, using the competitors' price index as the depended variable:

$$\log p_j = \gamma_0 + \gamma_1 \log p_i + \gamma_2 \log EURIB_{it} . \tag{1.27}$$

The competitive price elasticity will be estimated using parameter γ_1. Equation (1.27) also includes the Euribor rate ($EURIB_{it}$) as an independent variable, following the analysis of Heffernan (2002), which included the Libor rate in econometric models for lending interest rates in the UK. It is to be used as a proxy for the term $k_i + \mu_i$ in equation (1.25), which is the interest rate on minimum reserves plus the difference between the marginal cost of producing an additional unit of loans and the marginal cost of managing an additional unit of deposits placed with the monetary authorities.

1.5 An augmented conjectural variations model with opportunity cost

Consider the following profit maximization program of a representative bank that accepts deposits (D) and aims to maximize profits from loan (q^L) and security investment (I) operations

$$\max_{p^L} \Pi = p_i^L.q_i^L + p_i^I.I_i + k_i(q_i^D - q_i^L - I_i) - p_i^D.q_i^D - C_i(q_i^L, q_i^D, I_i, s_i, ...). \qquad (1.28)$$

The profit maximization program includes the following variables:

$q^D = deposits$
$q^L = loans$
$s =$ minimum $(or\ excess)$ $reserves$
$I = investments\ in$ securities
$k =$ interest $rate\ on$ minimum $reserves$
$p^L =$ nterest $rate\ on\ loans$
$p^D =$ nterest $rate\ on\ deposits$
$p^I =$ securities $market\ rate$
$OA = other\ assets$
$OL = other\ liabilities.$

The loan demand function facing the bank is $q_i^L = f(p_i^L, p_j^L, \chi_i)$, as defined in the previous sub-section. The balance sheet identity of the bank is $q^L + I + s + OA = q^D + OL$ and the reserves held with the monetary authorities are defined as $s = q^D - q^L - I$. Hence, reserves are defined as residual deposits after loan provisions and security investments and satisfy the minimum reserves requirements set by the monetary authorities. In this study, reserves held with the monetary authorities are considered as an alternative use of funds that are usually allocated to loans.

With regards to cost, an additive cost function is assumed. It consists of separate cost components for each principal bank operation plus fixed costs

$$C(q^L, q^D, I\ ,...) = C^L(q^L, s\ ,...) + C^D(q^D,...) + C^I(I\ ,...) + FC\,. \tag{1.29}$$

$C(q^L, q^D, I\ ,...) = \text{total cost}$
$C^L(q^L, s\ ,...) = \text{operating cost of loans}$
$C^D(q^D,...) = \text{operating cost of deposits}$
$C^I(I,...) = \text{operating cost of security investments}$
$FC = \text{fixed cost}$

Substituting the balance sheet identity ($I = -q^L - s - OA + q^D + OL$) into the profit function, the maximization program becomes

$$\begin{aligned}\max_p \Pi = {} & p_i^L.q^L + p_i^I.(-q_i^L - s_i - OA_i + q_i^D + OL_i) + k_i(q_i^D - q_i^L - I_i) - p_i^D.q_i^D \\ & - C_i(q_i^L, q_i^D{}_i, I_i, s_i,...).\end{aligned} \tag{1.30}$$

Subject to the profit and cost separability conditions stated above, Canhoto (2004) showed that the profit function can be rearranged into three components that can be maximised separately:

$$\pi_i^L = (p_i^L - p_i^I - k_i)q_i^L - C_i^L(q_i^L, s_i,...)$$
$$\pi_i^D = (p_i^I + k_i - p_i^D)q_i^D - C_i^D(D_i,...)$$
$$\bar{\pi} = p_i^I(OL_i - OA_i) - k_i I_i - C_i^I(I_i,...) - FC_i\,.$$

In order to maximise the profit function for loans, π^L, the partial derivative with respect to price is developed using the chain-rule

$$\frac{\partial \pi^L}{\partial p_i^L} = q_i^L + p_i^L \frac{\partial q_i^L}{\partial p_i^L} + p_i^L \frac{\partial q_i^L}{\partial p_j^L}\frac{\partial p_j^L}{\partial p_i^L} - k_i \frac{\partial q_i^L}{\partial p_i^L} - k_i \frac{\partial q_i^L}{\partial p_j^L}\frac{\partial p_j^L}{\partial p_i^L} - p_i^I \frac{\partial q_i^L}{\partial p_i^L} - p_i^I \frac{\partial q_i^L}{\partial p_j^L}\frac{\partial p_j^L}{\partial p_i^L}$$
$$- \frac{\partial C_i^L}{\partial q_i^L}\frac{\partial q_i^L}{\partial p_i^L} - \frac{\partial C_i^L}{\partial q_i^L}\frac{\partial q_i^L}{\partial p_j^L}\frac{\partial p_j^L}{\partial p_i^L}.$$

Setting this expression equal to zero and letting $\frac{\partial C_i^L}{\partial q_i^L} = MC_i^L$ provides the first order condition

$$q_i^L + \left[p_i^L - k_i - p_i^I - MC_i^L\right]\left[\frac{\partial q_i^L}{\partial p_i^L} + \frac{\partial q_i^L}{\partial p_j^L}\frac{\partial p_j^L}{\partial p_i^L}\right] = 0\,. \tag{1.31}$$

This can be written in a price-cost margin format

$$p_i^L - k_i - p_i^I - MC_i^L = -\frac{q_i^L}{\frac{\partial q_i^L}{\partial p_i^L} + \frac{\partial q_i^L}{\partial p_j^L}\frac{\partial p_j^L}{\partial p_i^L}}\,. \tag{1.32}$$

In order to obtain an equilibrium condition in terms of own and cross-price elasticities, both sides of the above equation are multiplied by q_i^L / p_i^L. In addition, the second term in the denominator of the right hand side is multiplied by $p_j^L / p_j^L = 1$. The equation then becomes

$$\frac{p_i^L - k_i - p_i^I - MC_i^L}{q_i^L}\frac{q_i^L}{p_i^L} = -\frac{1}{\frac{\partial q_i^L}{\partial p_i^L}\frac{p_i^L}{q_i^L} + \frac{\partial q_i^L}{\partial p_j^L}\frac{\partial p_j^L}{\partial p_i^L}\frac{p_i^L}{q_i^L}\frac{p_j^L}{p_j^L}}\,. \tag{1.33}$$

It can also be written in terms of the following elasticities: own-price elasticity $\left(\varepsilon_{ii}^{L} = -\frac{\partial q_i^L}{\partial p_i^L}\frac{p_i^L}{q_i^L}\right)$, cross-price elasticity $\left(\varepsilon_{ij}^{L} = +\frac{\partial q_i^L}{\partial p_j^L}\frac{p_j^L}{q_i^L}\right)$ and competitive price reaction elasticity $\left(\eta_{ji}^{L} = \frac{\partial p_j^L}{\partial p_i^L}\frac{p_i^L}{p_j^L}\right)$.

$$p_i^L - k_i - p_i^I - MC_i^L = -\frac{p_i^L}{\varepsilon_{ii}^L + \varepsilon_{ij}^L\,\eta_{ji}^L}\,. \tag{1.34}$$

Leaving only marginal cost on the left hand side and rearranging, provides the following equilibrium condition

$$MC = p_i^L\left(1 + \frac{1}{\varepsilon_{ii} + \varepsilon_{ij}\,\eta_{ji}}\right) - k_i - p_i^I\,. \tag{1.35}$$

The perceived marginal revenue function with the market power parameter h is

$$PMR_i^L = (1-h)\,p_i^L + h\,MR_i^L\,. \tag{1.36}$$

Profit maximization implies the equality between marginal revenue and marginal cost ($MR = MC$). Substituting (1.35) into (1.36), the perceived marginal revenue function becomes

$$PMR_i^L = p_i^L - h\,p_i^L + h\,p_i^L\left(1 + \frac{1}{\varepsilon_{ii}^L + \varepsilon_{ij}^L\,\eta_{ji}^L}\right) - h\,k_i - hp_i^I\,. \tag{1.37}$$

The partial derivative of the profit function with respect to the loan quantity demanded is as follows

$$\frac{\partial \pi^L}{\partial q_i^L} = p_i^L + \frac{\partial p_i^L}{\partial q_i^L} q_i^L - k_i - p_i^I - \frac{\partial C_i(q_i^L,.)}{\partial q_i^L} + \frac{\partial C_i^L(.,D_i - q_i^L - p_i^I)}{\partial q_i^L} .$$

Setting this first order condition equal to zero and substituting for the marginal revenue expression $MR_i^L = p_i^L + \frac{\partial p_i^L}{\partial q_i^L} q_i^L$ gives

$$MR_i^L = k_i + p_i^I + \mu \ . \qquad (1.38)$$

Similarly with Section 3, $\mu = \frac{\partial C_i(q_i^L,.)}{\partial q_i^L} - \frac{\partial C_i^L(.,D_i - q_i^L - p_i^I)}{\partial q_i^L}$ represents the difference in marginal cost associated with the two alternative lending options: provision of loans to households and enterprises or holding excess reserves with the monetary authorities. Equating the PMR_i^L in (1.37) with the MR_i^L in (1.38), the following equilibrium price equation is obtained

$$p_i^L\left(1 + \frac{h}{\varepsilon_{ii}^L + \varepsilon_{ij}^L\,\eta_{ji}^L}\right) = (1+h)(k_i + p_i^I) + \mu \, . \qquad (1.39)$$

This equation is similar to equation (1.25), however, it differs is one important aspect. The opportunity cost in this case consists of two sources: (i) the interest rate on minimum reserves (k_i) and (ii) the security market rate (p_i^I). Both can be considered as alternative uses of funds usually allocated to loans.

Following a similar line of reasoning it is possible to derive an equilibrium price equation for deposits rates. To see this consider the partial derivative

$$\frac{\partial \pi^{D}}{\partial p_i^{D}} = -D_i + p_i^{I}\frac{\partial q_i^{D}}{\partial p_i^{D}} + p_i^{I}\frac{\partial q_i^{D}}{\partial p_j^{D}}\frac{\partial P_j^{D}}{\partial p_i^{D}} + k_i\frac{\partial q_i^{D}}{\partial p_i^{D}} + k_i\frac{\partial q_i^{D}}{\partial p_j^{D}}\frac{\partial P_j^{D}}{\partial p_i^{D}} - p_i^{D}\frac{\partial q_i^{D}}{\partial p_i^{D}} - p_i^{D}\frac{\partial q_i^{D}}{\partial p_j^{D}}\frac{\partial P_j^{D}}{\partial p_i^{D}}$$
$$- \frac{\partial C_i^{D}}{\partial q_i^{D}}\frac{\partial q_i^{D}}{\partial p_i^{D}} + \frac{\partial C_i^{D}}{\partial q_i^{D}}\frac{\partial q_i^{D}}{\partial p_j^{D}}\frac{\partial P_j^{D}}{\partial p_i^{D}}$$

Setting this expression equal to zero and letting $\frac{\partial C_i^{D}}{\partial q_i^{D}} = MC_i^{D}$, provides the first order condition

$$p_i^{I} + k_i - p_i^{D} - MC_i^{D} = \frac{q_i^{D}}{\frac{\partial q_i^{D}}{\partial p_i^{D}} + \frac{\partial q_i^{D}}{\partial p_j^{D}}\frac{\partial p_j^{D}}{\partial p_i^{D}}} \,. \qquad (1.40)$$

Solving this equation in terms of $p_i^{I} + k$ and substituting into (1.32) provides the following equation for the interest rate spread

$$p_i^{L} - p_i^{D} = MC_i^{D} + \frac{q_i^{D}}{\frac{\partial q_i^{D}}{\partial p_i^{D}} + \frac{\partial q_i^{D}}{\partial p_j^{D}}\frac{\partial p_j^{D}}{\partial p_i^{D}}} + MC_i^{L} - \frac{q_i^{L}}{\frac{\partial q_i^{L}}{\partial p_i^{L}} + \frac{\partial q_i^{L}}{\partial p_j^{L}}\frac{\partial p_j^{L}}{\partial p_i^{L}}} \,. \qquad (1.41)$$

The lending-deposit interest rate spread depends on the following factors: (a) the total level of marginal cost in the financial intermediation process ($MC_i^{D} + MC_i^{L}$) and (b) the quantities demanded for loans (q_i^{L}) and deposits (q_i^{D}), adjusted by their interest rate sensitivities.

1.6 Econometric analysis

1.6.1 Data and preliminary analysis according to the S-C-P model

The empirical testing of the modelling framework proposed in Section 1.4 was conducted using data from the banking industry in Cyprus. The data were collected from the financial statements of twenty financial institutions that offer banking services in the country. Commercial banks can be separated in two groups: (a) those with an ultimate controlling affiliate in Cyprus and (b) those with an ultimate controlling affiliate abroad. In addition to traditional banks, which primarily offer retail, corporate and investment banking services, the sample includes cooperative societies (COOPs), which have developed similar products during the last decade and can be considered direct competitors of traditional banks in Cyprus. For simplicity, in the remainder of the analysis, all financial institutions in the sample will be referred to as banks.

Table 1.2 provides information on the lending market shares of the top six banks in Cyprus, which account for 80% of the market. Two major commercial banks in the market have market shares ranging from 16% to 24%. The total market share of COOPs ranges from 19% to 20%. Most COOPs are local in character, and, according to their constitutions, are allowed to provide services only within the boundaries of the municipality in which they are established. However, they have a central authority (the Cooperative Central Bank), which takes over their economic and financial policies towards competition and administers a common network of automatic teller machines (ATMs). For example, the ATM network of a particular COOP can be used by the subscribers of any other COOP, irrespective of geographical location.

Table 1.2: Lending shares of major financial institutions in Cyprus

Financial Institution	Dec 2008	Dec 2009	Dec 2010	Dec 2011
Bank of Cyprus Ltd	24%	24%	24%	24%
Cooperative Societies	19%	19%	20%	20%
Marfin Popular Bank Public Co Ltd	16%	16%	17%	16%
Hellenic Bank Ltd	7%	7%	7%	7%
Alpha Bank Cyprus Ltd	8%	8%	8%	7%
Eurobank EFG Cyprus Ltd	2%	5%	4%	3%
Total	76%	79%	80%	77%

Source: Central Bank of Cyprus

To estimate market power in the banking industry of Cyprus based on the modelling method outlined in Section 1.4, several variables were collected or constructed from the financial statements of the banks used in the sample: the total value of loans for each bank i in year t (Q_{it}); the interest rate on loans for each bank i in year t (P_{it}), obtained by dividing total interest income from loans by total loans, as suggested by Coccorese (2005); the (loan value based) weighted average lending interest rate of the competitors of bank i in year t (CP_{it}); and the total loans included in the assets of the competitors of bank i in year t (CQ_{it}).

Table 1.3: Summary statistics

Variable	Sample	Mean	St dev	Min	Max
Loans (million EUR)	216	1843	4991	31	27725
Interest rate on loans (%)	216	7.138	1.722	2.026	11.739
Competitors inter. rate on loans (%)	216	6.531	1.393	1.748	9.357
Interest rate on deposits (%)	216	3.952	1.103	1.624	7.010
Cost of physical capital (EUR)	216	1.246	1.323	0.017	9.136
Euribor (%)	216	2.893	1.263	0.931	4.571
Interest rate on minimum reserves (%)	216	2.395	1.064	0.546	3.689
Gross domestic product (million EUR)	216	13769	2931	9008	17761
Total cost (million EUR)	161	114.419	298.628	1.254	1864
Average cost (EUR)	161	0.062	0.023	0,03	0,214
Labour cost (EUR)	161	25573	18313	6965	72300
Total Revenues (million EUR)	161	865	2995	1046	19744
Deposits to total assets ratio (%)	161	81.053	17.074	6.726	98.937
Assets (million EUR)	161	3867	7559	79.783	42638
Herfindahl-Hirschman Index (%)	206	10.231	0.707	9.380	11.700

Banks with ultimate controlling affiliate in Cyprus:
Bank of Cyprus Ltd (2000-2011)
Marfin Popular Bank Public Co Ltd (1999-2011)
Hellenic Bank Ltd (1999-2011)
Cyprus Development Bank Ltd (1999-2010)

Banks with ultimate controlling affiliate abroad:
Universal Bank Public Ltd (1999-2011)
Alpha Bank Cyprus Ltd (1999-2011)
Eurobank EFG Cyprus Ltd (2008-2011)
Emporiki Bank Cyprus Ltd (2007-2010)
Piraeus Bank (Cyprus) Ltd (2008-2010)

Cooperative societies:
Strovolos (1999-2010)
Paralimni (1999-2010)
Makrasyka (2001-2010)
Lemesos (1999-2010)
Lefkosia (1999-2010)
Aradippou (2000-2010)
Nea Ledra (1999-2011)
Central Coop (1999-2011)
Lakatamia (1999-2009)
Latsia (2000-2011)
Stadyl (2000-2011)

For comparison purposes, measures of market power in the banking industry of Cyprus were also obtained by using two additional models. The first is the conjectural variations model proposed by Coccorese (2005). Because this method requires the full specification of a cost function, additional cost information was collected from financial statements: the total cost of each bank i in year t (C_{it}); the average cost of each bank i in year t (AC_{it}), calculated by dividing total cost by the total value of loans; a proxy variable for the per unit cost of physical capital of bank i in year t (ω_{1it}), calculated by dividing operating cost (excluding interest expenses and labour costs) by funds under management, as in Coccorese (2005); the average labour cost of each bank i in year t (ω_{2it}), calculated by dividing total labour expenses (salaries, social security and pension fund contributions plus any other staff expenses) by the total number of employees; and the interest rate paid on deposits (ω_{3it}), calculated by dividing total interest expenses on deposits by the total amount of deposits reported by each bank i in the financial statements of year t.

The second model considered in Section 1.6 is based on a revenue test model proposed by Delis et al. (2008) to estimate the Panzar-Rosse statistic, which additionally includes the following variables: total assets ($Assets_{it}$); the ratio of total deposits to total assets for each bank i and year t ($Deposits_{it} / Assets_{it}$) and total revenues ($TR_{it}$).

Additional variables obtained from other sources included the interest rate on minimum reserves in year t (R_t), calculated from data in the annual reports of the Central Bank of Cyprus by dividing the interest expenses of the Central Bank by the total deposits placed with the Central Bank by banks; the gross domestic product of the country in year t at current market prices (Y_t), published by the Cyprus Statistical Service; the Herfindahl-Hirschman index of market concentration (HHI_{it}) for the banking industry of Cyprus, published by the European Central Bank (ECB), based on the assets of the banks operating in the country from 1999 to 2010; and the Euribor rate (euro interbank offered rate) in year t ($EURIB_t$), published by the European Banking Federation and upon which banks in the European Union (EU) exchange funds. The financial statements used in the dataset covered the period from 1999 to 2011; however, the data were not available for all the periods in the analysis for several banks. In several cases, this was because some banks began operating in Cyprus after

2005. The available annual observations for each bank are included in Table 1.3 along with summary statistics for the variables used in the analysis.

To investigate the relationship between lending interest rates and market concentration according to the S-C-P model, a log-linear fixed effects model was estimated as follows:

$$\log P_{it} = \beta_0 + \sum_{i=1}^{19} \delta_i d_i + \beta_1 \log CP_{it} + \beta_2 \log EURIB_{it} + \beta_3 \log HHI_{it} + \beta_4 \log Y_{it} + \beta_5 \log T + \varepsilon \tag{1.42}$$

The lending interest rate of bank i in year t is a function of competitive interest rates, the Euribor rate, which is used as a proxy for the average competitive interest rate in the EU, the Herfindahl-Hirschman index of market concentration, the gross domestic product, and a time trend variable (T). In addition, model (1.42) includes indicator variables for each bank in the sample to be used as fixed effects.

Table 1.4: Log-linear interest rate model

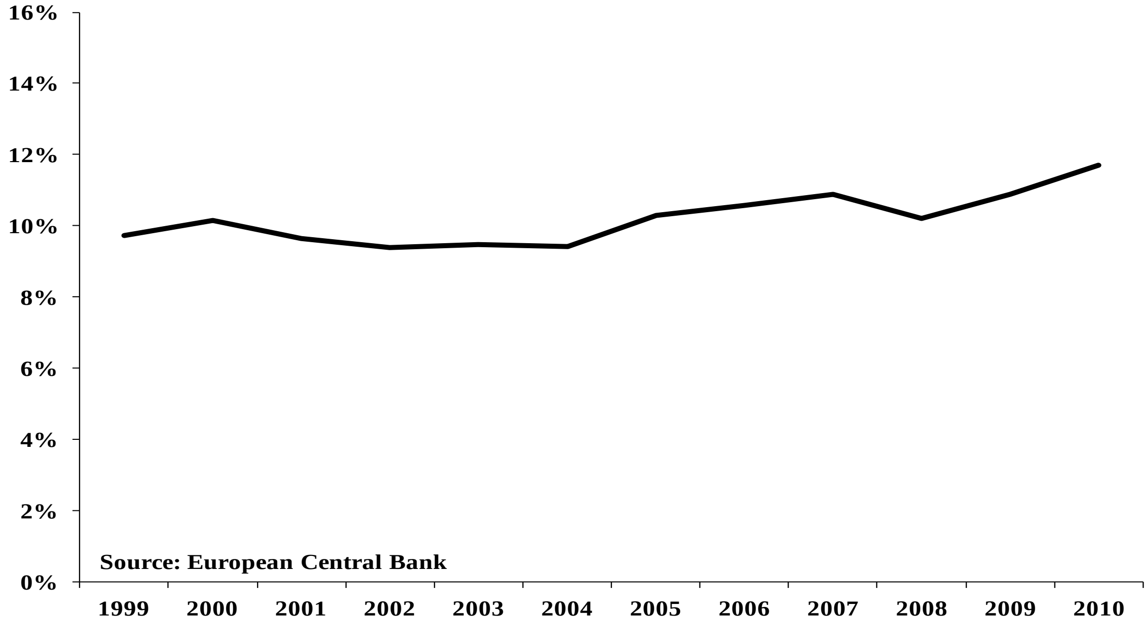
Coefficient	Estimate	St error	t test
Intercept	10.273	2.977	3.450*
log HHI	0.565	0.293	1.930*
log Y	-0.769	0.239	-3.210*
log EURIB	0.047	0.027	1.740**
log CP	0.160	0.078	2.050*
log T	0.055	0.063	0.870
R squared:	0.667		

Sample size: 206

*Significance level: *5%; **10%*

The coefficient estimates from the OLS method are included in Table 1.4. All the coefficients are statistically significant at the 5% or 10% level and have the expected sign, except for the coefficient of the trend variable. The coefficient for the Herfindahl-Hirschman index of market concentration is positive and suggests that a 1% increase in market concentration is associated with a 0.565% increase in lending interest rates. This finding is consistent with the S-C-P model, according to which increasing market concentration is associated with increasing market power, resulting in higher prices (lending interest rates) and higher profits. Conversely, increasing competition in the market (decreasing market concentration) is associated with declining interest rates.

Figure 1.1 exhibits the trend of the Herfindahl-Hirschman index during the period from 1999 to 2000. The overall market concentration level is low and has not changed significantly during the decade covered by the available data, remaining mostly around the 10% level. Small increases are observed after 2005, particularly in 2007 and 2010. This analysis, similarly to most empirical S-C-P tests, cannot offer clear conclusions with regards to the level of competition in the industry. It is therefore necessary to estimate a market power parameter as explained in Section 1.4.



