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Essays in interest rates, exchange rates and savings

Essays in Interest rates, Exchange rates and Savings

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Abstract

This thesis studies the behaviour of interest rates in government bonds markets, foreign exchange rates and national savings. There are three main chapters in the thesis. The first chapter consists of a comparative study of government securities and risk. It generates monthly interest rate risk premium data and examines their determinants. The results show that the risk premia are time varying and also vary considerably across sample countries. In particular, countries with better financial development and higher income generally have lower risk premia of government assets. Additionally, the risk premia are significantly affected by macroeconomic circumstances, especially economic growth and the real effective exchange rate.

The second chapter revisits the empirical literature testing the efficiency of forward markets for foreign exchange. According to the forward rate unbiasedness hypothesis, the forward premium should be an unbiased estimate of the subsequent exchange rate change. However, not only is this hypothesis rejected by standard regressions of the spot return on the forward premium but there are also puzzling negative coefficients from these regressions, which is referred to as the forward premium anomaly. This chapter addresses the forward premium anomaly first by examining statistical artifacts of the data; and secondly by considering the presence of a foreign exchange risk premium. This chapter finds the forward premium series to be fractionally integrated, which contributes to the forward premium anomaly. Structural breaks in the forward premium series are also found to increase the persistency of the series. Lastly, this chapter suggests a new methodology in estimating the exchange rate risk premium and finds that the foreign exchange risk premium is non-trivial.

The last main chapter models and simulates individuals' savings behaviour. The study extends the Diamond (1965) Overlapping Generation Model (OLG) to model more explicitly the difference between the accumulation and decumulation phases of private pensions. A similar question to that of Diamond is analysed, namely the effect of government debt, but changes the logic considerably by noting that government debt has different characteristics from private debt (i.e. equity) and that the two may be less close substitutes than is usually assumed. Rather than crowding out private investment, government bonds provide an important part of the funded system. The results show that it is socially optimal to rely entirely upon a funded pension system.

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I would also like to thank my family for their encouragement and support; my fiance, Sanha Hemvanich, for his care and morale support; my beloved friends who made everyday in Bristol meaningful; and my PhD colleagues for making my study an enjoyable one. Help from several support staff in the Department of Economics is also appreciated. I dedicate this dissertation to my parents.

Declaration

I declare that the work in this dissertation was carried out in accordance with the Regulations of the University of Bristol. The work is original, except where indicated by special reference in the text, and no part of the dissertation has been submitted for any other academic award. Any views expressed in the dissertation are those of the author.

SIGNED: [illegible] DATE: 15 June 2008.

Contents

1 This chapter is joint work with Edmund Cannon.

List of Abbreviation

Abbreviations	Definitions
ACF	Autocorrelation Functions
ARCH	Autoregressive Conditionally Heteroskedastic
ARCH-M	Autoregressive Conditionally Heteroskedastic in Mean
ARFIMA	Autoregressive Fractionally Integrated Moving Average
ARIMA	Autoregressive Integrated Moving Average
ARMA	Autoregressive Moving Average
CIP	Covered Interest Parity condition
CRRA	Constant Relative Risk Aversion
EMU	European Monetary Union
ERM	Exchange Rate Mechanism
FIGARCH	Fractionally Integrated Generalised Autoregressive Conditionally Heteroskedastic
FRU	Forward Rate Unbiasedness
G7	Group of Seven
GARCH	Generalised Autoregressive Conditionally Heteroskedastic
GDP	Gross Domestic Product
GMM	General Method of Moment
GNP	Gross National Product
GPH	Geweke and Potter-Hudak (1983)
HAC	Heteroskedasticity and Autocorrelation Consistent estimator
IMF	International Monetary Fund
IV	Instrumental Variable
IV-FE	Instrumental Variable - Fixed Effect estimation
JIT	Jensen Inequality terms
LP	Log periodogram regression
LSDV	Least Square Dummy Variable, Fixed Effect Estimation
LSDVC	Least Square Dummy Variable Corrected
ML	Maximum Likelihood
MLP	Modified log periodogram regression
OECD	Organisation for Economic Co-operation and Development
OLG	Overlapping generations model
OLS	Ordinary Least Square
PAYG	Pay As You Go pension scheme
RBD	Residual based diagnostic
UIP	Uncovered Interest Parity condition

Chapter 1
Introduction

1.1 Background:

Three major areas of interest for financial economics and macro-economics are interest rates, exchange rates and national savings.

The first main chapter of the thesis is called "Determinants of the time varying risk premia". It studies the behaviour of the risk premia of short term government assets (treasury bills). First, I generate monthly risk premia data using zero coupon government treasury bills. The risk premia series in this study are proxied by the time series volatility of the excess holding yields for short- and long-term treasury bills. The risk premia measure is based on the ARCH-in-Mean (ARCH-M) model introduced by Engle, Lilien and Robins (1987). The estimated risk premia are found to be time varying and also vary considerably across countries. Second, this chapter examines the macroeconomic and political determinants of government asset risk premia by using cross section and dynamic panel regression analyses. The results show that the risk premia are significantly affected by macroeconomic circumstances, especially economic growth and the real effective exchange rate. The results are robust across the majority of countries in our study.

The next chapter, entitled "Fractional integration and the forward premium puzzle", revisits the empirical literature testing the efficiency of forward markets for foreign exchange. In previous work, forward rate unbiasedness has been rejected. Generally, the literature has found that the future exchange rate change is negatively related to the forward premium. This chapter assumes rational expectations and attempts to explain the forward premium anomaly by firstly the presence of foreign exchange risk premia and secondly by examining statistical artifacts of the data.

The results show that the forward premium series are fractionally integrated while the return on the spot exchange rate is stationary. This yields an unbalanced test regression of Uncovered Interest Parity, which causes the violation of the hypothesis. However, there is also evidence that structural breaks cause spurious long memory in the forward premium series for some of the sample currencies analysed in this study. Nonetheless, the finding of a fractionally integrated forward premium series is still robust.

This chapter also pays attention to the relationship of the exchange rate risk premia and the test of forward market efficiency. Previous studies suggested that the time varying foreign exchange risk premium is extremely small at the monthly level. However, this chapter argues that the time varying foreign exchange risk premium is significant at the daily level. This chapter also suggests a new methodology to estimate the foreign exchange risk premium by modelling the time series volatility in the conditional variance of the forward rate forecast error series. The results show that such series appear to have long memory in the conditional variance for all of the sample currencies in this study.

The last main chapter of this thesis is entitled "Funded and PAYG Pensions when Annuities are backed by Bonds". This chapter studies optimal proportion of the private funded pensions and the state pay-as-you-go pension. Previous literature (such as Samuelson, 1958 and Aaron, 1966) suggested that funded pensions provides better incentives to save and results in higher capital. On the other hand, the state pension (funded by government debt and income tax) reduces the utility of individuals in the long run since it creates a further reduction in the productive capital stock arising from the substitution of government debt for physical capital in individual portfolios. However, the contribution of this chapter is motivated by the observation that government debt is both a complement and a substitute to physical capital.

The analysis in this chapter extends the Diamond (1965) Overlapping Generation Model (OLG) model to model more explicitly the difference between the accumulation and decumulation phases of private pensions. We observe that ownership of physical capital is mediated largely through equity, which provides a high risk high return financial asset, whereas government debt is mediated through bonds, which can insure against long term risk. So although government debt and physical capital compete for funds (suggesting they are substitutes), the financial assets which result are quite different (suggesting complementary).

The results show that the optimal government policy is to have people funded (via annuity market) for retirement consumption rather than public pension transfer. People are worse off in the long run after a sudden fall in national debts, while a baby-boom generates a small improvement to the welfare in the long run. Lastly, these two shocks do not only affect the current generations but also hurts the future generations.

1.2 Literature Review

Each separate main chapter contains review material of the principal literature. The purpose of this section is to provide an overview of the related literature

1.2.1 The interest rate risk premia and their determinants:

There is an abundance of work on the term structure of interest rates but this focuses mainly on the validity of the expectations hypothesis. Empirical evidence of time varying risk premia in government asset returns is frequently interpreted as evidence against the expectations hypothesis. However, we need a better understanding of the determinants of the term premia. This will in turn give a clearer explanation for the rejection of expectations hypothesis. This is the purpose of the research in chapter.

The literature has not yet fully identified the determinants of risk premia in government assets. There are a few works that attempt to relate the term structure to movements in macroeconomic variables such as Wu (2002), Hordahl, Tristani, and Vestin (2006), and Rudebusch and Wu (2003). These papers explain how macroeconomic factors (inflation, output gaps and the short term policy interest rates) drive movements in the term structure of interest rates and how they affect the behaviour of the risk premia embedded in observed yields. However, these works ignore the role of time-varying risk premia which is an important component in explaining movements in yields over time.

Previous literature addressing cross country comparison of risk premia includes Alesina, DeBroek, Prati, Tabellini, Obstfeld, and Rebelo, 1992; Lemmen and Goodhart , 1999; Giavannini and Piga, 1994; Favero, Giavazzi and Spaventa, 1997; IMF, 1997; Mc Cauley, 1996; Eijfinger, Huizinga and Lemmen, 1998. However, these works examine risk premia based on the credit risk of government debt and ignore the time variation in risk premium within countries obviously.

Alesina, DeBroek, Prati, Tabellini, Obstfeld and Rebelo (1992) study the default risk on government debt in OECD countries. The risk is derived by comparing the return from holding government debt with the return from holding corporate debt denominated in the same currency. However, the drawback is that the measure of default risk tends to be sensitive to significant changes in private risk. Additionally, they consider a variety of different maturities for both public and private yields.

However, differences in the maturity between the public and private yields may lead to inaccurate measurement of the magnitude of government default risk.

Lemmen and Goodhart (1999) find the determinants of credit risk in the European government bond markets using fixed effects estimation. The risk specified in their work is the default risk (credit risk) proxy by the spread of 10-year benchmark government bond yields over the corresponding swap yield of the same 10-year maturity denominated in the same currency. Although the risk specified represents an improvement to the one used by Alesina, DeBroek, Prati, Tabellini, Obstfeld and Rebelo (1992), the risk measure still has several problems. Firstly, the risk premia may not be a good proxy for country risk if the government and private bonds interact with each other. According to Lemmen and Goodhart (1999), uncertainty about government debt servicing will affect private sector risks particularly when bank or other financial institutions hold a large proportion of their assets in government debt, leading private and public risks to move in a lockstep fashion. Secondly, Lemmen and Goodhart (1999) consider government bond redemption yield data. The use of redemption yields introduces coupon reinvestment risk in the default risk measure. The redemption yield depends on the coupon size. To solve this problem, we use zero coupon yields data to calculate the risk premia.

Other works employ the credit risk of sovereign debt, which can be assessed by comparing yields on domestic government bonds with high quality private risk as represented by interest rate swap yields (see Giavannini and Piga, 1994; Favero, Giavazzi and Spaventa, 1997; IMF, 1997; Mc Cauley, 1996; Eijfinger, Huizinga and Lemmen, 1998). The risk premia measures in these papers cannot distinguish between credit risk and liquidity risk. The measure of the risk premia in these studies requires the assumption that variations in liquidity are negligible. However, liquidity effects may play a central role in government assets return. In the first main chapter of this thesis, the liquidity effect is automatically taken into account in the risk premia estimation.

From previous literature, there is a room for improvement in the estimation of the interest rate risk premia. The generated risk premia from the first main chapter of this thesis is suitable not only for cross country comparison but also capable to explain the time variation factor of risk premia within countries.

1.2.2 The forward premium anomaly and related literatures:

According to the forward rate unbiasedness hypothesis, the forward premium should be an unbiased estimate of the subsequent exchange rate change. However, previous literature (such as Frankel, 1980; Fama, 1984; Bekaert and Hodrick, 1993; and others) finds that not only is this hypothesis rejected by standard regressions of the spot return on the forward premium but there are also puzzling negative coefficients from these regressions. This is referred to as the forward premium anomaly. Froot and Thaler (1990) find that the average coefficient in the regression across some 75 published estimates is -0.88. Of these estimates, a few are positive, but none of them have the coefficient statistically greater than or equal to unity. There are two main explanations for the forward premium anomaly, which are the statistical artifacts of the data and the existence of the foreign exchange risk premium.

The statistical artifacts of the data view explains the forward premium anomaly based on 1) the long memory behaviour of the forward premium; and 2) the existence of structural breaks in the forward premium. Baillie and Bollerslev (2000) proposed that the anomaly is caused by a very persistent autocorrelation in the forward premium. Maynard and Phillips (2001) also found evidence of fractionally integrated behaviour in the forward premium. However, the returns on the spot exchange rate are widely accepted to be stationary (Cornell, 1977; Meese and Singleton, 1982; Corbae and Ouliaris, 1986, Baillie and Bollerslev, 1989; Baillie and Bollerslev, 1993; Maynard and Phillips, 2001; Choi and Zivot, 2005). This suggests that the traditional asymptotic regression for unbiasedness may not be suitable due to the difference in persistence between the two series mentioned above. Recently, Choi and Zivot (2005) pointed out the importance of structural breaks and confirmed that both explanations are important.

The foreign exchange risk premium explanation states that the forward rates are biased predictors of actual exchange rate movements because there exists a risk premium on one country's currency relative to another. Fama (1984) originally attempted to explain the anomaly by arguing that the time varying risk premium term can lead to bias and inconsistency in the OLS estimates of coefficient in the unbiasedness regression. There are two approaches used within the literature in interpreting the risk premia. The first approach estimate the risk premium by the Capital Asset Pricing Model (CAPM), the second approach specifies the statistical model of the risk premium.

Literature using the CAPM approach to estimate the risk premium include Hansen and Hodrick, 1980; Hodrick and Srivastava, 1984, 1986; Giovannini and Jorion, 1987; Bekaert and Hodrick, 1992 and Bekaert, 1994. In this approach, the risk premium is generally defined by the sum of the conditional variance of the spot exchange rates, the covariance of the spot exchange rates and inflation, and the covariance of the intertemporal marginal rate of substitution in consumption and exchange rate. However, this research in general finds that the model has limited success in explaining the risk premium. A key problem is that the risk premium model is not robust across different datasets and time periods (Lewis, 1995). There is also a limitation of the data, in particular, a small covariation between consumption and exchange rates since the data of the former is fairly smooth (Kaminsky and Peruga, 1990;Sarno and Taylor, 2002; Baillie and Bollerslev 1989, 1990; and Bekaert and Hodrick, 1993)), and a small covariation between exchange rate and inflation (Lewis, 1995; Engel, 1984; and Cumby, 1988) and the foreign exchange rate data show little evidence of conditional heteroskedasticity (Baillie and Bollerslev 1989, 1990; and Bekaert and Hodrick, 1993).

Literature that specifies "statistical" models of the risk premium include Domowitz and Hakkio (1985) and Baillie and Bollerslev (1990). This approach tests for certain patterns in or across excess exchange rate returns. Domowitz and Hakkio (1985) give an explanation of the risk premia through the conditional variance of the forward exchange rate forecast error using monthly exchange rate data. The forward rate forecast error is defined as the difference between the future spot rate and the forward rate. The estimation of the variance of the forward rate forecast error is by the ARCH-in-mean Model by Engle, Lilien and Robins (1987). However, the estimated risk premium is found to be insignificant in the unbiasedness regression. Baillie and Bollerslev (1990) use an alternative approach to estimate the risk premium, by specifying it as a function of the conditional covariance matrix for all currencies. They examine the conditional variance and covariance of the forward rate forecast error using a multivariate GARCH model in the weekly data. Unfortunately, the test for the presence of the time varying risk premium gains little support which confirms the finding in Domowitz and Hakkio (1985).

The "statistical" models of the risk premium have not been completely successful and there is room for new research. The test for the presence of a time varying risk premium gains little support when using monthly data in Domowitz and Hakkio (1985) and weekly data in Bailie and Bollerslev (1990). In contrast, my thesis chapter argues that there is a possibility that using higher frequency data and a longer time span would give more information to the model. Moreover, in estimating the conditional variance of the forecast error, one can find a model that fits the data better than the ARCH-in-mean model and the multivariate GARCH model. Thus, my thesis chapter employs the second approach in estimating the risk premium.

1.2.3 Pensions and related literature

The first and foremost research in pension modelling the difference between funded and unfunded schemes is the overlapping generations framework is by Samuelson (1958) and Aaron (1966). The research shows that with a Pay-As-You-Go (PAYG) scheme, it is possible in principle for every generation to receive more in pensions than it paid in contributions, provided that the rate of growth of total real earnings exceeds the interest rate indefinitely. However, this argument does not appear to be currently relevant. The old age dependency ratio in nearly all developed economies is substantially higher than it used to be. This has generated a large literature on the reform of the pension systems (such as Feldstein, 1996; Feldstein and Samwick, 1998; Mitchell and Zeldes, 1996; Disney, 1996; Kotlikoff, 1996; Huang, Imrohoroglu and Sargent, 1997; Miles and Timmerman, 1999; Sinn, 1999; and Campbell and Feldstein, 2001). These works mainly concentrate on the appropriate proportion of unfunded state pensions and private pension provision.

Among these, many researchers have expressed different opinions on the problem whether the pay-as-you-go state pension system should be replaced with a funded system. The first group suggested efficiency gains from a transition to a funded system (Diamond, 1965; Feldstein, 1977, 1995; Kotlikoff, Smetters, and Walliser, 1998; Feldstein and Samwickm 1998; Borsch-Supan, 1998; Homburg, 1990, 1997; and Mile, 1999). The second group argued that a Pareto improving transition to a funded system is not possible (Breyer, 1989, Fenge, 1995; Brunner, 1996; Sinn, 1997, 1998; and Geanakopolos, Mitchell, and Zeldes, 1998).

The first group argued that the pay-as-you-go schemes induce important labour market distortions due to the income tax. Moreover, such schemes diminish the capital stock because they are a special form of government debt (Feldstein, 1977). On the other hand, funded pensions are dynamically more efficient, since they encourage saving, which in turn raises the capital-labour ratio and income per head. As a result, phasing out state pensions completely generates higher saving, a higher capital stock and lower real rates of return.

The second group argued that this comparison of rates of return does not imply that the abolition of the state pay-as-you-go pension system would lead to an intergenerational Pareto improvement since it is impossible to compensate the losers of the transition (namely, the first generation which does not receive the state pension) without making at least one of the later generations strictly worse off. In addition to this, the argument is on income redistribution grounds, whilst the public pension systems often redistribute income from the rich to the poor (see Barr and Diamond, 2006).

Within the work of Diamond, 1965; Feldstein, 1977; and others, the pay-as-you-go schemes diminishes the capital stock because they rely on a special form of government debt. The logic of the models is that government debt will reduce welfare because it will divert savings away from productive capital, which is a typical example of crowding out.

The contribution of this chapter is motivated by the observation that government debt is both a complement and a substitute to physical capital. We observe that ownership of physical capital is mediated largely through equity, which provides a high risk high return financial assets, whereas government debt is mediated through bonds, which can insure against long term risk. So although government debt and physical capital compete for funds (suggesting they are substitutes), the financial assets which results are quite different (suggesting complementary).

Chapter 2
Determinants of the time varying risk premia

2.1 Introduction

This paper studies the behaviour of the risk premia of short term government assets (treasury bills). The paper makes 2 contributions to the literature. Firstly, we generate monthly risk premia data using zero coupon government treasury bills for 43 countries over the period of 1994-2006. The risk premia measure is based on the ARCH-in-Mean (ARCH-M) model introduced by Engle, Lilien and Robins (1987). The estimation of the risk premia in this paper can perform the same function as the agencies' credit ratings as it allows us to extract the market perceptions of the risk in holding government assets. Moreover, the risk premia data generated in this study are somewhat more continuous and more time varying measure of risk in holding government asset than the risk indices based on credit ratings. We find that the risk premia are time varying and also vary considerably across countries. The second contribution of this paper is that we examine the macroeconomic and political determinants of the risk premia by using cross section and dynamic panel regression analyses. The results show that the risk premia are significantly affected by macroeconomic circumstances, especially economic growth and the real effective exchange rate. The results are robust across the majority of countries in our study.

The risk premia series in this study are proxied by the time series volatility of the excess holding yields for short- and long-term treasury bills. Thus the risk premia in this study correspond to the term premia in the theory of the term structure of interest rates, I will use these two terms interchangeably. The process used to construct risk premia data follows the argument of Engle, Lilien and Robins (1987), and associates the mean of the excess returns on holding long-term comparing to short-term government bills to the volatility of the excess returns. It focuses on the fundamental trade-off between expected returns and their volatility. The theoretical appeal of this model is that it provides microeconomics foundations by measuring the response of risk averse economic agents to uncertainty using the time series data. Estimating the risk premia from the treasury bills data is relevant to previous studies which have documented that the treasury bills rates contain time varying term premia[2].

There is an abundance of work on the term structure of interest rate but this focuses mainly on the validity of the expectations hypothesis. Empirical evidence of time varying risk premia in government asset returns is frequently interpreted as evidence against the expectations hypothesis. However, we need a better understanding of the determinants of the term premia. This will in turn give clearer explanation for the rejection of expectations hypothesis.

The literature has not yet fully identified the determinants of risk premia in government assets. There are a few works that attempt to relate the term structure to movements in macroeconomic variables such as Wu (2002), Hordahl, Tristani, and Vestin (2006), and Rudebusch and Wu (2003). However, these works ignore the role of time-varying risk pre-

[2] Many papers provide evidence that the risk (term) premium in term structure of interest rate varies over time instead of being constant. Parts of this evidence consist of repeated rejection of the expectation hypothesis [Shiller, 1979; Startz, 1982; Shiller, Campbell and Schoenholtz, 1983; Fama, 1984a; Mankiw, 1986; Mankiw and Miron, 1986; Campbell, 1987; Engel, Lilien and Robins, 1987; Fama and Bliss, 1987; Shiller and McCulloch, 1987; Hardouvelis, 1988; Froot, 1989; Simon, 1989; Campbell and Shiller, 1991 and others]

mia which is an important component in explaining movements in yields over time. Ang and Piazzesi (2003) suggest that macroeconomic factors (inflation and economic growth factors) have an important explanatory role for the dynamics of the yield curve, and that including these variables in a term structure model can improve its one-step ahead forecasting performance[3]. They find that macro factors explain up to 85 percent of the observed variation in bond yields. Hordahl, Tristani and Vestin (2006) employ macroeconomic variables to indirectly explain the risk premia. Their paper explains how macroeconomic factors (inflation, output gaps and the short term policy interest rates) drive movements in the term structure of interest rates and how they affect the behaviour of the risk premia embedded in observed yields. Their paper utilises a dynamic term structure model based on macroeconomic factors, which allows for an explicit feedback from the short term policy rates to macroeconomic outcomes. At the same time, the explicit modelling of risk premia captures dynamics of the entire term structures. They conclude that the dynamics of risk premia can ultimately be attributed to underlying macroeconomic dynamics[4].

This paper can be divided into two main parts. In the first part, we generate measures of the risk premia of government securities for 43 countries over the period 1994-2006. In the latter part, we find the determinant of the risk premia using the data generated from the fist part. In examining the determinants of risk premia, we carefully deal with the characteristics of small sample sizes in our study. In the cross section regression analysis, we use the small sample version of the heteroskedasticity consistent covariance matrix

[3] Their two stage estimation methods is based on the asssumption that short term interest rates do not affect macroeconomic variables.