Source: European Central Bank

Figure 1.1: Herfindahl-Hirschman index for the banking industry in Cyprus (assets)

1.6.2 Estimation of the conjectural variations model

To measure market power in the banking industry of Cyprus, this section proceeds with the estimation of the nonlinear system consisting of equations (1.25) - (1.27). Using the variable notation introduced in sub-section 1.6.1, the system (model 1) to be estimated can be represented as follows:

$$\begin{aligned}\log Q_{it} &= a_0 + a_1 \log P_{it} + a_2 \log CP_{it} + a_3 \log Y_t + a_4 \log P_{it} \log CP_{jt} + a_5 Coop_{it} \\ &\quad + a_6 Foreign_{it} + a_7 \log T + \varepsilon_Q\end{aligned}$$

$$P_{it} = \beta_0 + hR_{it} - h\left(\frac{P_{it}}{\varepsilon_{ii} + \varepsilon_{ij}\eta_{ji}}\right) + \beta_1 EURIB_{it} + \beta_2 \log T + \varepsilon_P \tag{1.43}$$

$$\log CP_{it} = \gamma_0 + \gamma_1 \log P_{it} + \gamma_2 \log EURIB_{it} + \gamma_3 \log CQ_{it} + \gamma_4 \log T + \varepsilon_{CP}$$

The system is subject to the following parameter restrictions: $a_1 = \varepsilon_{ii}$, $a_2 = \varepsilon_{ij}$, $n_{ji} = \gamma_1$, $0 \le h \le 1$ and all interest rate variables have been deflated as suggested by Coccorese (2005). The first equation is similar to equation (1.26) except that it includes a time trend (T) variable and two indicator variables in the list of explanatory variables. The first indicator variable takes the value of 1 when the observations in the sample refer to a cooperative society ($Coop_i$). Similarly, the second indicator variable takes the value of 1 when the observations in the sample refer to a bank with a foreign ultimate controlling affiliate ($Foreign_{it}$).

Based on these specifications, the control group, which refers to banks with an ultimate controlling affiliate in Cyprus, is captured by the intercept. Because we are using a panel dataset in the estimation procedure, it is possible to use separate indicator variables (fixed effects) for each bank in the sample. However, due to the limited number of observations and the unbalanced nature of the panel dataset, it was not possible to obtain individual coefficient estimates for all the banks in the sample because the generalised method of moments (GMM) estimator did not converge[2]. The classification based on the control status of the banks is sensible because the banks within each group have historically been similar in terms of size and strategic orientation.

In the second equation, the Euribor rate has been included as a proxy for the term $k_i + \mu_i$ in equation (1.25), which is equal to marginal revenues. It consists of the interest rates

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[2] Estimation with the IGMM, N3SLS, N2SLS and FIML methods was also tested but without success.

on minimum reserves (k) and the difference in marginal cost (μ) associated with the two alternative uses of deposits (provision of loans or deposits at the Central Bank). In a competitive market, the marginal revenues of a bank should reflect movements in the Euribor rate, as suggested by Heffernan (2002) for the relationship between lending rates and the Libor rate in the UK economy.

The third equation has been included in the system in log-linear form to estimate the competitive price elasticity $\eta_{ji} = (\partial p_j / \partial p_i)(p_i / p_j)$ that enters the equilibrium price condition in the second equation. The weighted average index of the competitive interest rates is considered to be a function of four independent variables: the lending rate offered by bank i, which is not included in the calculation of the index (for each cross section of observations in the panel dataset); the Euribor rate, which is again used as a proxy for the competitive interest rate in the EU; and the total amount of loans offered by all banks included in the calculation of the index except for bank i. The equation also includes a trend variable. The terms $\varepsilon_Q, \varepsilon_P$ and ε_{CP} represent the errors in the system of equations.

It is possible that the CQ_{it} variable is correlated with the error term of the model in the third equation because a general decline in the total loans in the market might generate a reduction in the lending interest rates charged by banks to recover. However, because the total amount of loans comprises data collected from nineteen different banks, it is unlikely that it will exhibit variation that represents a common policy reaction by all banks. To test for the possibility of endogeneity bias in the results, a Haussman endogeneity test (see Cameron and Trivedi, 2005, p.271) was constructed and a second version of the model (model 2) was estimated without the use of this variable. In addition to the exogenous variables included in each equation, the per unit cost of physical capital was used as an instrumental variable for CQ_{it} in the estimation of model 1 and was also used as an additional instrumental variable in the estimation of model 2. This variable is expected to be correlated with the total value of loans in the market because it is one of the input factors used by banks in the production process.

Table 1.5: GMM estimation of nonlinear conjectural variations models

	Model 1			Model 2		
Coefficients	**Estimate**	**St error**	**t test**	**Estimate**	**St error**	**t test**
Equation 1: Demand for loans						
Intercept	-56.429	22.155	-2.550*	-42.881	21.409	-2.000*
log P	-2.087	0.976	-2.140*	-1.819	0.853	-2.130*
log CP	2.172	1.025	2.120*	1.953	0.923	2.120*
log Y	4.824	1.396	3.450*	3.962	1.349	2.940*
Coop	-2.880	0.295	-9.760*	-2.933	0.281	-10.430*
Foreign bank	-1.801	0.357	-5.050*	-1.826	0.351	-5.200*
log P × log CP	-0.003	0.030	-0.110	0.002	0.032	0.080
log T	-0.578	0.443	-1.300	-0.282	0.424	-0.670
R square	0.573			0.578		
Equation 2: Interest rate						
Intercept	-0.005	0.003	-1.690**	-0.009	0.002	-4.380*
log Euribor	-0.002	0.001	-3.060*	-0.002	0.001	-3.970*
log T	0.003	0.001	2.160*	0.003	0.001	4.210*
R square	0.990			0.990		
Equation 3: Competitors' interest rate index						
Intecept	2.781	0.982	2.830*	0.287	0.085	3.400*
log P	0.956	0.025	37.580*	0.925	0.021	44.730*
log CQ	-0.116	0.045	-2.570*	-	-	-
log Euribor	-0.154	0.048	-3.240*	-0.142	0.045	-3.160*
log T	-0.042	0.052	0.810	-0.086	0.022	-3.970*
h (market power)	0.010	0.003	2.980*	0.013	0.003	4.400*
R square	0.894			0.898		
Chi-square test	10.650			17.580		

Sample size:216

*Significance level: *5%, **10%*

The coefficient estimates for the two versions of the model using the GMM method are included in the Table 1.5. Heteroskedasticity and autocorrelation consistent estimates of the standard errors were obtained with the Newey and West (1987) method, using a lag truncation parameter of length 2. All coefficients have the expected sign and are statistically significant at the 5% or 10% level, except the coefficients for the CQ_{it} and $EURIB_{it}$ variables in the third equation. However, because the CQ_{it} variable in this model concerns average movements in 19 different banks, the sign of its coefficient cannot be known with certainty.

The demand for loans is elastic, and the elasticities that enter the price equation have the following values: $\varepsilon_{ii} = -2.087$, $\varepsilon_{ij} = 2.172$ and $\eta_{ji} = 0.956$. The estimated market power parameter is statistically significant at the 5% level in both models with a value close to zero, which suggests a negligible mark-up on competitive prices. Based on these values, the banking industry of Cyprus can be considered to have a (nearly) perfectly competitive market structure, and the monopoly hypothesis is rejected.

This conclusion is valid only with regards to the credit market for loans. It cannot be generalised to other banking services in Cyprus, like for example exchange rate services, underwriting and investment consulting. Further, it is valid only for the period covered by the dataset under consideration. In a rapidly changing banking environment, the level of competition can change significantly, especially if major mergers and acquisitions come to foreground.

1.6.3 Testing for exogenous and weak instruments

Based on the values of the Chi-square test for over-identifying restrictions (see Cameron and Trivedi, 2005, p.181-182), the validity of the additional instruments used in the GMM estimation procedures cannot be rejected at the 5% significance level. The associated p-values of the tests in models 1 and 2 were 0.385 and 0.227 respectively.

A Hausman endogeneity test (see Cameron and Trivedi, 2005, p.271) for the CQ_{it} variable (total amount of loans included in the assets of competitive banks) in model 1 was also constructed. This test is based on the equivalence of the OLS and 2SLS estimators, under the null hypothesis that the two estimators are asymptotically equivalent and there is no endogeneity bias. The estimates provided a value of 6.75 and a corresponding p-value of 0.748, according to which the equality of the two estimators is accepted at both the 5% and 10% significance levels.

Since the cost of physical capital was used as an instrumental variable in the estimation of model 1, the possibility of using a weak instrument was also tested. When the instrumental variables in a GMM estimation procedure are only weakly correlated with the endogenous variables, the respective moment conditions do not provide the necessary information in the estimation procedure in order to achieve identification. This is generally referred to as the "weak identification" problem (see, Verbeek, 2012, p. 171) and can have several negative consequences for GMM estimation procedures. These include inconsistency, sample size sensitivity and sensitivity with regards to the number of instruments used (see, Stock et. al., 2002).

In order to test the "per unit cost of physical capital" variable for the possibility of weak identification, the F-statistic was used as suggested by Staiger and Stock (1997). This statistic provides an indication regarding the information content of the instrumental variables. It consists of: (i) estimating a reduced form regression of the endogenous variable on the exogenous variables and the instruments and (ii) computing the F-statistic for zero instrumental variable coefficients. In the context of this study, the following reduced form regression was used for CQ_{it}

$$\begin{aligned} \log CQ_{it} &= \beta_0 + \beta_1 Coop_i + \beta_2 Foreign_i + \beta_3 \log EURIB_{it} + \beta_4 \log CP_{it} \\ &\quad + \beta_5 \log CPC_{it} + \varepsilon \end{aligned} \tag{1.44}$$

The four dependent variables consist of: one indicator variable for cooperative societies ($Coop_i$); one indicator variable for banks with a foreign controlling affiliate ($Foreign_i$); the Euribor rate ($EURIB_t$); the weighted average lending interest rate derived from the competitors of bank i (CP_{it}); and the per unit cost of physical capital (CPC_{it}) which is to be used as an instrument for CQ_{it}. The model was estimated with the OLS method. The coefficient estimates, together with their robust standard errors, are reported in Table 1.6.

Table 1.6: Reduced form regression for weak instruments

Coefficient	Estimate	St error	t test
Intercept	25.018	0.028	119.720*
Coop	0.048	0.018	0.350
Abroad	0.453	0.024	2.820*
log Euribor	-0.930	0.006	-8.660*
log CP	-0.212	0.007	-2.800*
log CPC	-0.437	0.013	-3.420*
R squared:	0.347		

Sample size: 216
F statistic for zero CPC coefficient: 11.700
*Significance level: *5%*

All the coefficients have the expected signs. The coefficient for the CPC_{it} variable is significantly different from zero, which provides a first indication that the proposed instrumental variable is not irrelevant. More importantly, the F-statistic for $\beta_5 = 0$ generated a value of 11.700. According to the Staiger and Stock criterion, there is no indication for a weak instrument when the F-statistic generates values greater than 10.

1.6.4 Results with alternative models

As already noted in Section 1.2, most empirical studies of market power in banking have concentrated on the conjectural variations and Panzar-Rosse approaches. To compare the results of the previous sub-section with alternative models that have been proposed in the banking literature, two further models were estimated and compared with model (1.43) using similar variable definitions as in the previous section.

The first is the conjectural variations model proposed by Coccorese (2005), who estimated a nonlinear system of three equations using a translog cost function as shown below (model 3). The first equation in this system represents loan demand and is similar to the first equation in system (1.43). In this case too the interest rate variables in the system have been deflated. The translog-cost equation includes three input factors: the interest rate paid on deposits $\omega_{1it} = DR_{it}$, the average price of labour $\omega_{2it} = LB_{it}$ and the per unit cost of physical capital $\omega_{3it} = PC_{it}$. The third equation represents the price-cost margin equilibrium condition derived from profit maximisation and includes the conjectural derivative $J = \partial p_j / \partial p_i$. The nonlinear system of equations has the following form:

$$\log Q_{it} = a_0 + a_1 \log P_{it} + a_2 CP_{it} + a_3 \log Y_{it} + a_4 Coop_i + a_5 Foreign_i + a_6 \log T + \varepsilon_Q$$

$$\begin{aligned} \log C_{it} &= d_0 + \beta_0 \log Q_{it} + \frac{\beta_1}{2}(\log Q_{it})^2 + \sum_{k=1}^{3} d_k \log \omega_{kit} + \sum_{k=1}^{3} \beta_{k+1}(\log Q_{it} \times \log \omega_{kit}) \\ &\quad + \sum_{k=1}^{3} d_{k+3}(\log \omega_{kit})^2 + d_7(\log \omega_{1it} \times \log \omega_{2it}) + d_8(\log \omega_{1it} \times \log \omega_{3it}) \\ &\quad + d_9(\log \omega_{2it} \times \log \omega_{3it}) + d_{10} Coop + d_{11} Foreign + d_{12} \log T + \varepsilon_C \end{aligned} \tag{1.45}$$

$$P_{it} = AC_{it}\left(\beta_0 + \beta_1 \log Q_{it} + \sum_{k=1}^{3} \beta_{k+1} \log \omega_{it}\right) - \frac{1}{(a_1 / P_{it}) + J(a_2 / CP_{jt})} + \beta_5 T + \varepsilon_p$$

To identify parameter J, five restrictions that are consistent with linear homogeneity in input prices were incorporated into the estimation procedure as follows: $d_1 + d_2 + d_3 = 1$, $\beta_2 + \beta_3 + \beta_4 = 0$, $d_6 = d_4 + d_5 - 2d_7$, $d_8 = -d_4 - d_7$ and $d_9 = -d_5 - d_7$. The conjectural derivative measures the degree of competition in the market and can take the following values:

(a) $0 < J < 1$: Competitors are expected to match price changes in the market (cooperative profit maximisation).

(b) $J = 1$: Collusion.

(c) $J = 0$: Non-cooperative Nash equilibrium (each firm acts alone in setting its own optimal price in the market).

(d) $J < 0$: Competition among firms (competitive price reductions expected from rivals).

(e) $J = -\infty$: Perfect competition

Table 1.7: GMM estimation of conjectural variations model 3

Coefficients	Estmate	St error	t test
Equation 1: Demand for loans			
Intercept	-1.036	0.009	-118.880*
log P	-1.459	0.355	-4.100*
log CP	1.498	0.377	3.970*
log Y	1.334	0.019	69.210*
Coop	-2.619	0.239	-10.970*
Foreign bank	-1.461	0.235	-6.230*
log T	0.364	0.136	2.680*
R square	0.450		
Equation 2: Total cost			
Intercept	-17.575	6.436	-2.730*
log Q	2.322	0.529	4.390*
log Q × log Q	-0.114	0.031	-3.660*
log DR	-0.321	0.703	-0.460
log LB	- 0.053	0.893	-0.060
log PC	1.374	0.743	1.850**
log Q × log DR	0.055	0.039	1.430
log Q × log LB	0.109	0.033	3.330*
log Q × log PC	-0.164	0.033	-4.930*
log DR × log DR	0.066	0.036	1.840**
log LB × log LB	-0.093	0.021	-4.420*
log PC × log PC	0.143	0.066	2.180*
log DR × log LB	-0.084	0.030	-2.810*
log DR × log PC	0.018	0.034	0.540
log LB × log PC	0.177	0.048	3.670*
Coop	-0.229	0.096	-2.400*
Foreign bank	-0.287	0.095	-3.010*
log T	-0.265	0.055	-4.860*
R square	0.975		
Equation 3: Interest rate			
J (conjectural derivative)	-0.023	0.007	-3.020*
T	-0.006	0.001	-10.910*
R square	0.990		
Chi-square test	53.904		

Sample size:161

*Significance level: *5%, **10%*

The system was estimated with the GMM method, using the competitors' weighted average interest rate on deposits as an instrumental variable for the average competitive interest rate on loans (CP_{it}) in the equilibrium interest rate equation, which could be endogenous. A reduction in the average competitive interest rate on loans can motivate an interest rate reduction response by firm i. However, an inverse relationship is also possible and can introduce endogeneity bias to the system. If firm i is associated with a sufficiently large market share and reduces its lending interest rate, competitors might follow by adopting similar interest rate reductions. It is therefore necessary to use instrumental variables to consistently estimate the parameters of the third equation. The results are presented in Table 1.7.

The interest rate elasticities in the first equation ($\varepsilon_{ii} = -1.459$, $\varepsilon_{ij} = 1.498$) are lower than the results obtained in model (1.43) and included in Table 1.5. In the cost equation, the coefficient of the loan variable (Q) indicates that the banking industry of Cyprus is characterised by diseconomies of scale. This finding is most likely due to the inclusion in the sample of COOPs, which generally exhibit lower efficiency and productivity levels than commercial banks. The market power parameter in the third equation has a statistically significant negative value that is close to zero. This result suggests the existence of a competitive market structure (compatible with a Nash non-cooperative equilibrium) and a rejection of the monopoly hypothesis.

The second model considered in this section is a log-linear econometric model for total revenues. As shown in the study by Delis et al. (2008), this model can be used to measure market power based on the Panzar-Rosse statistic (H) and has the following form:

$$\begin{aligned} \log TR_{it} &= \beta_0 + \sum_{i=1}^{19} b_i D_i + \beta_1 \log LB_{it} + \beta_2 \log DR_{it} + \beta_3 \log PC_{it} + \beta_4 \log\left(\frac{Depos._{it}}{Assets_{it}}\right) \\ &\quad + \beta_5 \log Assets_{it} + \beta_6 \log T + u_{it} \end{aligned} \tag{1.46}$$

The dependent variable represents the total revenues of bank i in period t (TR_{it}). The independent variables in the model include the average labour wage (LB_{it}); the interest paid on deposits (DR_{it}); the cost of physical capital (PC_{it}); bank-specific dummy variables (D_i); the ratio of total deposits to total assets for each bank and period in the sample ($Deposits_{it} / Assets_{it}$); the size of assets of each bank ($Assets_{it}$); a time trend variable (T) and an idiosyncratic disturbance term (u_{it}). Instead of using the deposits-to-assets ratio and total assets independent variables, Delis et al. (2008) used the equity-to-assets ratio. However, because the dataset used in this study includes unlisted COOPs that do not report equity capital in their financial statements, the deposits-to-assets ratio and total assets variables were preferred in this study. Furthermore, for the cost of physical capital, the authors used the ratio of administrative expenses to total fixed assets and the same variable was used for the estimation of model (1.46).

According to the intermediation approach in finance, banks are considered primarily as financial intermediaries whose output is measured by the total value of generated loans, investments and other income (see, Matthews and Thompson, 2008, p. 148). Deposits can be considered as a necessary input in the production process (see, Berger and Mester, 1997) and accordingly input costs consist of deposit interest rate payments and payments to other factors of production e.g. labour. Use of the deposits-to-assets ratio in model (1.46) is further justified by the fact that, during the last ten years, there has been a considerable inflow of foreign deposits in the Cyprus banking system. This has enabled the banking industry in Cyprus to expand the size of its assets considerably (primarily through loan provisions and investments in securities abroad), thus reaching 896% of the country's GDP in 2010.

The Panzar-Rosse statistic for measuring market power is equal to $H = \beta_1 + \beta_2 + \beta_3$ (the sum of the three input factor elasticities). According to this definition, market power is measured by the sensitivity of total revenues to changes in the three factor prices and can take the following values: (1) $H < 0$, the market structure is characterised by a monopoly or by collusion in an oligopoly; (2) $H = 1$, the market structure is characterised by perfect competition and (3) $0 < H < 1$, the market structure is characterised by monopolistic competition. Following Delis et al. (2008), two versions of the model were estimated. The first was based on the specification in equation (1.46), whereas in the second, time dummy variables were used in place of the time trend variable.

Table 1.8: Estimation of the revenue test (Panzar-Rosse) model

	with time dummies			with time trend		
Coefficients	**Estimate**	**St error**	**t test**	**Estmate**	**St error**	**t test**
Intercept	-7.151	2.908	-2.460*	-9.321	2.330	-4.000*
log LB	0.298	0.108	2.750*	0.368	0.069	5.360*
log DR	-0.025	0.124	-0.200	0.011	0.114	0.100
log PC	0.087	0.127	0.690	0.129	0.125	1.040
log (Deposits/Assets)	0.246	0.102	2.410*	0.290	0.103	2.810*
log (Assets)	0.901	0.113	7.950*	0.943	0.111	8.480*
log T	-	-	-	-0.062	0.077	-0.800
H	0.360			0.508		
R square	0.990			0.990		
Wald test (H=0)	2.750			6.160		
Wald test (H=1)	8.710			5.740		

Sample size:161
*Significance level: *5%*

The estimation was performed with the OLS method and the results are included in Table 1.8. In the model with the time dummies, all the coefficients have the expected sign except the coefficient of the deposit interest rate. The coefficients of the deposit interest rate and cost of physical capital variables are not statistically significant in either of the two models; however, the F-test for the joint significance of all the coefficients in the models indicated a valid specification at the 1% significance level. The positive coefficient of the total deposits-to-total assets ratio suggests that higher deposit rates provide a bank with the opportunity for additional loan provisions and investments (e.g., in government bonds), which can further increase interest rate revenues. Similarly, the positive coefficient of the total assets variable suggests that increases in bank size are associated with higher revenues.

The estimated values of the Panzar-Rosse statistic (0.360 and 0.508) suggest a market structure that is compatible with monopolistic competition and are similar to those calculated for Latvia, Greece and Spain by Delis et al. (2008). Wald tests were also conducted to test the hypotheses of perfect competition ($\beta_1 + \beta_2 + \beta_3 = 1$) and monopoly ($\beta_1 + \beta_2 + \beta_3 = 0$) and were rejected at the 5% significance level in both models.

1.7 Discussion and policy implications

During the last ten years the banking sector in Cyprus has grown enormously mainly due to a considerable inflow of foreign deposit funds. In 2010 alone, the assets of the industry (consisting mostly of loans and foreign securities) amounted to 896% of the country's GDP. In their efforts to attract foreign funds, Cypriot commercial banks increased deposit interest rates to unprecedented levels. This had the effect of increasing the primary input cost for loans and led to a situation where both the deposit and lending interest rates in Cyprus were by all means exceptionally high when compared with European standards.

At first glance the level of lending rates in the country, for the period under consideration, raises reasonable considerations for the possibility of monopoly pricing. However, if this was true, then one would expect deposit rates to be significantly lower. This was not the case in Cyprus. In order to measure market power in the Cyprus banking sector, a conjectural variations model was developed. Such an analysis is important for the monetary authorities in the country, in order to understand whether the industry operates efficiently and its implications for consumer welfare. Depending on the results, the following market structures - outcomes are possible (see, Van Hoose, 2010, p. 51):

- Perfect competition: Highest possible amount of loans at the lowest possible lending rates. Highest possible amount of deposits at the highest possible deposit rates. Consumer surplus is maximised.

- Monopoly: Lower amount of loans, higher interest rates on loans and a transfer of consumer surplus to the monopolist (the bank). Bank profits are maximized.

- Monopsonist: Lower amount of deposits, lower interest rates on deposits and a transfer of consumer surplus to the monopsonist (the bank).

- Oligopoly (Cournot-Nash): Equilibrium values between the two extremes of perfect competition and monopoly (monopsony).

- Monopolistic (monopsonistic) competition: As the number of banks in the industry increases, the equilibrium values tend to converge to the results of perfect competition. Increased product differentiation in the industry, increases the interest rate on loans and decreases the interest rate on deposists.

The preceding market power estimates for the banking industry in Cyprus suggest a market structure close to perfect competition. This is also supported by the fact that the spreads between the lending and deposit interest rates in the country were among the lowest in Europe during the period under consideration. As already explained above, deposit rates were mainly driven by the strong competition among Cypriot banks in their effort to attract foreign deposit funds.

The analysis in this chapter proposed an alternative way for estimating conjectural variations models. It will be of interest to practitioners working in the area of competition in banking. Most available banking micro-data (e.g. from financial statements) provide only aggregated cost information that cover all the operations and products offered by banks. The aggregated nature of most available cost data introduces several measurement problems in empirical applications and therefore complicates the validity of the results.