[4] Anyhow, the paper did not include the foreign variables or exchange rate, which will provide fully satisfactory account of macroeconomic dynamics in the country of study e.g. Germany.

estimates (HC3) suggested by MacKinnon and White (1985) to improve the performance of the analysis in small samples. In the dynamic panel regression, the determinants of risk premia are estimated using a Least Squares Dummy Variable Corrected (LSDVC) procedure proposed by Bruno (2005a, b). This estimator is a recently proposed panel data technique that is suitable for small samples in unbalanced panels.

The result from the cross section analysis can be briefly summarised as follows. On average, over the period 1994-2006, the risk premia for holding government assets required by risk averse investors is positively associated with the level of inflation and the budget deficit as a percentage of GDP (both variables are significant at the 1 percent level), and is negatively affected by the country's economic growth (significant at the 5 percent level). Additionally, low income countries are estimated to have risk premia about 19 percent higher than in the high income countries outside the Eurozone, holding other variables constant. In the high income countries outside the Eurozone, the risk premia on holding government assets is predicted to be 10 percent more than those in Eurozone.

Using panel data analysis, we found that economic growth and the volatility of real effective exchange rates are the main determinants of the risk premia in the full sample. Risk averse investors require lower risk premia for holding government assets in countries with good economic performance i.e. high economic growth and a stable external price competitive position i.e. low volatility of real effective exchange rate. If we split the sample by income group, economic growth remains the main determinant of the risk premia. However, we also find that the real effective exchange rate plays an interesting and important role: in high income countries, devaluations bring favorable results to the economy as consistent with the Mundell-Fleming model. There is a better price competitiveness which in turn reduces the country risk premia. The opposite relationship is found in the sample of low income countries. One possible mechanism explaining this may be that in financial vulnerable countries, weaker local currency can exacerbate the external debt service difficulties. Devaluations therefore raise the country risk premia. This corresponds to the results from Cespedes, Chang and Velasco (2004).

The paper is organised as follows. The following section first outlines the definition of the risk premia and the departure from the existing literature. We then present a theoretical model for the ARCH-M methodology of time varying risk premia following Engle, Lilien and Robins (1987). Next, we construct measures of the time varying risk premia for 43 countries over the period of 1994 to 2006. The results show that, in general, term premia exist, are time varying and different between countries. After deriving measures of the term premia, we then ask what factors determine the movements in risk premia and what makes it vary across countries and through time[5]. Using cross sectional and panel data regression analysis, our main aim is to establish how macroeconomic, financial and political conditions determine the differences in risk premia. The final section concludes.

2.2 Risk premia and related literatures

In this section, we first discuss the concept of the risk premia. We explain the rationale for estimating the risk premia from the term structure of interest rates. We also explain why

[5] Assuming that investors form their expectations concerning movements in the interest rate using all available information, perfect capital mobility no longer implies that interest rates on the same asset class are equal across countries. The widely used measure of risk in finance is the volatility, however, in an arbitrage free economy; the risk perceived corresponds to the relevant information available to an investor as well.

the risk premia estimated here provides an alternative to those used in previous literatures. Lastly, we briefly discuss the rationale for using the ARCH-M model to estimate the risk premia.

The risk premium is the differential in the expected rate of return on a risky asset as compared with a safe asset. The risk associated with holding government assets can be classified into 2 aspects; the pure time factor of the risk, and the risk of default.

The pure time factor of risk refers to the term or maturity risk, and this is directly related to the term structure of interest rates in monetary economics through the expectation hypothesis. This hypothesis states that the interest rate on the long term asset must equal the average of the expected future interest rate on short term assets plus the term premium (Campbell and Schiller, 1991). Hence, this term premium is simply an increment of return required to induce investors to hold longer term securities. The longer maturities entail greater risks for the investor. With longer maturities, more catastrophic events might occur that may impact the investment, hence the need for a risk (term) premium[6]. This pure time factor of risk premia series can be directly estimated by the ARCH-M model by Engle, Lilien and Robins (1987).

The risk of default refers to the likelihood of the loan not being repaid. Although it is generally recognised that securities issued by governments are relatively safer than other types of assets, the risk associated with holding them as perceived by international in-

[6] This explanation depends on the distant future being more uncertain than the near future, and risk of future adverse events (such as default and higher short-term interest rates) being higher than the chance of future positive events (such as lower short-term interest rates).

vestors, varies according to the economic and political conditions of the country of issuer[7]. This risk is thus country specific and is regarded as a country's credit-risk.

Previous literature addressing cross country comparison of risk premia includes Alesina, DeBroek, Prati, Tabellini, Obstfeld, and Rebelo, 1992; Lemmen and Goodhart , 1999; Giavannini and Piga, 1994; Favero, Giavazzi and Spaventa, 1997; IMF, 1997; Mc Cauley, 1996; Eijfinger, Huizinga and Lemmen, 1998 . However, these works examine risk premia based on the credit risk of government debt. I believe that the measure of the risk premia in my study is a somewhat better measure of risk in holding government assets than the government defaulted risk constructed by previous literatures in several ways as follows.

Alesina, DeBroek, Prati, Tabellini, Obstfeld and Rebelo (1992) study the default risk on government debt in OECD countries. The risk is derived by comparing the return from holding government debt with the return from holding corporate debt denominated in the same currency. However, the drawback is that the measure of default risk tends to be sensitive to significant changes in private risk. Additionally, they consider a variety of different maturities for both public and private yields. However, differences in the maturity between the public and private yields may lead to inaccurate measurement of the magnitude of government default risk.

Lemmen and Goodhart (1999) find the determinants of credit risk in the European government bond markets using fixed effects estimation. The risk specified in their work is the default risk (credit risk) proxy by the spread of 10-year benchmark government bond

[7] This concept is quite similar to the asset market and portfolio balance approach in international economics which states that domestic and foreign bonds are not perfect substitutes and foreign bonds carry some additional risk with respect to domestic bonds. However, in some countries with less financial stability, the domestic bonds may be relatively more risky than the government bonds in developed countries.

yields over the corresponding swap yield of the same 10-year maturity denominated in the same currency. Although the risk specified in Lemmen and Goodhart (1999) offer an improvement to the one used by Alesina, DeBroek, Prati, Tabellini, Obstfeld and Rebelo (1992), the risk measure still has several problems. Firstly, the risk premia may not be a good proxy for country risk if the government private bonds interact with each other. According to Lemmen and Goodhart (1999), uncertainty about government debt servicing will affect private sector risks particularly when bank or other financial institutions hold large proportion of their assets in government debt, leading private and public risks to move in a lockstep fashion. Secondly, Lemmen and Goodhart (1999) consider government bond redemption yield data. The use of redemption yields introduces coupon reinvestment risks in the default risk measure. The redemption yield depends on the coupon size. To solve this problem, I use zero coupon yields data to calculate the risk premia.

Other works employ the credit risk of sovereign debt, which can be assessed by comparing yields on domestic government bonds with high quality private risk represented by interest rate swap yields (see Giavannini and Piga, 1994; Favero, Giavazzi and Spaventa, 1997; IMF, 1997; Mc Cauley, 1996; Eijfinger, Huizinga and Lemmen, 1998). The risk premia measures in these literatures cannot distinguish between credit risk and liquidity risk. The measure of the risk premia in these studies requires the assumption that variations in liquidity are negligible. However, liquidity effects may play a central role in government assets return. In my study, the liquidity effect is automatically taken into account in the risk premia estimation.

2.3 Methodology: Measuring risk premia

This section describes the construction of our risk premia data. Section 2.3.1 describes the source of data for calculating the excess holding yield for 6 month treasury bills over 3 month treasury bills in 43 countries over the sample period of 1994:12 to 2006:2. Due to the limited availability of the zero coupon yield data, there are 43 countries in our studies. These include both developed and developing countries. See table 2.1 for a list of countries, data definition and the period of observation. Section 2.3.2 presents the theoretical derivation of the time varying risk premia. Section 2.4.1 uses the calculated excess holding return to generate the risk premia data by applying the ARCH-M methodology. The formulation closely follows Engle, Lilien and Robins (1987). The risk premia is the dependent variables in the cross section regression and the panel data analysis in sections 2.5 and 2.6, respectively.

2.3.1 The data

The term structure data available in each country start in different years and was collected from Bloomberg L.P. We use monthly observations[8] of the yield on short term assets, i.e. 3 month- and 6-month treasury bills, to calculate the excess holding yield. We use the volatility of excess holding yield to generate the risk premia.

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[8] This chapter closely follows Engle, Lilien and Robins (1987) in using the risk premia estimation technique and the choice of data frequency.

An additional reason is that the objective of this work is to explain the risk premia by the macro economic and institutional variables. The frequency of these data is naturally low (i.e. monthly, quarterly and annually). Thus, the monthly estimated of the risk premia is considered to be sufficient to fulfil the objective of the paper.

Instead of using the outstanding coupon treasury securities to calculate the excess holding yields, we use the calculated zero coupon instruments (fixed income) instead. This methodology is the same as Dotsey and Otrok (1995) and Harris (2004)[9]. The zero coupon instruments make a single payment at the maturity date. The size of the payment is the face value of the instruments. The advantage of the zero coupon bills is that it is free of liquidity and coupon effects that are common in outstanding treasury securities. The data is, therefore, of the same type as that is analysed in Campbell and Schiller (1991). This type of data is suitable for the analysis of term structure of interest rate since they have no effects from different coupons and compounding methods. To interpret a zero coupon yield index, the zero coupon yields are derived by stripping the par coupon curve. For example, the USD Government Agency (FMC84) Zero Coupon Yield is the zero coupon rate derived by stripping FMC[10] curve84. Most of the yield indices are denominated in national currencies except Turkey, Brazil and Uruguay. These 3 countries are denominated in US Dollars. The dataset we obtained here is daily reported, and the last trading day of the month is therefore chosen to serve as the end of month observation. Naturally, the 30th or 31st data of each month is used except for national Holidays or other non-trading days.

In estimating the term premia, we first define the excess holding yield. The formula for constructing the excess holding yield of 6-month over 3-month zero coupon treasury bills is analogous to Engle, Lilien and Robins (1987), Dotsey and Otrok (1995) and Harris

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9 Engle, Lilien and Robins (1987) uses the treasury bills rate to calculate the riks premia. However, the US treasury bills are zero coupon bills in that they do not pay interest prior to maturity; instead they are sold at a discount of the par value to create a positive yield to maturity.

10 FMC stands for Fair market value curve. The fair market value indices are derived from data points on Bloomberg's option free market curves. The yield at each maturity point represents the composite yield of securities around that maturity.

(2004). To set the notation, $y_t^{6,3}$ is defined as the excess holding return from holding a 6-month treasury bill compared to the return from holding consecutive 3-month treasury bills. The unit of time period in t stands for every 3 months. Thus, the time t is actually 3 months ahead of time $t+1$. R_t is the 6-month zero coupon treasury bill rate and r_t is the zero coupon yield of the treasury bill with maturity of 3 months. The excess holding period yield can therefore be calculated as:

$$y_t^{6,3} = \left[(1+R_t)^2 / (1+r_{t+1})\right] - (1+r_t)\,, \tag{2.1}$$

and following Engle, Lilien and Robins (1987), the linear approximation of equation (2.1) is used to calculate the excess holding yield as follow,

$$y_t^{6,3} = 2R_t - r_{t+1} - r_t.$$

Prior to generating the risk premia, it is useful to briefly explain the descriptive statistics of the excess holding yield in the different countries. This will help in visualising the expected characteristics of risk premia.

Table 2.3 illustrates descriptive statistics of the excess holding yield generated from equations (2.1). The number of observations is represented by number of months observed. The main findings are summarised as follows.

First, the mean of the excess holding yield for 6 months vs 3 months of our sample countries is positive in sign with value between 0 to 1 per cent per annum. Argentina and Uruguay are exceptions; the mean of the excess return is -3.95 and -0.99 percent per annum, respectively. This means that an investor would be better off if he keeps investing in a shorter term asset (3-month bill) for a year than buying a single 6-month treasury bill which gives less return in time $t+1$. Additionally, the excess holding yield of government securities in these 2 countries is extremely volatile with standard deviations[11] of 18.09 and 9.73 in Argentina and Uruguay[12], respectively. The data available for Argentina is from 1998:07 to 2002:03. Hence, it includes the time of economic crisis[13] in Argentina in 2001-2002. Over the period of study, the excess holding yield in Argentina hit its low at -71.716 per cent per annum in late 2001. This probably reflects the lack of confidence in economic prospects as investors do not want to take a risk in longer term assets. From Figure 2.1, thanks to the currency board, we can see a period of stability in the excess holding yield from late 1998 to late 2000. The volatility coincides with the time of crisis[14].

During year 2001-2002, Uruguay (see Figure 2.41) went through a similar economic and financial crisis[15] which developed mostly from external factors, not least the crisis in Argentina. As a result, there was considerable volatility in the excess holding period return during late-2001 to mid-2002. Although Uruguay's economy recovered in 2003 through improving its export performance and a more positive investment climate, the excess hold-

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11 The standard deviation of the return measures the average deviations of the return series from its mean, and is often used as a measure of risk. A large standard deviation implies that there have been large swings in the return series of assets.

12 The volatility of excess return is also increasing with maturity of longer term bonds.$s.d.(y_{12,3}) > s.d(y_{6,3})$. Appendix table A, the standard deviation of excess return in Argentina and Uruguay are 48.96 and 35.15, respectively. The excess holding yield in these 2 countries is also the most volatile among 43 countries in this study.

13 This entailed output falling by 20 percent over 3 years, high inflationary pressure, a severe devaluation of Argentine peso, government debt default, and lastly, a stagnant banking system.

14 Unfortunately, we cannot obtain the zero coupon yield data for Argentina after 2002:03.

15 The crisis started by the devaluation in Brazilian Reais in 1999 made Uruguayan exports relatively less competitive. In late 2000, the situation was exacerbated by the economic crisis in Argentina, which is Uruguay's major trading partner. Subsequently in mid 2002 there was a bank run due to massive withdrawals from Uruguayan banks. The bank run was unfortunately overcome by massive borrowing from international financial institutions which in turn, led to a serious debt sustainability problem.

ing yield swung wildly over the studied period. This reflects a persisting unstable financial system.

At the other extreme is the excess holding return of government securities in the Philippines (in table 2.3) which has a mean value of 1.91 percent per annum. It is also highly volatile with a standard deviation of 2.10. Figure 2.32 shows that the excess holding yield fluctuates wildly throughout the period of study. The excess holding yield is especially volatile with the sharp spikes in 1997-1998 and in late 2000 owing probably to the Asian financial crisis and oil price shocks[16], respectively.

Apart from the countries already mentioned, there have also been large swings in the excess return series in Brazil, with a standard deviation of 4.27 (see table 2.3). This is probably because Brazil was also affected by the South American economic crisis of 2002. Like other emerging market economies in general, Brazil was susceptible to contagion effects. In Brazil's case, it was contagion from Argentina's economic melt down causing a crisis of confidence among investors and lenders who were demanding higher interest rates. That put increasing pressure on the Brazilian economy to come up with those higher interest rates. Figure 2.5 shows that the excess holding yield series is again extremely volatile.

.Our second finding is that the less volatile excess holding yield series relate to economies with more stable financial systems and better economic development. For example, the

[16] Nevertheless, the Philippines was less severely affected by the Asian financial crisis of 1998 than its neighbours, aided in part by its high level of annual remittances from overseas workers, and no sustained run up in asset prices or foreign borrowing prior to the crisis. The impact from surging petroleum prices shock during late 2000 was more serious since the Philippines is an oil importer country. Overall, we find that the excess rates of return from holding Philippines' securities are highly erratic.

mean of the excess return is relatively lower but exhibits much less variation over time in the majority[17] of countries in the EU, compared to the rest of the world.

To test the robustness of the econometric results, the excess holding yield[18] of 12-month over 3-month zero coupon rate, $y_t^{12,3}$ is constructed. Following the same fashion as (2.1), the excess holding yield of 12-month versus 3-month Treasury bill can be generated as

$$y_t^{12,3} = \left[\frac{(1+R_t)^4}{(1+r_{t+3})(1+r_{t+2})(1+r_{t+1})} \right] - (1+r_t), \tag{2.2}$$

and the linear approximation is

$$y_t^{12,3} = 4R_t - r_{t+3} - r_{t+2} - r_{t+1} - r_t.$$

The descriptive statistics of the excess holding yield generated from equation (2.2) is presented in the Appendix Table A. The results show that mean and volatility of excess return are increasing with maturity of longer term bonds. There is higher uncertainty associating with the longer horizon, thus investors require more excess return. The excess return series are also more fluctuate with longer maturity spread. The standard deviation of excess return in Argentina and Uruguay are 48.96 and 35.15, respectively. The excess holding yield in these 2 countries is also the most volatile among 43 countries in this study.

The next section describes how the excess holding yield data can be used to construct risk premia.

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[17] Turkey, Poland and Hungary are exceptions. The excess return series of government assets in these three countries are relatively highly volatile with the standard deviation of 1.63, 1.48 and 1.40, respectively. The mean excess holding yield of these countries is in the range of 0.02 to 0.46 percent annually.

[18] It is defined as the excess returns from holding 12-month treasury bills for 3 months compared to the return from holding 3-month treasury bills. As mentioned earlier, the unit of time period in t stands for every 3 months. Thus, the time t is actually 3 months ahead of time $t + 1$.

2.3.2 The theoretical derivation of time varying risk premia

The estimation of the risk premia in my study bases on the model of the relation between risk and return in Engle, Lilien and Robins (1987). It is useful to discuss the theoretical derivation of the risk premia in their model as follows. The risk averse economic agents require compensation for holding risky assets. In this model, the risk is measured by the variance of return from holding assets and the compensation by the rise in the expectation of the return. The relation between the mean and the variance of returns which will insure that the asset is fully held in equilibrium will depend on the utility function of the agents and the supply condition of the assets.

The variance of the payoff of the risky assets is assumed to be able to change over time and consequently the price offered by risk averse agents will change over time. This equilibrium price determines the relation between mean and variance of excess returns from holding risky assets and therefore how the risk premium is related to the variance of returns.

The model assumes two assets economy (risky and safe assets). Assuming that r is the rate of return on the safe assets with the price of unity; q is a random total return on the risky assets with the price of p. The random return has mean θ and variance ϕ. Agent's wealth, W can be expressed as

$$W = ps + x,$$

where x is share of the safe assets and s is the share of the risky assets. The excess return per pound invest in shares of the risky asset is

$$y = (q/p) - r,$$

and the mean and variance of the excess return is given by

$$\begin{aligned} E(y) &= \mu = (\theta/p) - r, \\ V(y) &= \sigma^2 = \phi/p^2. \end{aligned}$$

Assuming constant absolute risk aversion, agents maximize expected utility of the end-of-period wealth. The expected utility of the agent is

$$EU \doteq 2E(qs + rx) - bV(qs + rx),$$

Agents maximise the utility by choosing

$$ps = \mu/(b\sigma^2).$$

The equilibrium can be written as

$$\mu = [-r + \sqrt{r^2 + 4bs\sigma^2\theta}]/2,$$

so that for a large variance, the mean is proportional to the standard devation. Equation above shows that there is a relation between observed means and variances of return which move them in the same direction but not proportionally.

The econometric model that best suits the risk return trade off is the ARCH-in-mean model. This model allows the conditional variance to affect the mean. In this way, changing conditional variances directly affect the expected return on a portfolio.

Engle, Lilien and Robins (1987) construct an ARCH-M model where the conditional variance of excess return determines the current risk premium. They then test their model by applying it to quarterly data on 3-month comparing to 6-month US Treasury bill rates from 1960:Q1 to 1984:Q2. The data are obtained from Salomon Brothers. The results imply that the risk premia vary systematically over time with agent's perceptions of underlying uncertainty.

In this section we generate measures of the term premium by estimating the ARCH-M model of excess holding yields for 6 month treasury bills over 3 month Treasury bills over the sample period of 1994:12 to 2006:2. The formulation closely follows Engle, Lilien and Robins (1987) and specifies that the contemporaneous expected conditional standard deviation of the error term be included in the mean equation of the excess holding yield. This specification follows from a micro-founded model with risk averse agents.

Firstly, the excess holding yield can be decomposed:

$$y_t = \mu_t + \varepsilon_t, \tag{2.3}$$

where (y_t) is the excess holding yield on 6 month zero coupon treasury bills. The non-stochastic term μ_t is the risk premium or the expected return that the risk averse investor would demand for holding the (riskier) long-term asset. In contrast, ε_t is the difference between the ex ante and ex post rate of return which is unforecastable in an efficient market. This means that the expected excess return from holding the longer-term asset is just equal to the risk premium $[E_{t-1}y_t = \mu_t]$.

The equation for risk premium is expressed as

$$\mu_t = \beta + \nu h_t, \qquad \nu > 0, \tag{2.4}$$

where h_t is the conditional standard deviation of the unforecastable shocks (ε_t) to the excess return on the long term asset. The term ν is the coefficient of relative risk aversion. The risk premium is assumed to be an increasing function of the conditional standard deviation of the unforecastable shocks (ε_t).

The conditional variance of the error term is h_t^2 and is a function of the information set available to investors.

$$h_t^2 = Var(\varepsilon_t | \text{ all available information}) \tag{2.5}$$

We note here that the model takes the mean as a linear function of the standard deviation (h_t) instead of the variance (h_t^2). This represents the assumption that changes in the variance are reflected less than proportionally in the mean. This specification has been widely used by other papers such as Domowitz and Hakko (1985), and Bollerslev, Engle and Wooldridge (1988).

Following Engle, Lilien and Robins (1987), it is assumed that the conditional variance is a weighted sum of past squared innovations, ε_{t-i}^2. This conditional variance follows an ARCH(P) process as follows:

$$h_t^2 = \alpha_0 + \alpha_1 \sum_{i=1}^{P} w_i \varepsilon_{t-i}^2. \tag{2.6}$$

Here, the variance of the error term depends on the intercept α_0 and the weighted average of past squared innovations, where w_i are the weighting parameters. Using monthly observations[19], the ARCH specification has 12 months lags[20] as we assume that information from the past year is useful for predicting the mean. We discount the older information

19 Engle, Lilien and Robins (1987) use quarterly formulation and use four lags.

20 The conditional variance follows a 12-order autoregressive process.

using a linearly declining weight scheme where $w_i = (13 - i)/78$, and $i = 1 - 12$. This declining weight scheme on lag structures also helps cope with the collinearity of the past square innovation terms, ε_{t-i}^2 [see Engle (1982)]. The equation can therefore be written as[21]

$$h_t^2 = \alpha_0 + \alpha_1 \left(\frac{12}{78} \varepsilon_{t-1}^2 + \frac{11}{78} \varepsilon_{t-2}^2 + .. + \frac{1}{78} \varepsilon_{t-12}^2 \right). \tag{2.7}$$

From the specification above (equation (2.3)-(2.5)), we can conclude that the conditional mean of the excess holding yield $E(y_t)$ depends on the conditional standard deviation of the unforecastable error term. Given that the variation of return measures riskiness, as $E_{t-1}y_t = \mu_t$, the risk premium is an increasing function of the conditional standard deviation of the returns.

The model specification above is used to generate risk premia[22] for our entire sample.

2.4 The variables

This section describes characteristics of the dependent variable, the risk premia and the explanatory variables.