Not surprisingly, many studies that estimate conjectural variations models use cost equations with several proxies in the context of systems of equations that include in addition demand and equilibrium price equations. As a result, many coefficient estimates are not consistent with economic theory and can change substantially with simple modifications in the specification of the model. In this study, a conjectural variations model for loans was developed that incorporates an opportunity cost concept in the form of the interest rate on minimum reserves. Two main advantages are associated with this approach in applied work:

1. Simultaneous equations systems can be estimated without any measurement errors in the aggregated cost variables.
2. By directly incorporating the conjectural price reaction elasticities in the modelling framework, the explanatory power of the system is significantly enhanced.

2 A market test for nonlinear pricing in European banking

2.1 Introduction

Price discrimination, the practice of setting different prices for the same product or service, is an important research area in the industrial organisation literature and a pricing tactic that is extensively used in many industries. This practice can be implemented based on observable consumer characteristics or through the self-selection of consumers who choose different versions of the same product. By exercising price discrimination, firms with monopoly power can increase profits by extracting additional consumer surplus. This is achieved by servicing more consumers based on their willingness to pay for a product or service (see, Cabral, 2000, p. 181).

There are three basic types of price discrimination (Tirole, 1987, p.135). First-degree price discrimination is rarely encountered in practice because it is based on setting different prices for each consumer, using knowledge of their individual characteristics. Second-degree price discrimination is based on offering different versions of the same product (e.g., different package sizes) and is a frequently employed pricing technique, especially in the fast moving consumer goods industry. Third-degree price discrimination is another frequently used pricing tactic that uses market segmentation or different versions of the same product in order to set different prices per segment or version. Under third-degree price discrimination, profit maximisation presupposes an inverse relation between price and market segment price elasticity (see Pepall et al., 2008, p. 98).

Nonlinear pricing is the most common form of second-degree price discrimination and describes any pricing schedule in which the prices paid by the consumers for a product or service are not proportional to the quantities they purchase. Iyengar and Gupta (2009) identify three main reasons for firms to adopt nonlinear pricing tactics: (a) consumer heterogeneity (consumers choose the quantities they prefer), (b) cost considerations (quantity discounts encourage stockpiling by consumers and hence reduce the inventory cost of the firm) and (c) competition (quantity discounts are used as a form of competitive response).

Most of the theoretical research on nonlinear pricing (as in for example Schmalensee, 1981a and Armstrong and Vickers, 2001) has concentrated on two-part tariffs that consist of a fixed fee (f), which must be paid regardless of the consumed quantity, and a variable fee (p), which is proportional to the consumed quantity (q). Although the total amount paid by a consumer ($c = f + p.q$) is a linear function of the consumed quantity (q), the associated price per unit ($p + f / q$) is not constant.

Nonlinear pricing can be associated with monopolistic or oligopolistic market structures. In a monopoly context, Oi (1971) demonstrated the optimality of the two-part tariff pricing scheme, which achieves both allocation efficiency and profit maximisation. Setting the price close to the marginal cost ensures efficiency while the fixed fee component achieves profit maximisation by extracting the consumer surplus. Schmalensee (1981b) and Varian (1985) further demonstrated that an increase in the total produced quantity is a necessary condition in order to increase the total welfare (the sum of the consumer and producer surpluses) in a price-discriminating monopoly.

In an oligopoly context, nonlinear pricing has been associated with increasing competition among firms that use quantity discounts as a form of competitive response. A related effect demonstrated by Stole (1995) is that the quality distortion induced by monopolists in order to increase profits, decreases with increasing competition in oligopolistic markets. Furthermore, Armstrong and Vickers (2001) and Rochet and Stole (2002) showed that, under specific conditions, two-part tariffs can be nearly optimal in oligopolistic markets. Iyengar and Gupta (2009) note that oligopolistic markets are especially complex to model and that examining the optimality of the different nonlinear pricing schemes under different market conditions remains an open research area. For example, Jensen (2006) demonstrated the optimality of two-tier tariffs in a duopoly setting.

Numerous empirical studies have investigated the existence of price discrimination in banking markets (to be reviewed in Section 2.2), particularly in the United States (US). However, given the size and importance of the global banking industry, it is surprising that the existence and causes of nonlinear pricing in credit markets have not been previously explored in the literature, despite the existence of a variety of products in the loan market and the substantial variation of interest rates among products. In this study, an econometric test is proposed to verify the existence of nonlinear pricing in European lending interest rates and its relationship with market structure. This study especially focuses on examining whether nonlinear pricing is associated with increasing market concentration (monopoly power) or increasing competition, as in oligopolistic markets. This type of analysis has important implications for financial regulation and banking competition policy in Europe.

The econometric test is based on the cash demand model originally developed by Baumol (1952) and Tobin (1956), which, in turn, is based on the economic order quantity (EOQ) inventory model. This model was later extended to firms by Miller and Orr (1966), while modern versions of the EOQ model have incorporated quantity discounts in the optimal inventory demand function (see for example Brigham and Ehrhardt, 2011, ch. 28). Using a panel dataset consisting of seven European countries and eight annual time periods, the differences in lending interest rates, with and without the existence of quantity discounts, are modelled as a function of market concentration and other country- and market-specific characteristics that likely influence interest rates in Europe. The results, based on a fixed effects model specification, suggest that the use of nonlinear pricing in the European market for loans is associated with increasing market concentration.

The rest of the chapter is organised as follows. Section 2.2 reviews the literature on price discrimination in banking and Section 2.3 discusses the theoretical cash demand model upon which the market test for nonlinear pricing is based. The dataset used in the analysis is discussed in Section 2.4 and the empirical results of the study are presented in Section 2.5. In Section 2.6, some alternative panel data estimators are used in order to evaluate the robustness of the results. Section 2.7 concludes with a discussion.

2.2 Literature review

The existence of various forms of price discrimination has been the subject of many empirical studies in the banking literature. Most of these studies examined topics like: monopoly pricing, racial discrimination, credit card price discrimination mechanisms, ATM surcharging and spatial price discrimination. In contrast, there are very few theoretical studies on price discrimination on banking. Two such studies were provided by Gorton and Kahn (2000) for nonlinear pricing and Gary-Bobo and Larribeau (2004) for first-degree price discrimination.

An important gap in the existing literature refers to the lack of any empirical support regarding the use of non-linear pricing tactics by banks. This chapter aims to address this gap by developing and executing a test for the existence of nonlinear pricing in European banking. Before proceeding with the test, this section presents the main findings of all the aforementioned studies. Nearly all of these studies suggest the existence of some form of monopolistic price-discriminatory behaviour in the banking industry.

Theoretical studies

Gorton and Kahn (2000) examined a theoretical model where banks can renegotiate or liquidate (e.g., end the contract early and seize the collateral) loan contracts as new information about the borrowers becomes available. In this way, banks can monitor borrowers (control risk) by either choosing to liquidate inefficient projects or renegotiating lower interest rates to prevent risk taking on the part of the borrower. In this framework, the authors showed that renegotiation has efficiency considerations and the characteristics of loan contracts endogenously emerge to enhance efficiency.

The initial interest rates on loans result primarily from the renegotiation process and are set to minimise expected asset substitution by borrowers. To a lesser extent, they reflect a default risk premium. Also, the interest rate that minimises the expected agency cost in a loan contract, is not guaranteed to result in zero expected profits for banks. In these types of settings, competition among banks can result in nonlinear pricing schedules for loans in the form of origination fees or cross-subsidisations with other products.

Gary-Bobo and Larribeau (2004) presented a model with two important features: (a) it simultaneously explains the size of a loan and its associated interest rate and (b) it explains how the lender can price discriminate (in the first-degree sense) based on observable borrower characteristics. Two versions of the model were examined: a competitive equilibrium model and a model for a price-discriminating monopolist with market power. Based on French mortgage data, the authors discovered the presence of social discrimination in which blue-collar workers pay more than executives. The empirical results were more consistent with the monopolistic version of the model because the estimated price mark-ups were too high to reflect default risks only.

Monopoly pricing

Earlier studies on monopoly power in banking, such as Meyer (1967), investigated the influence of market concentration on the lending interest rates offered to small firms in the US. The author concluded that the lack of competition in small markets resulted in higher lending interest rates and therefore in discrimination against small firms. This finding has important implications for credit markets because it suggests a misallocation of resources among different types of firms. Another interesting conclusion of this study is that small banks gradually reduce interest rates as they grow due to economies of scale. However, after a certain level of market concentration, monopoly power is created, and the associated increase in interest rates offsets their initial decline.

Berger and Hannan (1989) examined the relationship between market structure and the deposit interest rates offered by banks across a large sample of metropolitan areas in the US. Their results were consistent with the structure-conduct-performance hypothesis in the industrial organisation literature, according to which the concentration resulting from anti-competitive behaviour (e.g., collusion) results in lower deposit rates.

Neumark and Sharpe (1992) extended the work of Berger and Hannan by considering the impact of market structure on the price rigidity of consumer deposits. The authors found that, when general market interest rates rise, banks in concentrated markets are slower to raise interest rates on deposits but are faster to reduce them when general market interest rates decline. This price rigidity associated with market concentration enables banks to extract additional consumer surplus on movements in both directions.

Heffernan (1994) estimated a generalised pricing model in order to examine the pricing behaviour of financial institutions in Canada related to mortgages, term deposits, fixed rate registered retirement saving plans and registered retirement income funds. The results suggested a Cournot-type market structure where increases in the number of competing firms were found to be associated with lower prices in the market. Furthermore, the author found evidence of price discrimination in the market. The presence of transaction and switching costs encouraged: (a) consumers to purchase all their financial products from a single firm and (b) firms to adopt price discrimination tactics (see, Heffernan, 2005, p. 516).

Prager and Hannan (1998) examined the deposit rates offered by banks that participated in substantial horizontal mergers in the US. They were compared with those of their rivals, as well as with those offered by banks operating in markets where no substantial mergers had taken place. The authors found that the deposit rates offered by banks participating in mergers, as well as those of their rivals, declined by a greater percentage than the rates offered by banks operating in markets with no merger activity. These results suggest that the market power effect of the mergers dominated any positive effects on deposit interest rates that might have resulted from improved efficiency.

Racial discrimination

Another topic that has received considerable attention in the banking literature is the existence of racial discrimination in the United States credit market. Berkovec et al. (1998) used a large sample of mortgage loans insured by the US Federal Housing Administration in order to study the relationship between market concentration and discriminatory behaviour against minority borrowers. According to the authors, in the presence of noneconomic discrimination, minority borrowers will tend to have better loan performance than white borrowers in more concentrated markets. This is because financial institutions will ask them to meet higher underwriting standards. A detailed empirical test conducted by the authors failed to reject the null hypothesis of no noneconomic discrimination.

The relationship between market structure and racial discrimination was also examined by Cavalluzzo and Cavalluzzo (1998) with regards to the credit market for small businesses in the US. Based on their empirical results, the authors concluded that there was clear evidence of discrimination against small businesses owned by Hispanics and weaker evidence of discrimination against small businesses owned by Asians. Furthermore, the authors did not find any evidence of discrimination against white female owners. In addition, no evidence was found that the Black disadvantage in loan provision arises from prejudice.

A general conclusion drawn from these studies is that banks operating in markets characterised by high concentration, acquire market power that enables them to exercise price discrimination. Banks appear to have used this market power in order to offer more favourable lending rates to female business owners and to charge Hispanic and, to a lesser extent, Asian small business owners, more.

Credit card price discrimination mechanisms

In addition to the loan and deposit markets, several researchers have examined the credit card market for the possibility of price-discriminatory behaviour on the part of the banks. Murphy and Ott (1977) provided several theoretical arguments according to which credit cards can be used as a price discrimination mechanism by firms. By separating consumers into high- and low-elasticity groups and simultaneously providing discounts to high-elasticity cardholders, firms can realise profits as suggested by third-degree price discrimination. The authors also provide several examples of goods, services and firm types that should be expected to profit from the use of credit cards as a price discrimination mechanism.

One distinct feature of the credit card market is that it is characterised by different equilibrium conditions than those encountered in the loan and deposit markets. For example, Stango (2000) estimated a (non-cooperative) game theoretic model of credit card competition in the US market. His results suggest that firms that are ex-ante identical commit to different variable-rate pricing structures for their homogeneous products that smooth the level of competition in the market. In addition, the author found evidence that volatile movements in the prime rate index of the market (which represents the cost of funds) increase the prices and profits of variable-rate firms more than those of fixed-rate firms.

ATM surcharging

Studies of the impact of market structure on the prices of banking services have also examined ATMs. Hannan et al. (2003) investigated the impact of various institutional and market-specific factors on the decision of banks to impose surcharges and identified both direct and indirect effects of this decision on profitability. The direct effect is related to the demand of non-depositors to use the bank’s ATMs and the indirect effect is related to the influence of surcharging on the decision of non-depositors to change banks. The authors found that, among other factors (e.g., demographic), the profitability of surcharging is higher in less concentrated markets and increases with the banking institution’s share of ATMs.

Spatial price discrimination

Another study that suggests the existence of monopoly power in a credit market was conducted by Degryse and Ongena (2005). The authors found that there is substantial evidence of spatial price discrimination in Belgium’s credit market. Lending rates were found to decrease with the distance between the firm and the lending bank and increase with the distance between the firm and competing firms. This spatial discrimination policy by banks is caused by transportation costs because borrowers encounter higher transportations costs when visiting competing banks. The resulting market power of the lenders leads to monopolistic pricing tactics, in which, banks charge higher lending rates to closer borrowers who have lower transportation expenses. Furthermore, the authors found lending interest rates to be positively related with market concentration and negatively related with the size of the loan, which suggest the existence of monopolistic nonlinear pricing strategies.

2.3 Theoretical framework

2.3.1 The Baumol model of cash demand

To develop a formal test for the relationship between nonlinear pricing and market concentration in the banking industry, two versions of the Baumol model of cash demand will be used. This model, originally developed by Baumol (1952) based on the EOQ model of inventory management, can be used in order to estimate the optimal cash balances of a firm through the minimisation of an appropriate cost function. It relies on the assumptions that firms use cash at a steady rate and that cash inflows also occur at a steady rate; therefore, net cash inflows are also characterised by a steady rate.

A further assumption adopted in this study is that cross boarder provision of loans - where financial institutions in one country provide loans to entities based in other countries - exists only to a limited extend and represents only a small fraction of the aggregated amount of loans provided by the financial institutions of a country. Further, distance is negatively related with the value of loans provided by these institutions. This assumption is supported by both theoretical and empirical evidence in the banking literature.

One rationale put forward in the literature concerns the travel cost faced by borrowers when considering the choice of bank for deposit or lending services (see for example Chiappori et. al., 1995). When this cost is high, borrowers will prefer to do business with banks located close to their premises. A second rationale concerns the high monitoring cost faced by lenders when providing loans to entities located far away. When a prospective borrower is located in another country, credit management and the process of screening loan applications requires the collection of a considerable amount of information (in addition to financial statements) that is frequently associated with multiple site visits, knowledge of foreign industry characteristics and knowledge of the legal and corporate environment associated with a foreign country. Since the collection of this information can prove costly, banks will be less willing to lend funds to entities located in other countries.

Further, entities operating in close proximity to the premises of a bank are more likely to expand their cooperation with this particular bank (e.g. by placing deposits or purchasing additional non-loan related services) which provides an additional source of information and reduces the cost of monitoring for the lender (see Almazan, 2002). In an empirical study of the US banking industry Brevoort and Hannan (2006) found that, indeed, distance operates as a deterrent to lending.

Brigham and Ehrhardt (2011, 28E-3) identify three main reasons for firms to hold cash balances: (a) to carry out their everyday payments and transactions, which are necessary for the smooth operation of the firm; (b) to maintain the compensating balances that are frequently required by financial institutions in return for the provision of a loan; and (c) as a precaution in the event of unexpected cash flow volatility or to take advantage of trade discounts. However, because cash balances do not earn any interest, an important opportunity cost is associated with holding them, which Baumol proposed could be minimised by choosing the optimal cash balance quantity.

The first version of the model used in this study is based on the augmented cost model presented in Lal and Staelin (1984) and Dada and Srikanth (1987), according to which the total cost consists of three elements: the purchase cost, which is equal to the annual sales value ($p_0 D$), the holding cost ($(Q/2).H$) and the transaction cost ($(D/Q)/A$). The model has the following form:

$$TC = p_0 D + \frac{D}{Q} A + \frac{Q}{2} H \, . \tag{2.1}$$

In inventory models, the quantity demanded, D, represents the period's sales in units and $p_0 D$ is the total value of sales. In a cash demand context, the amount of net new cash needed for the transactions of a period is D. The associated end-of-period cost of this amount is $(1+p_0)D$, which includes the amount of net new cash needed and can be obtained with a loan (D) plus, the end of period interest expense ($p_0 D$) associated with the loan. The remaining terms follow Baumol's model and can be written in augmented form as follows:

$$TC_1(Q, p_0) = (1 + p_0)D + \frac{D}{Q}A + \frac{Q}{2}H \quad . \tag{2.2}$$

The holding cost ($(Q/2).H$) is the product of the average cash balance ($Q_1/2$) and the opportunity cost rate (H). The opportunity cost rate is usually considered to be equal to the rate of return forgone on marketable securities such as treasury bills or the cost of borrowing to hold cash (Brigham and Ehrhardt, 2011, 28E-4). The transactions cost ($(D/Q).A$) is the product of the number of transactions (D/Q) and the cost per transaction (A). Therefore, D is the total amount of net new cash needed for the transactions of a period and Q is the amount of cash raised by borrowing (the cash balance). Closing costs, such as loan origination and underwriting fees that are usually associated with loans, are not included in the Baumol model.

The optimal borrowing quantity is found by partially differentiating equation (2.1) with respect to the required cash balance (the first-order condition) and setting the derivative equal to zero:

$$\frac{\partial TC_1(Q, p_0)}{\partial Q} = \frac{H}{2} - \frac{(D)(A)}{Q^2} = 0 \, .$$

The associated second-order condition that ensures minimisation is:

$$\frac{\partial TC_1^2(Q, p_0)}{\partial Q^2} = \frac{2DA}{Q^3} > 0 \, .$$

Solving the first-order condition for Q gives:

$$Q^2 = \frac{2DA}{H} \tag{2.3}$$

and taking the square root on both sides provides the optimal cash borrowing quantity (balance),

$$Q^* = \sqrt{\frac{2DA}{H}}\,. \tag{2.4}$$

This formula suggests that the optimal cash balance increases less than proportionately with increases in the total amount of net new cash needed for the transactions of a period (D). Consequently, economies of scale are associated with holding cash balances that are greater, the larger the firm is (Brigham and Ehrhardt, 2011, 28E-3).

Lal and Staelin (1984) extended the EOQ modelling framework by including quantity discounts and examining their impact on the optimal inventory balance. Their model is based on six assumptions: (a) the inventory policies of sellers and buyers can be described by EOQ models; (b) the seller is assumed to have knowledge of the buyer's holding and ordering costs; (c) the seller is price competitive and offers the same price to all buyers; (d) the annual demand rate of the buyer is considered exogenous and fixed and is independent of the seller's discount policy; (e) the buyer faces greater holding costs than the benefits accruing to the seller from early payments; and (f) the seller is assumed to face homogeneous buyers.

The proposed quantity discount policy is based on the following schedule. Let $Q_0 < Q_1 < Q_2 < ... < Q_n$ and $p_0 > p_1 > p_2 > ... > p_n$. For orders $Q \le Q_0$, the price charged is p_0; for orders between $Q_0 < Q < Q_1$, the price charged is p_1; for orders between $Q_1 < Q < Q_2$, the price charged is p_2 and so on up to quantity n. Lal and Staelin (1984) suggested approximating such a discrete discount pricing schedule with a continuous price-quantity relationship $p(Q)$ that satisfies the following criteria:

1. $p'(Q)$, $p''(Q)$ both exist.

2. Given the price list $p(Q)$, a cost-minimising buyer benefits by ordering the quantity Q^{**} that minimises the joint cost function of the buyer and the seller.

3. The seller's augmented cost function, evaluated at the jointly optimal quantity Q^{**} and given the discount schedule $p(Q)$, should be less than the augmented cost function with no quantity discounts. This is equal to requiring the following condition for the seller's annual cost $S(Q^{**}) < \bar{S}(Q^{*})$, where $S(Q) = (p_0 - p(Q))D + \bar{S}(Q)$, $Q > Q^{*}$ and $\bar{S}(Q) = \frac{D}{Q}A - \frac{Q}{2}H$.

Based on the above requirements, the authors proposed the cost function:

$$TC_2(Q, \bar{p}) = (1 + \bar{p})D + \frac{D}{Q}A + \frac{Q}{2}H \tag{2.5}$$

that incorporates quantity discounts according to the following schedule:

$$\begin{aligned} \bar{p} &= p_0 \exp[-a(Q - Q^{*})] && \text{for} \quad Q \geq Q^{*} \\ \bar{p} &= p_0 && \text{for} \quad Q \leq Q^{*}. \end{aligned} \tag{2.6}$$

The constant a is greater than zero and is a decision variable for the seller. Lal and Staelin (1984) experimented with other (one- or two-parameter) monotonically decreasing functions but failed to find any other function that satisfies the listed requirements.

Minimisation of the cost function in (2.5) yields the following first-order condition:

$$\frac{\partial TC_2(Q,\bar{p})}{\partial Q} = -a^* D\, p_0 \exp[-a^*(Q-Q^*)] - \frac{(D)(A)}{Q^2} + \frac{(H)}{2} = 0 .$$

The associated second-order condition which ensures minimisation in this case is:

$$\frac{\partial TC_2^2(Q,p_0)}{\partial Q^2} = a^{*2} D\, p_0 \exp[-a^*(Q-Q^*)] + \frac{2DA}{Q^3} > 0 .$$

Letting $L = \exp[-a^*(Q-Q^*)]$ and rearranging the first-order condition provides:

$$\frac{(D)(A)}{Q^2} = -a^* D\, p_0\, L + \frac{(H)}{2} . \tag{2.7}$$

Multiplying both sides with Q^2 gives:

$$(D)(A) = Q^2\left(-a^* D\, p_0\, L + \frac{H}{2}\right) . \tag{2.8}$$

Dividing both sides by 2 and solving for Q provides the optimal cash borrowing quantity in the presence of quantity discounts:

$$Q^{**} = \sqrt{\frac{2DA}{-2a^* D\, p_0\, L + H}} . \tag{2.9}$$

2.3.2 A market test for nonlinear pricing

It is assumed that two distinct groups of borrowers exist in the market:

(a) Group A: Consists of borrowers that demand small cash balances and do not qualify for quantity discounts. They are represented by the first Baumol model in equation (2.1).

(b) Group B: Consists of borrowers that demand large cash balances and qualify for quantity discounts. They are represented by the second Baumol model in equation (2.5).

To construct a market test for nonlinear pricing, it is necessary to explicitly model the interest rate spread: the difference between the interest rate without quantity discounts and the interest rate with quantity discounts. To achieve this, the following four step procedure is used:

Step1: Derivation of the inverse cost functions from the Baumol models in equations (2.1) and (2.5).

Step2: An expression for the interest rate spread is formed based on the difference between the two inverse cost functions.

Step3: The optimal cash balances in the interest rate spread equation are replaced with their equilibrium conditions in (2.4) and (2.9).

Step4: Using algebra, the interest rate spread equation is set in a format that can be estimated with econometric methods.

To see this, consider taking the difference between equations (2.1) and (2.5) evaluated at the optimal quantities Q^* and Q^{**}, respectively. The following expression is obtained:

$$TC_1 - TC_2 = [p_0 D - p_0 D L] + \left[\frac{D}{Q^*} A - \frac{D}{Q^{**}} A\right] + \left[H \frac{Q^*}{2} - H \frac{Q^{**}}{2}\right]. \qquad (2.10)$$

Rearranging the equation and dividing both sides by D results in the following form:

$$p_0 - p_0 L = \left(\frac{TC_1 - TC_2}{D}\right) + A\left[\frac{1}{Q^{**}} - \frac{1}{Q^*}\right] + H\left(\frac{Q^{**} - Q^*}{2D}\right). \qquad (2.11)$$

The left-hand side of the equation represents the spread in interest rates as a result of quantity discounts. This is also the quantity that will be used as the dependent variable in the econometric models of Section 2.5. The first term on the right side of the equation is the additional total cost faced by the borrower when no quantity discounts are offered by the bank as a percentage of the annual cash demand. Because this term is expected to vary by country and year in a panel data context, in addition to country specific variables, cross-sectional and time-period fixed effects will be used in the econometric models of Section 2.5.

The third term on the right side of the equation also represents a difference between the two pricing policies. It is the product of the opportunity cost of holding cash (H) and the additional optimal cash balance requested by the borrower when quantity discounts are offered, expressed as a percentage of the annual cash demand ($(Q^{**}-Q^{*})/D$). In the econometric applications of Section 2.5, proxy variables will be used for these quantities. The short-term interest rate (3-month Treasury bill) will be used as a proxy for the opportunity cost of holding cash. For the additional optimal cash balance, the proxy variable will be the annual flows of loans (new loans provided by the financial institutions in each country and year) because it should be expected that large increases in the amounts of loans provided every year will be correlated with higher differences between the optimal cash balances obtained with and without quantity discounts.

The second term on the right side of equation (2.11) is more complicated and requires further analysis. By taking the difference of the inverse squared quantities in (2.4) and (2.9), the following equality is formed:

$$\frac{1}{Q^{**2}}-\frac{1}{Q^{*2}}=\frac{-2a^{*}D\,p_0\,L+H}{2DA}-\frac{H}{2DA}\,. \tag{2.12}$$

Taking $2DA$ as a common denominator and cancelling out common terms gives:

$$\frac{1}{Q^{**2}}-\frac{1}{Q^{*2}}=\frac{-a^{*}\,p_0\,L}{A}\ . \tag{2.13}$$

Substituting the square root of this expression into (2.11) gives the following equation:

$$p_0 - p_0\ L = \left(\frac{TC_1 - TC_2}{D} \right) - A^{1/2} \left(a * p_0\ L\right)^{1/2} + H \left(\frac{Q^{**} - Q^{*}}{2D} \right). \tag{2.14}$$

The second term on the right side of the equation is the square root of the discounted price ($a * p_0\, L$) associated with a quantity discount pricing policy on the part of the seller multiplied by the square root of the cost per transaction (A). According to the industrial organisation literature, the exercise of nonlinear pricing by firms depends on the market structure that they encounter and can be associated with a monopolistic or an oligopolistic market, as explained in Section 2.2. Consequently, the average discounted interest rate in each country and year will depend on the level of market concentration in the country. For this reason, the Herfindahl-Hirschman index of market concentration will be used as an explanatory variable for the interest rate spread in the econometric models of Section 2.5. For the cost per transaction, the harmonised index of consumer prices will be used as a proxy variable.

2.4 Data description

2.4.1 Sources

Several different sources of information were combined in order to construct the panel dataset used in this study. The cross-sectional units in the panel consisted of Austria, Finland, France, Germany, Italy, the Netherlands, and Spain and the time period was from 2003 until 2007. The first source of information was the statistical data warehouse of the European Central Bank (ECB), from which the following variables were obtained for each country and year:

(1) The percentage change in the M3 monetary aggregate index.

(2) The Herfindahl-Hirschman index of market concentration.

(3) The actual changes (flows) in the end-of-year stocks of loans, as recorded in the aggregated balance sheets of the monetary and financial institutions in each country (MFIs, excluding money market funds and central banks).

(4) The average annual interest rate for “Loans other than revolving loans and overdrafts, convenience and extended credit card debt” as reported by the MFI sector in each country. Specifically, separate interest rates were obtained for the following cases (using the ECB definitions):
   - Loans of “Up to and including EUR 1 million” with maturities of “Up to 1 year”.
   - Loans of “Up to and including EUR 1 million” with maturities of “Over 1 and up to 5 years”.
   - Loans of “Up to and including EUR 1 million” with maturities of “Over 5 years”.
   - Loans of “Over EUR 1 million” with maturities of “Up to 1 year”.
   - Loans of “Over EUR 1 million” with maturities of “Over 1 and up to 5 years”.
   - Loans of “Over EUR 1 million” with maturities of “Over 5 years”.

The second source of information was the EUROSTAT statistical database, from which the gross domestic product (GDP, at market prices) and the harmonised index of consumer prices were obtained for each country and year. Finally, the average annual short-term interest rate for each country and year was obtained from the OECD statistical database. According to the OECD, “short-term rates are usually either the three-month interbank offer rate attaching to loans given and taken amongst banks for any excess or shortage of liquidity over several months or the rate associated with treasury bills, certificates of deposit or comparable instruments, each of three-month maturity”.

Table 2.1: Summary statistics

Variable	Sample	Mean	St dev	Min	Max
Interest rate spread	168	0.783	0.505	-0.159	2.709
Herfindalh-Hirschman Index	168	980	940	170	3550
Short term interest rate (%)	168	2.582	1.266	0.811	4.634
Gross domestic product (EUR trillions)	168	1.061	0.758	0.147	0.246
M3 monetary aggregate (% change)	168	7.244	4.231	-3.111	17.949
Change in the stock of loans (EUR 00000s)	168	0.200	0.798	-2.074	3.215
Harmonized index of consumer prices (2005=100)	168	103.224	4.790	93.860	112.900

2.4.2 Descriptive statistics

In the econometric models of Section 2.5, the dependent variable will be the interest rate spread, calculated as the difference between loans of “Up to and including EUR 1 million” and those of “Over EUR 1 million” for three maturity segments: “Up to 1 year”, “Over 1 and up to 5 years” and “Over 5 years”. This variable will be used in order to test for the presence of quantity discounts in the provision of loans and how this is influenced by changes in the market structure. Table 2.1 presents summary statistics for all the variables including the interest rate spread. It can be observed that the average value of the interest rate spread, 0.783, is positive, which suggests that interest rates on loans of up to and including 1 million are generally higher than interest rates on loans that exceed the 1 million benchmark.

Figure 2.1 illustrates the distribution of interest rate spreads across all countries, years and loan maturities. The interest rate on loans exceeding the 1 million benchmark is higher than the interest rate on loans below the 1 million benchmark in only 3% of the cases. Another interesting finding from Table 2.1 concerns the maximum and minimum values of the Herfindahl-Hirschman index. The large difference between the two values suggests that there are considerable differences in market concentration among the examined countries. This is of particular importance for the econometric analysis of Section 2.5 because market concentration will be the main explanatory variable in explaining movements in the interest rate spread. Figure 2.2 illustrates the distribution of values for the Herfindahl-Hirschman index.

To further investigate the variability inherent in the interest rate spread variable, Table 2.2 presents analytical values of this variable per loan duration segment, country and year. From the 168 values included in the table, the average interest rate on loans exceeding the 1 million benchmark was higher than the average interest rate on loans lower than or equal to the 1 million benchmark in only 5 cases. All 5 cases concerned loans with durations of more than 5 years and were from Austria (2003, 2008 and 2009), Germany (2008) and France (2007). Consequently, it can generally be inferred that large loans that exceed the 1 million benchmark are consistently associated with lower interest rates, which suggests the existence of quantity discounts in the loan market. It is therefore of interest to examine the factors influencing the level of these quantity discounts, particularly focusing on changes in the market structure as suggested in the industrial organisation literature.

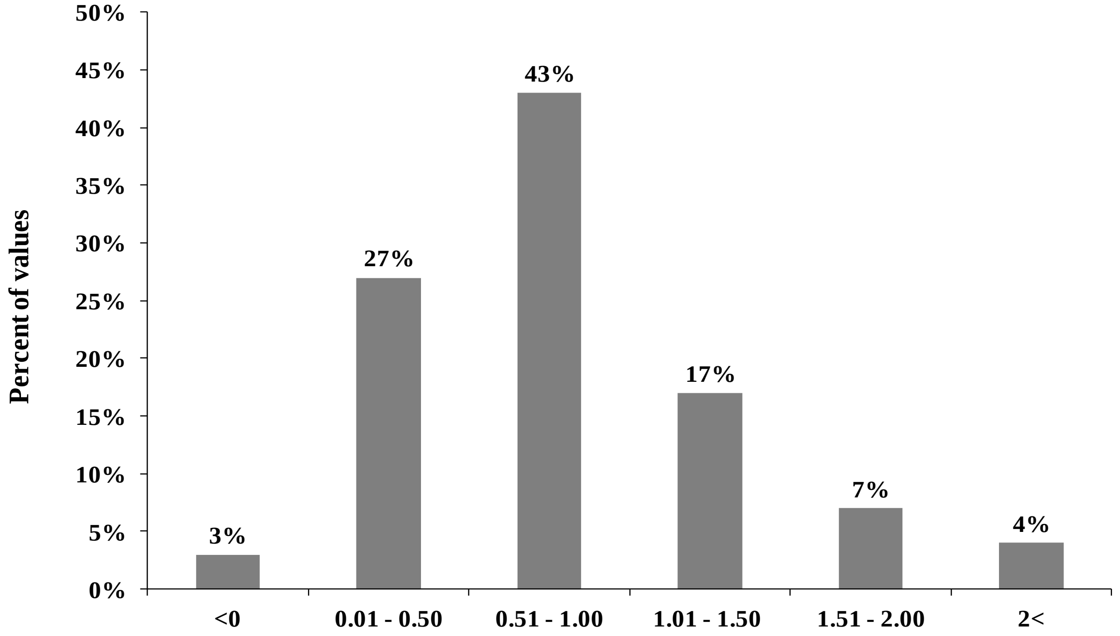

Figure 2.1: Distribution of the interest rate spread variable

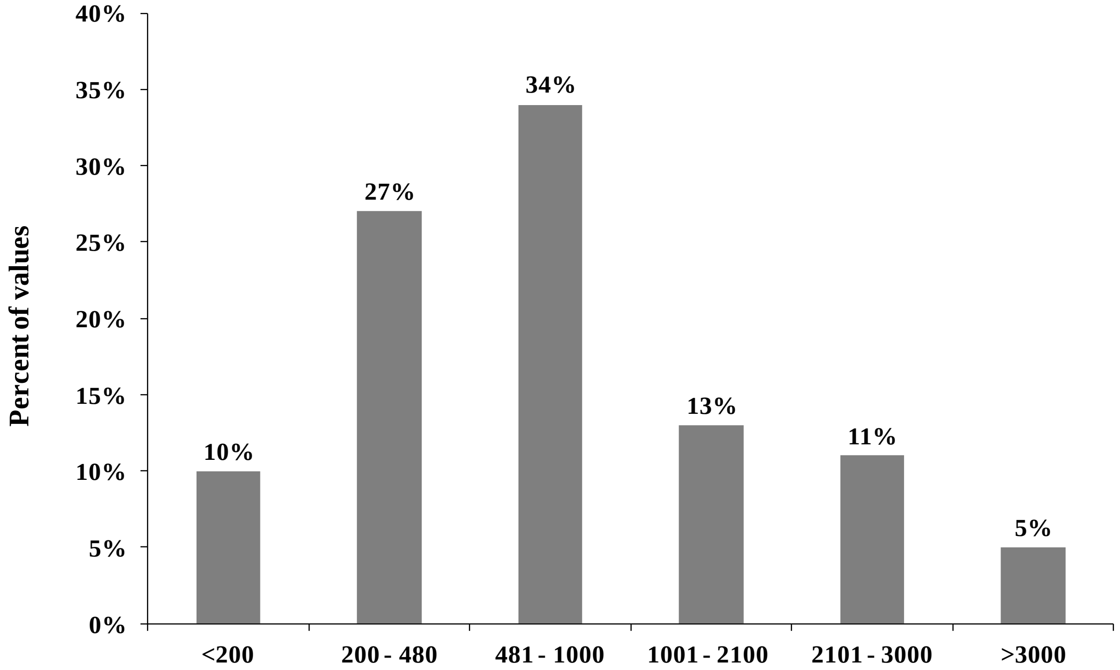

Figure 2.2: Distribution of the Herfindahl-Hirschman index variable

Table 2.2: Interest rate spreads

Year	AUS	GER	SPA	FIN	FRA	ITA	NED
Spreads for loans with duration up to 1 year							
2003	0.890	1.190	1.120	0.600	0.310	1.110	0.720
2004	0.770	1.270	0.960	0.580	0.610	1.140	0.690
2005	0.590	1.160	0.800	0.460	0.780	0.940	0.480
2006	0.480	1.110	0.610	0.560	0.540	0.760	0.410
2007	0.430	1.040	0.540	0.550	0.710	0.930	0.350
2008	0.440	0.950	0.800	0.500	0.710	0.850	0.600
2009	0.530	0.710	1.630	0.670	0.970	1.130	1.510
2010	0.480	0.990	1.470	0.810	0.630	1.100	1.090
Spreads for loans with duration between 1-5 years							
2003	0.820	1.010	1.230	0.930	2.020	1.630	0.360
2004	0.350	1.130	1.710	0.900	1.820	1.530	0.610
2005	0.420	0.850	1.190	1.010	1.580	1.330	0.190
2006	0.570	0.570	1.090	1.040	1.050	1.190	0.260
2007	0.540	0.380	0.790	0.970	0.740	1.060	0.310
2008	0.320	0.200	0.990	0.800	0.870	1.010	0.530
2009	0.730	0.600	2.710	1.520	2.030	2.270	1.300
2010	0.090	0.870	2.180	1.570	1.750	2.460	0.830
Spreads for loans with duration more than 5 years							
2003	0.380	0.460	0.470	0.490	0.450	0.830	0.630
2004	-0.160	0.340	0.550	0.270	0.390	0.780	0.640
2005	0.150	0.360	0.620	0.500	0.350	0.930	0.380
2006	0.260	0.200	0.690	0.210	0.100	0.910	0.430
2007	0.160	0.030	0.440	0.520	-0.150	0.580	0.540
2008	-0.080	-0.120	0.890	0.730	0.020	0.330	0.780
2009	-0.070	0.090	1.300	1.210	0.820	1.090	0.530
2010	0.070	0.180	1.810	0.970	0.590	1.150	0.990

2.5 Econometric analysis

In this section, a panel data econometric model is estimated in order to examine the determinants of interest rate spreads, with particular emphasis on market concentration. In order to measure the difference between interest rates on loans below the €1 million benchmark (p_S) and interest rates on loans above the €1 million benchmark (p_L) the following two distance metrics were defined for the interest rate spread ($d = p_S - p_L$):

1. $S^a = \operatorname{sgn}(d).|\, d \,|$ 2. $S^b = \operatorname{sgn}(d).(|\, d \,|)^2$.

The sign (sgn) expression in the two distance metrics is defined according to the following rule:

$$\operatorname{sgn}(d) = \begin{cases} +1 \ if \ \ d > 0 \\ \ \ 0 \ if \ \ d = 0 \\ -1 \ if \ \ d < 0 \end{cases} \tag{2.15}$$

Both metrics will be used as dependent variables in econometric models. The sign definition, $\operatorname{sgn}(d)$, is included because the difference between the two interest rates $d = p_S - p_L$ can also result in negative values, when the interest rate on loans below the €1 million benchmark is lower than the interest rate on loans above the €1 million benchmark. Based on the descriptive statistics for the dataset reported in Section 2.3, negative values of the interest rate spread were realized in 3% of the sample cases. Equation (2.16) is the basic form of the model that will be estimated:

$$S_{it}^{a,b} = \beta_0 + \sum_{i=1}^{6} \delta_i^C C_i + \sum_{t=1}^{7} \delta_t^D D_t + \sum_{j=1}^{2} \delta_j^M M_j + \beta_1 R_{it} + \beta_2 L_{it} + \beta_3 \log P_{it} + \beta_4 \log HHI_{it}$$

$$+ \beta_5 M3_{it} + \beta_6 \log GDP_{it} + \varepsilon_{it}. \tag{2.16}$$

The model's variables are as follows: C_i represents 6 country dummy variables, except for Spain, which is the country captured by the intercept β_0; D_i represents 7 annual time period dummy variables, except for year 2010, which is captured by the intercept; M_i represents 2 dummy variables for the different loan maturities in the dataset, loans of "Up to 1 year" and of "Over 1 and up to 5 years" (loans with maturity of "Over 5 years" are captured by the intercept). The remaining variables vary by country and time period as follows: R_{it} is the short-term interest rate as published by the OECD; L_{it} is the annual change (flow) in the total stock of loans; P_{it} is the harmonised index of consumer prices; HHI_{it} is the Herfindahl-Hirschman index of market concentration and ε_{it} is the idiosyncratic disturbance.

Two additional explanatory variables were incorporated into the model, which, according to economic theory, can influence lending interest rates in an economy. These variables are not under the control of the firms but they represent the general macroeconomic environment in which firms operate in each country. The first is the percentage change in the M3 monetary aggregate ($M3_{it}$) and the second is the gross domestic product (GDP_{it}) of each country. The M3 monetary aggregate is a measure of the money supply, which, according to the liquidity preference theory of interest rate determination, is inversely related to the interest rates of an economy. In contrast, the gross domestic product is positively related to the level of interest rates of an economy because increases in income during business cycle expansions positively influence interest rates through an increase in the demand for money (see, Mishkin, 1998, p. 118 - 119).

The estimation results based on the first metric are presented in Table 2.3. The fixed-effects model in equation (2.16), estimated with the OLS method, provided a good fit for the data, as shown by the corresponding R-square value. Heteroskedasticity and autocorrelation consistent estimates of the standard errors were obtained with the Newey and West (1987) method using a truncation parameter of length 2. The Herfindahl-Hirschman index has a positive coefficient and is statistically significant at the 5% significance level. This suggests that higher concentration (increasing monopoly power) in the market is associated with higher interest rate spreads.

Table 2.3: Estimation results for absolute metric

	OLS			2SLS		
	Coefficients	**St error**	**t-value**	**Coefficients**	**St error**	**t-value**
Intercept	-3.077	11.952	-0.260	-4.459	12.637	-0.350
R	0.709	0.278	2.550*	0.628	0.275	2.290*
L	0.027	0.043	0.620	0.046	0.047	0.980
log P	-4.115	3.041	-1.350	-2.719	2.944	-0.920
log HHI	0.761	0.234	3.250*	1.586	0.477	3.320*
M3	-0.021	0.011	-1.870**	-0.026	0.012	-2.230*
log GDP	5.458	1.385	3.940*	4.875	1.213	4.020*
R-squared	0.649			0.610		

*Significance level: *5%; **10%*

The short-term interest rate (a proxy for the opportunity cost of holding cash) is statistically significant and positive, which suggests that the opportunity cost of holding cash is positively related to the level of interest rate spreads in the market. This could be due to the use of quantity discounts by financial institutions as a means of increasing loan demand when short-term interest rates increase. Because an increase in the opportunity cost of holding cash reduces the optimal demand for cash balances, as shown in equation (2.9), a corresponding increase in quantity discounts offered by financial institutions could be a reasonable price strategy for these institutions in order to increase the demand for loans. The coefficients of the harmonised index of consumer prices (a proxy for the transaction cost) and the change in the actual demand for loans (a proxy for the difference in the optimal cash balances with and without quantity discounts) are not statistically significant but their signs are consistent with equation (2.14).

The coefficient of the percentage change in the M3 monetary aggregate is negative, close to zero and statistically significant at the 5% level. An increase in the money supply leads to a marginal decline in the interest rate spread. According to the liquidity preference theory of interest rate determination, an increase in the money supply lowers market interest rates. In this case, it is possible that financial institutions consider that the reduced interest rates are already attractive for firms and are therefore less willing to further reduce the lending interest rate through the provision of quantity discounts. In contrast, the coefficient for the GDP is positive and statistically significant at the 5% level. Consequently, an increase in the GDP of a country is associated with an increase in the level of interest rate spreads. This could be due to the business cycle effect on the demand for loans. A growing GDP leads to increases in investment activity and consumption and therefore in the demand for cash, which is associated with increases in the interest rate spreads (quantity discounts).

The GDP and L variables in model (2.16) could be endogenous and correlated with the error term of the model. An increase in the GDP and the total demand for loans in a country increases the demand for cash by firms and, consequently, the provision of quantity discounts by banks. However, an inverse relation where an increase in the level of quantity discounts offered by banks lowers the interest rate per currency unit available to firms and therefore provides motivation for increased demand for loans is also possible. The associated increases in investment activity and consumption will also influence the GDP level of the economy. To address this endogeneity issue, model (2.16) was also estimated with the two-stage least squares (2SLS) method using one-period lagged values of the GDP, L and HHI variables as instruments. Some useful graphs and model fit diagnostics associated with the OLS and 2SLS estimation procedures are included in Appendix A.1. The coefficient estimates are included in the second part of Table 2.3 and have the same signs as in the OLS case. The coefficient of the Herfindahl-Hirschman index is higher than in the OLS case and higher absolute values are also observed for the intercept and the M3 and L variables.

Estimation results based on the second metric are presented in Table 2.4. The standard errors were again estimated with the Newey and West (1987) method using a lag truncation parameter of length 2. The signs and the statistical significance of the coefficients are the same as in Table 2.3, however their magnitudes differ since different dependent variables were used in the two models. Both estimation methods provided a lower R-squared value when the squared metric was used and the coefficient for the M3 monetary aggregate was not statistically significant when the OLS method was used. However, in general, both metrics provided similar results.

Table 2.4: Estimation results for squared metric

	OLS			2SLS		
	Coefficients	**St error**	**t-value**	**Coefficients**	**St error**	**t-value**
Intercept	-34.210	30.237	-1.130	-34.685	31.666	-1.100
R	1.727	0.704	2.450*	1.696	0.688	2.460*
L	0.017	0.110	0.160	0.064	0.117	0.550
log P	-1.245	7.692	-0.160	-0.723	7.378	-0.100
log HHI	1.453	0.592	2.450*	3.485	1.196	2.910*
M3	-0.041	0.028	-1.460	-0.051	0.029	-1.760**
log GDP	9.666	3.505	2.760*	10.534	0.039	3.470*
R-squared	0.542			0.501		

*Significance level: *5%; **10%*

2.6 Results with alternative panel data estimators

In this section, some additional panel data estimators are considered in order to evaluate the robustness of the preceding analysis. The emphasis is on using alternative estimation methods and model specifications that can possibly improve the accuracy of the results in Section 2.5. The analysis is restricted to the absolute metric version of model (2.16), which provided the best fit to the data. A preliminary analysis based on the Hausman test for fixed effects (see, Cameron and Trivedi, 2005, p. 717) provided a value of 19.73, which suggests correlation between the errors and the exogenous variables. Consequently, the null hypothesis of equality between the fixed effects and random effects estimators is rejected in favour of the fixed effects estimator.

The possibility of serial correlation in the errors was tested using the generalized Durbin-Watson test for panel data proposed by Bhargava et. al. (1983) and Wooldridge's test for AR(1) errors in fixed effects models (see Wooldridge, 2002, p. 274). The Durbin-Watson test generated a value of 1.648 and the Woodridge test a value of 55.922. Based on these values both tests reject the null hypothesis of no serial correlation in the errors. Consequently, in the analysis that follows the emphasis is on using heteroskedasticty and autoccorelation consistent covariance matrix estimators. For the sake of completeness, two additional random effects estimators are considered, which explicitly allow for dynamics in the specification of the error. Specifically, the following estimators are considered:

1. The first difference GMM estimator with two step estimation of the optimal weighting matrix. This procedure produces panel-robust standard errors that allow for both heteroskedasticity and serial correlation over time (see, Cameron and Trivedi, 2005, p. 746).