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[21] We use monthly data and assume that the useful information for predicting the mean comes from the past year. Thus, in the conditional variance equation, we specify the declining weight on lag structure of past square innovations as in equation (2.7). However, Engle, Lilien, and Robins (1987) use quarterly data, the lag structure is instead characterised by $h_t^2 = \alpha_0 + \alpha_1 \left(\frac{4}{10} \varepsilon_{t-1}^2 + \frac{3}{10} \varepsilon_{t-2}^2 + \frac{2}{10} \varepsilon_{t-3}^2 + \frac{1}{10} \varepsilon_{t-4}^2 \right)$.

[22] Initially, I attempted to use a more general specification i.e. the GARCH-M model but (presumably because of small samples) results were quite noisy. There were problems of flat log-likelihood.

Generally the GARCH-M methods are very data- intensive. The ARCH-in-mean model with the declining weighted lag structure makes minimal demands of the data.

There's a trade-off between bias and efficiency. ARCH-M is better from the point of view of efficiency.

2.4.1 Dependent variable: Risk premia

This section describes the risk premia data which is the dependent variable in the cross section regression and the panel data analysis sections. The risk premia generated from volatility of excess holding yield is referred to as the ex-post term premia or liquidity premia since the excess holding yield represents the realised or expost premium from holding the long-term as compared to short-term securities.

In this section, we present the estimation of the risk premia for 43 countries[23] derived from the ARCH-M model in equations (2.3)-(2.6). The time series plots of estimated risk premia (together with the excess holding yield) are presented in figures 2.1 to 2.43. This is to illustrate their characteristics over time and across countries. Figure 2.44 gives broader view; it shows average risk premia over the period of 1994-2006 for all 43 countries.

It is useful to first consider the descriptive statistics of the estimated risk premia. Table 2.4 gives descriptive statistics of the risk premia of 6 month versus 3 month treasury bills across the sample period of 1994-2006. The risk premia appear to be highest in the

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[23] The data is country level and expressed in different currencies. One may argue that all of the excess holding yield and risk premia data should be in a common currency.

Ideally we would like to be able to separate out risk from return and risk from exchange rate volatility.

Although there is an issue of currency risk, the reality is rather problematic. Sovereign risk, currency risk, and interest rate risk are all highly correlated in many emerging countries (which are the majority of the sample countries considered in my thesis) and hence by gaining exposure to currency and interest rate risk one is still primarily taking one decision: a decision about the macro risk, just as one does with dollar-denominated debt.

Also, there is an issue of differences in maturity. The objective of this study is to measure the risk of government short term assets. Dollar-denominated debt has an average duration of 4 to 5 years, whereas local currency debt typically comes with interest rate duration of as little as 6 weeks. Moreover, dollar denominated debt is not widely available in developing countries.

One may argue that the local currency debt is not protected against sudden currency devaluation. However, a sudden devaluation is rarely an altogether unforeseen event, not least as the macro-economic factors that eventually lead to such drastic policy action builds up over a period of time and can be foreseen.

Philippines with average value of 1.98 percent annually. The risk premia in this country are also highly volatile with standard deviation of 0.58.

The risk premia are also highly volatile in the Latin American countries. The standard deviations of risk premia within this country group is in the range of 0.52 (Mexico) to 1.94 (Uruguay). On the other hand, the risk premia is relatively low in almost all European countries and the series are much less volatile. Excluding the Czech Republic, the average risk premia in the EU is in the range of 0.06 to 0.27 percent annually with standard deviations ranging from 0.10 to 0.23. Hence, there seems to be a relationship between economic as well as financial development and the risk premia.

Table 2.5 illustrates estimated coefficients from the ARCH-M estimations in equations (2.3)-(2.6) and their t-statistics for each of the 43 countries. The results can be summarised as follows. Firstly, there is an ARCH in mean relationship in 16 out of the 43 countries. The ARCH in mean relationship exists when the disturbances are heteroscedastic and the standard deviation of each observation is found to affect significantly the mean of that observation ($\alpha_1 \neq 0$ and $\nu \neq 0$). Additionally, the ARCH-M coefficient shows the correct sign ($\nu > 0$) in 34 out of the 41 countries; the risk premia is an increasing function of the conditional variance of returns[24].

Secondly, from the result of ARCH-M estimation in table 2.5, the conditional variance of ARCH (12) process is constant (i.e. $\alpha_1 = 0$ and thus $\nu = 0$) in China, Hungary,

[24] We can conduct the sign test to see whether there is a significant positive relationship between the risk premia and the conditional variance of return. The null hypothesis to be tested here is that there is no significant positive relationship between them. This hypothesis implies that both the positive and negative of ν in equation (2.4) are equally likely to be larger than the other. The results show zero p-value, which indicates that there is a strong positive relationship between the risk premia and the volatility [$Pr(k >= 34) = 0.000013$, $Pr(k <= 34) = 0.999998$, given $N=41, k=34$].

Indonesia, Korea, and Sri Lanka. The models show relatively flat and less volatile risk premia in Indonesia and Sri Lanka as are illustrated in figures 2.19 and 2.26, respectively. However, this does not imply that the risk premia of government assets in these countries are constant.

From the plots of the excess holding yields and estimated risk premia, the series of excess holding yield in these five countries are so noisy[25] that a systematic pattern of conditional heteroscedasticity does not hold given the quite short time-horizon under consideration. Thus, the conditional variance cannot be predicted by the past squared innovations as is suggested by Engle, Lilien and Robins (1987). We also find that the excess return series shows extreme volatility in Hungary (Figure 2.18) and Korea (Figure 2.25). The excess return swings wildly (with periods of both negative and positive excess return) without any systematic pattern in Indonesia and Sri Lanka. We cannot find information for the risk premia in China (Figure 2.8) and Slovak Republic (Figure 2.36). Again, this can be attributes to the short horizon of the observations in China and Slovak Republic (see table 2.1 for data sources, definitions and period of observations).

Lastly, for some countries, although the disturbance is heteroscedastic ($\alpha_1 \neq 0$), the data are not suggestive of an ARCH-M process i.e. the conditional standard deviation does not affect the mean. These countries are Norway, Sweden, Finland, Greece, Ireland, Turkey, South Africa, Argentina, Uruguay, Israel, Hong Kong and Hungary. From figures, there is no period of stability in the excess holding yield in any of these countries. Hence, the estimated risk premium is characterised by a relatively flat line. Good examples here are

[25] There is no variation in volatility of the excess holding yield. In other words, the series are constantly highly-volatile.

the excess holding return series in Sweden, South Africa, Israel and Ireland. In Sweden, the variance of the excess return is very stable as illustrated in Figure 2.37. The excess return series in South Africa (see Figure 2.43) fluctuates around the constant mean with a brief shock in 1998. In Ireland, the excess return is also volatile throughout (see Figure 2.21). The excess return in Israel is severely volatile around the constant mean (see Figure 2.22), the series distributed evenly between positive and negative values. This reflects a fairly unstable financial condition in this country. The risk premia is unsurprisingly high throughout. The problem therefore is that the time period under consideration is not long enough to observe both periods of stability and volatility e.g. Engle, Lilien and Robins (1987) look at the risk premia in USA during 1960-1985, wherein there is a period of stability followed by a volatile period. In order to find an ARCH-M process, the samples must contain both.

The excess holding yield in some other countries swings unsystematically and the past innovation does not contain information of the risk premia such as Turkey and Uruguay (see Figure 2.39 and 2.41). For Argentina (see Figure 2.1), there is too large shock in 2001 following period of stability, thus it mimics the predictive ability of the past innovations. Similarly, surrounded by periods of stability in excess holding yield, there is a large shock 1997-1998 in Hong Kong (see Figure 2.17) according to the Asian financial crisis.

In Finland, there is a negative time trend during late 20th century (see Figure 2.14). The mean and variance of the excess return are trending downward over the period of studies. On the other hand, there is no trend in the excess return in Greece and Norway, but the series is highly volatile that the risk premia is unpredictable.

As mentioned above, there is a significant ARCH in mean relationship in 26 countries ($\alpha_1 > 0$ and $\nu > 0$) in our study. The characteristics of the excess return are quite similar to the case of the USA during 1960-1985. From Engle, Lilien and Robins (1987)'s work, over the period of analysis there are a few interesting shocks in the US economy. There was an oil price shock in 1973 and 1980,. and the severe economic recessions in early 1982. During these periods, there was instability in financial and economic conditions, and people lost confidence in the assets markets. They were unable to forecast future returns and demanded more return from holding long-term assets. The volatility in the excess holding yield produces a higher risk premium in these periods. However, during the more stable period (1960-1967), we find that the risk premium is quite low and the long run value of the excess return is constant. In our work, the excess returns of 6 month treasury bills in France (Figure 2.15), Mexico (Figure 2.27), Malaysia (Figure 2.28), and New Zealand (Figure 2.31) follow the same pattern as the USA case in Engle, Lilien and Robins (1987): there is a period of tranquillity followed by a period of volatility. Brazil (Figure 2.5) also follows this pattern, but the volatility in the excess holding return is more drastic.

In Australia (Figure 2.2), Austria (Figure 2.3), Belgium (Figure 2.4), Czech Republic (Figure 2.10), the excess holding return is characterised by a negative time trend in short run (during late 20th century) and fluctuates around the constant mean in the long run. In Spain (Figure 2.13), the mean of excess return fluctuates up and down but the variances have large swing. There are time trends in the excess return and its variance is not constant throughout the period of studies with shocks in some periods in Germany (Figure 2.11), Switzerland (Figure 2.7), Canada (Figure 2.6), Colombia (Figure 2.9), Denmark (Figure 2.12), and India (Figure 2.20).

Figures 2.1-2.4 illustrates the average risk premia for all 43 countries over the period of 1994-2006. Figure 2.44 is the average risk premia for holding 6 month treasury bills (comparing to 3-month treasury bills). Figure 2.45 is the average risk premia for holding 12 month treasury bills (comparing to 3-month treasury bills). The purpose of figure 2.45 is to show that the difference in average risk premia across countries is consistent across maturities. We find that the risk premia is generally low in countries with better financial development and economics conditions. Government assets in Singapore, Australia and Japan are relatively less risky compared to other countries in the study. Government assets in the Philippines and all Latin American countries are considered to be more risky than the rest. We can also perform country comparison of the risk premia by considering the countries' income and economic development. Figure 2.46 presents risk premia (for holding 6 month treasury bills) comparisons by country group. We find that the risk premia of government assets in the non-OECD countries are relatively higher than the OECD country group. Figure 2.47 presents risk premia (for holding 6 month treasury bills) comparisons by country's income. The higher income countries have relatively safer government assets.

From a rough comparison of risk premia in 43 countries in this study, it is useful to extend an analysis by doing the cross section and panel data analysis. In sections 2.5 and 2.6, we examine whether the country's macroeconomic variables affect the risk premia.

2.4.2 Explanatory variables:

This section defines our control and explanatory variables used in the risk premia regression and discusses the expected sign of relationships with the risk premia. The macroeconomic variables we examine are economic growth ($GGDP$), the inflation rate ($INFL$), the real effective exchange rate ($REER$), and the volatility of real effective exchange rate, ($VREER$). The government fiscal variables pertain to government debt as a percentage of GDP ($DEBTGDP$) and the fiscal deficit as a percent of GDP ($DEFGDP$). The institutional variables consist of political constraints ($POLCON5$) and a political risk index ($ICRG$). These variables will be defined subsequently. The sources and definition of data are detailed in the data appendix in table 2.2.

A preliminary examination of these relationships is presented by using the bar charts of the explanatory variables and bivariate regression plots of the risk premia and explanatory variables. The bar charts of average value of each explanatory variables are presented in figures 2.48A-2.48I. The bivariate regression plots of the mean value of country's risk premia and explanatory variables are presented in figures 2.49A-2.49I.

The initial income level ($GDP94$) is our control variable for differences in initial development levels. The initial level of income is derived from the natural log of real gross domestic product per capita in year 1994 of each country. Initial income also is a proxy for the financial development. We might expect that there is less risk premia in holding government assets in countries with higher initial income and better financial development.

To control for heterogeneity among groups of economies, the regression analysis also include 3 groups of dummies, namely, EMU, $NEMU\_RICH$ and $POOR$. The dummy variable EMU stands for member countries of the European Monetary Union (EMU). We can refer to these countries as the Eurozone[26]. The second dummy variable, $NEMU\_RICH$ stands for other high income countries outside the Eurozone such as Denmark, Sweden, United Kingdom[27], USA, Canada, Japan, etc. Lastly, the dummy variable $POOR$ stands for the low to middle income countries such as Czech Republic, Slovak Republic, Hungary, Poland[28], Malaysia and Thailand, etc. The partitioning of these three groups is presented in the variable list in table 2.2. The definition of high/low income countries is obtained from the World Bank (2006). Using dummy variables also allow us to compare these 3 countries groups in the regression analysis. We discuss the reason for adding these three dummy variables in paragraphs below.

In our context, the inclusion of a Euro-zone dummy variable could be particularly relevant. The inflation and exchange rate risk associated with their government assets are closely aligned, given their common currency. We begin our analysis in 1994 which is the second stage of the implementation of the European Economic and Monetary Union (EMU)[29]. At this stage, economic convergence criteria among member countries had been

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[26] The Eurozone (also called Euro Area, Eurosystem or Euroland) is the subset of European Union member states which have adopted the euro, creating a currency union. The European Central Bank is responsible for the monetary policy within the eurozone.

[27] Denmark, Sweden and the UK are countries in the European Union that do not use the Euro.

[28] Czech republic, Slovak republic, Hungary and Poland joined the EMU on 1 May 2004. However, we do not include them in the group of Eurozone due to the early stage of membership and their income level.

[29] The first stage on the EMU (started on 1 July 1990) was to provide complete freedom for capital transactions, to improve economic convergence and to raise co-operation between central banks. There was also a free use of the European Currency Unit (a forerunner of the Euro currency) [European Central Bank, 2006].

The second stage (1 January 1994) is to strengthen co-ordination and economic convergence, to establish European Monetary Institute and to foster the process leading to the independence of the national central bank.

The last stage (1 Jan 1999) is to officially introduce Euro, to conduct the single monetary policy by the European System of Central Banks and entry into effect of the intra-EU exchange rate mechanism ($ERM\ II$) and into force of the Stability and Growth pact.

in process, although the official launch of the Euro was not until 1 January 1999. The EMU had a major impact on the European financial markets and the management of the economic policies. It was argued that the currency risk would be reduced following EMU. Government assets will instead be subjected just to the default risk.

> "Government assets among EMU member countries would mainly differ with respect to their credit worthiness, liquidity and tax treatment since intra-EMU exchange risk should be zero and inflation risk would be the same for every country in the Euro zone" [Lemmen and Goodhart, 1999].

Thus the principal source of relative risk in government debt markets in EMU is credit risk. The variation in interest rates and exchange rates, which we regard as the market risk is no longer involved at least in intra-EMU [IBCA, 1996]. We thus may expect no significant difference between the exchange rate and inflation risk among EMU member countries in our regression[30].

Basically, the initial income and these dummies are similarly functioning as control variables. They are employed to control for the financial development in general. The countries' initial incomes take the economic convergence into account when we measure the economic growth. The dummy variables help enhance the predictability of the model by taking into account the income difference and the inflation and exchange rate agreements[31]. An interesting research question is to examine whether EMU member countries have lower risk premia as a result of their exchange rate arrangement. This issue will be unfolded in cross section and panel data analysis section.

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30 We note that, however, the exchange rate risk still exists externally. The EMU member that trades externally has more risk than a member that does not i.e. it depends on extent of external trade.

31 Including income dummies tends to enhance the predictability of the model. Figures 2.47A and 2.47B show that countries with high incomes tend to have lower risk premia. We partially control for income by using dummy variables, $NEMU\_RICH$ and $POOR$.

Next, we discuss the characteristics of the explanatory variables. Countries with superior macroeconomic conditions, less exchange rate volatility, better fiscal conditions and more reliable political conditions, are expected to have lower risk premia. The superior macroeconomic conditions are characterised by low inflation rate and high output growth. The government will have a good fiscal condition if it has low debt and budget deficit in proportion with the gross domestic product. The political conditions are relatively more reliable if there is less political risk in the country and more stable government policy.

The percentage increase in gross domestic product (GDP) during one year defines economic growth, $GGDP$. Economic growth is defined as

$$GGDP_{it} = \frac{1}{4}\log(GDP_{it}/GDP_{it-4}).$$

The GDP data are available on a quarterly basis. $GGDP_{it}$ is the rate of change in the gross domestic product of country i at quarter t comparing to the same quarter last year, $t-4$. In the risk premia regression, we use the natural log of the average GDP growth of each particular country over 1994 to 2006 as an explanatory variable. We expect that a good economic performance comes along with stable financial market conditions. Alternatively slow economic growth might make the government asset in that country is more risky.

The GDP growth data suggests that there tends to be convergence across the economies in our sample. Figure 2.48B is bar chart of economic growth on average over 1994-2006. It suggests that lower income or developing countries (labelled by $POOR$) experience significantly higher growth rates than the higher income group (labelled by EMU and $NEMU\_RICH$). Comparing this figure with the bar chart of each country's initial level of income measured by the gross domestic product in 1994 (figure 2.48A), it suggests that the less advanced economies with lower value of initial income (and initial capital) have higher growth rate of income (and capital).

In the bivariate regression in figure 2.49A, there is a strongly negative relationship between initial level of income ($GDP94$) and the risk premia as suggested earlier. On the other hand, the bivariate regression in figure 2.49B shows a strongly positive relationship between the risk premia and economic growth. This relationship is somewhat contradict to our prior that the better economic growth leads to less risk premia required. Referring back to the chart of average risk premia over 1994-2006 (figures 2.44 and 2.45), the estimated risk premia for the developing countries are quite high. However, during this period the more backward economies have higher economic growth rate than developed countries as suggested by the convergence. This shows the importance of including the initial level of income variable to control for other factors determining the risk premia apart from the economic growth.

Inflation is also a potential determinant of risk premia. Investors protect themselves by requiring nominal interest rates that compensate them for expected inflation as well as for the risk that the inflation deviates from their expectations. The higher prices rise, the lower will be the purchasing power of the principal and nominal interest payments correspondingly must be higher. Not only do investors want to be compensated for the inflation they expect, they also want to be compensated for the risk that inflation could increase during the term of their loan. Inflation ($INFL$) is defined as the percentage change of consumer price index over the corresponding period of previous year. In the cross section regression, we use the natural log of the mean inflation for each country over 1994-2006. We expect a positive relationship between the inflation rate and the risk premia.

The data suggest that the attempt to stabilise inflation among member countries in EMU seems to be successful. This can be seen in the charts of average country's inflation over 1994-2006 in figure 2.48C. Within the Eurozone (excluding Greece[32]), the country's average inflation over the period varies between the minimum value of 1.88 percent[33] in France to maximum value of 4.89 percent in Italy (excluding Greece, the mean inflation of this group is 2.83 percent).

As mentioned earlier, the inflation levels of the Eurozone members tend not to be different from each others possibly due to the single currency convergence criteria. The higher income countries (both inside and outside Eurozone) have lower inflation rate than the lower income group. Comparing inflation level between countries in EMU and $NEMU\_RICH$, the difference between these 2 groups is not obvious[34]. However, there is slightly higher variation in inflation rates in the latter group. The developing countries group ($POOR$) has highest levels of inflation and the variation of inflation rates is quite substantial.

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32 The average inflation over 1994-2006 of Greece is 8.21 percents which is substaintially higher than the rest of countries in the Eurozone . This is partly because Greece if the last country that join this group. Greece was qualified as an EMU memeber state in 2000 and was admitted on 1 January 2001.

33 In the cross section regression, we use the natural log of this value instead.

34 Additionally, we find that the mean inflation in the UK, Denmark and Sweden are not very much different from the Eurozone (see figure 2.48C). This is reasonable. These three countries are reluctant to join the Eurozone on political ground; it is not because these three countries have problem qualifying for membership.

However, one might think of the Black Wednesday. It is useful to note that the period of study in this chapter is from 1994-2006. The Black Wednesday refers to 16 Sept. 1992 when the conservative government in the UK was forced to withdraw the Pound from the European Exchange rate mechanism (ERM) due to pressure by currency speculators; as a result the UK economy went through recession. However, it began a sustained recovery few years later and the economy has been significantly stronger than that of the Eorozone, despite the damage caused to the economy in the short term.

The scatter plots illustrating the relationship between risk premia and the inflation are presented in figure 2.49C. From the figure, the EMU members are clustered around one another. The majority of countries in the $POOR$ group are more dispersed in terms of both the risk premia and inflation. Overall, the fitted line shows a clear upward trend, which reflects a strongly positive relationship between the risk premia and the level of inflation. The t statistics from the single regression in both figures are significant at the 1 percent level.

The real effective exchange rate ($REER$) provides a measure of a country's competitive position over time by taking the effect of price movements into account[35]. Movements in real effective exchange rates provide an indication of the evolution of a country's aggregate external price competitiveness since it measures the currency's appreciation/depreciation against a weighted basket of foreign currencies and adjusts for relative prices between countries. The goods and services produced in particular country may not find buyers in both foreign and domestic markets if there is a fall in competitiveness. An improvement/fall in international price competitiveness affects the country's international trade position, national production, employment and income. We might expect that a rise in the $REER$ (a fall in international competitiveness) results in an economic contraction as

[35] To explain the concept of real effective exchange rate, we first refer to the real exchange rate. The real exchange rate is the nominal exchange rate adjusted for relative prices between the countries under consideration. It is expressed as:

$$E_{\text{real}} = \frac{EP}{P^*}$$

where E_{real} is the index of the real effective exchange rate, E is the nominal exchange rate (foreign currency per unit of domestic currency) in index form, P is the index of the domestic price level, and P^* is the index of the foreign price level. Instead of using a single foreign currency, the real effective exchange rate is concerned with what is happening to it against a basket of foreign currencies with whom the country trades [Pilbeam, 1998, pp.13-16].

suggested in the Mundell-Fleming model. This in turn might be expected to be associated with a rise in the risk premia for holding government bonds in that country.

We also link real effective exchange rate volatility ($VREER$) to the risk premia of government treasury bills. We measure real exchange rate volatility as the natural log of the standard deviation of the real effective exchange rate over 1994-2006. Using monthly data (t) of $REER$ in country i, we define the annual standard deviation of the real effective exchange rate as

$$VREER = \sigma_i^{REER} = \left[\frac{1}{T}\sum_{t=1}^{T}(REER_{it} - \overline{REER_i})^2\right].$$

In this analysis, more volatile real effective exchange rates imply more uncertainty in the country's competitiveness position. Thus, we would expect a positive relationship between real effective exchange rate volatility and the risk premium.

Differences in the country's competitive position, as measured by the real effective exchange rate ($REER$), between the three countries groups is less clear-cut in the data. The charts of the country's average real effective exchange rate over 1994-2006 are presented in figure 2.48D. On the other hand, the exchange rate volatility ($VREER$) over the period is generally higher in the $POOR$ group than the higher income group (EMU and $NEMU\_RICH$). Additionally, the majority of countries in the EMU group have relatively lower exchange rate volatility than the rest. The charts of the real effective exchange rate volatility are presented in figure 2.48E.