2. The first difference GMM estimator with two step estimation of the optimal weighting matrix and 3 additional instrumental variables. Since the variables GDP and L could be endogenous and correlated with the error term in model (2.16), the GMM estimator described in point 1 was also used with 3 additional instrumental variables. These referred to one-period lagged values of the variables GDP, L and HHI. The lagged value of the HHI variable was used as an additional instrument in order to improve the efficiency of the GMM estimator (see, Verbeek, 2012, p. 399).

3. The feasible GSL estimator proposed by Park (1967). This estimator allows for first-order serial correlation in the errors ($\varepsilon_{it} = \rho\varepsilon_{it-1} + u_{it}$), group wise heteroskedasticity ($E(\varepsilon_{it}^2) = \sigma_{ii}$) and time-invariant cross sectional dependence ($E(\varepsilon_{it}\varepsilon_{jt}) = \sigma_{ij}$). A recent

study by Reed and Ye (2011) used extensive simulations in order to examine the efficiency properties of a number of commonly used panel data estimators. For panel datasets where the number of time periods exceeds the number of cross-sectional units, the authors found that Park's estimator tends to provide the best results in terms of efficiency.

4. The two step GSL estimator proposed by Da Silva (1975). This estimator is based on a variance component model where the errors follow a moving average process $\varepsilon_{it} = a_0 e_t + a_1 e_{t-1} + ... + a_m e_{t-m}$ and e_t is a white noise process.

Table 2.5: GMM estimation results for absolute metric

	GMM			GMM with 3 additional IVs		
	Coefficients	St error	t-value	Coefficients	St error	t-value
Intercept	0.001	0.009	-0.110	-0.009	0.009	-0.980
R	0.505	0.077	6.600*	0.480	0.071	6.780*
L	0.033	0.010	3.350*	0.022	0.009	2.490*
P	-5.355	0.868	-6.170*	-5.017	0.812	-6.180*
log HHI	0.563	0.186	3.030*	0.715	0.125	5.720*
M3	-0.022	0.005	-4.490*	-0.020	0.005	-4.190*
log GDP	6.507	0.972	6.690*	6.210	0.904	6.870*
Sargan test	23.12			26.590		
AR(1) test	-1.430			-1.520		

*Significance level: *5%*

The results based on the two GMM procedures are presented in Table 2.5. With the exception of the intercept, all the coefficients are statistically significant at the 5% level and have similar signs and magnitudes with the coefficients in Table 2.3. Based on the GMM estimates, two panel data specification tests have been included in Table 2.5 (see, Verbeek, 2012, p. 409). The first is the Sargan test of over-identification that follows the chi-square distribution with degrees of freedom equal to the number of moment conditions minus the number of parameters. The values of this test confirm the validity of the moment conditions used in the estimation procedures. The second test is the standard normal test for first order serial correlation in the residuals. Based on the values of this test, there is no strong evidence of autocorrelation in the residuals.

The results from the two random effects GLS estimators are included in Table 2.6. None of these estimators provided a good fit to the data, as shown by the low R-squared values. Park’s method did not provide any statistically significant coefficients and the coefficients of the variables R and L do not have the expected signs. The Da Silva method with a first-order moving average specification (MA(1)) provided some statistically significant coefficients and a different sign only in the case of one variable (R). Nevertheless, both estimators provided a positive coefficient for the HHI variable, in line with the analysis in Section 2.5. In the case of the Da Silva method, this coefficient is statistically significant at the 5% level.

Table 2.6: Estimation results with dynamic error specifications

	Park's method - AR(1) errors			Da Silva's method - MA(1) errors		
	Coefficients	**St error**	**t-value**	**Coefficients**	**St error**	**t-value**
Intercept	-0.725	16.939	-0.040	0.642	3.546	0.180
R	-0.106	0.106	-1.000	-0.076	0.035	-2.160*
L	-0.032	0.128	-0.250	0.034	0.027	1.260
P	0.317	3.628	0.090	0.013	0.775	0.020
log HHI	0.073	0.262	0.280	0.223	0.098	2.270*
M3	-0.005	0.030	-0.180	-0.031	0.007	-4.330*
log GDP	0.099	0.328	0.300	0.258	0.117	2.200*
R-squared	0.271			0.215		
Av. AR(1) coef.	0.652			-		
MA(1) coef.	-			0.040		

*Significance level: *5%*

Overall, the results in this section provide further support to the conclusions reported in Section 2.5. With regards to the relationship between market concentration and nonlinear pricing, it can be inferred that higher concentration (increasing monopoly power) in the European banking industry tends to be associated with higher interest rate spreads.

2.7 Discussion and policy implications

The econometric results of the previous two sections are consistent with the theoretical work of Oi (1971) who demonstrated the optimality of the two-part tariff pricing scheme. It achieves both allocation efficiency and profit maximisation. Banks in Europe appear to use nonlinear pricing tactics in order to extract additional consumer surplus and maximize profits in their loan operations.

Given the crucial role of banking in every economy and the increasing emphasis that is nowadays placed on financial regulation and supervision, the results in this chapter raise some interesting questions and policy implications for the monetary authorities in Europe. The European Central Bank and the monetary authorities in member states are nowadays closely supervising commercial banks and their interest rate strategies. A major scope of this supervision is to minimise the possibility of collusion and prevent the application of monopoly pricing tactics. The empirical literature regarding market power in European banking, as reviewed in Chapter 1, suggests that to some satisfactory extend, there are no indications of monopoly pricing in European banking.

However, monopoly behaviour can manifest itself in many different formats and the Industrial Organization literature can provide useful guidance in this direction. The possibility of adopting a non-linear pricing tactic with regards to loan provisions is one such possibility that was demonstrated theoretically and tested empirically in this chapter. Non-linear pricing tactics have a long standing presence in fast moving consumer goods markets and are still being heavily used by consumer goods manufacturers. Given the wide range of products offered by banks these days, there is no reason why it should not be adopted by commercial banks too.

It is also important to emphasize at this point that non-linear pricing is not necessarily bad for an economy, even though it can be associated with a monopolistic market structure. Schmalensee (1981b) showed that total welfare can increase with nonlinear pricing, provided that the total quantity produced increases too. However, an important distinction holds. Despite the increase in total welfare, consumer welfare in particular decreases. On the other hand more consumers are served under non-linear pricing (see, Cabral, 2000, p. 181) which is important for entrepreneurship and small business development in an economy. Cabral (2000), provides a valid point in this direction: price discrimination should not be of concern, as long as there are no distribution issues between firms and consumers. It is therefore important for the monetary authorities to ensure that liquidity and project financing, in the market for loans, is equally accessible to the whole distribution of consumers and firms.

There are many other possibilities for monopoly behaviour in banking, beyond pricing strategies, which constitute fruitful areas for further research. Some notable examples are:

- <u>Vertical restrains</u>: Many commercial banks are nowadays vertically linked with insurance companies under a common group structure. Can vertical restrains be used as a collusion device when bank and insurance operations exist under the same group?
- <u>Product differentiation</u>: It is well known in the Industrial organization literature that greater product differentiation tends to be associated with greater market power. What product differentiation strategies are nowadays used by banks with substantial market power? Do banks use tactics such as product bundling?
- <u>Countervailing buyer power</u>: In many retail markets, large retailers (buyers) with substantial bargaining power are able to obtain discounts from suppliers when more options are available to them. To what extend are large depositors able to exert bargaining power in banking?

3 Multiscale analysis of the liquidity effect in the UK economy

3.1 Introduction

The study of interest rate determinants is important to financial institutions because changes in interest rates can greatly influence the profitability of investments. For this reason, many institutions use sophisticated forecasting methods in order to accurately predict changes in equilibrium rates (Mishkin and Eakins, 1998, p.128). The liquidity preference theory for the determination of interest rates suggests that changes in equilibrium rates result from changes in the demand for and supply of money. Increases in the money supply by monetary authorities lead to a reduction in interest rates (i.e., the liquidity effect). On the other hand, changes in demand result from income and price effects that are positively related to interest rates.

The liquidity effect generally has an immediate influence on interest rates, while income and price effects can involve substantial time lags. This is because increases in the money supply take time before they increase prices and income in the economy. Expectations of inflation also have an important influence on interest rates. It is important for monetary policy to identify the magnitude and speed of adjustment associated with each effect (Mishkin and Eakins, 1998, p.123).

Following an increase in the money supply, interest rates can rise or fall depending on the size of each effect, and results can vary by time-scale or cycle (e.g., short-term, medium-term and long-term). In order to analyse the liquidity effect, this study estimates regression models at different frequency bands as proposed by Cochrane (1989). However, instead of using a band-pass filter the model's variables are analysed with a wavelet multiresolution analysis, as suggested by Ramsey and Lampart (1998a,b), in order to examine the liquidity, income and price effects on interest rates in the UK economy by time-scale.

Ramsey and Lampart's methodology is also extended by standardising the time-scale regression models so that all of the variables in a given time-scale are set on a common scale of measurement. This allows for a comparison of the three effects (liquidity, price and income). Results for the UK economy suggest that the magnitude of the liquidity effect varies by time-scale. When only short-term cycles are examined, the liquidity effect is higher than the income and price effects, but in the medium and long-term, the price and income effects become more important.

The rest of this article is organised as follows: Section 3.2 reviews the main empirical evidence for the liquidity effect, and Section 3.3 provides a short introduction to wavelets and presents the standardised time-scale regression models. The data used in the analysis are explained in Section 3.4, and the results for the UK economy are presented in Section 3.5. In Section 3.6 two alternative modelling approaches are considered. Section 3.7 concludes with a discussion and some policy implications.

3.2 Empirical evidence for the liquidity effect

3.2.1 Definition and identification

There is a considerable amount of empirical literature on the liquidity effect, with different authors emphasizing different problems associated with its empirical identification. A particularly lucid explanation of the liquidity effect can be found in Leeper and Gordon (1992) who relied on the analysis provided by Friedman (1968) and Cagan (1972). This study adopts the definition provided by Lastrabes and McMillin (2004): "the temporary, but persistent, negative response of interest rates to nominal money supply shocks". However, as documented by Friedman and Schwartz (1963), money supply shocks are also associated with long-term effects on prices and income in the economy.

According to Thornton (1996), the main difficulty in identifying the liquidity effect lies with the interest rate targeting policies of the monetary authorities. With the beginning of the 21st century, many Central Banks abandoned the conduct of monetary policy through changes in the money supply and focused instead on interest rate targeting (Friedman, 2008). When such policies are in effect, innovations in the monetary aggregates are mainly reactions to changes in money demand, not money supply. As a result, it becomes difficult to correctly identify the policy actions of the monetary authorities and therefore consistently estimate the liquidity effect.

Since in equilibrium money supply equals money demand, it is not straightforward to quantify the money supply-price and money supply-income relationships by merely looking at changes in the quantity of money. This is because many hypothetical changes in the supply of money are in reality changes in money demand triggered by changes in prices and income. As a result, the statistical identification of the money supply-price and money supply-income relationships becomes difficult. This problem is further complicated by the increasing evidence that money demand is unstable (see, Friedman, 2008). For these reasons, nowadays Central Banks are mainly focusing on interest rate targeting as a way to stabilize prices and income. The purpose of changes in the money supply has been confined to accommodate changes in money demand that are not related with changes in prices and income.

Some authors, like Strongin (1995), have suggested using non-borrowed reserves (total reserves held as deposits with the monetary authorities minus any funds that may have been borrowed through the discount window) as an accurate reflection of the policy actions implemented by the monetary authorities. This is because in modern banking systems, the money supply process consists mostly of deposits created by banks for the general public and the reserves that these banks are required by law to hold with the monetary authorities. The rest of the money supply concerns currency in circulation.

However, this approach was also criticised (see for example, Pagan and Robertson, 1998) since in many cases it is not clear how changes in interest rates are linked with policy changes regarding non-borrowed reserves. For example, Friedman (2008) notes that Central banks apart from open market operations, can also create reserves through lending to commercial banks. In many monetary systems the distinction between borrowed reserves and reserves held as deposits with the Central bank (non-borrowed reserves) is not always clear. This further complicates the identification of monetary policy changes. From an econometric perspective, if interest rates are influenced by more factors (other than money supply, prices and income), then it is very important to include all the interest rate determinants in the model specification. In a different case there is a high possibility of omitted variable bias.

3.2.2 Empirical studies

Mishkin (1982) used a linear rational expectations model to study the liquidity effect in the US and found that increases in money supply are not correlated with declines in interest rates. A different conclusion was reached by Urich and Watchel (1984); they identified a liquidity effect using actual money supply announcements and survey data on expectations of the money supply and the producer price index. Cochrane (1989) used band-pass filters in the context of a regression model to isolate short-term movements in money growth and interest rates. His results suggest the existence of a short-term negative relationship between money growth and interest rates.

Subsequent studies by Leeper and Gordon (1992), Strongin (1995) and Serletis and Chwee (1997) used structural models to show that increased money growth can lower interest rates in the short-term, but the correlation often becomes positive and results are not as robust across sub-periods. Boyd and Caporale (1997) used the Kalman Filter to also demonstrate that monetary innovations do not always impact interest rates in the same direction.

Pagan and Robertson (1998) emphasise that the results from structural models are much less certain due to several econometric problems related to the choice of instrumental variables and the sample period used in the analysis. More recent studies have examined alternative explanations of the liquidity premium puzzle in the US. For example, Thornton (2004) introduced the idea of interest rate smoothing, which suggests that the US Federal Reserve aims to smooth the transition of rates to new equilibrium levels following economic shocks. As a result, the relationship between interest rates and monetary aggregates can be positive. Another interesting aspect is discussed by Kelly et al. (2011), who showed that monetary aggregates are characterised by measurement errors that can distort estimates of the liquidity effect.

Kopchak (2011), building on earlier work by Hamilton (1997), found that deviations of the federal funds rate from its target level are frequently the result of demand-side shocks. In order to identify the liquidity effect, Kopchak modelled demand-side forecast errors with a Kalman Filter model consisting of permanent and transitory components. Hamilton used a different identification approach by using forecast errors as an instrumental variable for exogenous changes in the supply of reserves. Kopchak found a considerably larger liquidity effect than Hamilton.

All of the above-mentioned studies concentrate on the US economy. Studies investigating the liquidity effect in other countries (mainly the G-7 group) include Sims (1992), Grilli and Roubini (1996), Cushman and Zha (1997), Fung and Kasumovic (1998), Lastrapes (1998), Kim (1999), Halabi and Lastrapes (2003) and Lastrapes and McMillin (2004). All of these studies report evidence of a short-term liquidity effect.

3.2.3 Information loss due to data filtering

The empirical studies presented so far, point to the existence of a short-term impact of money supply changes on interest rates. This chapter proposes that to a large extend these results are biased due to the filtering procedures imposed on time series used in econometric models. Frequently, these procedures involve first-differencing of the data as in Lastrapes and Selgin (1995). When time series contain unit roots (and are therefore non-stationary), econometric methods generate inconsistent estimates. For this reason, in applied work, economic time series are almost always subjected to a filtering procedure like first-differencing, before they are used in econometric models.

While the use of first differencing is sensible in order to remove stochastic trends, it also implies a cost that confines the practitioner's ability to accurately measure the liquidity effect. First differencing re-weights the information content in the time series towards higher frequencies and short-term cycles. At the same time it down-weights the lower frequencies that are associated with the long term cyclical components in the time series (see, Baxter and King, 1999). As a consequence, long-term causal effects cannot be estimated accurately and short-term effects are over-emphasized. This information loss can be easily demonstrated with a spectral analysis (see, Wei, 2006, p. 284).

A fundamental tool in spectral analysis is the spectral density function of a stochastic process that decomposes its variance into different frequencies. Adding the generated components across frequencies results in the total variance of the series. In order to show the information loss associated with first-differencing, two empirical estimates of the spectral density function were generated using the short-term interest rate variable: (1) based on the actual data and (2) based on first-differences of the actual data. The population spectral density function of a time series process y can be represented as follows (see, Greene, 2003, p. 625)

$$S_y(\omega) = \frac{1}{2\pi}\left[\lambda_0 + 2\sum_{k=1}^{\infty}\lambda_k \cos(\omega k)\right] \tag{3.1}$$

where $\omega \in [0, \pi]$ and $\lambda_k = Cov[y_t, y_{t-k}]$. The spectral density function is continuous and the cyclical features in the series are captured by the cosine functions which have periods 2π. The spectral density function and the autocovariances are linked with the following relationship

$$\lambda_k = \int_{-\pi}^{\pi} S_y(\omega) \cos(\omega) d\omega . \tag{3.2}$$

It suggests that a time series process can be represented in two equal ways: (a) in the time domain with the autocovariance function and (b) in the frequency domain with the spectral density function.

There are several ways to consistently estimate the spectral density function using sample data. For the purposes of this study, the lag window estimator with Bartlett weights (see, Greene, 2003, p. 628) was used

$$\hat{S}_y(\omega) = \frac{1}{2\pi}\left[c_0 + 2\sum_{k=1}^{L} w(k, L) c_k \cos(\omega k) \right] . \tag{3.3}$$

The weights are defined as $w(k, L) = 1 - \frac{k}{L+1}$ and the sample autocovariances have the following form $c_k = \frac{1}{T}\sum_{t=k+1}^{T} (y_t - \bar{y})(y_{t-k} - \bar{y})$.

The estimated spectral densities for the actual and first-differenced data are included in Figure 3.1. The horizontal axis depicts the different cycles (in number of weeks) in the data. The horizontal axis can also be represented in terms of frequencies however the representation in terms of cycles is intuitively more appealing and was preferred (see, Greene, 2003, p. 627). The vertical axis provides the value of the estimated spectral density. The spectrum estimates based on the actual data show that most variability is concentrated in cycles exceeding 50 weeks (low frequencies).

A different result is obtained once first differencing has been applied to the data. Most variability is now located in medium (less than 100) and short-term (less than 50) cycles. Evidently, first differencing eliminated most long-term cyclical variation in the data and shifted contributions towards short-term and medium-term cycles. As a result, any analysis of the liquidity effect based on first differences of the data will tend to provide misleading results, especially with regards to the long-term effects of money supply increases.

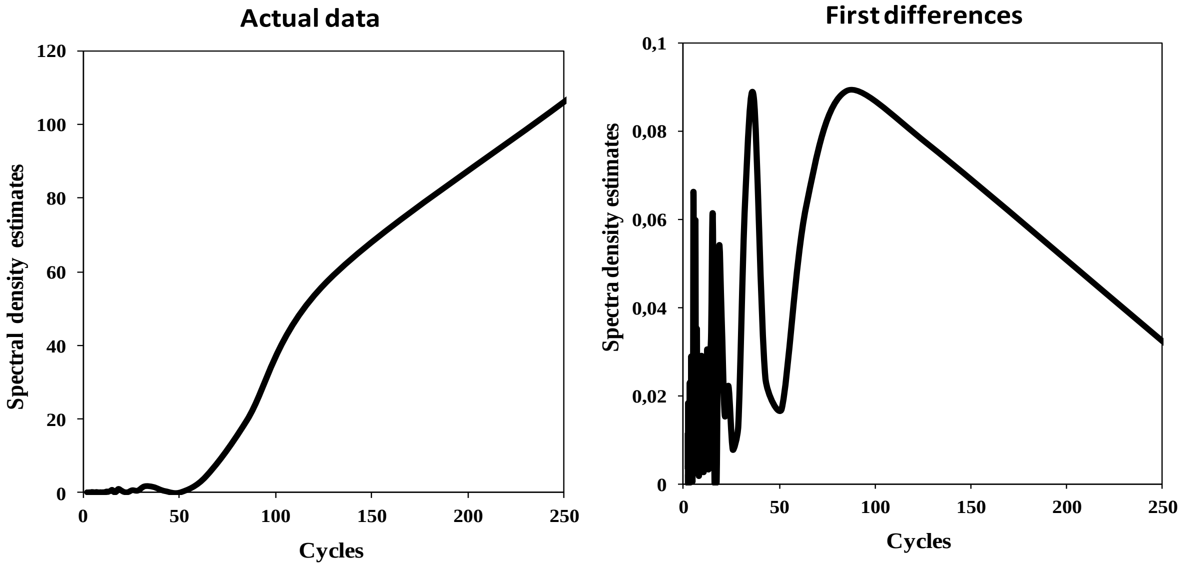

Figure 3.1: Spectral density estimates for interest rate data

3.3 Methodology

3.3.1 Wavelet analysis of macroeconomic data

Ramsey and Lampart (1998b) first proposed the use of wavelets in order to analyse the effect of monetary policy on income over different time-scales. The authors identified four key properties of wavelet analysis that are useful for analysing macroeconomic and financial data: the ability to analyse non-stationary time series, localisation in time, orthogonal time-scale decomposition of the data and the fact that pre-filtering is not needed when analysing time series with wavelets. Using Granger causality tests, the authors concluded that, in short-term cycles (or high frequencies), income Granger-causes money; at business cycle frequencies, money Granger-causes income.

Ramsey and Lampart recognised that when examining relationships between macroeconomic variables, the value of the parameters involved may be different across time-scales. Even if the direction of the relationship is the same, the intensity may be different. When the effects vary by time-scale, important insights are lost by concentrating only on the observed sampling rate of the data that averages, and therefore masks, the individual time-scale effects.

Cochrane (1989) also used a frequency domain framework to examine the liquidity effect but concentrated only on short-term movements in interest rates. By running a regression of filtered interest rates on filtered money growth, he was able to find a short-term negative relationship for periods of up to one year. He also showed that regressions with band-pass filtered data are equal to placing restrictions on the lag coefficients of the actual data and the frequency response of the lag coefficients is constant within each band. In addition, his analysis is based on the assumption of stationary time series.

In this study, the liquidity effect is studied across a range of cyclical movements (i.e., frequencies) but without the stationarity requirement because wavelets can also handle nonstationary time series. The regression framework is similar to the one proposed by Cochrane (1989) but differs in two important aspects: (a) a wavelet filter is used in order to extract the frequency components of the model’s variables and (b) the model’s variables are standardized so that they are measured on a common scale.

3.3.2 Time-scale decompositions

Wavelets are functions defined within the set of square integrable functions $L^2(R)$. They have compact local support but decay quickly to zero elsewhere. Any function of time can be represented by a sequence of projections onto a basis of orthonormal elements called the father (ϕ) and mother wavelets (ψ). They are defined as follows (see, Ramsey and Lampart, 1998b):

$$\phi_{j,k} = 2^{-j/2}\phi\left(\frac{t-2^j k}{2^j}\right) \quad \text{and} \quad \psi_{j,k} = 2^{-j/2}\psi\left(\frac{t-2^j k}{2^j}\right).$$

The mother wavelet can be dilated (expansion in the range) proportional to $2^j \{j = 1,2,3,...,J\}$ and translated (shift in the range) by changing the parameter $k = \{0,1,2,...\}$. A unit decrease in j packs the wavelet oscillations of $\psi_{j,k}$ closer (i.e., reduces their width and doubles their frequency). A unit increase in k shifts the location of the wavelet by an amount proportional to its width. The father wavelet remains unaffected by changes in j but is also translated by changes in k.

In a wavelet multiresolution analysis, the wavelet series approximation of a time series χ_t can be represented as follows:

$$\chi_t = \sum_k s_{J,k}\,\phi_{J,k}(t) + \sum_k d_{J,k}\,\psi_{J,k}(t) + \sum_k d_{J-1,k}\,\psi_{J-1,k}(t) + ... + \sum_k d_{1,k}\,\psi_{1,k}(t). \qquad (3.4)$$

The coefficients associated with the father and mother wavelets are, respectively:

$$s_{J,k} = \int \chi_t\,\phi_{J,k}(t)\,dt \quad \text{and} \quad d_{j,k} = \int \chi_t\,\psi_{j,k}(t)\,dt.$$

The father wavelet coefficients capture the low-frequency trend behaviour in the data, and the mother wavelet coefficients capture all deviations from the trend. The wavelet series approximation, can be written more compactly as

$$\chi_t = D_J + D_{J-1} + ...D_j + ... + D_1 \qquad (3.5)$$

with the following orthogonal components:

$$D_J = \sum_k s_{J,k}\,\phi_{J,k}(t) \quad \text{and} \quad D_j = \sum_k d_{j,k}\,\psi_{j,k}(t).$$

Large values of j refer to low-frequency variation (in long-term cycles) of the time series χ_t, and small values refer to high frequency variation (in short-term cycles). Consequently, wavelets provide an orthogonal decomposition of the data into J time-scales, with each component providing a resolution of the data at scale 2^j. In the next section, monthly interest rate time series with a length of $2^8 = 256$ are decomposed into 8 time-scales. The coefficients of the first time-scale component (D_1) capture frequency variation over durations of 2-4 months, and the coefficients of the second component (D_2) capture frequency variation over durations of 4-8 months. Accordingly, the coefficients of the third component (D_3) capture frequency variation over durations of 8-16 months and so on up to level J[3].

In the empirical applications of this study the maximal overlap discrete wavelet transform (MODWT) was used. Unlike the discrete wavelet transform (DWT), the MODWT does not provide an exactly orthogonal decomposition but it provides efficiency gains and generates wavelet coefficient vectors of equal length with the original time series that can be used in a time-scale regression analysis. In practice, both the DWT and the MODWT are computed using a pyramid algorithm that starts by filtering the time series with a high and a low-pass filter in order to produce the wavelet and scaling coefficients respectively. This procedure is then iteratively repeated up to J times on the scaling coefficients. In the case of the DWT the algorithm involves a down-sampling step, as a result of which only half the coefficients are retained after each iteration. Consequently, the resulting time series do not have equal length with the original time series as in the case of the MODWT. More details on the pyramid algorithm are provided in Appendix A.2.

3.3.3 Standardization of time-scale regressions

Instead of looking at the relationship between money and income averaged over all time-scales, Ramsey and Lampart (1998b) examined the relationship at each time-scale by

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[3] The complete list of monthly cycles covered by each time-scale component is as follows: $D_1 : 2-4$, $D_2 : 4-8$, $D_3 : 8-16$, $D_4 : 16-32$, $D_5 : 32-64$, $D_6 : 64-128$, $D_7 : 128-256$, $D_8 : Trend$.

estimating separate regression models using the wavelet orthogonal components as variables. In Section 3.4, the same method is adopted in order to examine the effects of the money supply (M), inflation (I) and the industrial production index (P) on short-term interest rates (R) in the UK economy. The industrial production index is used as a proxy for real income. These variables are commonly used in Keynensian liquidity preference models of interest rate determination (see, for example, Mishkin, 1982). Using the wavelet orthogonal components, the following regression models are estimated:

$$R[D_J]_t = a + \beta_1 M[D_J]_t + \beta_2 I[D_J]_t + \beta_3 P[D_J]_t$$

$$R[D_j]_t = a + \beta_1 M[D_j]_t + \beta_2 I[D_j]_t + \beta_3 P[D_j]_t, \; j = 1,2,...,J-1. \tag{3.6}$$

In the first regression equation, the low-frequency (trend) behaviour of short-term interest rates ($R[D_J]_t$) is a function of the low-frequency behaviour of the money supply ($M[D_J]_t$), inflation ($I[D_J]_t$) and industrial production ($P[D_J]_t$). In the second regression equation, the same relationships are estimated for $J-1$ time-scales, with each scale covering a different short-term frequency band of the data as described in the previous sub-section.

This second step in the analysis extends the time-scale regression framework developed by Ramsey and Lampart (1998b) by standardising the coefficients in each time-scale. The usefulness of this step derives from the complex data transformations generated by wavelets, as a result of which the variables are difficult to interpret within each time-scale. By subtracting off their mean and dividing by their standard deviation, all the variables in a particular time-scale are standardised so that they are measured on a common scale. The corresponding coefficients are called standardised coefficients (see Wooldridge, 2003, p. 186). It is therefore possible to examine whether the income and price effects are more important than the liquidity effect within a specific time-scale.

Standardisation of the regression equation at time-scale j proceeds as follows:

$$\frac{R[D_j]_t - \overline{R}}{\hat{\sigma}_R} = \beta_1 \frac{M[D_j]_t - \overline{M}}{\hat{\sigma}_M} + \beta_2 \frac{I[D_j]_t - \overline{I}}{\hat{\sigma}_I} + \beta_3 \frac{P[D_j]_t - \overline{P}}{\hat{\sigma}_P}, \; j = 1,2,...,J \tag{3.7}$$

where $\overline{R}$ and $\hat{\sigma}_R$ are the sample mean and standard deviation, respectively, of the time series $R[D_j]_t$, and the intercept of the regression model (a) is eliminated from the transformation. The notation for the other variables is similar. The standardised regression can also be written more compactly as:

$$Z_R = b_1 Z_M + b_2 Z_I + b_3 Z_P \qquad (3.8)$$

where the standardised coefficients are

$$b_1 = \frac{\hat{\sigma}_M}{\hat{\sigma}_R}\beta_1, \quad b_2 = \frac{\hat{\sigma}_I}{\hat{\sigma}_R}\beta_2 \quad \text{and} \quad b_3 = \frac{\hat{\sigma}_P}{\hat{\sigma}_R}\beta_3 .$$

These coefficients are measured in terms of standard deviations (e.g., a one standard deviation increase in the money supply changes interest rates by b_1 standard deviations), and their magnitudes can be compared in order to identify which has the biggest effect on interest rates at the specific time-scale j.

3.3.4 Data description

Monthly data for short-term interest rates, the M3 monetary aggregate, inflation and the industrial (manufacturing) production index in the UK were used in the analysis covering the period from April 1987 until July 2008. The data were obtained from the OECD statistical database. Based on the OECD definition, short-term interest rates in the UK refer to the 91-day interbank equilibrium rate. For the M3 monetary aggregate the OECD publishes an index with 2005 as the base year. Following Mishkin (1982), the industrial production index was used as a proxy for real income, which is not available with monthly regularity for the UK economy. In addition to the above, data for two further variables were collected: the first is the short-term interest rate in the US economy as published by the OECD and the second is the monthly average effective exchange rate index of the sterling against the US dollar (1990 average = 100) published by the Bank of England. Table 3.1 includes summary statistics for all variables.

Table 3.1 Summary statistics

	Mean	Std Dev	Min	Max
Short term inter. rate (%)	7.175	3.131	3.420	15.286
M3 index (2005=100)	62.592	31.847	20.245	144.977
Inflation (%)	0.230	0.442	-1.000	3400
Industrial production index	96.756	5.432	83.858	104.835
US short term inter. rate (%)	4.950	2.128	1.040	10.090
Sterling/USD exch. rate	932338	274032	832	1240789

Charts of the first four variables are included in Figure 3.2. Each variable exhibits different characteristics and cyclical behaviour. For example, inflation is characterised by strong short-term and seasonal variability, while M3 is characterised by an upward trend. Interest rates and industrial production exhibit different cyclical movements, structural changes and trend characteristics.

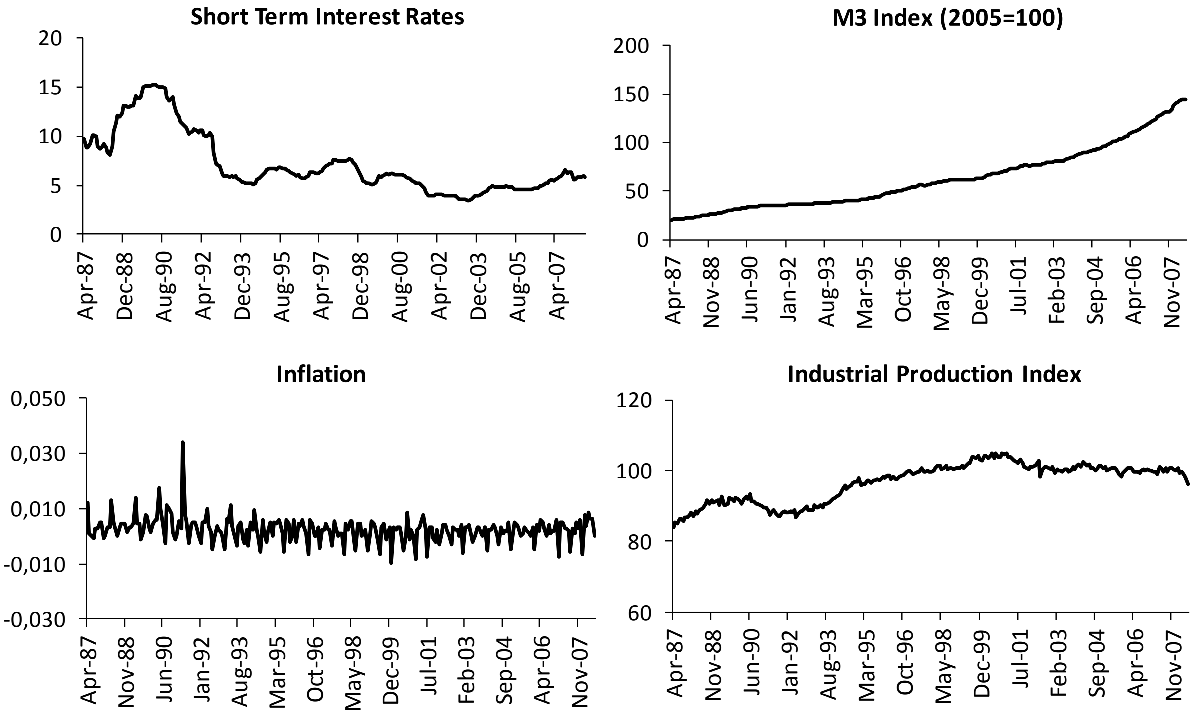

Figure 3.2: UK data (source OECD).

Wavelet decompositions of all the time series were performed using the Daubechies least asymmetric family of wavelets with a filter length of 8 in the context of the MODWT. Because the time series consist of 256 observations, eight time-scales were generated from the MODWT. This wavelet family was also used by Gencay et al. (2001, p. 159) to analyse the money supply in G-7 countries. It was also found to provide good resolutions for the data used in this study.

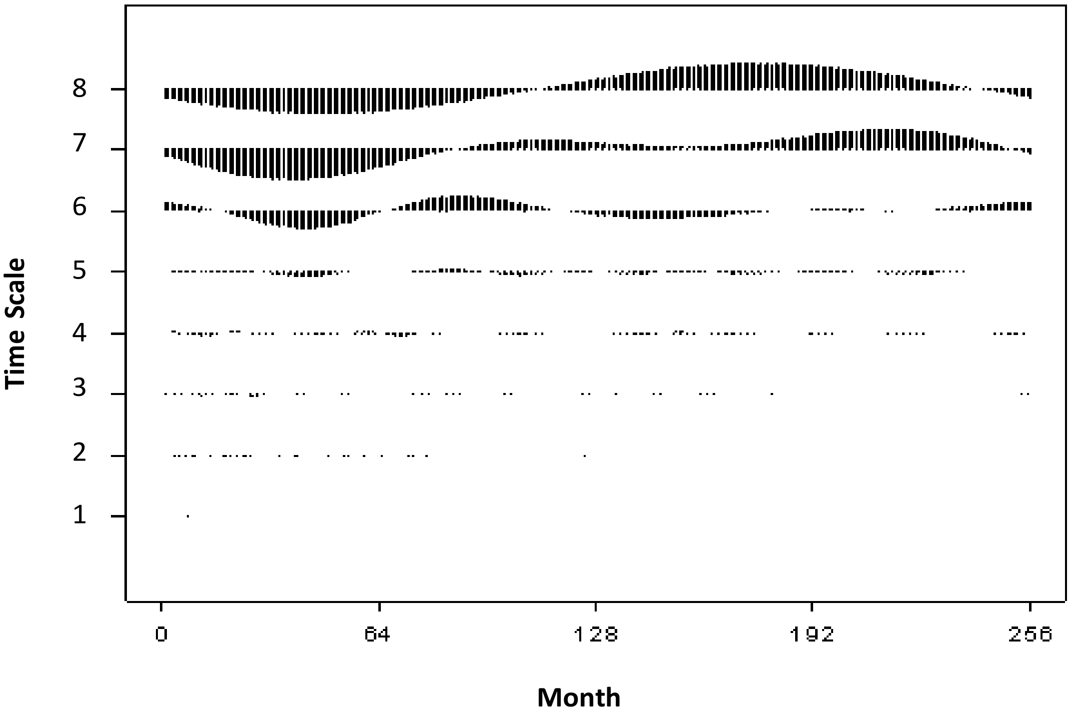

Figure 3.3: Wavelet coefficients for UK interest rates generated using the Daubechies least asymmetric family of wavelets with a filter length of 8 in the context of the MODWT.

Figure 3.3 shows the wavelet coefficients for interest rates generated by the MODWT. Time-scale 1 corresponds to cyclical movements of 2-4 months, and time-scale 8 corresponds to the long-term trend in the data. It can be observed that interest rates exhibit variation at all time-scales, particularly in time-scales 6 to 8. Furthermore, the cyclical variation in each time-scale differs according to the time location (i.e., month). Unlike the Fourier transform, wavelets are flexible enough to provide a frequency decomposition of a time series while also preserving the time location dimension of the data. For example, although there is considerable variability during the first 60 months of the analysis at all time-scales, the situation is different during the last 60 months of the sample period when variability is high at time-scales 7 and 8. Also, the long-term trend of the series exhibits variability between months 128 and 256, a period when variability in time-scales 1-4 is not high.

Because variability differs by time-scale, it is interesting to examine how the liquidity, income and price effects differ (and their relative importance changes) by time-scale. This difference is the subject of Section 3.4.

3.4 Estimation results

In order to examine the existence of a liquidity effect by time-scale, separate regression models were estimated using the wavelet components of the variables. Two stage least squares (2SLS) estimates of the coefficients, together with their robust standard errors, are included in Table 3.2 (see, Appendix A.3 for the relevant R code). The instrumental variables in the 2SLS procedure consisted of two lags for each one of the regressors, since it is expected that the three regressors are jointly determined (endogenous) and are therefore correlated with the error term of the model. The standard errors were estimated using the heteroskedasticity and autocorrelation consistent covariance matrix estimation method proposed by Newey and West (1987) with a truncation parameter of length 2. The first part of the table includes estimates of the model (3.6) coefficients, and the second part (standardised coefficients) includes estimates of the model (3.7) coefficients.

The first column, f(t), includes coefficient estimates from a regression using the actual data for each variable (without the application of a wavelet transform) and is included for comparison with the time-scale regressions. Only the coefficients for M3 and inflation have the expected sign and only the coefficient for M3 is statistically significant at the 5% level in this case. The standardised coefficient estimates in the second part of Table 3.2 show that M3 has the second biggest effect on interest rates which suggests the existence of a liquidity effect in the UK economy. Based on these results it is not possible to draw clear conclusions concerning the liquidity effect because the observed time series provide only a mixture of the different frequency variations that characterize the phenomena. In order to decompose the variability into different frequencies it is necessary to use wavelets.

Table 3.2 UK time scale regressions

	f(t)	**D1**	**D2**	**D3**	**D4**	**D5**	**D6**	**D7**	**D8**
Intercept	32.985*	-3.759**	-1.205	-4.107	8.402**	0.964	1.762	128.643*	-26.887
	(7.515)	(1.906)	(2.572)	(2.862)	(3.408)	(3.266)	(3.490)	(1.756)	(18.812)
M3	-0.028*	-0.035*	-0.035*	-0.046*	-0.047*	-0.030*	-0.024*	-0.053*	-0.037*
	(0.006)	(0.007)	(0.007)	(0.005)	(0.004)	(0.002)	(0.002)	(0.001)	(0.001)
Production	-0.253*	0.003	0.073	0.032	0.137*	0.058**	0.055*	0.440*	0.081*
	(0.075)	(0.051)	(0.048)	(0.037)	(0.030)	(0.020)	(0.016)	(0.005)	(0.017)
Inflation	177.705	-6.142	97.516**	844.136*	1727.09*	1707.02*	1782.66*	2871.14*	1736.43*
	(275.400)	(36.375)	(45.670)	(69.241)	(80.800)	(59.906)	(42.003)	(17.079)	(51.943)
R-square	0.53	0.201	0.158	0.57	0.855	0.96	0.921	0.99	0.99
Standardized Coefficients									
M3	-0.285	-0.454	-0.442	-0.478	-0.434	-0.299	-0.247	-0.414	-0.368
Production	-0.439	0.006	0.131	0.048	0.209	0.113	0.086	0.488	0.168
Inflation	0.251	-0.019	0.141	0.523	0.803	0.911	0.895	1.043	0.948

*Note: Standard errors in parenthesis; Significant at: *1% level; **5% level*

The regression results for the eight time-scales are included under columns D1-D8. All the coefficients have the expected sign except for the coefficient for inflation in time-scale D1, which is negative and could be associated with noise in the data. All the coefficients for M3 are statistically significant at the 5% significance level. Most of the other coefficients are statistically significant at the 5% significance level, except for the coefficients for production in time-scales D1-D3 and the coefficient for inflation in time-scale D1. The R-square values are higher for time-scales D4-D8, which capture the low-frequency, long-term cyclical movements in the data. The coefficient for the money supply (M3) is higher in time-scales D1-D4 (higher frequencies) and D7, while the coefficients for industrial production and inflation vary by time-scale. Accordingly, interest rates can rise or fall when examined within a specific time-scale, depending on whether the income and price effects are higher than the liquidity effect.

To be able to compare the three effects within each time-scale, it is necessary to standardise the coefficients as described in sub-section 3.3.3. When only time-scales D1-D2 are examined, the standardised coefficient of the liquidity effect is higher than the standardised coefficients of the income and price effects. It should therefore be expected that interest rates will decrease in the short-term following an increase in the money supply.

When considered together, the total income and price effects are higher than the liquidity effect in all the other time-scales. Time-scale D8 captures the long-term trend in the data, and the coefficients at this scale suggest that, in the long-term, the income and price effects become more important than the liquidity effect. Therefore, interest rates rise as suggested by Friedman (1968).

The short-term liquidity effect identified in time-scales D1-D2 is consistent with the work of Cochrane (1989), who estimated a negative relationship between money growth and interest rates in the US when only the high-frequency data are considered. Also, the results suggest that the inflation effect becomes more important than the liquidity effect at time-scale D3, which covers cyclical movements with a duration of 8-16 months. This result is consistent with the empirical studies of Mishkin (1982), Leeper and Gordon (1992) and Strongin (1995), who reported a positive relationship between money supply and interest rates in the long-term in the US.

3.5 Results with alternative approaches

In this section the previous analysis is compared with two alternative approaches. The first examines the liquidity effect using an alternative model specification with two additional explanatory variables: the short-term interest rate in the US economy (US int. rate -1) and the Sterling/USD effective exchange rate (US ex. rate -1). According to the real interest rate differentials model of exchange rate determination proposed by Frankel (1979), a short-run change (deviations from its long-run equilibrium) in the exchange rate of an economy reflects a disequilibrium between the real domestic and real foreign interest rates. If the real domestic interest rate is lower than the real foreign interest rate, then the real exchange rate of the domestic currency will be undervalued from its long-run equilibrium and should be expected to appreciate to compensate (see, Pilbeam, 2010, p.315).

The two additional variables were included in the model specification with a one period time lag since their contemporaneous values did not provide statistically significant coefficients. Additional variables were also considered for inclusion in the model (e.g. the exchange rate against the euro) but did not provide statistically significant coefficients. Estimation was performed with the 2SLS method using as instrumental variables two lags for each one of the original regressors (M3, production and inflation) and the two new lagged variables. The results are presented in Table 3.3.

Table 3.3 UK time scale regressions with additional variables

	f(t)	**D1**	**D2**	**D3**	**D4**	**D5**	**D6**	**D7**	**D8**
Intercept	30.592*	11.051*	21.032*	-10.898**	29.332*	51.448*	-15.146**	355.077*	-572.234*
	(4.939)	(2.503)	(3.372)	(4.168)	(5.976)	(6.039)	(8.528)	(16.469)	(35.811)
M3	0.002	-0.043*	-0.040*	-0.453*	-0.036*	-0.017*	-0.030*	0.018*	-0.018
	(0.006)	(0.006)	(0.006)	(0.096)	(0.004)	(0.003)	(0.002)	(0.006)	(0.002)
Production	-0.292*	0.214*	0.281*	2.698**	0.220*	0.275*	0.040***	0.576*	0.465*
	(0.053)	(0.053)	(0.048)	(0.792)	(0.039)	(0.027)	(0.024)	(0.003)	(0.018)
Inflation	-114.145	-5.599	70.613***	664.221*	1706.44*	1772.11*	1858.87**	1876.18*	2151.47*
	(207.01)	(32.991)	(40.354)	(115.860)	(79.710)	(52.747)	(47.429)	(97.529)	(28.771)
US int. rate -	0.906*	0.277*	0.312*	0.403*	0.092*	0.145*	-0.034***	0.461*	0.189
	(0.086)	(0.033)	(0.035)	(0.094)	(0.023)	(0.017)	(0.019)	(0.036)	(0.016)
US ex. rate -	0.001	0.001	0.0001	-2.743*	0.004*	0.001	0.001*	0.001*	0.001**
	(0.001)	(0.001)	(0.001)	(0.801)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
R-square	0.746	0.372	0.36	0.612	0.868	0.972	0.933	0.99	0.99
Standardized Coefficients									
M3	0.020	-0.565	-0.504	-4.719	-0.335	-0.175	-0.307	0.144	-0.185
Production	-0.507	0.396	0.501	4.020	0.335	0.531	0.062	0.639	0.967
Inflation	-0.161	-0.017	0.102	0.411	0.791	0.945	0.933	0.681	1.174
US Int. Rate	0.060	0.040	0.046	0.403	0.025	0.050	-0.008	0.102	0.071
US Ex. Rate	0.028	0.038	0.048	-2.743	0.097	0.007	-0.088	0.680	-0.009

*Note: Standard errors in parenthesis; Significant at: *1% level; **5%; ***10% level*

The coefficient estimates are similar to the estimates in Table 3.2 except now in time-scale D7 the coefficient for M3 is positive. The coefficients for the industrial production index and inflation are statistically significant in all time-scales except for inflation in time-scale D1 which could be due to the existence of noise in the data. With regards to the two new variables, the coefficients for the US interest rate are statistically significant in all time-scales except from time-scale D8, while for the exchange rate variable the coefficients are statistically significant only in time-scales D4 and D6-D8. The standardized coefficients provide similar conclusions as in the case of Table 3.2. The liquidity effect is higher and dominates the income and price effects in time-scales D1-D3. The situation is reversed in time-scales D4-D8 with the price and income effects dominating the liquidity effect when considered jointly.

The second approach considered in this section is similar to the one proposed by Cochrane (1989) but uses the Baxter-King (BK) filter in order to isolate the different frequency components in the variables of model (3.6). Band-pass filters like the Baxter-King and Hodrick-Prescott filters are frequently used in economics for de-trending time series and approximating their business cycle components (see, Gencay et. al., 2001, pp. 44-47). Eight different filtering procedures were performed using a BK filter of length 2, each time using a frequency band that covered the same frequencies as in the case of the wavelet filter. Filter BK1 isolates the same frequencies as captured by time-scale component D1 and similarly filter BK8 isolates the long-term trend in the data as captured by time-scale D8. Using the BK filtered data eight different regression models were estimated that corresponded to the frequencies covered by the time-scale regressions D1-D8. Estimation was again performed with the 2SLS method using two lags from each variable as instruments. The results are presented in Table 3.4.