The plots of the relationship between the risk premia and the real effective exchange rate are presented in figure 2.49D. The impact of the country's competitive position on the risk premia on holding 6-month treasury bills is unclear. Figure 2.49E presents data for the relationship between the risk premia and the volatility of the real effective exchange rate. There is a strongly positive relationship between the risk premia and the volatility of the real effective exchange rate which is consistent with our prior. The t statistics from the single regression is significant at the 1 percent level.

Government debt as a percentage of gross domestic products can be considered as a determinant of government default risk. The higher the existing debt stock to GDP ratios, the greater the debt service obligations and the lower the government's capacity to borrow and roll over debt declines. This ultimately may result in an increase in the risk of default. We thus might expect a positive relationship between the risk premia and the government debt. The regression uses the natural log of the mean government debt as a percentage of GDP over 1994-2006.

An increase in the fiscal deficit might impact the risk premium for two reasons. Firstly, fiscal expansion may worsen future public debt and increase the probability of a debt crisis. Secondly, it affects public trust and investors' expectations. The ability to control fiscal deficits reveals information about government preferences, the importance of lobbies (which expect tax cuts or expenditure increases) and the degree of reform implementation (i.e. future public deficits.). Hence, we might expect the risk premia is increasing with the government budget deficit. In the regression, we use the mean of the deficit as a percentage of GDP for country i over 1994-2006, $\overline{DEFGDP_i}$.

The data for government budget deficit and debt as a percentage of GDP over 1994-2006 are presented in figures 2.48F and 2.48G. There is not much different across the groups. In figure 2.48F, the negative value represents the government budget deficit[36]. On average of 1994-2006, majority of sample countries have government budget deficit. The exceptions are Ireland, New Zealand, Brazil, Hong Kong, Singapore, Thailand, and Slovak Republic, which have government budget surplus. Due to the high variation among samples, we normalize this variable by taking the natural log of $(1 + 0.1 * \overline{DEFGDP_i})$ in the regression.

A scatter plot of the risk premia and the government budget deficit data is presented in figure 2.49F. There is no significant relationship between these two variables. We suspect, however, that the budget deficit does not strongly drive risk due to the existence of the outliers e.g. Norway, Sri Lanka, India, Philippines and Singapore. We will leave this issue until the next section.

Figure 2.49G contains data on the risk premia and government debt. The predicted coefficient of government debts is not statistically significant. Surprisingly, the plots show negative relationship between government debts and the risk premia. It can be argued that government debts are not always bad. Debts reflect the demand for government assets by investors. The greater demand for them (given that there is no constraint on the supply side) may also mean that they are safer bet than private assets or foreign assets. Thus, it doesn't always mean that countries with high proportion of public debt will have inferior fiscal stance and economic condition. For example, Belgium and Philippines both have high government debt (see figure 2.48G) but the risk premia for holding securities in the

36 This rule applies for figure 2.48F only, to give clearer illustration. In the analysis beyond this point, such as in the bivariate regression plots, cross section and panel regression analysis; the government budget deficit has positive sign.

former is less than the latter country (see figure 2.44). On the other hand, there are low government debts in Australia and Colombia. Unsurprisingly, the risk premia in Australia is lower.

The political variables used in this paper are the political risk index ($ICRG$) created by the PRS group and the political constraints index ($POLCON5$) by Henisz (2000, 2002).

The political risk index ($ICRG$) measures the political stability of countries on a comparable basis. The index is based on 100 points. The higher number of points indicates lower potential political risk e.g. 80-100 points represent very low risk and 0-49.5 points represent very high risk. In the political risk assessment, the number of points depends on the fixed weight of the political risk components. The political risk components and their weights in the parentheses are Government stability (12), Socioeconomic Conditions (12), Investment Profile (12), Internal Conflict (12), External Conflict (12), Corruption (6), Military in Policies (12), Religion in Policies (6), Law and Order (6), Ethnic Tensions (6), Democratic Accountability (6) and Bureaucracy Quality (4). The data for $ICRG$ are available annually. In the regression, we take natural logs of the mean of the political risk index over 1994-2006. We might expect a negative relationship between $ICRG$ and $RP3\_6$. In other words, lower risk premia for holding government assets should be positively related to the ICRG rating.

The $POLCON5$ measures the effective political restrictions on executive behaviour. It accounts for the veto powers of the executive whether or not there are, two legislative chambers, sub national entities and an independent judiciary. The index ranges from zero to one, where the higher value indicates stronger political constraints on the government. We take the natural log of the average values of $POLCON5$ over 1994-2006. The stronger political constraint reflects a more stable government policy, which may in turn result in reduced risk premia.

Higher income countries tend to have lower political risk ratings (higher score) and stronger political constraints than the lower income group, as shown in figure 2.48H and 2.48I. From the scatter plots in figures 2.49H and 2.49I, the risk premia exhibit negative correlations with both political variables as expected. The scatter plot of the risk premia and the political risk rating is presented in figure 2.49H. The political risk index negatively determines the risk premia as we expected. The predicted coefficient is highly significant (at the 1 percent level). The scatter plot of the risk premia and the political constraint is illustrated in figure 2.49I. The determinant of the political constraint index on the risk premia is less strong but the sign of the predicted coefficient is correct. The predicted coefficient is significant at the 12 percent level.

The next section is to present the result from the cross section regression analysis.

2.5 The cross section regression

This section examines the determinants of risk premia on holding 6-month treasury bills in 43 countries using cross section regression analysis. We test whether macroeconomic variables, government fiscal variables and political variables determine the risk premia. The dependent variable in the regression is the average risk premia for holding 6-month treasury bills comparing to 3 month treasury bills ($RP3\_6$) for different countries over the period 1994-2006 (as depicted in figure 2.44). In general, investors who hold these assets are mainly financial institutions. These financial institutions are assumed to minimize investment risks by spreading assets among different investments both nationally and internationally. The difference between these 2 assets is that holding shorter term treasury bills is less subjected to liquidity risk. In other words, the ability to sell or convert a security into cash is obviously greater for the shorter term treasury bills.

A small sample version of heteroskedasticity consistent covariance matrix estimator, HC3 proposed by MacKinnon and White (1985)[37] is applied to correct for heteroskedasticity in the cross-sectional data analysis[38]. The following paragraphs present the results of the risk premia cross-section regression on the macroeconomic and political variables.

The starting point for the risk premia cross-section regression[39] is to regress the risk premia on the macroeconomic variables, initial level of income and the country's economic and income group dummies. The results are presented in column (1) of table 2.7. The results show that inflation ($INFL$) and the economic growth ($GGDP$) are significant at the 5 percent level[40]. The budget Deficit as a percentage of GDP ($DEFGDP$) has predictive power at the 10 percent level. Initial level of income is significant at the 15 percent level. Central government debt as a percentage of GDP ($DEBTGDP$) and the

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37 Long and Ervin (2000) produced an extensive study of small sample behaviour and arrive at the conclusion that HC3 provides the best performance in small samples (less than 250 observations) as it gives less weight to influential observations.

38 When the variance of the errors varies across observations, OLS becomes inefficient and the estimates of the standard errors are inconsistent. This results in incorrect inferences. For a careful data analysis, we thus correct for heteroskedasticity in the cross sectional data analysis by using MacKinnon and White (1985)'s HC3.

39 The regression is based on the heteroscedasticity consistent covariance matrix (HCMM) version HC3 by Mackinnon and White 1985. This helps correct heteroscedasticity in the small sample size model ($n \leq 250$).

40 The magnitude will be presented in the preferred model. It will be discussed in the latter paragraphs.

real effective exchange rate volatility ($VREER$) do not statistically determine the risk premia. Approximately 74 percent of the variability of the risk premia is accounted for by the explanatory variables in the model.

Column (2) of table 2.7 adds the political variables, $POLCON5$ and $ICRG$ to the model. The economic factors are robust to the inclusion of additional explanatory variables. However, the economic factors highly dominate in the risk premia regression, thus the political variables have limited explanatory power[41]. The sign of the predicted coefficients are as expected although are not significant. We can conclude from the regression in column (2) that the short run macroeconomic circumstances do most of the work in explaining the risk premia e.g. the higher inflation and government budget deficit, and the lower economic growth lead to higher risk premia. In contrast, the level of long run development as illustrated by the institutional variables, i.e. the political risk index and the political constraint index, and the public debt[42] do not determine the risk premia.

Column (3) of table 2.7 excludes the insignificant explanatory variables. The results from the previous section are unchanged. The effect of the deficit (as a percentage of GDP), $DEFGDP$ become stronger and is significant at the 5 percent level. The variables economic growth ($GGDP$) and inflation ($INFL$) are once again significant at the 5 percent level[43]. The standardized coefficient[44] (beta value) of this model is also presented in table

41 Adding political variables $POLCON5$ and $ICRG$ separately into the model in column (1) of table 2.7 also does not improve the explanatory power of each political variable in the regression.

42 A good example is again in Belgium. The average government debt as a percentage of gross domestic product over 1994-2006 is high in this country (as illustrated in figure 2.48G). However, the risk premia for holding government asset is quite low (see figure 2.44). For the case of this country, high debt may be a sign that a country is a safe bet.

43 Note that in column (3), omitting $DEBTGDP$ and REER volatility, $VREER$ yields 9 more observations which are Argentina, Brazil, Greece, Mexico, Sri Lanka, India, Indonesia, Korea and Thailand.

44 The standardised regression coefficients (Beta value) are computed by STATA to compare the relative

2.7. It indicates the size of the change in the risk premia, $RP3\_6$ (in term of its standard deviation) with respect to a one standard deviation in the explanatory variable. For example, based on the estimates in column (3), a one standard deviation increase in $INFL$ (from Germany to Portugal's level) raises the risk premium by 1.29 of a standard deviation (from Germany to Indonesia's level[45]).

Finally, it is possible that these outlying observations might skew our test for heteroscedasticity in column (3). We thus identify influential observations by DFITS measure[46] of Welsch and Kuh (1977). The measure suggests removing observations in Argentina, Brazil, Norway, Sri Lanka, Indonesia, Philippines and Singapore[47]. We omit these 7 countries from regression in column (3) and present the result in column (4). Comparing the previous column with the latter, dropping observations reduces the variation and standard errors of all estimated coefficients. Additionally, column (4) suggests the model

strength of various predictors within the model. These beta coefficients are measured in standard deviations instead of units of variables. The Beta values are presented for models in column (3) and (4) of table 2.7. The results are present at the bottom part of the table.

45 The rank of countries by the average risk premia for holding 6 month treasury bill (comparing to those with 3 month maturity) over 1994 to 2006 can be found in figure 2.44.

46 We assess "Influence" of the observations by DFITS measure by Welsch and Kuh (1977). An observation is said to be influential if removing the observation substantially changes the estimates of coefficients. Influence can be thought of as a product of leverage and outlier. The former measure how far an independent variable deviates from its mean. These leverage points can have an effect on the estimate of regression coefficient. The latter is an observation with large residual which may indicate a sample peculiarity or may indicate a data entry error. The DFITS measure summarise the information in the size of the residuals and the size of leverages (h_i). Following Bollen and Jackman (1990), the equation for DFITS is

$$DFITS_i = r_i\sqrt{\frac{h_i}{1-h_i}},$$

where r_i are the studentized residuals. Large residuals or large leverage increases the value of DFITS. Belsley, Kuh and Welsch (1980) suggest that DFITS values greater than $2\sqrt{k/n}$ deserve further investigation (where k = number of independent variables and n = number of observations).

47 Omitting observations from these 7 countries are reasonable. Firstly, there are limited observations in deriving risk premia for Argentina (from 1998:07-2002:01), Sri Lanka (from 1994:12-2001:01) and Indonesia (1994:12-2001:01). Lastly, the excess holding yield series in these 7 countries show the statistically insignificant ARCH-M. This is partly due to the economic crisis which generates large shocks to the excess holding yield series, for example, the large economic shock in Argentina in 2002.

does not suffer from heteroskedasticity (based on Breusch-Pagan and White tests[48]) and omitted variable bias (based on Ramsey's RESET statistics)[49]. Column (4) is thus the preferred model. In this regression, the power of $INFL$ and $DEFGDP$ becomes stronger and are both significant at the 1 percent level. $GGDP$ is once again significant at the 5 percent level. A one-standard-deviation increase in $INFL$ would raise the risk premia by 71.92 percentage points[50] (or a 1.23 standard deviation increase in the predicted risk premia). Additionally, a one standard deviation increase in an economic growth would yield a 0.91 standard deviation decrease (or 53 percent reduction) in the predicted risk premia. Lastly, a one standard deviation increase in the deficit as a percentage of GDP would yield a 0.55 standard deviation increase (or 32.3 percent increase) in the predicted risk premia.

The scatter plots of the risk premia regression of the preferred model (column (4) of table 2.7) are presented in figure 2.50. These figures show scatter plots of natural log of inflation, natural log of government budget deficit (% GDP) and economic growth, conditional on the natural log of initial level of income, and other control variables. All the three explanatory variables are in correct sign and are statistically significant. Empirically, risk averse investors appear to require less risk premia for holding government securities in countries with a sound and stable financial market condition, i.e., the lower level of inflations and the government deficits and higher economic growth. Referring back to section

[48] The test on heteroskedasticity given by the Breusch-Pagan test and the White's test. Both test the null hypothesis that the variance of the residuals is homogenous. From table 2.7, column (4), there is no evidence against the null hypothesis.

[49] The omitted variable bias test (ovtest) command performs a regression specification error test (RESET) for omitted variables under the null hypothesis that model has no omitted variables. From table 2.7, column (4), the null hypothesis is not rejected.

[50] This figure is also another form of the standardised regression coefficients. It represents the unit change in the dependent variable (natural log of risk premia) with a one standard deviation change in the explanatory variable. The figure is not shown in the table 2.7.

2.4.2, the bivariate regression plots of risk premia and economic growth (figure 2.49B) suggested a strongly positive relationship between the two variables. It is interesting to note here that after controlling for the initial level of income, there is a negative relationship between the risk premia and economic growth as suggests in our priors. The regression plots are presented in the upper right panel of figure 2.50.

We can also compare the scatter plots from a single cross section regression of risk premia on the $DEFGDP$ (in figure 2.49F) with its conditional plots (in second picture of figure 2.50). We observe that after cutting outliers (Argentina, Brazil, Norway, Sri Lanka, Indonesia, Philippines and Singapore), the $DEFGDP$ is statistically significant in the risk premia regression.

We can also undertake the risk premia analysis by country group from column (4) of table 2.7. Lower income countries[51] are estimated to have risk premia about 19 percent[52] more than in the high income countries outside the Eurozone, holding other variables constant. In the high income countries outside the Eurozone, the risk premia on holding government assets is predicted to be 10 percent more than those in Eurozone.

The results of the standard cross section regression tell us the relationship between the risk premia and the macroeconomic and political variable on average of time during 1994-2006. In the next section, we consider how changes in the macroeconomic and political variables over time affect the change in the risk premia over the same time period. This can be done by the panel estimation of the risk premia.

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[51] In the preferred model (column 4 of table 2.7), we include dummy variables EMU and $POOR$ in the model. The coefficient (and t-statistics) of dummy variables are $-0.10(0.20)$ and $0.19(0.38)$, respectively.

[52] The dependent variable (the risk premia) is measured in natural logs, thus we can interpret the coefficients of the dummy variables in percentage.

2.6 Panel Data Analysis

2.6.1 Methodology

In this section, we employ panel data analysis to study the behaviour and determinants of government asset risk premia in 43 countries over the period 1994-2006. In the panel regression, we examine annual data[53]. The risk premia and explanatory variables data are annualised by taking average value of the monthly observations.

In the panel regression analysis, there are 3 critical methodological considerations. Firstly, the panel regression analysis allow us to take into account the arguments that the risk premia is time varying (as stated in sections 2.1 to 2.4). Additionally, it accounts for omitted variables and unobserved heterogeneity by incorporating the fixed country effect into the model. Econometrically, the Hausman test indicates that the fixed effects model are more suitable for the data i.e. there is a systematic difference in the coefficients between the random effects and the fixed effects models (p=0.00).

The second methodological consideration concerns how the risk premia is modelled. Choosing the dynamic panel model by taking the lagged dependent variable as an additional regressor is appealing in econometric sense. The Augmented Dickey Fuller test reveals that the risk premia $RP3\_6_{it}$ series follow a stationary first order autoregressive process. Intuitively, the behaviour of the current risk premia partly depends on the measured value in

[53] We move to the use of annual data in the panel analysis because many series of the explanatory variables are only available annually such as political constrain index (POLCON), international country risk guide index (ICRG), government debt (%of GDP) and government deficit (% of GDP).

the recent past[54]. Thus, including the lag dependent variable accounts for partial adjustment of risk premia behaviour over time[55].

To examine the determinant of the risk premia, the following model is estimated:

$$y_{it} = \gamma y_{it-1} + x_{it}'\beta + \eta_i + \varpi_t + \epsilon_{it}, \tag{2.8}$$

given $|\gamma| < 1; i = 1, ..., N$ and $t = 1, ..., T$. The time effect is ϖ_t. The unobserved individual and time invariant country's fixed effect is η_i. An unobserved white noise disturbance is ϵ_{it}. The subscripts i and t represent country and annually observed time period from 1994-2006, respectively. Following Bond (2002), we assume that the disturbances ϵ_{it} are serially uncorrelated and are independent across individuals;

$$\begin{aligned} \sigma_\epsilon^2 &> 0, \\ E(\epsilon_{i,t}, \epsilon_{j,s}) &= 0; i \neq j \text{ or } t \neq s, \\ E(x_{i,t}, \epsilon_{j,s}) &= 0; \forall i, j, t, s. \end{aligned}$$

The term y_{it} is the dependent variable $RP3\_6_{it}$, the risk premia for holding 6 month treasury bills (compared with 3 month treasury bills). We normalise[56] it by taking natural

[54] Intuitively, people form their expectation about the risk premia in the future for holding government bonds based on its past. For example, if the risk premia has been higher than expected in the past, people would revise expectations for the future. This can be referred to the theory of adaptive expectations.

[55] Another motivation for including lags would be to account for exogenous shocks that are believed to have continual effects over time. The coefficients on lagged dependent variables imply whether these factors have a greater impact over time or whether their impact decays and the rate at which it decays. Including lags of dependent variables as regressors is a parsimonious way of accounting for the persistent effects of explanatory variables in the past and can also help eliminate serial correlations in the disturbance term (Beck and Katz, 1996, and Wawro, 2002).

[56] There is high variation in the samples. The risk premia data and other explanatory variables are thus normalised to correct for the relatively favourable and unfavourable economic conditions and other influences, which affect the risk premia difference among countries. The normalisation is implemented by taking natural log to the variables. However, in the raw data, some observations have negative value such as risk premia (minimum value is -1.94), inflation rate (minimum value is -3.96) and deficit as a percentage of GDP (minimum value is -20.79), the normalisation for such case is to take natural log to (1+0.1$RP3_6$), (1+0.1$INFL$) and (1+($DEFGDP$/30)).

log to $(1 + 0.1RP3\_6)_{it}$. The vector of strictly exogenous explanatory variables is x_{it} ,which consists of natural log of inflation $[\ln(1 + 0.1INFL)_{it}]$, natural log of real effective exchange rate $[\ln REER_{it}]$ and its annualised standard deviation, the economic growth $[GGDP_{it}]$, natural log of debt and deficit as a percentage of GDP $[\ln DEBTGDP_{it}$ and $\ln(1 + (DEFGDP/30))_{it}]$, natural log of the political risk index $[\ln ICRG_{it}]$ and the natural log of the political constraint index $[\ln POLCON_{it}]$. The descriptive statistics of these variables after normalisation is presented in table 2.8.

The last methodological consideration concerns the choice of estimators to accommodate the joint presence of dynamics and unobserved heterogeneity in individual countries. We employ Bruno's (2005a,b) bias-corrected least squares dummy variable (LSDVC) approach to model the risk premia. The rationale for using this estimator over the rests is presented in the following paragraphs.

Although the autoregressive panel data model helps account for dynamic partial adjustment of the dependent variable, it also introduces bias into the model (Nickell, 1981 and Bond, 2002). According to the standard results for omitted variable bias, the OLS estimator of γ (in equation (2.8)) is inconsistent and biased upwards since the lagged dependent variable is positively correlated with the error term due to the presence of the fixed effects. The Within group or fixed effect estimator (LSDV) is instead biased downwards in case of small T panel even when N is large (Bond, 2002). This is because the within group transformation induces a correlation between the transformed lagged dependent variable and the transformed error term in the case of small time period data (Nickell, 1981). In this study the time dimension of the panel is small $(T = 11)$ thus estimating the least square dummy variable model with a lagged dependent variable results in biased estimates. In estimating the dynamic panel data model, Judson and Owen (1999) found that the bias of the LSDV can be large even when $T = 20$.

The candidate consistent estimator will lie between the OLS and LSDV estimates. In previous literature, the first difference-IV estimators (Anderson and Hsiao, 1981 and 1982), the General Method of Moments (GMM) estimators (Arellano 1989; Arellano and Bond, 1991; Arellano and Bover, 1995) and system GMM (Blundell and Bond, 1998) are usually applied to solve the first order dynamic panel data models. However, these methods are only efficient asymptotically and thus are not suitable for small sample data. Bruno (2005a,b) pointed out that the weakness of these estimators is that their properties hold for large N, so they can be severely biased and imprecise in panel data with a small number of cross-sectional units, such as most macro panels.

A method for implementing the corrected least square dummy variable (LSDVC) gained popularity in recent literature and was introduced by Kiviet (1995 and 1999) for balanced panels. Bruno (2005a, b) extended the LSDVC estimation to unbalanced panels with a strictly exogenous selection rule. The LSDVC offers a method to correct the bias in LSDV estimator for samples where N is small or only moderately large. The Montecarlo evidence in Judson and Owen (1999)[57] showed that the LSDVC estimator is preferred to the GMM estimators when N is small or only moderately large. This argument is supported by Kiviet (1995) and Bun and Kiviet (2001).

[57] Judson and Owen (1999) use an RMSE criterion to evaluate different techniques for estimating dynamic panel models in macroeconomic balanced panel datasets. The study found that for panels of all sizes, a corrected LSDV (LSDVC) is the most preferred estimator since it generally has the lowest RMSE compared with OLS, LSDV, GMM (both one-step and two-step estimators by Arellano and Bond, 1991), Instrumental variables (by Anderson and Hsiao; 1981) estimators.

There are three consistent estimators available to initialise the bias correction in the LSDVC estimation, which are as follows. The first one is the Anderson and Hsiao estimator (AH), with the dependent variable lagged twice used as an instrument for the first difference model with no intercept. The second estimator is a standard one step Arellano and Bond's estimator (AB) with no intercept. Lastly, the standard Blundell and Bond estimator (BB) with no intercept. Considering the nature of the risk premia data in this study, the AH estimator is chosen to initialise the correction procedure. The data are characterised by small cross section observations, the BB estimator tends to perform badly since BB imposes more instrument and more moment conditions. The AB estimator performs better than AH if the estimated coefficient of the lagged dependent variable, γ (in equation (2.8)) in the LSDV estimation is persistent. However, from table 2.9 columns (3) and (4), γ is only approximately 0.26 in this study. Thus the AH estimator is the best choice[58]. Additionally, the statistical significance of the LSDVC coefficients has been tested using bootstrapped standard errors (with 200 iterations).

It is useful to point out that in the corrected least square dummy variable (LSDVC), Kiviet's bias correction assumes strict exogeneity in the explanatory variables, x_{it}. If the the explanatory variables are not strictly exogenous, the bias correction term is invalid[59]. In the risk premia measures, there is one concern about the exogeneity of the right hand side variables i.e. economic growth, $GGDP_t$[60]. We can do a robustness check for the correct

[58] However, we note that based on the finding by Bun and Kiviet (2001), differences in the initial estimators have only a marginal impact on the LSDVC performance.

[59] Huang (2005) correct the weakness of this methodology by using the lag of explanatory variables instead in the regression. However, this case cannot be applied to the risk premia measures. Intuitively, the risk premia is sensitive and reacts quickly to the shock in macroeconomic circumstances.

[60] If the $GGDP$ varible is proved to be endogenous, it is more proper to apply instrumental variable regres-

LSDV estimators by implementing Instrumental Variables estimation of the fixed effects panel data models (IV-FE), allowing possibility of endogenous regressors. The rest of the variables are treated as strictly exogenous.

We perform instrumental variables regression (or two stage least squares) to estimate the structural model for the risk premia, $RP3\_6_t$ (equation (2.8)). In the structural model, $RP3\_6$ is the endogenous dependent variable, $GGDP_t$ is an endogenous regressor, and the rest are exogenous variables. The first stage regression is modelled as

$$GGDP_t = \alpha + \beta_1 DUBI_t + \beta_2 DUBI_t^2 + \beta_3 Y_t + v_{i,t} \tag{2.9}$$

where $DUBI_t$ is the natural log of crude oil price[61] (Arab Gulf Dubai) in US dollars per barrel, $DUBI_t^2$ is the square of natural log of crude oil price and Y_t is the real Gross Domestic Product per capita relative to the United States[62]. We postulate that economic growth variable ($GGDP_t$) is a function of $DUBI_t$, $DUBI_t^2$, and Y_t. The variable Y_t reflects the degree of economic convergence. The change in oil price has short run impact[63] on the economy. A significant increase in oil price can slow down the economic growth in oil importer countries through its effects on spending, or aggregate demand. It also simultaneously create inflationary pressures through increased prices of oil products used by

sion. The limitation of the Kiviet's correction in the LSDVC measures is that it does not allow instruments.

61 The oil price data are obtained from Datastream, 2006.

62 The current per capita GDP expressed relative to the United states (US=100) is obtained from the Penn World Table, 2006.

63 Under the assumption that the oil prices do not increase sharply and become higher than their already high levels, their long run effect can be manageable. The impact of higher oil prices in the long run is that they possibly reduce the production capacity. However, dealing with the higher oil price in the long run can take place in many ways such as developing alternative energy sources and conserving the oil. Moreover, productivity gains from diverse sources, including technological improvements and a more highly educated workforce, are likely to exceed by a significant margin the productivity losses created by high oil prices (Bernanke, 2004).

consumers, such as gasoline and heating oil and prices of alternatives such as natural gas. Thus $GGDP_t$ is expected to have significant negative relationships with the oil price. The nonlinear relationships are examined as well. $GGDP_t$ is also expected to have inverse relationship with the economic convergence variable, Y_t. The country's real output relative to the United States implies the rate at which the economy catches up the United States. Countries with lower GDP per capita relative to the United States are expected to grow significantly faster than rich countries and they tend to catch up or converge to those with higher real per capita output in a faster speed.

The IV-FE is introduced as a robustness check estimator instead of being the best estimator because the IV-FE estimator also has a weakness. Its properties hold for large number of cross sectional units (N), so it can be biased and imprecise in panel data with small number of N, such as most macro panels including this work. In conclusion, the Kiviet corrections address the problem of small sample bias, but it is invalid if there is endogeneity problem. In contrast, the IV can correct the endogeneity, but it is problematic in the small cross sectional samples. The rationale for choosing best estimators here is to compare the results of the IV-FE and the fixed effect regressions (LSDV). If the estimated coefficients and standard errors are not systematically different, we can emphasise the Kiviet approach based on the bias correction of the LSDV estimator. Then the initial guess that the economic growth variable is endogenous is proved to be invalid as it does not change the results in the fixed effect estimations.

In the next section, we present the results of the risk premia regression using OLS, LSDV, IV-FE and LSDVC estimators. In the OLS, IV-FE and LSDV estimators, the standard errors computed are asymptotically robust to heteroskedasticity and serial correlation. The results will be presented in aggregate and subgroup estimates. The subgroup is determined by the country's income level. To pick up unobserved time effects, year dummies (ϖ_t) are included in all regressions in this study.

2.6.2 The Regression results

2.6.2.1 Whole sample results

The panel regression results for the whole sample are presented in table 2.9, including estimation by OLS, LSDV, IV-FE and LSDVC. Estimated p-values are given in parentheses below point estimates of parameters. For each estimation procedure, the first column is the baseline specification. In these, we control for the impact of macroeconomic variables in the risk premia regression. In the second column, we add political variables to the baseline model.

The results from the pooled regression (OLS) [in columns (1) and (2)] and the LSDV or the dynamic fixed effects estimator [columns (3) and (4)] are presented for comparisons with the result from the best estimates, LSDVC [(columns (7) and (8)]. Before discussing the result from the LSDVC estimates, we check the robustness of these measures by IV-FE.

The risk premia regression by IV-FE can be presented by columns (5) and (6). We test the validity of instruments in equation (2.9) by the Sargan-Hansen test for over-identifying restriction[64]. The results does not reject the null hypothesis that the instruments ($DUBI_t$, $DUBI_t^2$, and Y_t) in equation (2.9) are valid instruments (p-value =0.973 and 0.960 in mod-

64 See Hayashi (2000) page 227-228, 407, 417 for further discussion.

els in columns (5) and (6), respectively). Thus the instruments are indeed exogenous and correctly excluded from the estimated equation. We also perform the Anderson (1984) canonical correlations test. It is a likelihood ratio test of whether the equation is identified, i.e. that the excluded instruments are relevant, meaning correlated with the endogenous regressors. We reject the null hypothesis that the equation is under-identified (p-value =0.00 in both models in columns (5) and (6)). Thus, it indicates that the model is identified.

We then compare the coefficients of parameter in LSDV and IV-FE regressions. The results suggest the effect of the economic growth is weaker after being instrumented. However, the resulting coefficients of the lag dependent variables and the explanatory variables, and the standard errors in both measures are unchanged. We then employ the Hausman test to check whether there is a sufficient difference between the coefficients of the instrumental variables regression (IV-FE) and those of the standard fixed effect (LSDV). The Hausman test clearly indicates that coefficients estimated by IV-FE are not statistically different from those estimated by LSDV (the null hypothesis that different in coefficients is not systematic is not rejected, with $\chi^2(7) = 0.45$ and prob. > χ^2= 0.9996). This suggests that we can emphasise on the results of the LSDVC estimates. We can now proceed the analysis of the estimated results from the preferred estimators, LSDVC.

The estimated results from the LSDVC are as follows. First, the estimated coefficients of the lagged dependent variable estimated by LSDVC lie between the OLS and LSDV estimates as proposed in the methodology section.

Secondly, the greater real effective exchange rate volatility ($VREER$) and the lower economic growth ($GGDP$) are strongly suggestive of the higher the risk premia for holding government's short term assets ($RP3\_6$). In both the baseline model and the second model specifications (columns (7) and (8) of table 2.9, respectively), the real exchange rate volatility ($VREER$) is a highly significant determinant of the risk premia and is significant at the 1 percent level. The results are robust across all estimates. In both the baseline specification and the second model of the LSDVC estimates, the coefficients of the volatility of real effective exchange rate ($VREER$) in the risk premia ($RP3\_6$) regression is 0.01. The interpretation is that as $VREER$ increases by 1 unit, $RP3\_6$ rises by 0.0372 percent annually[65].

Economic growth ($GGDP$) significantly determines the risk premia ($RP3\_6$) at the 7 percent and the 5 percent levels in the first and second models, respectively. The results are consistent with the LSDV estimates. The coefficients of economic growth ($GGDP$) in the risk premia ($RP3\_6$) regressions are -0.461 and -0.481 in the baseline specifications and in the second model of LSDVC estimates, respectively. In the baseline specification, as the economic growth increase by 1 percent annually[66], the risk premia decline by 1.4642 percent annually. In the second model, the risk premia decline by 1.5442 percent annually with respect to 1 percent increase in the economic growth[67].

The finding that economic growth has an important explanatory role for the dynamics of the yield curve corresponds to the work of Ang and Piazzesi (2003) which suggest that

[65] This figure is obtained by anti-log procedure. The detailed calculation is presented in the Appendix to Thesis; Appendix A, section A.1.1.

[66] We calculate the economic growth variable by initially using quarterly data. The unit of growth is percent quarterly. The economic growth at quarter t is $GGDP_t$. It is defined as $(\ln GDP_t - \ln GDP_{t-4})/4$. This is to avoid the seasonal effects. Intuitively, this chapter calculates the economic growth by comparing GDP at a particular quarter this year with GDP at the same quarter in the following year. To convert it to annual data, we average the GDP growth of each quarter within a year.

[67] These figures are obtained by anti-log procedure. The detailed calculation is presented in the Appendix to Thesis; Appendix A, section A.1.2.

macro factors explain up to 85 percent variation in bond yields[68], and Hordahl, Tristani and Vestin (2006).

It is interesting to note that coefficients of the economic growth are not statistically significant in the OLS regressions (columns (1) and (2) of table 2.9). However, they are statistically significant in the regressions that include country fixed effects, η_i (in equation (2.8)), which can be seen in columns (3) to (8) of table 2.9. Referring back to figures 2.48B and 2.49B, countries with higher average risk premia are growing faster perhaps due to convergence. In contrast, countries with lower risk premia tend to have slower economic growth. The economic growth rate thus correlates to the fixed country effects. For example, assets in the US have low risk because of a highly developed financial system, economic stability and relatively high market confidence. However, the economic growth rate in the US is not as high as in Mexico, which is less financially developed. The OLS fails to distinguish the country specific factors, it thus bias coefficients of the economic growth in the risk premia regression back down to zero. This explains why the coefficients of $GGDP$ are not statistically significant in the OLS regression. In contrast, fixed effect estimation distinguishes institutional features of high and low income countries. Fixed effect estimation allows the influence of economic growth on the risk premia, holding country effects constant.

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68 Considering a reverse relationship between the risk premia and the economic growth, some other works use the yield curve to predict the macroeconomic conditions such as Ang, Piazzesi and Wei (2006). These studies find that the term spread has limited power in forecasting GDP growth but the short term interest rates perform better in predicting GDP growth. Accordingly, we check for the causal relationships between the risk premia and the economic growth and found that the risk premia does not determine the risk premia. Thus, we can be safe from the endogeneity problem.

A key finding is the strong statistical relationship between exchange rate volatility ($VREER$) and risk premia ($RP3\_6$). Intuitively, an investment decision is made under uncertainty over the economic environment such as exchange rates, and future tax and regulatory policy. An uncertainty in assets return and foreign exchange is generally captured by their series volatility. The risk averse investors tend to require higher risk premia for holding assets denominated in the higher volatile currency.

We found that government debt ($DEBTGDP$) and the fiscal deficit ($DEFGDP$) do not determine the risk premia in the panel regression[69]. This corresponds to Lamfalussy (1989) who argued that the fiscal stance of governments is often insufficiently reflected in risk premia.

In the OLS regression, the coefficients on $DEBTGDP$ are significant at the 6 percent and the 7 percent levels, respectively (see columns (1) and (2) of table 2.9). However, the fixed country effects eliminate the importance of $DEBTGDP$. From these results, we can infer that $DEBTGDP$ may be important but it correlates with the fixed effects. It is likely to be the data problem since the series observed are quite short and do not vary much over time.

Finally, the political variables ($POLCON$ and $ICRG$) have limited explanatory power to the risk premia. In a preliminary test using a simple pair-wise correlation, both variables appear to be individually significant at the 1 percent level. However, the effect of these two variables is weak in the panel regression. This is partially because of the

[69] Even though the risk premia appear to be significantly positively relate to the government debt as a percentage of GDP ($DEBTGDP$) at the 10 percent level in OLS regression (columns (1) and (2) of table 2.9). However, the estimates do not wipe out all the country's time invariant fixed effects that can influence the determinant of the risk premia. Thus the results from the OLS regressions are subjected to bias.

time dummies (ϖ_t) in the regression. Adding time dummies is a conventional way to pick up unobserved time effect. However, it is important to note that with the time dummies, we cannot identify variables whose change across time is common to each country. It is possible that the $ICRG$ index is collinear with the time dummies. As a result, the political risk index does not significantly determine the risk premia when the time dummies are included (Note that removing the time effects from the LSDV and LSDVC estimations in table 2.9, we found that the $ICRG$ index become significant at 5 and 12 percent level, respectively[70].).

We can conclude[71] that risk-averse investors tend to require less risk premia for holding government assets in countries with good economic performance e.g. high economic growth and stable external price competitive position e.g. low volatility of real effective exchange rate. Although with caveats, the political variable and government fiscal conditions have limited ability to explain the risk premia.

2.6.2.2 Sub-samples

In this subsection, we split the data according to income groups (high income and lower income groups) to restrict the income heterogeneity across countries. The definition of high/lower income countries is according to the World Bank (2006) country classification[72]. The details of this classification are presented in the last column of table 2.1. This

[70] Removing the time dummies from LSDV and LSDVC models, the coefficients (and probability) of $ICRG$ are -0.068 ($p = 0.04$) and -0.062 ($p = 0.12$), respectively.

[71] Lastly the results for a reduced panel where the countries with no time-variation in volatility are omitted is presented in appendix to table 2.9. Comparing with the panel regression with full sample in table 2.9, the results are quite similar. To maintain the sample size, we keep results in table 2.9 as the main finding.

[72] For operational and analytical purposes, the World Bank's main criterion for classifying economies is gross national income (GNI) per capita. Based on its GNI per capita, every economy is classified as low

definition is consistent with the cross section regression in the previous part. We start with the panel estimations of the countries in the high income group.

High income group

The panel regression results for the high income countries are presented in table 2.10. The data set consist of 21 countries. We find that the country's real effective exchange rate ($REER$) highly positively determines the risk premia ($RP3\_6$) across all estimates and model specifications. The estimated coefficient of this variable is significant at the 1 percent level in LSDV and LSDVC estimators. Comparing with the full sample regression in previous part, the impact of the real effective exchange rate volatility ($VREER$) become less strong here and is significant at about the 10 percent level in LSDVC (see columns (5)-(6) of table 2.10).

The strong positive relationship between real effective exchange rate and the risk premia in the sub-sample is intuitively reasonable[73]. As mentioned earlier, $REER$ measures the currency appreciation/ depreciation against weighted basket of foreign currencies and adjusts for relative prices between countries. A real depreciation lowers the country risk premia in financial robust country by shifting demand toward domestic goods as in the Mundell and Fleming model. This in turn raises output and the return earned by entrepreneurs. This also corresponds to Cespedes, Chang and Velasco (2004) which suggest that in the financial vulnerable countries, a real depreciation raises the country risk premium; in

income, middle income (subdivided into lower middle and upper middle), or high income.

[73] The real effective exchange rate ($REER$) does not appear to significantly determine the risk premia in the full sample regression. This is possibly due to the income heterogeneity across samples as we primarily find that initial income is an important determinant of the risk premia in the cross section regression.

contrast, country with financial robustness, the opposite happens. In the LSDVC estimates, the coefficients of $REER$ in the risk premia regressions are 0.059 and 0.058 in the first and second models, respectively (see columns (5) and (6) of table 2.10). The interpretation is as follows, a basis point increase in $REER$ index associates with 0.224 percent increase annually in the risk premia in the baseline model (and 0.220 percent increase annually in the risk premia in the second model)[74].

The effect of $VREER$ is weaker in this sub-sample. The coefficient of $VREER$ in the risk premia regression is 0.005 in both the first and second models using LSDVC estimates (see columns (5) and (6) of table 2.10). Thus, we can infer that as the $VREER$ increases by 1 unit, the risk premia increase by 0.018 units[75] (the size of coefficients on $VREER$ is half of those in the whole sample regression).

Other macroeconomic variables also determine the risk premia such as economic growth ($GGDP$), inflation ($INFL$) and government budget deficit as a percentage of GDP ($DEFGDP$).

Economic growth ($GGDP$) negatively determines the risk premia and is significant at the 10 percent level in the LSDV and LSDVC models (columns (3)-(6) of table 2.10). In the LSDVC estimates, the coefficients of $GGDP$ in the risk premia regression are -0.337 and -0.319 in the first and second models, respectively[76] (columns (5) and (6) of table 2.10). As economic growth increase by 1 percent per annum, the risk premia decline by

[74] These figures are obtained by anti-log procedure. The detailed calculation is presented in the Appendix to Thesis; Appendix A, section A.2.1.

[75] These figures are obtained by anti-log procedure. The detailed calculation is presented in the Appendix to Thesis; Appendix A, section A.2.2.

[76] These figures are obtained by anti-log procedure. The detailed calculation is presented in the Appendix to Thesis; Appendix A, section A.2.3.

1.002 percent annually in the first model and 0.940 percent annually in the second model, respectively.

In the OLS regression using a whole sample, the coefficient of $GGDP$ is found to be statistically insignificant (as shown in table 2.9 columns (1) and (2)). However, using the sample of high income country group, the coefficient of $GGDP$ become significant at the 5 percent level in OLS regressions (see columns (1) and (2) in table 2.10). This is because using the sub-sample helps restricting the income heterogeneity across countries. Since high income countries tend to have high economic growth and vice-a-versa, the growth data is also less heterogenous here[77]. The dividing samples by income group helps partially control for the country fixed effect and thus it allows the data to explain more variation in the risk premia.

Inflation ($INFL$) and government budget deficit as a percentage of GDP ($DEFGDP$) positively determine the risk premia and are significant at the 10 percent level in the LS-DVC estimates. The coefficients of $INFL$ in the risk premia regression are 0.029 and 0.027 in the first and second models, respectively (see columns (5) and (6) of table 2.10). As the inflation increase by 1 percent per annum, the risk premia increase by 0.017 percent annually in the first model and 0.016 percent annually in the second model[78]. The coefficients of DEF in the risk premia regression are 0.021 in both the first and second

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[77] This support our argument earlier that the OLS regressions fail to distinguish the association of the low risk and low growth countries in the full samples.

[78] These figures are obtained by anti-log procedure. The detailed calculation is presented in the Appendix to Thesis; Appendix A, section A.2.4.

models. As deficit increase by 1 percentage of GDP, the risk premia increase by 0.004 percent annually[79].

Medium to Low income group

The panel regression results for Medium to low income groups are presented in table 2.11. The sample size in this group is very small[80]. We thus consider dropping the dynamic analysis and the time dummies in the regression of the lower income group (as presented in table 2.11). This section therefore only roughly explains the relationship between risk premia and explanatory variables in these countries. The results from the OLS regression and the LSDV estimates are presented in table 2.11.

The fixed effect estimations in column (4) of table 2.11 show that the volatility of real effective exchange rate ($VREER$) and inflation ($INFL$) are the main determinants of the risk premia. The coefficients of $VREER$ and $INFL$ are significant at 1 percent and 5 percent level respectively. Although the real effective exchange rate ($REER$) does not significantly determine the risk premia, it is interesting to discuss the sign of the coefficient of this variable (see column (4) of table 2.11). There is a negative relationship between the risk premia and $REER$ which is in contrast to the results from the high income country group. One possible explanation here is that the medium to low income countries are countries with vulnerable financial systems. According to Cespedes, Chang

79 These figures are obtained by anti-log procedure. The detailed calculation is presented in the Appendix to Thesis; Appendix A, section A.2.5.

80 Initially, there are 13 countries in medium to low income group. However, 5 countries such as Turkey, South Africa, Uruguay, Hungary, and Poland are outliers. We drop observations of these 5 countries. After cutting outliers, there are 8 countries left for examinations which are Brazil, Colombia, Czech Republic, India, Malaysia, Mexico, Philippines and Thailand.

and Velasco (2004) a real depreciation has positive relationship with the country risk premium in financial vulnerable countries. In conventional textbook, expansionary monetary policy and depreciation of the currency are optimal in response to an adverse foreign shock. In practice, if an economy has a large debt denominated in foreign currency then a weaker local currency can also exacerbate debt service difficulties and wreck the balance sheets of domestic banks and firms. This channel may cause devaluations to be contractionary, not expansionary. As documented by Hausmann, Panizza and Stein (2001) and Calvo and Reinhart (2002), balance sheet effects have emerged as a prime reason why many central banks are reluctant to allow their currencies to devalue in response to external shocks.

Although the coefficient of $GGDP$ is significant at the 10 percent level in the fixed effects estimation in column (4), its predictive power is not strong after omitting other insignificant variables. The model in column (5) is the result of omitting all insignificant variables; $VREER$ and $INFL$ are still significantly determine the risk premia and are significant at 1 percent and 10 percent level respectively. In the regression of columns (4) and (5), there are only 4 countries observed here which are Colombia, Malaysia, Philippines and Czech Republic. This is because the data for $REER$ are not available in Brazil, Mexico, India, and Thailand.

In column (6), if we omit $REER$ from the regression, we can observe data for 8 countries which are Brazil, Colombia, Czech Republic, India, Malaysia, Mexico, Philippines and Thailand. The coefficient of $INFL$ is significant at 5 percent level.

2.7 Conclusion

This study generates monthly risk premia data using zero coupon government treasury bills for 43 countries over the period of 1994-2006. The measure of risk premia is based on the ARCH-in-Mean (ARCH-M) model introduced by Engle, Lilien and Robins (1987). We show that the risk premia are time varying and also vary considerably between countries. This study also examines the macroeconomic and political determinants of the risk premia by using cross section regressions and dynamic panel regression analysis.

The cross section regression shows that on average through 1994-2006, the risk premia for holding government assets required by risk averse investors is positively influenced by the level of inflation and the deficit as a percentage of GDP and is negatively determined by the country's economic growth. Additionally, lower income countries are estimated to have risk premia about 19 percent more than in the high income countries outside the Eurozone, holding other variables constant. In the high income countries outside the Eurozone, the risk premia on holding government assets is predicted to be 10 percent more than those in Eurozone.

Using panel regression analysis, we found that economic growth and the volatility of the real effective exchange rate are the main determinants of risk premia in the full sample regression. Risk averse investors require lower risk premia for holding government assets in countries with good economic performance e.g. high economic growth and stable external price competitive position e.g. low volatility of real effective exchange rate. If we split the sample by income group, the real effective exchange rate which reflects country's external price competitiveness plays important role in high income countries. In the high income countries, the devaluation of currency brings in the favorable result to the economy. This is consistent with the Mundell-Fleming model. There is a better price competitiveness which in turn reduces the country risk premia. The opposite relationship is found in the regression of lower income countries. The possible explanation is that in financial vulnerable countries, weaker local currency can exacerbate the external debt service difficulties which result in economic contraction. This in turn raises the country risk premia. However, the impact of the level of real effective exchange rate is less strong in the low income group. For lower income countries, the volatility of the real effective exchange rate which reflects uncertainty in the exchange rate market plays important role in determining the risk premia. The higher real exchange rate volatility, the greater risk premia require for holding government assets in that country.

The institutional variables and the government fiscal conditions have limited power in explaining the risk premia in this study. This is possibly due to the measurement errors.