The coefficient estimates for the money supply variable are not statistically significant in three time-scales, BK1, BK2 and BK7 and have a positive sign after time-scale BK4. The coefficient for inflation in time-scale BK1 is again not statistically significant and the same is true for the coefficient of the industrial production index in time-scales BK1 and BK2. Further, the inflation and production coefficients do not have the expected sign in several time-scales. The inflation coefficient is negative in time-scales BK2-BK4 and BK8 and the production coefficient is negative in time-scale BK6. In terms of the standardized coefficients, the liquidity effect is higher than the price and income effects only in time-scale BK1. In all the other time-scales the results are not conclusive because some of the coefficients do not have the expected sign.

When compared with wavelets, band-pass filters have two main disadvantages: (i) they are less suitable for analyzing non-stationary time series and (ii) they provide inaccurate estimates of the cyclical components when the data are characterized by low-frequencies (see, Guay and St-Amman, 2005). As evident from Figures 1 and 2 the short-term interest rate is characterized by both non-stationary behaviour and considerable low-frequency variation. As a result, it should be expected that wavelet analysis can provide better estimates of the frequency components in the time series and a more accurate measurement of the liquidity effect than the BK filter.

Table 3.4 UK time scale regressions with Baxter-King filtered data

	BK1	**BK2**	**BK3**	**BK4**	**BK5**	**BK6**	**BK7**	**BK8**
Intercept	-0.001	0.001	0.003	-0.036***	0.017	-0.028	0.013	-4.652
	(0.005)	(0.010)	(0.012)	(0.019)	(0.024)	(0.022)	(0.008)	(5.962)
M3	-0.094	-0.042	-0.083**	0.168*	0.168*	0.119*	0.002	0.156*
	(0.095)	(0.048)	(0.036)	(0.051)	(0.024)	(0.011)	(0.006)	(0.015)
Production	-0.003	0.001	0.195*	0.303*	0.485*	-0.083*	0.383*	0.161**
	(0.016)	(0.036)	(0.047)	(0.057)	(0.034)	(0.027)	(0.009)	(0.065)
Inflation	2.091	-9.675**	-29.00*	-282.312*	450.184*	1824.022*	2251.24*	-101.757*
	(2.472)	(4.148)	(8.766)	(48.309)	(61.400)	(54.882)	(44.845)	(17.693)
R-square	0.007	0.02	0.133	0.269	0.606	0.9	0.963	0.792
Standardized Coefficients								
M3	-0.137	-0.062	-0.140	0.181	0.293	0.242	0.008	0.981
Production	-0.013	0.002	0.251	0.296	0.620	-0.064	1.060	0.224
Inflation	0.071	-0.152	-0.199	-0.343	0.314	0.810	1.749	-0.216

*Note: Standard errors in parenthesis; Significant at: *1%; **5% level; ***10% level*

3.6 Discussion and policy implications

The analysis of the liquidity effect for the UK economy, presented in this chapter, underlines an important recommendation for practitioners examining the impact of money supply changes on interest rates. The application of band-pass filters to time series data like interest-rates, money supply, inflation and income should be used with extreme caution. This is because macroeconomic and financial variables typically evolve over several different cycles in time. As a result, the observed sampling rate of the data provides only a mixture of the different cycles that characterize a phenomenon.

When the purpose of the econometric analysis is to understand the short-term, medium-term and long-term effects of money supply changes on interest rates, it is important to ensure that the application of any filtering procedures on the variables will not distort their cyclical characteristics. Most frequently, studies of the liquidity effect involve first-differencing of the data (see, for example Lastrabes and Selgin, 1995) in order to remove stochastic trends from non-stationary time series. Although sensible from an estimation point of view, first-differencing distorts the cyclical characteristics of the data by re-weighting the information content in the series towards higher frequencies and short-term cycles. As a consequence, long-term relationships between variables cannot be estimated accurately and short-term effects are over-emphasized.

The wavelet methodology proposed in this chapter is a modern mathematical tool that can effectively address several of the issues that complicate the analysis of the liquidity effect. The wavelet technology is especially designed in order to analyse non-stationary time series without the need to pre-filter the data. It is therefore able to retain and extract the cyclical characteristics of the data. By decomposing and isolating the different cycles in the data, it permits the examination of the liquidity effect on a scale by scale basis. This is an important distinction in the examination of the liquidity effect because both theoretical and empirical studies suggest that different factors influence interest rates in the short-term than in the long-term.

With regards to the UK economy examined in this chapter, the results suggest that the impact of the liquidity, price and income effects vary by time-scale: in the short-term, interest rates are influenced primarily by money supply changes, but in the medium and long- terms, income and price effects become more important. The average duration of the liquidity effect is estimated to be less than 8 months.

These results are more closely related with the quantity-theory monetarism view in economics. According to this, due to the long-term effects imposed on interest rates by the money supply-prices and money supply-income relationships, it is recommended to keep the supply of money growing at a constant rate over short-term and medium-term horizons. As noted by Friedman (1956), this is recommended because money supply changes exert influence on prices and the economy over long time horizons and quite often in complicated and unpredictable ways. Consequently, the use of money supply changes as an economic policy tool can often have undesirable and destabilising effects.

CONCLUSIONS

In Chapter 1 a conjectural variations model for measuring competition in banking was developed using an opportunity cost concept. In contrast, most of the existing research in the competitive banking literature has relied on complete functional form specifications for the total cost. The adoption of an opportunity cost concept is sensible in banking because the available banking micro-data, used in most empirical studies, do not provide detailed cost information on the different financial products offered by banks. The interest rate offered by the monetary authorities on minimum reserves was used as the opportunity cost variable. It was incorporated in the equilibrium price condition through the inclusion of interest income from minimum reserves in the profit maximisation problem facing a representative bank.

The proposed model has two important advantages relative to the existing literature: (i) Conjectural price reaction elasticities are incorporated in the modeling framework. As a result, the explanatory power of the simultaneous equations system, derived from the theoretical model, is significantly enhanced. (ii) Estimates of market power are based on an opportunity cost concept and not on a complete cost function specification. This is desirable in banking because most empirical studies estimate complete cost function specifications using micro-data from financial statements. These data are frequently characterized by measurement errors and provide biased coefficient estimates. Estimates of the market power parameter, with the proposed model, rejected the monopoly hypothesis for the banking industry in Cyprus.

In Chapter 2 a method for testing for the existence of nonlinear pricing in the European credit market for loans was proposed. The test is based on a Baumol model of cash demand that includes quantity discounts. Such discounts are frequently used in the EOQ model literature, upon which the Baumol model was based. The interest rate spread associated with cash demand with and without quantity discounts, can be modelled as a function of market concentration in econometric models. It provides a direct test for the relationship between market structure and quantity discounts. The results based on a panel dataset consisting of seven European countries and eight annual time periods suggest that higher market concentration (increasing monopoly power) is associated with higher interest rate spreads (quantity discounts) in the credit market for loans.

Another implication of this finding is that monopoly behaviour can manifest itself in many different formats in banking. The ability of banks to adopt non-linear pricing tactics with regards to loan provisions is one such case that was demonstrated theoretically and tested empirically in Chapter 2. From a policy perspective, it is important for the monetary authorities to ensure that liquidity and project financing, in the market for loans, is equally accessible to the whole distribution of consumers and firms.

Chapter 3 concentrated on the examination of the liquidity effect using wavelet analysis. Wavelets can decompose time series into different time-scales, which allows us to identify how the money supply, inflation and income influence interest rates throughout the cycle. This influence is not evident when only the observed sampling rate of the data is studied because sampling provides a mixture of the different frequencies and masks differences between short-term and long-term relationships. This is an important distinction in the examination of the liquidity effect since both theoretical and empirical studies suggest that different factors influence interest rates in the short-term than in the long-term.

Another problem identified in the existing literature concerning the measurement of the liquidity effect, refers to the filtering procedures applied on time series used in econometric models, most commonly first differencing. First differencing re-weights the information content in the time series towards higher frequencies and short-term cycles. At the same time it down-weights the lower frequencies that are associated with the long-term cyclical components in the time series. As a consequence, long-term causal effects cannot be estimated accurately and short-term effects are over-emphasized. The wavelet technology is especially designed in order to analyse non-stationary time series without the need to pre-filter the data. It is therefore able to retain and extract the cyclical characteristics of the data, which enables accurate measurements of the liquidity effect.

Using data for the UK economy, the empirical analysis in Chapter 3 examined the liquidity, income and price effects on short-term interest rates over different time-scales. The results suggest that the impact of each effect varies by time-scale: in the short-term, interest rates are influenced primarily by the liquidity effect, but in the medium and long-terms, income and price effects become more important. The average duration of the liquidity effect is estimated to be less than 8 months.

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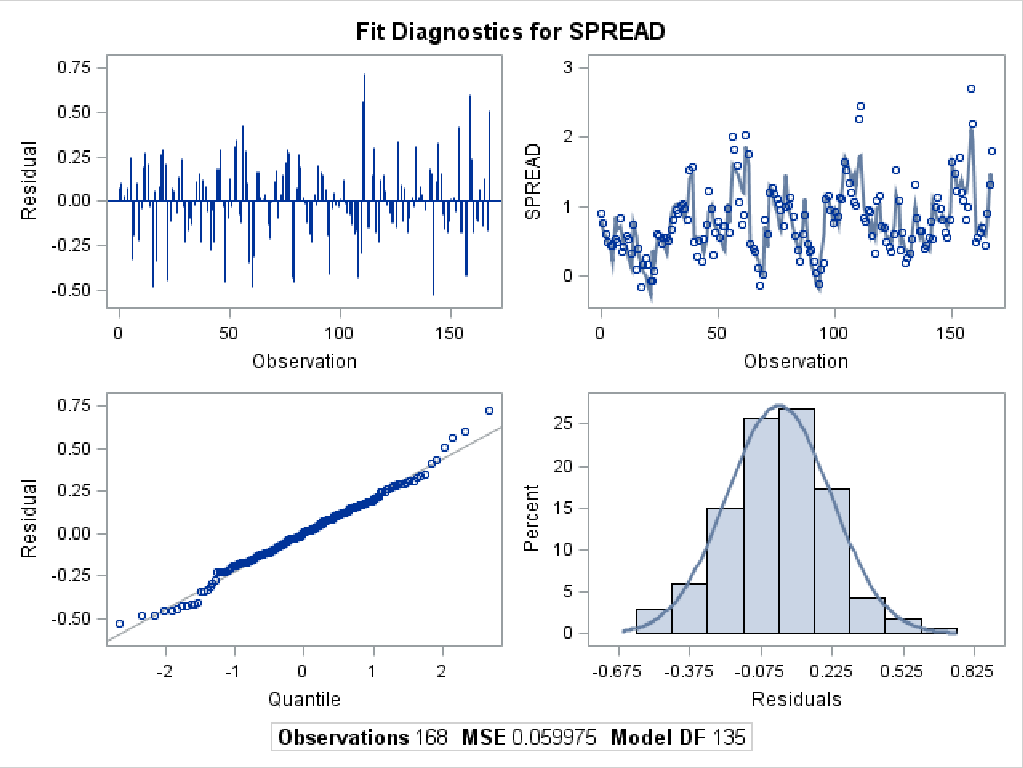
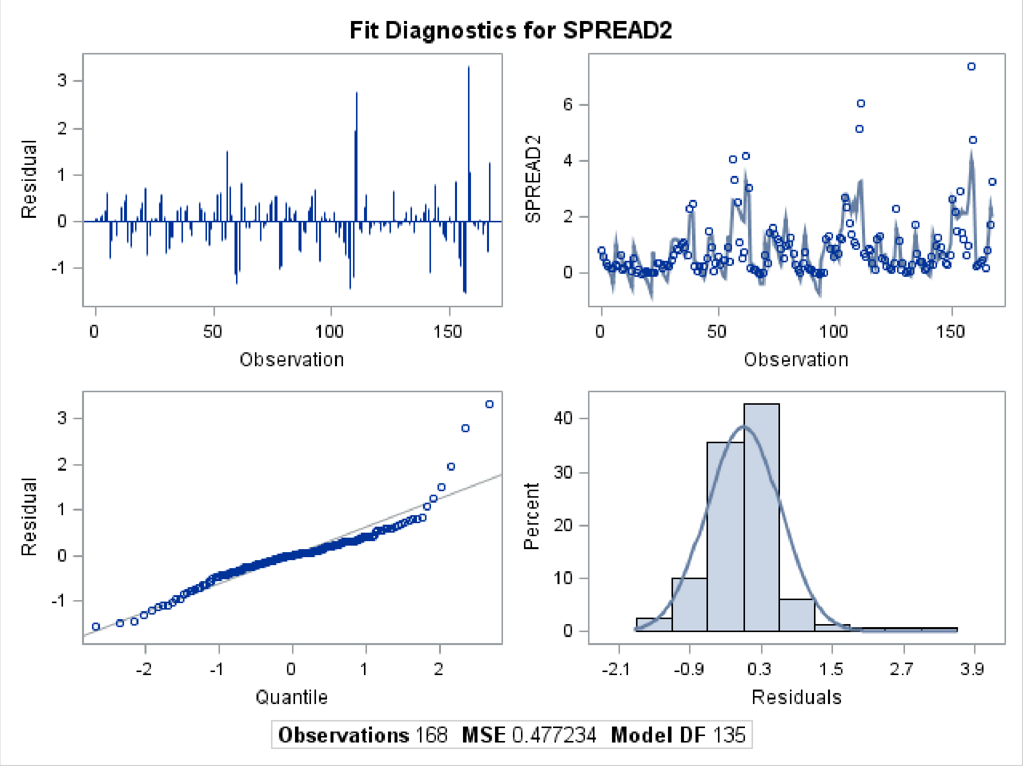
APPENDICES

A.1 Chapter 2: model fit diagnostics

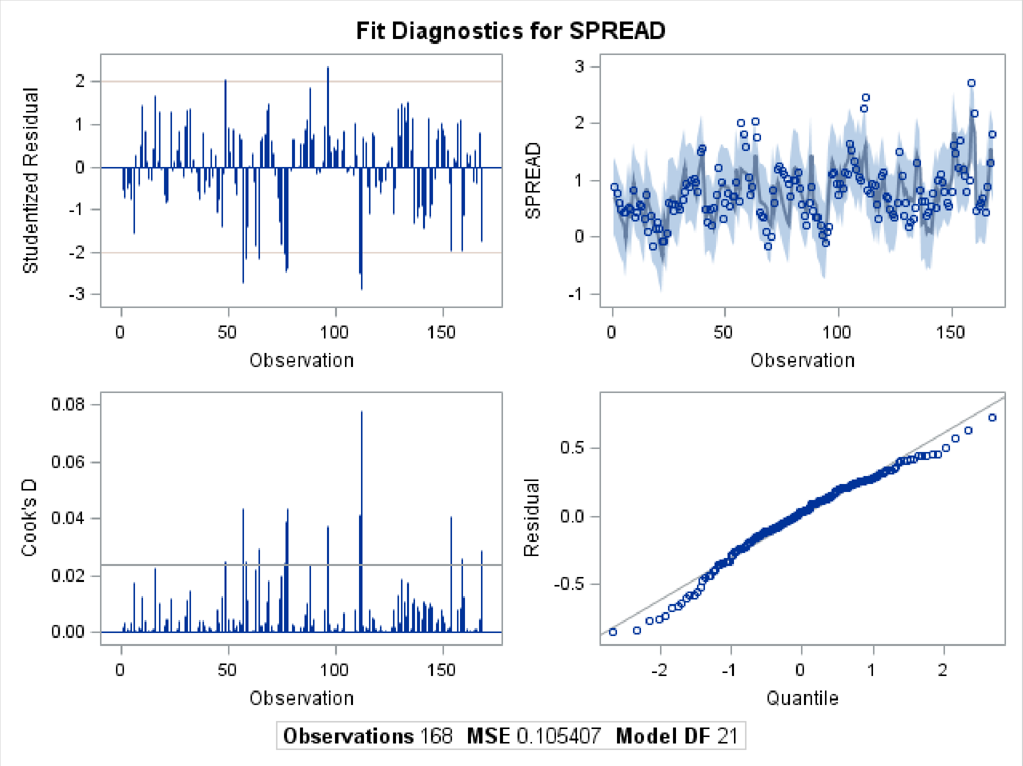
Dependent variable: SPREAD – OLS fit diagnostics

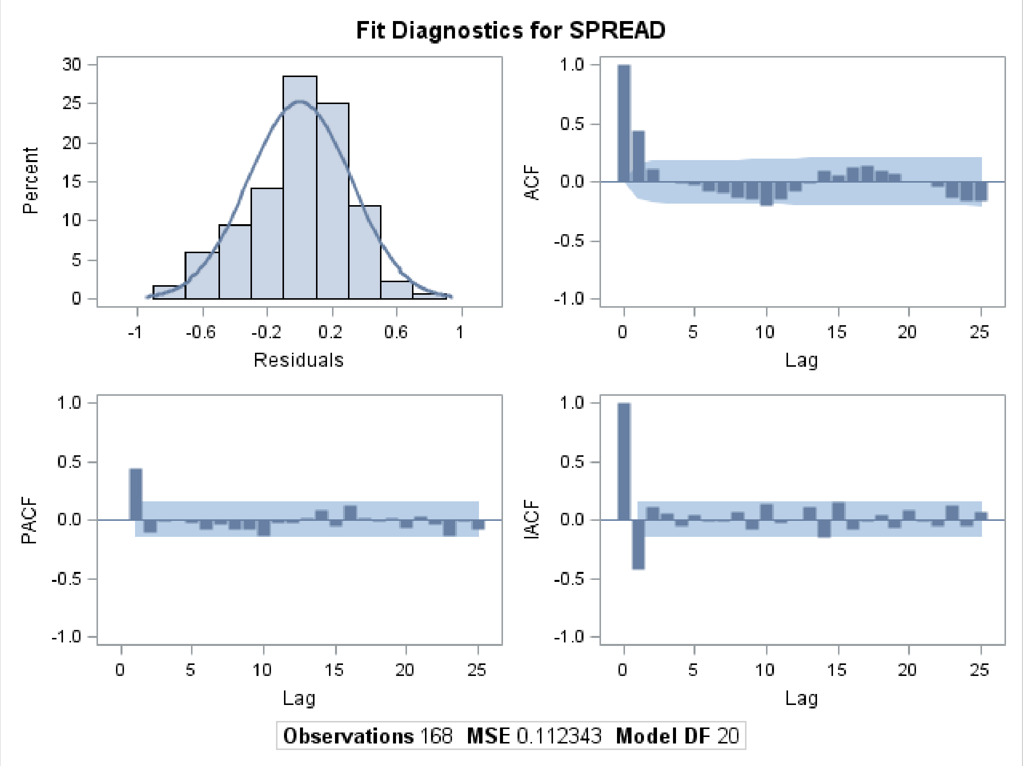


Dependent variable: SPREAD2 – OLS fit diagnostics

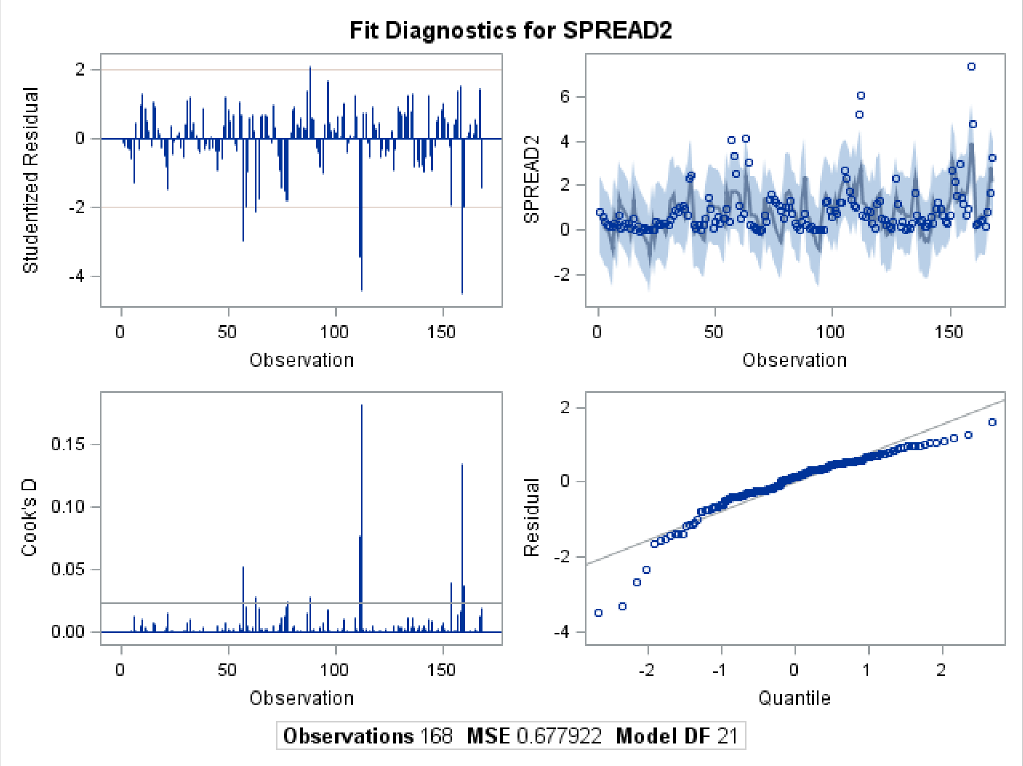


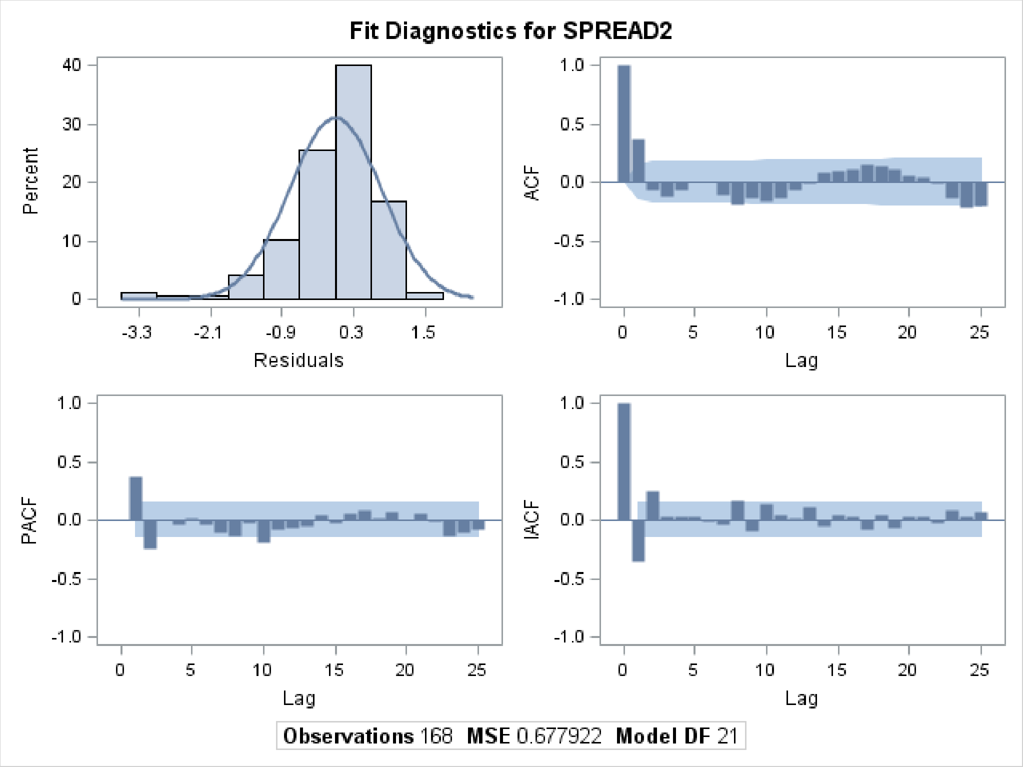
Dependent variable: SPREAD – 2SLS fit diagnostics





Dependent variable: SPREAD2 – 2SLS fit diagnostics





A.2 Chapter 3: computation of wavelet transforms

As noted in Section 3.3, in empirical applications both the DWT and the MODWT are computed using a pyramid algorithm. The algorithm utilizes two filters in order to iteratively filter the data: the first is a high pass (wavelet) filter with coefficients $(h_0,\ldots,h_{L-1})$ and the second is a low pass (scaling) filter with coefficients $(g_0,\ldots,g_{L-1})$. In the case of the DWT, the first stage of the algorithm consists of filtering an observed time series χ_t (of dyadic length $T=2^J$) with the two filters and downsampling the resulting outcome such that every other value of the filtered data is removed. The resulting wavelet and scaling coefficients can be represented as follows:

$$w_{1,t} = \sum_{l=0}^{L-1} h_l \, \chi_{2t+1-l \bmod N} \qquad \text{and} \qquad v_{1,t} = \sum_{l=0}^{L-1} g_l \, \chi_{2t+1-l \bmod N} \; .$$

The downsampling operation has been incorporated in the filtering procedure though the subscript of χ_t (see, Gencay et. al., 2000, p. 122), as a result of which the two generated vectors of coefficients are restricted to have length $T/2$. In the second stage of the algorithm, the two filters are applied to the vector of scaling coefficients v_1, in order to generate the second level of wavelet and scaling coefficients vectors w_2 and v_2. Due to the downsampling procedure incorporated in the filters, these vectors will now have length $T/4$. The third stage of the algorithm proceeds in a similar way by applying the two filters to the second stage vector of scaling coefficients v_2, in order to generate the third level of wavelet and scaling coefficient vectors w_3 and v_3. Accordingly, due to the downsampling procedure, these vectors will now have length $T/8$. This filtering procedure is repeated up to J times and the resulting vector of DWT coefficients is $w = [w_1,..., w_J, v_J]^T$.