Lastly, it is useful to discuss the policy recommendations as follows. The membership of the European Monetary Union is proved to reduce the risk premia in this study. The economic growth is good as it associates with lower risk premia. On the average of time (using cross section regression), the inflation and the budgetary positions tend to have strong effect on the economy which is on contrary to the IMF conventional wisdom.

Table 2.1: Data Sources for Treasury bills rate and definitions of country income group

Country Code	Country	Descriptions for the Treasury Bills rates	Currency	Period of observation	Income group
111	USA	USD Government Agency (FMC84) Zero Coupon Yield	US Dollar	1994:12-2006:01	High income
112	UK	GBP United Kingdom (IYC22) Zero Coupon Yield	British Pound	1994:12-2006:01	High income
122	AUSTRIA	EUR Austria Sovereign (IYC63) Zero Coupon Yield	Austrian Schilling	1994:12-2006:01	High income
124	BELGIUM	EUR Belgium Sovereign (IYC6) Zero coupon Yield	Belgian Franc	1994:12-2006:01	High income
128	DENMARK	DKK Denmark Sovereign (IYC11) Zero coupon Yield	Danish Krone	1994:12-2006:01	High income
132	FRANCE	EUR France Sovereign (IYC14) Zero Coupon Yield	French Franc	1994:12-2006:01	High income
134	GERMANY	EUR Germany Sovereign (IYC16) Zero Coupon Yield	German Mark	1994:12-2006:01	High income
136	ITALY	EU Italy Sovereign (IYC40) Zero Coupon Yield	Italian Lira	1994:12-2006:01	High income
138	NETHERLANDS	EUR Netherlands Sovereign (IYC20) Zero coupon yield	Dutch Guilder	1994:12-2006:01	High income
142	NORWAY	NOK Norway Sovereign (IYC78) Zero coupon yield	Norwegian Krone	1994:12-2006:01	High income
144	SWEDEN	SEK Sweden Sovereign (FMC259) Zero coupon Yield	Swedish Krona	1994:12-2006:01	High income
146	SWITZERLAND	CHF Switzerland Sovereign (IYC82) Zero coupon Yield	Swiss Franc	1994:12-2006:01	High income
156	CANADA	CAD Canada Sovereign (UYC7) Zero Coupon Yield	Canadian Dollar	1994:12-2006:01	High income
158	JAPAN	JPY Sovereign (IYC18) Zero Coupon Yield	Japanese Yen	1990:01-2006:01	High income
172	FINLAND	EUR Finland Sovereign (IYC81) Zero Coupon Yield	Finnish Markka	1994:12-2006:01	High income
174	GREECE	EUR Greece Sovereign (FMC904) Zero coupon Yield	EURO	2000:08-2006:01	High income
178	IRELAND	EUR Ireland Sovereign (IYC62) Zero coupon Yield	Irish Punt	1994:12-2006:01	High income
182	PORTUGAL	EUR Portugal Sovereign (IYC84) Zero coupon Yield	Portuguese Escudo	1994:12-2006:01	High income
184	SPAIN	EUR Spain Sovereign (IYC61) Zero coupon Yield	Spanish peseta	1994:12-2006:01	High income
186	TURKEY	US dollar Turkey Sovereign (IYC I249) Zero coupon yield	US Dollar	2002:12-2006:01	Upper middle income
193	AUSTRALIA	AUD Australia Sovereign (IYC1) Zero Coupon Yield	Australian dollar	1994:12-2006:01	High income

Note: [1]The Treasury bills rates data are obtained from Bloomberg.
[2]The classification for Income group is calculated by World Bank country classification, 2006: Economies are divided according to 2005 GNI per capita, calculated using the World Bank Atlas method. The groups are: low income, \$875 or less; lower middle income, \$876 - \$3,465; upper middle income, \$3,466 - \$10,725; and high income, \$10,726 or more.
[3]FMC and IYC stand for Fair market value curve and International Yield curve, respectively. The fair market value indices are derived from data points on Bloomberg's option free market curves. The yield at each maturity point represents the composite yield of securities around that maturity.

Table 2.1: Data Sources for Treasury bills rate and definitions of country income group (cont.)

Country Code	Country	Descriptions for the Treasury Bills rates	Currency	Period of observation	Income group
196	NEW ZEALAND	NZD New Zealand (IYC49) Zero Coupon Yield	New Zealand Dollar	1994:12-2006:01	High income
199	SOUTH AFRICA	ZAR South Africa Sovereign (FMC262) Zero Coupon Yield	South African Rand	1994:12-2006:01	Upper middle income
213	ARGENTINA	Argentina Sovereign (FMC801) Zero Coupon Yield	Argentine Peso	1998:07-2006:01	Upper middle income
223	BRAZIL	USD Brazil Sovereign (FMC802) Zero coupon yield	US Dollar	1998:06-2006:02	Lower middle income
233	COLOMBIA	USD Colombia Sovereign (FMC803) Zero Coupon Yield	Colombian Peso	1998:06-2006:01	Lower middle income
273	MEXICO	USD Mexico Sovereign (FMC804) Zero coupon yield	Mexican Peso	1998:06-2006:02	Upper middle income
298	URUGUAY	BFV USD Uruguay Sovereign	US Dollar	2000:07-2006:01	Upper middle income
436	ISRAEL	Israel Makam Bond	Israel Shekel	1996:11-2006:01	High income
524	SRI LANKA	LKR Sri Lanka Sovereign (FMC133) Zero Coupon Yield	Sri Lankan Rupee	1994:12-2001:01	Lower middle income
532	HONG KONG	Hong Kong Sovereign (IYC95) Zero Coupon Yield	Hong Kong Dollar	1994:12-2006:01	High income
534	INDIA	India Sovereign (FMC123) Zero Coupon Yield	Indian Rupee	1998:11-2006:01	Low income
536	INDONESIA	IDR Indonesia Sovereign (FMC132) Zero Coupon Yield 3 Month	Indonesian Rupiah	1994:12-2001:01	Lower middle income
542	KOREA	KRW Korea Treasury (FMC232) Zero Coupon Yield	South Korean Won	1999:09-2006:01	High income
548	MALAYSIA	MYR Malaysia Sovereign (FMC128) Zero Coupon Yield 3 Month	Malaysian Ringgit	1999:09-2006:01	Upper middle income
566	PHILIPPINES	Philippines treasury Bill Generic yield	Philippines Peso	1995:10-2006:01	Lower middle income
576	SINGAPORE	Singapore Sovereign (IYC107) Zero Coupon Yield	Singapore Dollar	1994:12-2006:01	High income
578	THAILAND	THB Thailand Sovereign (FMC122) Zero Coupon Yield 3 Month	Thai Baht	1994:12-2006:01	Lower middle income
924	CHINA	CNY China Sovereign (FMC20) Zero Coupon Yield	China Renminbi	2003:09-2006:01	Lower middle income
935	CZECH REPUBLIC	CZK Chech Republic (FMC480) Zero Coupon Yield	Czech Koruna	1997:07-2006:01	Upper middle income
936	SLOVAK REPUBLIC	SKK Slovakia Swap rate Zero coupon yield	Slovakia Koruna	2003:02-2006:01	Upper middle income
944	HUNGARY	HUF Hungary Sovereign (FMC114) Zero Coupon Yield	Hungarian Forint	1998:06-2006:01	Upper middle income
964	POLAND	PLN Poland Sovereign (FMC119) Zero coupon Yield	Polish Zloty	1998:05-2006:01	Upper middle income

Note: [1]The Treasury bills rates data are obtained from Bloomberg.

[2]The classification for Income group is calculated by World Bank country classification, 2006: Economies are divided according to 2005 GNI per capita, calculated using the World Bank Atlas method. The groups are: low income, $875 or less; lower middle income, $876 - $3,465; upper middle income, $3,466 - $10,725; and high income, $10,726 or more.

[3]FMC and IYC stand for Fair market value curve and International Yield curve, respectively. The fair market value indices are derived from data points on Bloomberg's option free market curves. The yield at each maturity point represents the composite yield of securities around that maturity.

[4]Makam is a short-term (up to one year) zero-coupon bond. The Makam market is the most sensitive barometer of expected changes in interest rates in the monetary auction, and is generally the first to respond to interest rate changes.

Table 2.2: The Variables for the Cross Section and Panel Regression (Data from 1994-2006)

Variable	Description	Source
RP3_6	Risk premia for holding 6 month treasury bills (comparing to 3 month treasury bills)	Calculation
GDP94	The initial level of income or the Gross domestic product in 1994	IFS, 2006
INFL	CPI % CHANGE over corresponding period of previous year (Percent per annum)	IFS, 2006
REER	Real Effective Exchange Rate (CPI Based) (REER Based on REL.CP)	IFS, 2006
VREER	The Volatility of Real Effective Exchange Rate (CPI Based)	Calculation
GGDP	Growth rate of GDP over corresponding period of previous year [g_t= (log (GDP_t)-log(GDP_{t-4}))/4]. The unit is in percent quarterly.	Calculation
	Raw data source: Gross Domestic Product (National Currency), IFS, 2006	
DEFGDP	The government deficit as a percentage of GDP	Calculation
	Raw data source: DEFICIT (-) OR SURPLUS, IFS and Gross Domestic Product (National Currency), IFS	
DEBTGDP	Total central government debt % of GDP (AMT: Stocks: Outstanding amounts)	Calculation
	Raw data source: For OECD countries: International Comparisons - Central Government Debt, statistical yearbook, 1980-2003, OECD	
	: For non-OECD countries: World Development Indicators	
POLCON5	Extent of political constraints in policy-making process. Higher value implies stronger constraints and more stability in the policy.	Heinsz, 2005.
ICRG	The political risk rating. The lower the risk point assigned, the higher the political risk.	The PRS group, 2006
EMU	Dummy for member countries joining the European Economic and Monetary Union (EMU)	EU, 2006
	Country list: Austria, Belgium, France, Germany, Italy, Netherlands, Finland, Greece, Ireland, Portugal and Spain	
NEMU_RICH	Dummy for high income countries which do not belong to the EMU	EU, 2006
	Country list: Canada, Denmark, Norway, Switzerland, Sweden, UK, Japan, USA, Australia, New Zealand, Israel, Hong Kong, Korea, and Singapore	
POOR	Dummy for low income to middle income countries.	World Bank, 2006
	Country list: Turkey, South Africa, Argentina, Brazil, Colombia, Mexico, Uruguay, Sri Lanka, India, Indonesia, Malaysia, Philippines, Thailand, China, Czech republic, Slovak republic, Hungary and Poland	
	Note that Czech republic, Slovak republic, Hungary and Poland are EMU member on 1 May 2004. They are included in this group due to their income level and stage of EU membership.	

Table 2.3: Descriptive Statistics for Excess holding yield for the 3 month comparing to 6 month Treasury bills rates.

Country	Mean	Std. Dev.	Min	Max	Obs*
USA	0.35	0.40	-0.33	1.90	130
UK	0.20	0.38	-0.56	1.38	131
AUSTRIA	0.15	0.28	-0.46	0.95	131
BELGIUM	0.16	0.37	-1.42	1.57	131
DENMARK	0.22	0.37	-0.70	1.35	131
FRANCE	0.14	0.41	-1.49	1.74	131
GERMANY	0.09	0.27	-0.46	0.91	131
ITALY	0.23	0.48	-1.81	1.74	131
NETHERLANDS	0.16	0.29	-0.46	1.09	131
NORWAY	0.03	0.67	-3.15	1.66	131
SWEDEN	0.28	0.39	-0.83	1.50	131
SWITZERLAND	0.21	0.40	-0.88	1.20	131
CANADA	0.34	0.56	-0.95	2.68	131
JAPAN	0.11	0.39	-1.73	1.57	190
FINLAND	0.16	0.43	-0.73	2.02	131
GREECE	0.11	0.22	-0.47	0.83	63
IRELAND	0.10	0.43	-1.45	1.44	131
PORTUGAL	0.20	0.50	-1.11	1.97	131
SPAIN	0.17	0.32	-0.58	1.11	131
TURKEY	0.11	1.63	-5.03	5.33	35
AUSTRALIA	0.12	0.46	-0.75	2.17	131
NEW ZEALAND	0.14	0.74	-1.44	3.49	131
SOUTH AFRICA	0.79	1.58	-7.49	5.88	131
ARGENTINA	-3.95	18.09	-71.72	52.48	42
BRAZIL	1.00	4.27	-9.80	22.75	89
COLOMBIA	0.57	1.73	-4.49	5.69	89
MEXICO	0.56	1.23	-2.42	7.91	89
URUGUAY	-0.99	9.73	-38.55	20.22	64
ISRAEL	0.03	1.26	-3.67	2.62	108
SRI LANKA	0.14	0.37	-0.53	1.13	71
HONG KONG	0.34	1.06	-3.51	4.93	131
INDIA	0.36	0.56	-1.54	1.81	84
INDONESIA	0.17	0.36	-0.47	1.18	71
KOREA	0.59	0.44	-0.39	1.56	74
MALAYSIA	0.18	0.37	-1.04	1.56	74
PHILIPPINES	1.91	2.10	-7.36	9.33	121
SINGAPORE	0.11	0.77	-1.87	2.64	131
THAILAND	0.19	0.44	-1.20	1.27	131
CHINA	0.02	0.03	-0.04	0.07	26
CZECH REPUBLIC	0.46	0.73	-0.51	4.18	90
SLOVAK REPUBLIC	0.28	0.64	-0.56	1.42	23
HUNGARY	0.02	1.40	-3.36	3.62	89
POLAND	0.17	1.48	-4.28	6.86	90

* Obs stands for number of months observed.

Table 2.4: Descriptive Statistics for the estimated risk premia for holding 3 month comparing to 6 month Treasury bills (RP3_6)

Country	Mean	Std. Dev.	Min	Max	Obs*
USA	0.24	0.21	-0.04	0.87	133
UK	0.18	0.20	-0.07	0.78	133
AUSTRIA	0.13	0.06	0.03	0.27	133
BELGIUM	0.17	0.10	0.05	0.40	133
DENMARK	0.20	0.15	0.00	0.59	133
FRANCE	0.14	0.23	-0.09	0.84	133
GERMANY	0.06	0.10	-0.09	0.34	133
ITALY	0.19	0.10	0.06	0.41	133
NETHERLANDS	0.14	0.05	0.07	0.30	133
NORWAY	0.02	0.02	0.01	0.08	133
SWEDEN	0.25	0.08	0.15	0.50	133
SWITZERLAND	0.27	0.13	0.09	0.56	133
CANADA	0.25	0.13	0.09	0.65	133
JAPAN	0.09	0.08	0.01	0.32	192
FINLAND	0.13	0.05	0.05	0.31	133
GREECE	0.10	0.04	0.06	0.22	65
IRELAND	0.12	0.06	0.05	0.33	133
PORTUGAL	0.16	0.10	0.04	0.42	133
SPAIN	0.14	0.10	0.00	0.36	133
TURKEY	-	-	-	-	38
AUSTRALIA	0.10	0.25	-0.24	0.93	133
NEW ZEALAND	0.11	0.16	-0.05	0.65	133
SOUTH AFRICA	0.70	0.23	0.50	1.62	133
ARGENTINA	0.80	1.01	0.03	3.33	45
BRAZIL	0.88	0.85	0.08	3.07	91
COLOMBIA	0.61	0.52	-0.03	1.90	91
MEXICO	0.53	0.52	0.00	2.13	91
URUGUAY	0.59	1.94	-4.32	2.57	66
ISRAEL	0.15	0.16	-0.23	0.39	110
SRI LANKA	0.16	0.03	0.07	0.20	74
HONG KONG	0.42	0.11	0.32	0.75	133
INDIA	0.36	0.29	0.06	1.31	86
INDONESIA	0.17	0.00	0.16	0.17	74
KOREA	0.54	0.21	0.22	1.05	76
MALAYSIA	0.23	0.14	0.08	0.57	76
PHILIPPINES	1.98	0.58	1.49	3.60	123
SINGAPORE	0.09	0.16	-0.10	0.52	133
THAILAND	0.21	0.12	0.08	0.66	133
CHINA	-	-	-	-	28
CZECH REPUBLIC	0.41	0.38	0.03	1.63	92
SLOVAK REPUBLIC	-	-	-	-	24
HUNGARY	0.16	0.56	-1.14	0.85	91
POLAND	0.26	0.41	-0.88	0.68	92

* Obs stands for number of months observed.

Table 2.5: Results from ARCH-M regression using the excess holding yield of 3 month Comparing to 6 month.

		ARCH-M				ARCH			
No.	Country	β		v		α_0		α_1	
1	USA	-0.32	(-2.02)**	1.61	(3.49)***	0.02	(0.80)	0.95	(3.79)***
2	UK	-1.00	(-1.78)*	3.35	(1.99)**	0.06	(1.92)*	0.55	(1.86)*
3	AUSTRIA	-0.07	(-1.07)	0.75	(2.45)*	0.01	(2.13)*	0.84	(4.10)***
4	BELGIUM	0.01	(0.28)	0.42	(3.02)***	0.00	(0.32)	1.52	(9.27)***
5	DENMARK	-0.24	(-2.32)**	1.39	(3.63)***	0.03	(2.57)***	0.79	(3.78)***
6	FRANCE	-0.20	(-5.27)***	1.06	(5.91)***	0.01	(1.24)	1.07	(6.69)***
7	GERMANY	-0.26	(-2.92)***	1.29	(3.11)***	0.01	(1.38)	0.87	(3.89)***
8	ITALY	0.00	(0.03)	0.40	(1.90)*	0.01	(1.32)	1.27	(7.45)***
9	NETHERLANDS	-0.01	(-0.17)	0.57	(1.67)*	0.01	(1.03)	0.89	(3.59)***
10	NORWAY	0.00	(0.03)	0.03	(0.26)	0.00	(0.15)	1.34	(9.45)***
11	SWEDEN	-0.32	(-0.66)	1.58	(1.12)	0.08	(4.53)***	0.36	(1.70)*
12	SWITZERLAND	0.03	(1.17)	0.60	(6.02)***	0.01	(1.38)	1.21	(7.13)***
13	CANADA	-0.01	(-0.09)	0.50	(1.83)*	0.02	(0.90)	1.08	(7.12)***
14	JAPAN	0.00	(1.30)	0.28	(3.69)***	0.00	(1.22)	1.13	(11.29)***
15	FINLAND	0.02	(0.49)	0.28	(1.46)	0.01	(1.27)	1.03	(5.52)***
16	GREECE	-0.03	(-0.28)	0.60	(1.25)	0.01	(1.46)	0.77	(2.67)***
17	IRELAND	0.02	(0.53)	0.25	(1.51)	0.01	(1.13)	1.08	(5.68)***
18	PORTUGAL	-0.01	(-0.19)	0.39	(1.86)*	0.01	(2.09)**	1.04	(6.13)***
19	SPAIN	-0.11	(-1.80)*	0.88	(2.91)***	0.01	(0.80)	0.97	(4.31)***
20	TURKEY	-0.30	(-1.47)	0.20	(0.96)	-0.11	(-0.57)	1.05	(5.05)***
21	AUSTRALIA	-0.67	(-3.90)***	1.94	(4.42)***	0.04	(1.47)	0.82	(4.11)***
22	NEW ZEALAND	-0.14	(-1.99)**	0.42	(2.29)**	0.03	(1.46)	0.96	(5.16)***
23	SOUTH AFRICA	0.38	(1.30)	0.23	(0.81)	0.15	(0.99)	1.10	(4.95)***
24	ARGENTINA	-0.05	(-0.09)	0.05	(0.26)	0.05	(0.03)	2.45	(4.88)***
25	BRAZIL	-0.03	(-0.23)	0.26	(2.72)***	0.05	(0.60)	1.36	(7.62)***
26	COLOMBIA	-0.29	(-1.28)	0.59	(2.69)***	0.01	(0.07)	1.15	(5.21)***
27	MEXICO	-0.15	(-2.11)**	0.69	(3.94)***	0.01	(0.41)	1.12	(6.31)***
28	URUGUAY	4.16	(1.49)	-0.39	(-0.90)	14.79	(1.79)*	0.96	(4.47)***
29	ISRAEL	0.68	(1.13)	-0.43	(-0.85)	0.42	(2.14)**	0.72	(3.74)***
30	SRI LANKA	0.32	(1.12)	-0.44	(-0.55)	0.06	(1.15)	0.61	(1.31)
31	HONG KONG	0.28	(2.23)**	0.16	(0.74)	0.04	(1.12)	1.14	(7.66)***
32	INDIA	-0.47	(-2.78)***	1.64	(4.03)***	0.10	(3.24)***	0.68	(3.42)***
33	INDONESIA	0.18	(0.39)	-0.04	(-0.03)	0.07	(1.14)	0.49	(0.81)
34	KOREA	-0.66	(-0.86)	3.01	(1.53)	0.05	(0.87)	0.67	(1.55)
35	MALAYSIA	0.04	(1.25)	0.43	(2.12)**	0.00	(-0.14)	1.94	(5.07)***
36	PHILIPPINES	-0.01	(-0.01)	1.02	(2.50)**	2.03	(7.60)***	0.50	(3.27)***
37	SINGAPORE	-0.26	(-2.07)**	0.50	(2.27)**	0.06	(1.43)	0.92	(4.34)***
38	THAILAND	-0.12	(-1.12)	0.81	(2.64)***	0.04	(2.79)***	0.88	(3.84)***
39	CHINA	2.31	(5.68)***	-82.78	(.)	0.00	(2.71)***	0.04	(1.29)
40	CZECH R.	-0.12	(-1.15)	0.98	(3.19)***	0.01	(0.74)	1.19	(7.05)***
41	SLOVAK R.	Flat Log likelihood							
42	HUNGARY	3.80	(1.54)	-2.93	(-1.39)	0.80	(1.85)*	0.49	(1.44)
43	POLAND	0.86	(4.29)***	-0.49	(-2.02)**	0.03	(0.32)	1.07	(6.97)***

Notes: [1]Figures in parenthesis () are t-ratios. *** indicates that a coefficient is significant at the 1% level, ** significant at the 5% level, and * significant at the 10% level.

The parameters correspond to equation (2.3)-(2.7) in Section 2.3.2: Theoretical Derivation of time varying risk premia.

Table 2.6: The descriptive statistics for variables in the cross section regression (1994-2006)

Table 2.6A: Summary Statistics for risk premia and their determinants

Variable	Obs	Mean	Std.Dev.	Min	Max
RP3_6	40	-1.5328	0.8325	-3.7883	0.6808
GGDP	43	-3.8833	0.8013	-6.0142	-1.5871
INFL	43	1.5289	0.8927	-0.5443	4.1087
DEFGDP	42	-0.2599	0.4146	-1.6359	0.6707
DEBTGDP	41	3.7382	0.6488	2.0096	4.7878
POLCON5	42	-0.3917	0.2945	-1.7863	-0.1134
ICRG	43	4.3107	0.1489	3.9604	4.4848
GDP94	42	8.9157	0.7357	7.2464	9.8349
VREER	32	2.0933	0.5419	1.1106	2.9326

Table 2.6B: Correlations between risk premium and macroeconomic and political variables

	RP3_6	GGDP	INFL	DEFGDP	DEBTGDP	POLCON5	ICRG	GDP94	VREER
RP3_6	1								
GGDP	0.4313	1							
INFL	0.5700	0.9069	1						
DEFGDP	-0.1669	-0.0511	-0.3336	1					
DEBTGDP	-0.1518	-0.2136	-0.2458	-0.1253	1				
POLCON5	-0.5805	-0.5743	-0.6047	0.1159	0.2220	1			
ICRG	-0.6275	-0.6109	-0.6733	0.2486	0.0602	0.7843	1		
GDP94	-0.7674	-0.7118	-0.7334	0.2221	0.2222	0.6631	0.7217	1	
VREER	0.5015	0.5284	0.5689	-0.1844	-0.2585	-0.4178	-0.4773	-0.6448	1

Table 2.7: The cross section regression: determinants of the risk premia for holding 6-month treasury bills (RP3_6)

	(1)		(2)		(3)		(4)	
Inflation	1.41	(0.60)**	1.35	(0.61)**	1.18	(0.47)**	0.85	(0.31)***
Deficit (%GDP)	1.62	(0.87)*	1.67	(0.94)*	1.12	(0.54)**	1.15	(0.40)***
Economic Growth	-1.73	(0.75)**	-1.73	(0.78)**	-1.33	(0.61)**	-0.76	(0.38)**
Debt (%GDP)	0.17	(0.22)	0.13	(0.25)				
REER volatility	0.22	(0.39)	0.24	(0.44)				
ICRG			-1.19	(2.08)				
POLCON5			0.06	(1.11)				
Control Variable								
Intial Income (GDP94)	-1.09	(0.73)	-0.93	(0.74)	-0.82	(0.61)	-0.61	(0.30)*
Group dummies								
EMU	0.28	(0.42)	0.36	(0.52)	0.04	(0.24)	-0.1	(0.20)
POOR	-0.1	(1.00)	0.00	(1.02)	0.16	(0.76)	0.19	(0.38)
R^2	0.7354		0.7462		0.5461		0.6445	
Adjusted R^2	0.6295		0.6052		0.4522		0.5557	
Number of Countries	29		29		38		31	
P-value for httest	0.23		0.28		0.07		0.37	
P-value for white test	0.07		0.03		0.00		0.31	
P-value for Ovtest	0.01		0.02		0.03		0.44	

Beta value	Col (3)	Col (4)
Inflation	1.29	1.23
Deficit (%GDP)	0.53	0.55
Economic Growth	-1.15	-0.91
Initial Income (GDP94)	-0.76	-0.65
EMU	0.02	-0.08
POOR	0.10	0.16

Notes: [1]The dependent variable is natural log of the risk premia for holding 6-month treasury bills (comparing to 3-month treasury bills), RP3_6.
[2]Numbers shown in parentheses are MacKinnon and White (1985) heteroskedasticity-consistent (hc3) standard errors. All regressions have a constant.
[3]The hettest performs the Breusch-Pagan test for heteroskedasticity in the independent variables. The whitetst performs a variant of the White test for heteroskedasticity that uses the predicted values from the original regression and their squared values. The ovtest performs the regression specification error test (RESET) for omitted variables. The corresponding numbers shown are p-values. ***, ** and * denote significance at 1%, 5% and 10%, respectively.
[4]The group dummies consist of; 1) Eurozone (*EMU*); 2) High income countries outsie Eurozone (*NEU_RICH*); 3) The Medium to Low income group (*POOR*). In the regression, the omitted category is *NEU_RICH*.

Table 2.8: Descriptive Statistics for variables in the panel regression1994-2006

Table 2.8A: Summary Statistics for risk premium and its determinants

Variable	Obs	Mean	Std. Dev.	Min	Max
RP3_6	391	0.028	0.041	-0.215	0.226
GGDP	454	0.024	0.107	-1.439	1.486
DEBTGDP	324	3.805	0.630	1.687	4.932
REER	360	4.626	0.109	4.093	4.915
VREER	360	0.625	0.606	-0.991	2.410
INFL	451	0.355	0.389	-0.504	2.429
DEFGDP	350	-0.054	0.147	-1.181	0.446
POLCON5	409	-0.337	0.227	-1.903	-0.113
ICRG	418	4.353	0.138	3.883	4.565

Table 2.8B: Correlations between risk premium and macroeconomics and political variables

	RP36	GGDP	DEBTGDP	REER	VREER	INFL	DEFGDP	POLCON	ICRG
RP3_6	1								
GGDP	0.236	1							
DEBTGDP	-0.028	-0.008	1						
REER	-0.037	-0.119	-0.099	1					
VREER	0.427	0.250	-0.280	0.100	1				
INFL	0.343	0.705	-0.112	-0.072	0.374	1			
DEFGDP	-0.121	-0.049	-0.064	-0.032	-0.139	-0.213	1		
POLCON5	-0.416	-0.256	-0.006	0.169	-0.163	-0.242	0.143	1	
ICRG	-0.498	-0.407	-0.068	0.112	-0.415	-0.538	0.318	0.471	1

Table 2.9: The Panel Regression: Determinants of the risk premia for holding 6-month treasury bills (RP3_6) (Using Whole Sample)

Dependent variable: RP36_(i,t)	OLS		LSDV		IV- FE		LSDVC	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
RP36_(i,t-1)	0.824	0.830	0.257	0.255	0.236	0.236	0.416	0.421
	(0.076)***	(0.060)***	(0.126)**	(0.125)**	(0.119)**	(0.116)**	(0.092)***	(0.093)***
GGDP_(i,t)	-0.254	-0.262	-0.502	-0.520	-0.978	-0.906	-0.461	-0.481
	(0.219)	(0.237)	(0.242)**	(0.247)**	(0.816)	(0.755)	(0.255)*	(0.250)**
DEBTGDP_(i,t)	0.004	0.005	0.000	0.000	0.001	0.001	0.004	0.004
	(0.002)*	(0.003)*	(0.009)	(0.009)	(0.005)	(0.006)	(0.009)	(0.010)
REER_(i,t)	-0.005	-0.007	0.015	0.020	0.012	0.018	0.012	0.019
	(0.025)	(0.025)	(0.020)	(0.017)	(0.029)	(0.026)	(0.022)	(0.024)
VREER_(i,t)	0.008	0.009	0.010	0.010	0.009	0.009	0.010	0.010
	(0.002)***	(0.003)***	(0.003)***	(0.003)***	(0.003)***	(0.003)***	(0.003)***	(0.003)***
INFL_(i,t)	0.021	0.024	0.010	0.013	0.017	0.019	0.014	0.019
	(0.015)	(0.013)*	(0.018)	(0.021)	(0.023)	(0.026)	(0.016)	(0.016)
DEFGDP_(i,t)	0.016	0.014	0.020	0.020	0.024	0.022	0.022	0.022
	(0.012)	(0.013)	(0.012)	(0.011)*	(0.017)	(0.016)	(0.018)	(0.019)
POLCON5_(i,t)		-0.004		-0.003		0.001		-0.005
		(0.010)		(0.011)		(0.009)		(0.015)
ICRG_(i,t)		0.015		-0.034		-0.048		-0.031
		(0.015)		(0.036)		(0.043)		(0.040)
Anderson cannon p-value					0.000	0.000		
Hansen J statistics p-value					0.973	0.960		
Countries	27	27	27	27	27	27	27	27
Observations	186	186	186	186	186	186	186	186

Notes: [1]Standard errors are reported in parenthesis below point estimates.
***, ** and * denote significance at 1%, 5% and 10%, respectively.
[2]Year dummies are included in all models.
[3]The LSDV estimator is the fixed effect estimator. In the OLS and LSDV estimators, the standard errors computed are asymptotically robust to heteroskedasticity and serial correlation. The LSDVC estimator is the corrected LSDV estimator by Bruno (2005).

Table 2.10: The Panel Regression: Determinants of the risk premia for holding 6-month treasury bills (RP3_6) (Using Sub-Sample: Rich Countries)

Dependent variable: RP36_(i,t)	OLS		LSDV		LSDVC	
	(1)	(2)	(3)	(4)	(5)	(6)
RP36_(i,t-1)	0.493	0.481	0.229	0.237	0.426	0.436
	(0.055)***	(0.051)***	(0.050)***	(0.050)***	(0.101)***	(0.107)***
GGDP_(i,t)	-0.350	-0.340	-0.354	-0.340	-0.337	-0.319
	(0.155)**	(0.145)**	(0.183)*	(0.176)*	(0.185)*	(0.187)*
DEBTGDP_(i,t)	0.003	0.004	0.008	0.008	0.008	0.009
	(0.002)	(0.002)**	(0.008)	(0.008)	(0.008)	(0.008)
REER_(i,t)	0.020	0.021	0.065	0.064	0.059	0.058
	(0.010)*	(0.011)*	(0.019)***	(0.018)***	(0.023)***	(0.023)***
VREER_(i,t)	0.003	0.004	0.005	0.006	0.005	0.005
	(0.002)	(0.002)*	(0.004)	(0.004)	(0.003)	(0.003)*
INFL_(i,t)	0.020	0.024	0.027	0.025	0.029	0.027
	(0.014)	(0.015)	(0.019)	(0.018)	(0.015)*	(0.016)*
DEFGDP_(i,t)	0.006	0.005	0.022	0.022	0.021	0.021
	(0.006)	(0.006)	(0.010)**	(0.010)**	(0.012)*	(0.013)*
POLCON5_(i,t)		0.008		0.000		-0.001
		(0.008)		(0.006)		(0.015)
ICRG_(i,t)		0.019		0.031		0.038
		(0.009)*		(0.032)		(0.039)
Countries	21	21	21	21	21	21
Observations	151	151	151	151	151	151

Notes: [1]Standard errors are reported in parenthesis below point estimates.
***, ** and * denote significance at 1%, 5% and 10%, respectively.
[2]Year dummies are included in all models.
[3]The LSDV estimator is the fixed effect estimator. In the OLS and LSDV estimators, the standard errors computed are asymptotically robust to heteroskedasticity and serial correlation. The LSDVC estimator is the corrected LSDV estimator by Bruno (2005).

Table 2.11: The Panel Regression: Determinants of the risk premia for holding 6-month treasury bills (RP3_6) (Using Sub-Sample: Medium to Low Income Countries)

Dependent variable: RP36_(i,t)	OLS			LSDV		
	(1)	(2)	(3)	(4)	(5)	(6)
GGDP_(i,t)	2.855			-1.808		
	(3.544)			(0.694)*		
REER_(i,t)	-0.160			-0.034		
	(0.077)			(0.063)		
VREER_(i,t)	0.043	0.050		0.036	0.030	
	(0.011)**	(0.018)*		(0.007)***	(0.005)***	
INFL_(i,t)	0.016	0.081	0.107	0.111	0.078	0.079
	(0.079)	(0.059)	(0.015)***	(0.030)**	(0.031)*	(0.028)**
DEFGDP_(i,t)	0.315			-0.073		
	(0.096)**			(0.053)		
Countries	4	4	8	4	4	8
Observations	34	34	70	34	34	70

Notes: [1]Standard errors are reported in parenthesis below point estimates.
***, ** and * denote significance at 1%, 5% and 10%, respectively.
[2]Year dummies are included in all models.
[3]The LSDV estimator is the fixed effect estimator. In the OLS and LSDV estimators, the standard errors computed are asymptotically robust to heteroskedasticity and serial correlation.

Appendix Table A: Descriptive Statistics for Excess holding yield for the 3 month comparing to 12 month Treasury bills rates.

Country	Mean	Std. Dev.	Min	Max	Obs
USA	1.69	2.03	-1.88	7.26	124
UK	0.83	1.44	-1.88	4.54	125
AUSTRIA	1.10	1.35	-1.61	4.59	125
BELGIUM	1.07	1.70	-2.73	6.58	125
DENMARK	1.90	1.70	-1.44	7.34	125
FRANCE	1.04	1.41	-1.43	6.10	125
GERMANY	0.76	1.45	-2.17	5.18	125
ITALY	1.63	1.89	-3.73	7.11	125
NETHERLANDS	1.07	1.45	-1.61	5.77	125
NORWAY	0.47	2.84	-9.02	6.37	125
SWEDEN	2.10	2.00	-1.57	7.40	125
SWITZERLAND	1.31	1.65	-2.89	5.90	125
CANADA	1.98	2.08	-1.04	7.16	125
JAPAN	0.89	1.61	-3.63	6.36	184
FINLAND	1.19	1.98	-2.13	7.71	125
GREECE	0.98	1.10	-1.56	3.84	57
IRELAND	0.81	1.86	-3.46	5.95	125
PORTUGAL	1.22	2.17	-4.16	8.04	125
SPAIN	1.33	1.68	-1.55	6.02	125
TURKEY	1.67	4.95	-5.59	20.51	29
AUSTRALIA	0.90	1.92	-2.28	7.28	125
NEW ZEALAND	0.31	2.82	-5.76	9.65	125
SOUTH AFRICA	2.06	5.44	-17.63	17.56	125
ARGENTINA	-19.79	48.96	-189.72	22.96	36
BRAZIL	5.07	14.68	-35.00	80.73	83
COLOMBIA	3.51	5.42	-4.39	26.22	83
MEXICO	3.22	3.81	-2.07	22.94	83
URUGUAY	-10.97	35.15	-112.44	47.69	58
ISRAEL	1.21	4.37	-11.92	9.46	102
SRI LANKA	0.82	1.69	-2.09	5.59	65
HONG KONG	1.98	3.60	-7.11	16.71	125
INDIA	1.96	1.99	-2.51	7.35	78
INDONESIA	0.99	1.77	-1.79	6.27	65
KOREA	3.04	2.10	-1.26	7.89	68
MALAYSIA	1.03	1.05	-1.88	3.77	68
PHILIPPINES	7.42	7.40	-14.38	28.98	115
SINGAPORE	0.74	1.89	-3.75	7.96	125
THAILAND	1.39	1.74	-3.79	6.75	125
CHINA	0.14	0.12	-0.06	0.32	20
CZECH REPUBLIC	2.74	3.31	-0.74	15.80	84
SLOVAK REPUBLIC	2.19	2.07	-2.02	4.12	17
HUNGARY	0.40	5.38	-16.06	8.81	83
POLAND	0.63	7.13	-17.17	18.11	84

Appendix Table B: Descriptive Statistics for the estimated risk premia for holding 3 month comparing to 12 month Treasury bills (RP3_12)

Country	Mean	Std. Dev.	Min	Max	Obs
USA	0.81	1.77	-2.06	5.96	133
UK	-	-	-	-	-
AUSTRIA	0.93	0.27	0.53	1.57	133
BELGIUM	0.87	0.33	0.40	1.87	133
DENMARK	1.87	0.90	0.84	4.56	133
FRANCE	0.72	0.54	-0.11	2.46	133
GERMANY	0.84	0.58	0.13	2.58	133
ITALY	1.22	0.52	0.43	2.45	133
NETHERLANDS	0.89	0.32	0.41	1.74	133
NORWAY	0.76	0.01	0.72	0.77	133
SWEDEN	1.74	0.56	1.12	3.39	133
SWITZERLAND	1.16	0.32	0.73	1.94	133
CANADA	1.25	0.79	0.29	3.14	133
JAPAN	0.71	0.65	0.04	2.36	192
FINLAND	0.90	0.40	0.39	2.18	133
GREECE	-	-	-	-	-
IRELAND	0.73	0.17	0.50	1.20	133
PORTUGAL	1.07	0.55	0.46	2.35	133
SPAIN	1.02	0.42	0.45	2.05	133
TURKEY	-	-	-	-	-
AUSTRALIA	0.50	1.23	-1.19	4.45	133
NEW ZEALAND	0.08	0.87	-1.10	2.82	133
SOUTH AFRICA	2.57	0.07	2.49	2.79	133
ARGENTINA	3.42	3.03	0.69	9.86	45
BRAZIL	5.05	5.16	0.75	20.15	91
COLOMBIA	2.66	2.33	-0.36	9.85	91
MEXICO	2.48	2.30	-0.33	9.68	91
URUGUAY	-1.34	22.07	-51.18	22.46	66
ISRAEL	2.38	1.12	-0.40	3.84	110
SRI LANKA	0.90	0.37	-0.03	1.42	74
HONG KONG	2.45	0.53	1.84	3.70	133
INDIA	1.36	1.68	-1.61	5.85	86
INDONESIA	1.16	0.25	0.58	1.53	74
KOREA	2.43	1.74	-0.08	6.89	76
MALAYSIA	1.25	0.87	0.36	3.47	76
PHILIPPINES	7.46	1.54	5.77	12.25	123
SINGAPORE	0.46	1.10	-1.22	3.96	133
THAILAND	1.71	2.09	-0.16	10.51	133
CHINA	-	-	-	-	-
CZECH REPUBLIC	-	-	-	-	-
SLOVAK REPUBLIC	-	-	-	-	-
HUNGARY	-	-	-	-	-
POLAND	1.60	3.90	-8.43	5.48	92

Appendix Table C: Results from ARCH-M regression using the excess holding yield of 3 month comparing to 12 month.

No. Country	ARCH-M				ARCH			
	β		ν		α_0		α_1	
1 USA	-14.07	(-1.76)*	10.60	(1.89)*	0.93	(1.51)	0.54	(1.76)*
2 UK	Flat Log likelihood							
3 AUSTRIA	0.35	(2.58)***	0.47	(3.24)***	·0.11	(1.17)	0.97	(4.28)***
4 BELGIUM	0.28	(2.37)**	0.39	(3.14)***	0.04	(0.48)	1.11	(5.74)***
5 DENMARK	0.39	(2.69)***	1.01	(7.19)***	0.12	(1.54)	1.01	(4.93)***
6 FRANCE	-0.47	(-2.26)**	1.04	(4.83)***	0.07	(0.65)	0.99	(4.33)***
7 GERMANY	-0.10	(-0.98)	0.69	(6.59)***	0.04	(0.38)	1.06	(4.81)***
8 ITALY	0.21	(1.61)	0.62	(4.84)***	0.06	(0.58)	1.07	(5.81)***
9 NETHERLANDS	0.27	(2.39)**	0.48	(4.2)***	0.06	(0.88)	1.05	(5.81)***
10 NORWAY	0.77	(2.01)**	-0.01	(-0.04)	0.91	(3.03)***	0.92	(5.19)***
11 SWEDEN	0.41	(0.75)	0.86	(2.06)**	0.46	(1.29)	0.80	(3.32)***
12 SWITZERLAND	0.54	(2.80)***	0.41	(2.27)**	0.18	(1.86)*	1.03	(5.10)***
13 CANADA	-0.40	(-0.82)	0.97	(2.36)**	0.24	(0.97)	0.92	(3.51)***
14 JAPAN	0.02	(1.04)	0.68	(7.23)***	0.00	(1.36)	0.99	(8.41)***
15 FINLAND	0.19	(1.14)	0.44	(3.02)***	0.11	(0.80)	1.00	(4.40)***
16 GREECE	Flat Log likelihood							
17 IRELAND	0.41	(1.74)*	0.20	(1.04)	0.15	(1.27)	0.98	(4.79)***
18 PORTUGAL	0.30	(1.76)*	0.43	(2.90)***	0.10	(1.68)*	1.03	(6.06)***
19 SPAIN	0.22	(1.30)	0.57	(3.15)***	0.12	(0.97)	0.98	(4.13)***
20 TURKEY	-150.39	(-0.05)	34.07	(0.05)	19.32	(2.15)**	0.03	(0.05)
21 AUSTRALIA	-4.20	(-2.89)***	3.16	(3.24)***	0.65	(1.57)	0.77	(3.50)***
22 NEW ZEALAND	-1.59	(-4.91)***	0.73	(4.20)***	0.27	(1.08)	1.01	(5.44)***
23 SOUTH AFRICA	2.43	(2.53)**	0.03	(0.11)	·2.35	(0.88)	0.97	(3.98)***
24 ARGENTINA	0.47	(0.13)	0.05	(0.15)	-2.16	(-0.08)	2.55	(2.83)***
25 BRAZIL	-1.84	(-1.09)	0.65	(2.71)***	13.57	(3.51)***	0.99	(5.37)***
26 COLOMBIA	-2.13	(-1.79)*	1.08	(3.16)***	1.44	(0.91)	1.08	(5.43)***
27 MEXICO	-1.12	(-3.68)***	1.10	(6.52)***	-0.18	(-0.49)	1.39	(7.88)***
28 URUGUAY	60.57	(2.10)**	-2.35	(-1.97)**	145.22	(0.59)	0.78	(1.99)**
29 ISRAEL	4.77	(4.79)***	-0.58	(-1.92)*	2.23	(1.70)*	1.02	(4.27)***
30 SRI LANKA	1.83	(3.97)***	-0.57	(-1.87)*	0.30	(0.80)	0.95	(2.93)***
31 HONG KONG	1.65	(5.58)***	0.23	(1.48)	0.43	(0.87)	1.26	(5.64)***
32 INDIA	-27.63	(-0.75)	21.71	(0.78)	1.40	(2.73)***	0.22	(0.77).
33 INDONESIA	1.87	(2.98)***	-0.42	(-1.11)	0.46	(1.01)	0.91	(2.85)***
34 KOREA	-17.13	(-0.85)	15.74	(0.98)	0.95	(1.42)	0.39	0.97)
35 MALAYSIA	0.22	(2.12)**	0.93	(4.15)***	0.01	(0.55)	1.55	(5.16)***
36 PHILIPPINES	4.31	(3.41)***	0.46	(2.02)**	7.39	(2.55)**	0.91	(4.74)***
37 SINGAPORE	-3.19	(-3.32)***	2.46	(3.44)***	0.46	(1.72)*	0.82	(3.87)***
38 THAILAND	-1.90	(-3.61)***	2.66	(6.63)***	0.18	(1.10)	1.23	(7.24)***
39 CHINA	Flat Log likelihood							
40 CZECH R.	Flat Log likelihood							
41 SLOVAK R.	Flat Log likelihood							
42 HUNGARY	Flat Log likelihood							
43 POLAND	7.55	(7.46)***	-1.20	(-4.65)***	·1.90	(1.03)	0.90	(3.90)***

Appendix Table D: Variables and Definitions used in the Cross section regression for the risk premia (RP3_6).

Variable	Description	Source
RP3_6	Natural log of country's mean risk premia, 1994-2006	Calculation
GDP94	Natural log of Gross domestic product in 1994 (national currency)	IFS, 2006
INFL	Natural log of mean inflation, 1994-2006.	IFS, 2006
REER	Natural log of mean REER, 1994-2006.	IFS, 2006
VREER	Natural log of standard deviation of Real effective exchange rate over 1994-2006.	Calculation
GGDP	Natural log of mean GDP growth, 1994-2006.	Calculation
DEFGDP	Natural log of (1+0.1DEFIGDP).	Nalculation
DEBTGDP	Natural log of mean debt as a percentage of GDP, 1994-2006	Calculation
POLCON5	Natural log of mean POLCON5, 1994-2006.	Heinsz, 2005.
ICRG	Natural log of mean ICRG, 1994-2006	The PRS group, 2006

Appendix Table E: Variables and Definitions used in the Panel regression analysis for the risk premia (RP3_6).