In the case of the MODWT, downsampling is removed and the algorithm starts by filtering the observed time series χ_t with the rescaled wavelet and scaling filters $\tilde{h}_j = h_j / 2^j$ and $\tilde{g}_J = g_J / 2^J$. The first stage of the pyramid algorithm generates the following vectors of coefficients:

$$\tilde{w}_{1,t} = \sum_{l=0}^{L-1} \tilde{h}_l \, \chi_{t-l \bmod N} \qquad \text{and} \qquad \tilde{v}_{1,t} = \sum_{l=0}^{L-1} \tilde{g}_l \, \chi_{t-l \bmod N} .$$

Since no downsampling procedure is involved in this case, the resulting vectors of wavelet and scaling coefficients will now have length T. In the second stage of the algorithm the rescaled filters are applied to the vector of scaling coefficients $\tilde{v}_1$, in order to generate the second level of wavelet and scaling coefficients vectors $\tilde{w}_2$ and $\tilde{v}_2$. In the absence of downsampling both vectors will have length T. The third stage of the algorithm proceeds in a similar way by applying the rescaled filters on the vector $\tilde{v}_2$. This procedure is repeated up to J times in order to produce the matrix of MODWT coefficients vectors $\tilde{w} = [\tilde{w}_1,..., \tilde{w}_J, \tilde{v}_J]^T$ and where each vector has length T.

A.3 Chapter 3: R code for the computation of standardized time-scale regressions

```
library(wavethresh)

library(zoo)

library(QuantPsyc)

library(sandwich)

library(gmm)

APT <- read.csv2("C:UK.CSV")

attach(APT)

detach("APT")


INTUK <- APT$INTUK

M3UK  <- APT$M3UK

PRUK  <- APT$PRUK

INUK  <- APT$INUK

INTUS <- APT$INTUS

EXUS  <- APT$EXUS

EXEUR <- APT$EXEUR

M3UK1  <- APT$M3UK1

PRUK1  <- APT$PRUK1

INUK1  <- APT$INUK1

INTUS1 <- APT$INTUS1

EXUS1  <- APT$EXUS1

EXEUR1 <- APT$EXEUR1
```

```
M3UK2  <- APT$M3UK2

PRUK2  <- APT$PRUK2

INUK2  <- APT$INUK2

INTUS2 <- APT$INTUS2

EXUS2  <- APT$EXUS2

EXEUR2 <- APT$EXEUR2

res <- lm(INTUK ~ M3UK + PRUK + INUK + INTUS + EXUS + EXEUR)

BETA <- lm.beta(res)

MODWT0 <- wd(INTUK, type="station", filter.number=8, family="DaubLeAsymm")

plot.wd(MODWT0)

SHORT1 <- accessC.wd(MODWT0, level=1)

SHORT2 <- accessC.wd(MODWT0, level=2)

SHORT3 <- accessC.wd(MODWT0, level=3)

SHORT4 <- accessC.wd(MODWT0, level=4)

SHORT5 <- accessC.wd(MODWT0, level=5)

SHORT6 <- accessC.wd(MODWT0, level=6)

SHORT7 <- accessC.wd(MODWT0, level=7)

SHORT8 <- accessC.wd(MODWT0, level=8)

MODWT1 <- wd(M3UK, type="station", filter.number=8, family="DaubLeAsymm")

M1 <- accessC.wd(MODWT1, level=1)

M2 <- accessC.wd(MODWT1, level=2)

M3 <- accessC.wd(MODWT1, level=3)

M4 <- accessC.wd(MODWT1, level=4)

M5 <- accessC.wd(MODWT1, level=5)
```

```
M6 <- accessC.wd(MODWT1, level=6)

M7 <- accessC.wd(MODWT1, level=7)

M8 <- accessC.wd(MODWT1, level=8)

MODWT2 <- wd(M3UK1, type="station", filter.number=8, family="DaubLeAsymm")

M11 <- accessC.wd(MODWT2, level=1)

M12 <- accessC.wd(MODWT2, level=2)

M13 <- accessC.wd(MODWT2, level=3)

M14 <- accessC.wd(MODWT2, level=4)

M15 <- accessC.wd(MODWT2, level=5)

M16 <- accessC.wd(MODWT2, level=6)

M17 <- accessC.wd(MODWT2, level=7)

M18 <- accessC.wd(MODWT2, level=8)

MODWT3 <- wd(M3UK2, type="station", filter.number=8, family="DaubLeAsymm")

M21 <- accessC.wd(MODWT3, level=1)

M22 <- accessC.wd(MODWT3, level=2)

M23 <- accessC.wd(MODWT3, level=3)

M24 <- accessC.wd(MODWT3, level=4)

M25 <- accessC.wd(MODWT3, level=5)

M26 <- accessC.wd(MODWT3, level=6)

M27 <- accessC.wd(MODWT3, level=7)

M28 <- accessC.wd(MODWT3, level=8)

MODWT4 <- wd(PRUK, type="station", filter.number=8, family="DaubLeAsymm")

PR1 <- accessC.wd(MODWT4, level=1)

PR2 <- accessC.wd(MODWT4, level=2)
```

```
PR3 <- accessC.wd(MODWT4, level=3)
PR4 <- accessC.wd(MODWT4, level=4)
PR5 <- accessC.wd(MODWT4, level=5)
PR6 <- accessC.wd(MODWT4, level=6)
PR7 <- accessC.wd(MODWT4, level=7)
PR8 <- accessC.wd(MODWT4, level=8)
MODWT5 <- wd(PRUK1, type="station", filter.number=8, family="DaubLeAsymm")
PR11 <- accessC.wd(MODWT5, level=1)
PR12 <- accessC.wd(MODWT5, level=2)
PR13 <- accessC.wd(MODWT5, level=3)
PR14 <- accessC.wd(MODWT5, level=4)
PR15 <- accessC.wd(MODWT5, level=5)
PR16 <- accessC.wd(MODWT5, level=6)
PR17 <- accessC.wd(MODWT5, level=7)
PR18 <- accessC.wd(MODWT5, level=8)
MODWT6 <- wd(PRUK2, type="station", filter.number=8, family="DaubLeAsymm")
PR21 <- accessC.wd(MODWT6, level=1)
PR22 <- accessC.wd(MODWT6, level=2)
PR23 <- accessC.wd(MODWT6, level=3)
PR24 <- accessC.wd(MODWT6, level=4)
PR25 <- accessC.wd(MODWT6, level=5)
PR26 <- accessC.wd(MODWT6, level=6)
PR27 <- accessC.wd(MODWT6, level=7)
PR28 <- accessC.wd(MODWT6, level=8)
```

```
MODWT7 <- wd(INUK, type="station", filter.number=8, family="DaubLeAsymm")
IN1 <- accessC.wd(MODWT7, level=1)
IN2 <- accessC.wd(MODWT7, level=2)
IN3 <- accessC.wd(MODWT7, level=3)
IN4 <- accessC.wd(MODWT7, level=4)
IN5 <- accessC.wd(MODWT7, level=5)
IN6 <- accessC.wd(MODWT7, level=6)
IN7 <- accessC.wd(MODWT7, level=7)
IN8 <- accessC.wd(MODWT7, level=8)
MODWT8 <- wd(INUK1, type="station", filter.number=8, family="DaubLeAsymm")
IN11 <- accessC.wd(MODWT8, level=1)
IN12 <- accessC.wd(MODWT8, level=2)
IN13 <- accessC.wd(MODWT8, level=3)
IN14 <- accessC.wd(MODWT8, level=4)
IN15 <- accessC.wd(MODWT8, level=5)
IN16 <- accessC.wd(MODWT8, level=6)
IN17 <- accessC.wd(MODWT8, level=7)
IN18 <- accessC.wd(MODWT8, level=8)
MODWT9 <- wd(INUK2, type="station", filter.number=8, family="DaubLeAsymm")
IN21 <- accessC.wd(MODWT9, level=1)
IN22 <- accessC.wd(MODWT9, level=2)
IN23 <- accessC.wd(MODWT9, level=3)
IN24 <- accessC.wd(MODWT9, level=4)
IN25 <- accessC.wd(MODWT9, level=5)
```

```
IN26 <- accessC.wd(MODWT9, level=6)
IN27 <- accessC.wd(MODWT9, level=7)
IN28 <- accessC.wd(MODWT9, level=8)
MODWT10 <- wd(INTUS, type="station", filter.number=8, family="DaubLeAsymm")
USIN1 <- accessC.wd(MODWT10, level=1)
USIN2 <- accessC.wd(MODWT10, level=2)
USIN3 <- accessC.wd(MODWT10, level=3)
USIN4 <- accessC.wd(MODWT10, level=4)
USIN5 <- accessC.wd(MODWT10, level=5)
USIN6 <- accessC.wd(MODWT10, level=6)
USIN7 <- accessC.wd(MODWT10, level=7)
USIN8 <- accessC.wd(MODWT10, level=8)
MODWT11 <- wd(INTUS1, type="station", filter.number=8, family="DaubLeAsymm")
USIN11 <- accessC.wd(MODWT11, level=1)
USIN12 <- accessC.wd(MODWT11, level=2)
USIN13 <- accessC.wd(MODWT11, level=3)
USIN14 <- accessC.wd(MODWT11, level=4)
USIN15 <- accessC.wd(MODWT11, level=5)
USIN16 <- accessC.wd(MODWT11, level=6)
USIN17 <- accessC.wd(MODWT11, level=7)
USIN18 <- accessC.wd(MODWT11, level=8)
MODWT12 <- wd(INTUS2, type="station", filter.number=8, family="DaubLeAsymm")
USIN21 <- accessC.wd(MODWT12, level=1)
USIN22 <- accessC.wd(MODWT12, level=2)
```

```
USIN23 <- accessC.wd(MODWT12, level=3)

USIN24 <- accessC.wd(MODWT12, level=4)

USIN25 <- accessC.wd(MODWT12, level=5)

USIN26 <- accessC.wd(MODWT12, level=6)

USIN27 <- accessC.wd(MODWT12, level=7)

USIN28 <- accessC.wd(MODWT12, level=8)

MODWT13 <- wd(EXUS, type="station", filter.number=8, family="DaubLeAsymm")

USEX1 <- accessC.wd(MODWT13, level=1)

USEX2 <- accessC.wd(MODWT13, level=2)

USEX3 <- accessC.wd(MODWT13, level=3)

USEX4 <- accessC.wd(MODWT13, level=4)

USEX5 <- accessC.wd(MODWT13, level=5)

USEX6 <- accessC.wd(MODWT13, level=6)

USEX7 <- accessC.wd(MODWT13, level=7)

USEX8 <- accessC.wd(MODWT13, level=8)

MODWT14 <- wd(EXUS1, type="station", filter.number=8, family="DaubLeAsymm")

USEX11 <- accessC.wd(MODWT14, level=1)

USEX12 <- accessC.wd(MODWT14, level=2)

USEX13 <- accessC.wd(MODWT14, level=3)

USEX14 <- accessC.wd(MODWT14, level=4)

USEX15 <- accessC.wd(MODWT14, level=5)

USEX16 <- accessC.wd(MODWT14, level=6)

USEX17 <- accessC.wd(MODWT14, level=7)

USEX18 <- accessC.wd(MODWT14, level=8)
```

```
MODWT15 <- wd(EXUS2, type="station", filter.number=8, family="DaubLeAsymm")
USEX21 <- accessC.wd(MODWT15, level=1)
USEX22 <- accessC.wd(MODWT15, level=2)
USEX23 <- accessC.wd(MODWT15, level=3)
USEX24 <- accessC.wd(MODWT15, level=4)
USEX25 <- accessC.wd(MODWT15, level=5)
USEX26 <- accessC.wd(MODWT15, level=6)
USEX27 <- accessC.wd(MODWT15, level=7)
USEX28 <- accessC.wd(MODWT15, level=8)

MODWT16 <- wd(EXEUR, type="station", filter.number=8, family="DaubLeAsymm")
EUEX1 <- accessC.wd(MODWT16, level=1)
EUEX2 <- accessC.wd(MODWT16, level=2)
EUEX3 <- accessC.wd(MODWT16, level=3)
EUEX4 <- accessC.wd(MODWT16, level=4)
EUEX5 <- accessC.wd(MODWT16, level=5)
EUEX6 <- accessC.wd(MODWT16, level=6)
EUEX7 <- accessC.wd(MODWT16, level=7)
EUEX8 <- accessC.wd(MODWT16, level=8)
MODWT17 <- wd(EXEUR1, type="station", filter.number=8, family="DaubLeAsymm")
EUEX11 <- accessC.wd(MODWT17, level=1)
EUEX12 <- accessC.wd(MODWT17, level=2)
EUEX13 <- accessC.wd(MODWT17, level=3)
EUEX14 <- accessC.wd(MODWT17, level=4)
```

```
EUEX15 <- accessC.wd(MODWT17, level=5)

EUEX16 <- accessC.wd(MODWT17, level=6)

EUEX17 <- accessC.wd(MODWT17, level=7)

EUEX18 <- accessC.wd(MODWT17, level=8)

MODWT18 <- wd(EXEUR2, type="station", filter.number=8, family="DaubLeAsymm")

EUEX21 <- accessC.wd(MODWT18, level=1)

EUEX22 <- accessC.wd(MODWT18, level=2)

EUEX23 <- accessC.wd(MODWT18, level=3)

EUEX24 <- accessC.wd(MODWT18, level=4)

EUEX25 <- accessC.wd(MODWT18, level=5)

EUEX26 <- accessC.wd(MODWT18, level=6)

EUEX27 <- accessC.wd(MODWT18, level=7)

EUEX28 <- accessC.wd(MODWT18, level=8)

DATA1 <-
data.frame(Y=SHORT1,M3=M1,LM3=M11,LLM3=M21,PROD=PR1,LPROD=PR11,LLPR
OD=PR21,INF=IN1,LINF=IN11,LLINF=IN21,USIN=USIN1,LUSIN=USIN11,LLUSIN=U
SIN21,
USEX=USEX1,LUSEX=USEX11,LLUSEX=USEX21,EUEX=EUEX1,LEUEX=EUEX11,L
LEUEX=EUEX21)

DATA2 <- data.frame(Y=SHORT2-SHORT1,M3=M2-M1,LM3=M12-M11,LLM3=M22-
M21,PROD=PR2-PR1,LPROD=PR12-PR11,LLPROD=PR22-PR21,INF=IN2-
IN1,LINF=IN12-IN11,LLINF=IN22-IN21,USIN=USIN2-USIN1,LUSIN=USIN12-
USIN11,LLUSIN=USIN22-USIN21,USEX=USEX2-USEX1,LUSEX=USEX12-
USEX11,LLUSEX=USEX22-USEX21,EUEX=EUEX2-EUEX1,LEUEX=EUEX12-
EUEX11,LLEUEX=EUEX22-EUEX21)

DATA3 <- data.frame(Y=SHORT3-SHORT2,M3=M3-M2,LM3=M13-M12,LLM3=M23-
M22,PROD=PR3-PR2,LPROD=PR13-PR12,LLPROD=PR23-PR22,INF=IN3-
IN2,LINF=IN13-IN12,LLINF=IN23-IN22,USIN=USIN3-USIN2,LUSIN=USIN13-
USIN12,LLUSIN=USIN23-USIN22,USEX=USEX3-USEX2,LUSEX=USEX13-
USEX12,LLUSEX=USEX23-USEX22,EUEX=EUEX3-EUEX2,LEUEX=EUEX13-
EUEX12,LLEUEX=EUEX23-EUEX22)
```

```
DATA4 <- data.frame(Y=SHORT4-SHORT3,M3=M4-M3,LM3=M14-M13,LLM3=M24-
M23,PROD=PR4-PR3,LPROD=PR14-PR13,LLPROD=PR24-PR23,INF=IN4-
IN3,LINF=IN14-IN13,LLINF=IN24-IN23,USIN=USIN4-USIN3,LUSIN=USIN14-
USIN13,LLUSIN=USIN24-USIN23,USEX=USEX4-USEX3,LUSEX=USEX14-
USEX13,LLUSEX=USEX24-USEX23,EUEX=EUEX4-EUEX3,LEUEX=EUEX14-
EUEX13,LLEUEX=EUEX24-EUEX23)

DATA5 <- data.frame(Y=SHORT5-SHORT4,M3=M5-M4,LM3=M15-M14,LLM3=M25-
M24,PROD=PR5-PR4,LPROD=PR15-PR14,LLPROD=PR25-PR24,INF=IN5-
IN4,LINF=IN15-IN14,LLINF=IN25-IN24,USIN=USIN5-USIN4,LUSIN=USIN15-
USIN14,LLUSIN=USIN25-USIN24,USEX=USEX5-USEX4,LUSEX=USEX15-
USEX14,LLUSEX=USEX25-USEX24,EUEX=EUEX5-EUEX4,LEUEX=EUEX15-
EUEX14,LLEUEX=EUEX25-EUEX24)

DATA6 <- data.frame(Y=SHORT6-SHORT5,M3=M6-M5,LM3=M16-M15,LLM3=M26-
M25,PROD=PR6-PR5,LPROD=PR16-PR15,LLPROD=PR26-PR25,INF=IN6-
IN5,LINF=IN16-IN15,LLINF=IN26-IN25,USIN=USIN6-USIN5,LUSIN=USIN16-
USIN15,LLUSIN=USIN26-USIN25,USEX=USEX6-USEX5,LUSEX=USEX16-
USEX15,LLUSEX=USEX26-USEX25,=EUEX6-EUEX5,LEUEX=EUEX16-
EUEX15,LLEUEX=EUEX26-EUEX25)



DATA7 <- data.frame(Y=SHORT7-SHORT6,M3=M7-M6,LM3=M17-M16,LLM3=M27-
M26,PROD=PR7-PR6,LPROD=PR17-PR16,LLPROD=PR27-PR26,INF=IN7-
IN6,LINF=IN17-IN16,LLINF=IN27-IN26,USIN=USIN7-USIN6,LUSIN=USIN17-
USIN16,LLUSIN=USIN27-USIN26,USEX=USEX7-USEX6,LUSEX=USEX17-
USEX16,LLUSEX=USEX27-USEX26,EUEX=EUEX7-EUEX6,LEUEX=EUEX17-
EUEX16,LLEUEX=EUEX27-EUEX26)

DATA8 <- data.frame(Y=SHORT8-SHORT7,M3=M8-M7,LM3=M18-M17,LLM3=M28-
M27,PROD=PR8-PR7,LPROD=PR18-PR17,LLPROD=PR28-PR27,INF=IN8-
IN7,LINF=IN18-IN17,LLINF=IN28-IN27,USIN=USIN8-USIN7,LUSIN=USIN18-
USIN17,LLUSIN=USIN28-USIN27,USEX=USEX8-USEX7,LUSEX=USEX18-
USEX17,LLUSEX=USEX28-USEX27,EUEX=EUEX8-EUEX7,LEUEX=EUEX18-
EUEX17,LLEUEX=EUEX28-EUEX27)

TABLE1 <- data.frame(DATA1)

TABLE2 <- data.frame(DATA2)

TABLE3 <- data.frame(DATA3)

TABLE4 <- data.frame(DATA4)

TABLE5 <- data.frame(DATA5)

TABLE6 <- data.frame(DATA6)
```

```
TABLE7 <- data.frame(DATA7)

TABLE8 <- data.frame(DATA8)

REGRES1 <- tsls(Y ~ M3 + PROD + INF, ~ LM3 + LPROD + LINF, data=DATA1)

BETA1 <- lm.beta(REGRES1)

summary(REGRES1)

vcovHAC(REGRES1,weights=weightsAndrews)

ts.plot(DATA1)

REGRES2 <- tsls(Y ~ M3 + PROD + INF, ~ LM3 + LLM3 + LPROD + LLPROD + LINF +
LLINF, data=DATA2)

BETA2 <- lm.beta(REGRES2)

summary(REGRES2)

vcovHAC(REGRES2,weights=weightsAndrews)

ts.plot(DATA2)

REGRES3 <- tsls(Y ~ M3 + PROD + INF, ~ LM3 + LLM3 + LPROD + LLPROD + LINF +
LLINF, data=DATA3)

BETA3 <- lm.beta(REGRES3)

summary(REGRES3)

vcovHAC(REGRES3,weights=weightsAndrews)

ts.plot(DATA3)

REGRES4 <- tsls(Y ~ M3 + PROD + INF, ~ LM3 + LLM3 + LPROD + LLPROD + LINF +
LLINF, data=DATA4)

BETA4 <- lm.beta(REGRES4)

summary(REGRES4)

vcovHAC(REGRES4,weights=weightsAndrews)

ts.plot(DATA4)
```

```
REGRES5 <- tsls(Y ~ M3 + PROD + INF, ~ LM3 + LLM3 + LPROD + LLPROD + LINF +
LLINF, data=DATA5)

BETA5 <- lm.beta(REGRES5)

summary(REGRES5)

vcovHAC(REGRES5,weights=weightsAndrews)

ts.plot(DATA5)

REGRES6 <- tsls(Y ~ M3 + PROD + INF, ~ LM3 + LLM3 + LPROD + LLPROD + LINF +
LLINF, data=DATA6)

BETA6 <- lm.beta(REGRES6)

summary(REGRES6)

vcovHAC(REGRES6,weights=weightsAndrews)

ts.plot(DATA6)

REGRES7 <- tsls(Y ~ M3 + PROD + INF, ~ LM3 + LLM3 + LPROD + LLPROD + LINF +
LLINF, data=DATA7)

BETA7 <- lm.beta(REGRES7)

summary(REGRES7)

vcovHAC(REGRES7,weights=weightsAndrews)

ts.plot(DATA7)

REGRES8 <- tsls(Y ~ M3 + PROD + INF, ~ LM3 + LLM3 + LPROD + LLPROD + LINF +
LLINF, data=DATA8)

BETA8 <- lm.beta(REGRES8)

summary(REGRES8)

vcovHAC(REGRES8,weights=weightsAndrews)

ts.plot(DATA8)
```