Variable	Description	Source
RP3_6	Natural log of (1+0.1RP3_6)	Calculation
GDP94	Natural log of Gross domestic product in 1994 (national currency)	IFS, 2006
INFL	Natural log of (1+0.1INFL)	IFS, 2006
REER	Natural log of REER	IFS, 2006
VREER	The annually observed VREER is calculated by taking natural log to standard deviation of REER over 12 month periods.	Calculation
GGDP	Annually observed GDP growth.	Calculation
DEFGDP	Natural log of (1+(DEFGDP/30))	Calculation
DEBTGDP	Natural log of government debt (%GDP)	Calculation
POLCON5	Natural log of POLCON5	Heinsz, 2005.
ICRG	Natural log of ICRG	The PRS group, 2006

Appendix to Table 2.9: The Panel Regression: Determinants of the risk premia for holding 6-month treasury bills (RP3_6)
(Using Whole Sample and good quality data)

Dependent variable: RP36_(i,t)	OLS		LSDV		IV- FE		LSDVC	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
RP36_(i,t-1)	0.862	0.882	0.202	0.195	0.238	0.234	0.401	0.408
	(0.060)***	(0.057)***	(0.091)**	(0.096)**	(0.069)***	(0.069)***	(0.114)***	(0.113)***
GGDP_(i,t)	-0.057	-0.110	-0.742	-0.738	-0.132	-0.153	-0.674	-0.688
	(0.326)	(0.364)	(0.326)**	(0.329)**	(0.288)	(0.309)	(0.339)**	(0.356)**
DEBTGDP_(i,t)	0.006	0.006	0.001	0.000	0.004	0.003	0.004	0.005
	(0.003)*	(0.003)**	(0.013)	(0.013)	(0.006)	(0.006)	(0.012)	(0.013)
REER_(i,t)	0.028	0.028	0.041	0.040	0.045	0.048	0.049	0.050
	(0.016)*	(0.013)**	(0.025)*	(0.024)*	(0.018)**	(0.019)**	(0.030)*	(0.031)*
VREER_(I,t)	0.006	0.006	0.007	0.007	0.010	0.010	0.007	0.007
	(0.004)	(0.004)	(0.005)	(0.005)	(0.004)***	(0.004)***	(0.006)	(0.006)
INFL_(i,t)	0.006	0.008	0.011	0.014	0.004	0.003	0.011	0.012
	(0.029)	(0.028)	(0.018)	(0.020)	(0.011)	(0.011)	(0.023)	(0.026)
DEFGDP_(i,t)	0.006	0.002	0.027	0.027	0.024	0.024	0.029	0.029
	(0.012)	(0.011)	(0.015)*	(0.015)*	(0.013)*	(0.013)*	(0.026)	(0.027)
POLCON5_(i,t)		0.006		-0.008		-0.005		-0.006
		(0.008)		(0.006)		(0.004)		(0.018)
ICRG_(i,t)		0.013		0.038		0.018		0.041
		(0.030)		(0.037)		(0.031)		(0.054)
Anderson cannon p-value					[0.000]	[0.000]		
Hansen J statistics p-value					[0.636]	[0.608]		
Countries	19	19	19	19	19	19	19	19
Observations	128	128	128	128	128	128	128	128

Notes: [1]Standard errors are reported in parenthesis below point estimates.
***, ** and * denote significance at 1%, 5% and 10%, respectively.
[2]Year dummies are included in all models.
[3]The LSDV estimator is the fixed effect estimator. In the OLS and LSDV estimators, the standard errors computed are asymptotically robust to heteroskedasticity and serial correlation. The LSDVC estimator is the corrected LSDV estimator by Bruno (2005).

Figure 2: Excess holding yield of 6 month Treasury Bills and estimated risk premia for 43 countries

Figure 2.1: Argentina

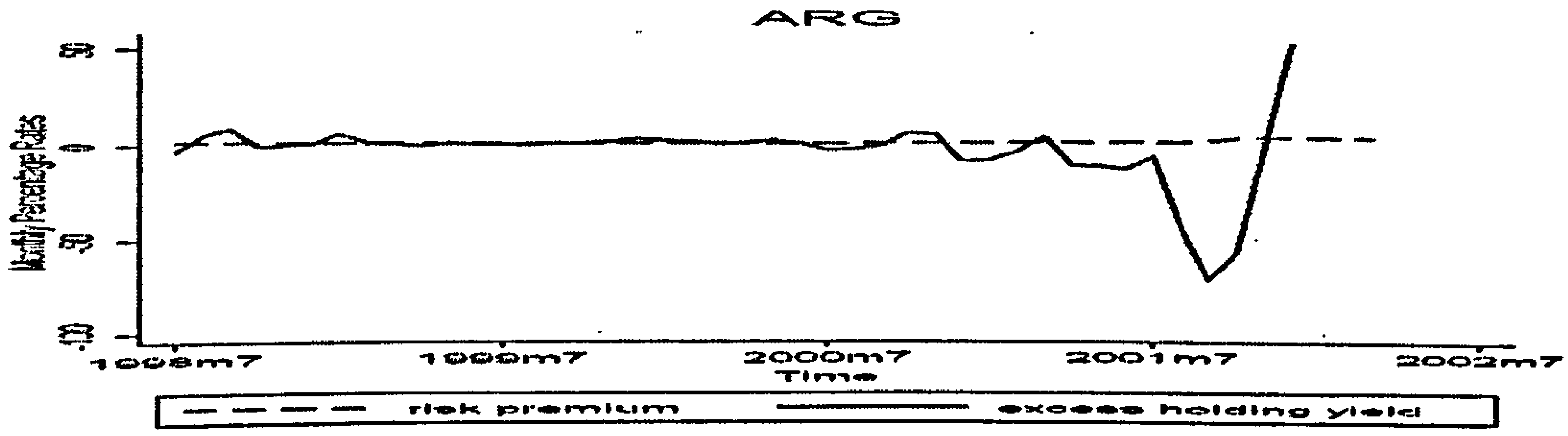


Figure 2.2: Australia

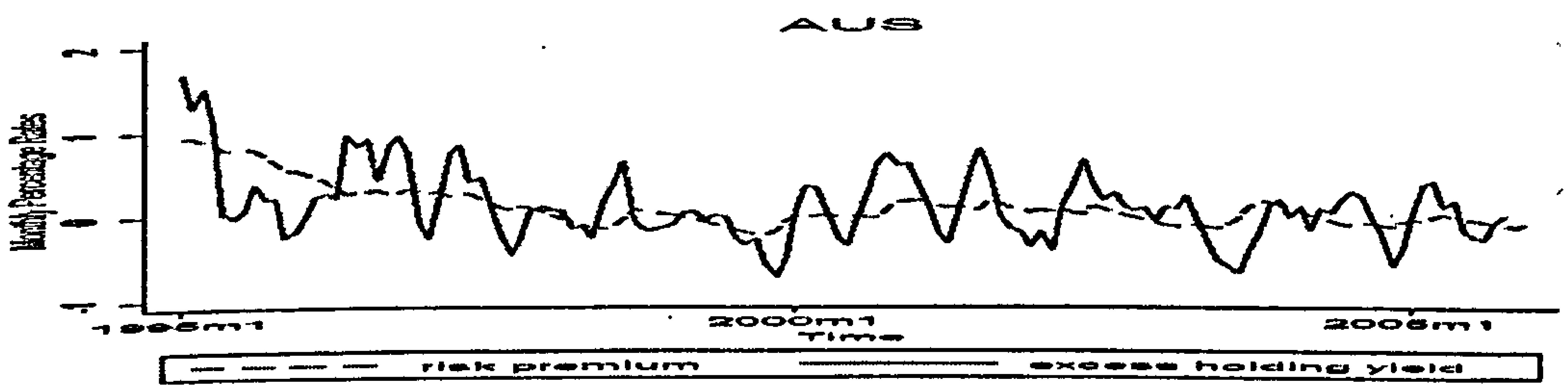


Figure 2.3: Austria

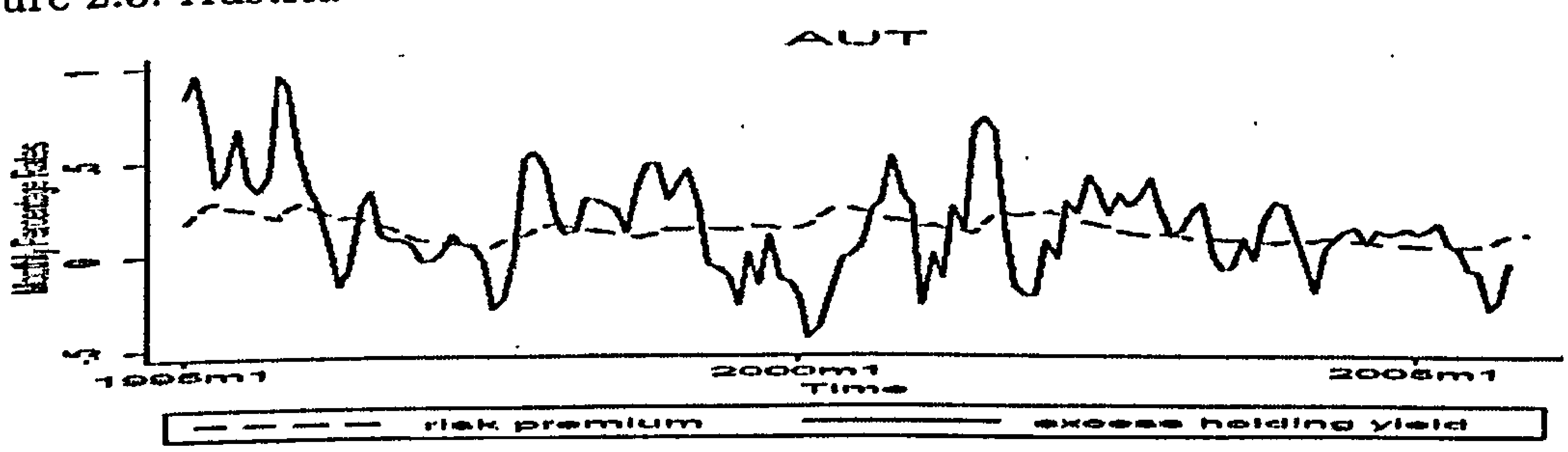


Figure 2.4: Belgium

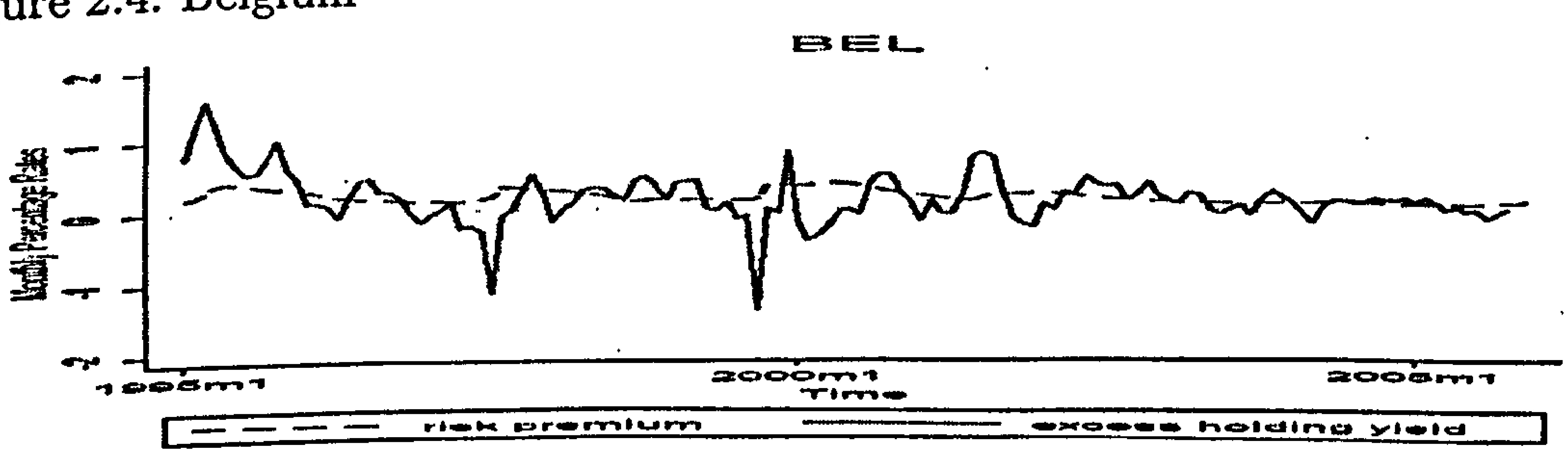


Figure 2.5: Brazil

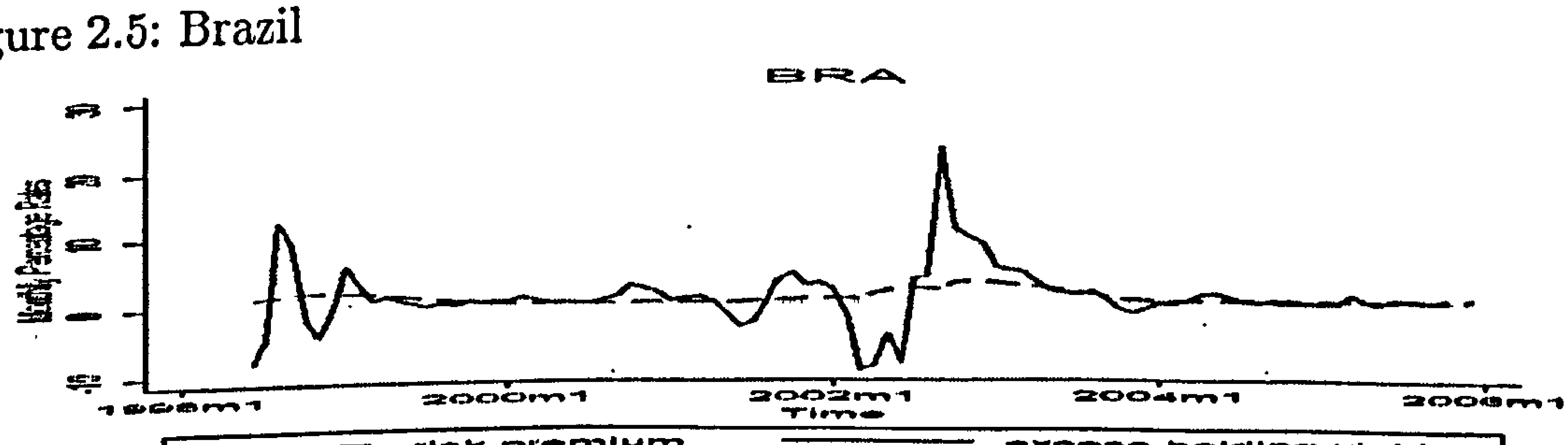


Figure 2.6: Canada

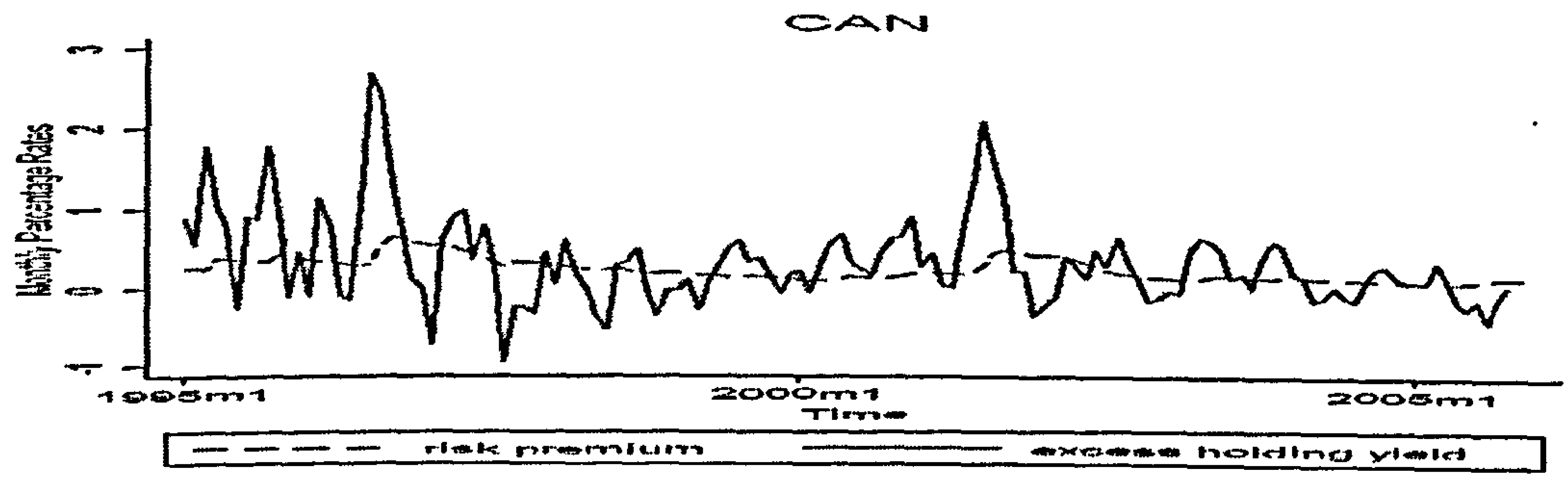


Figure 2.7: Switzerland

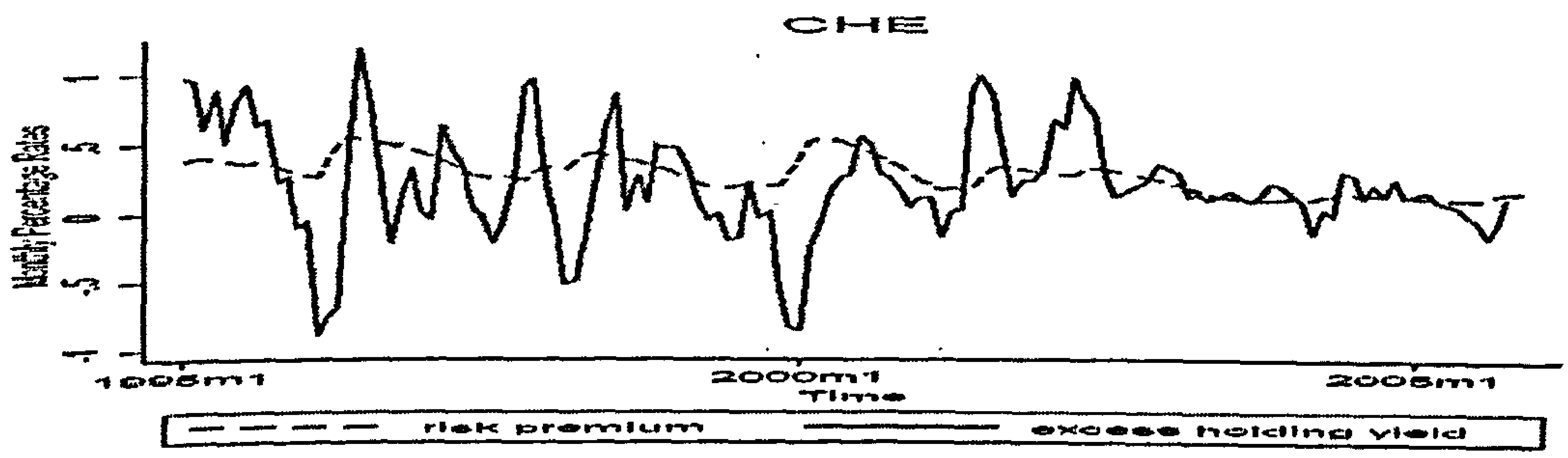


Figure 2.8: China

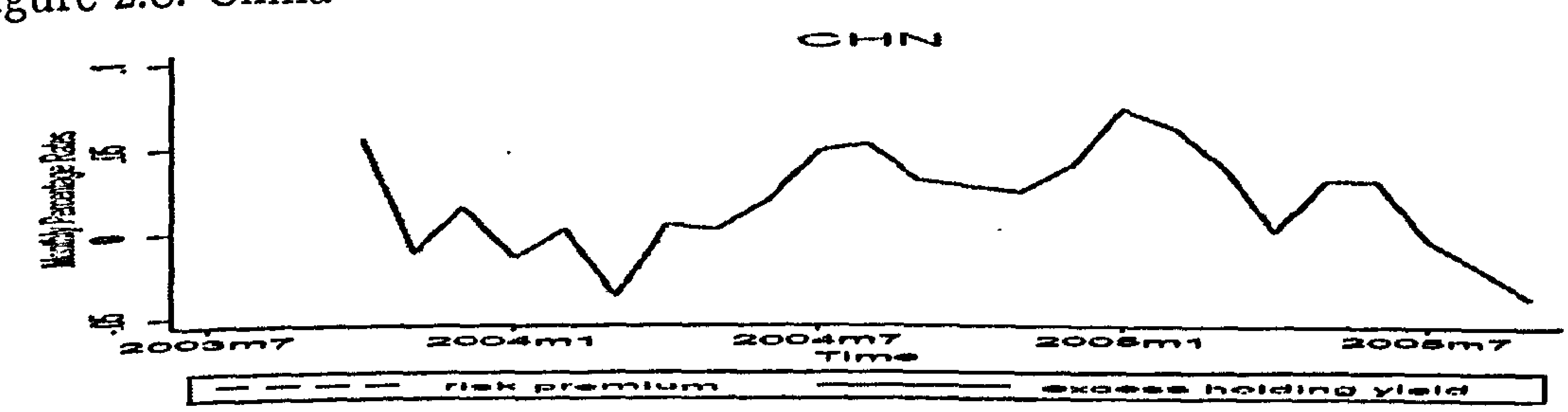


Figure 2.9: Colombia

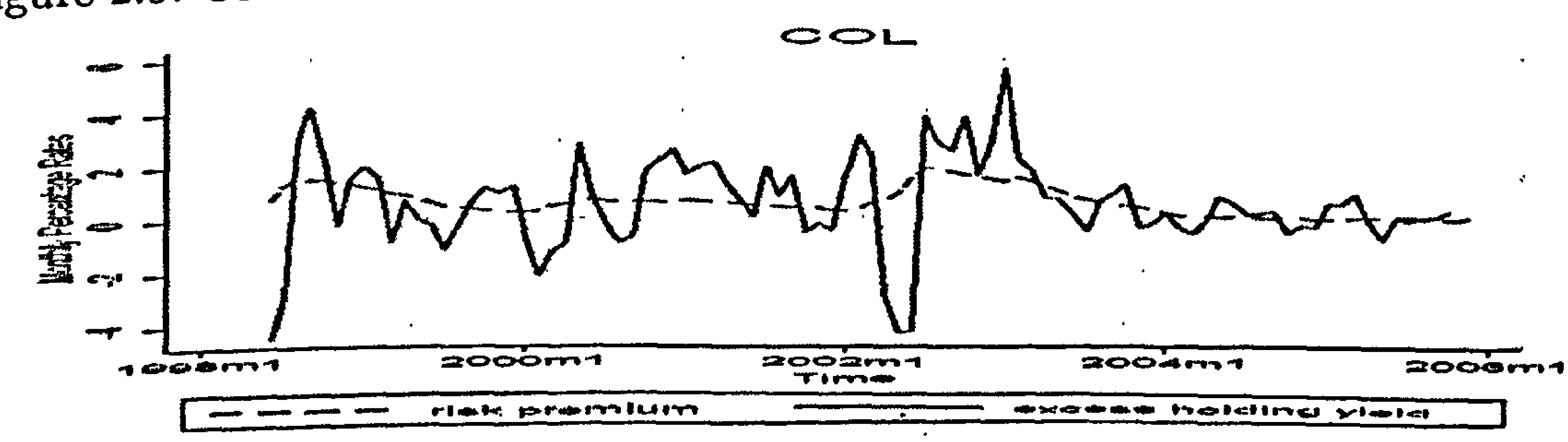


Figure 2.10: Czech Republic

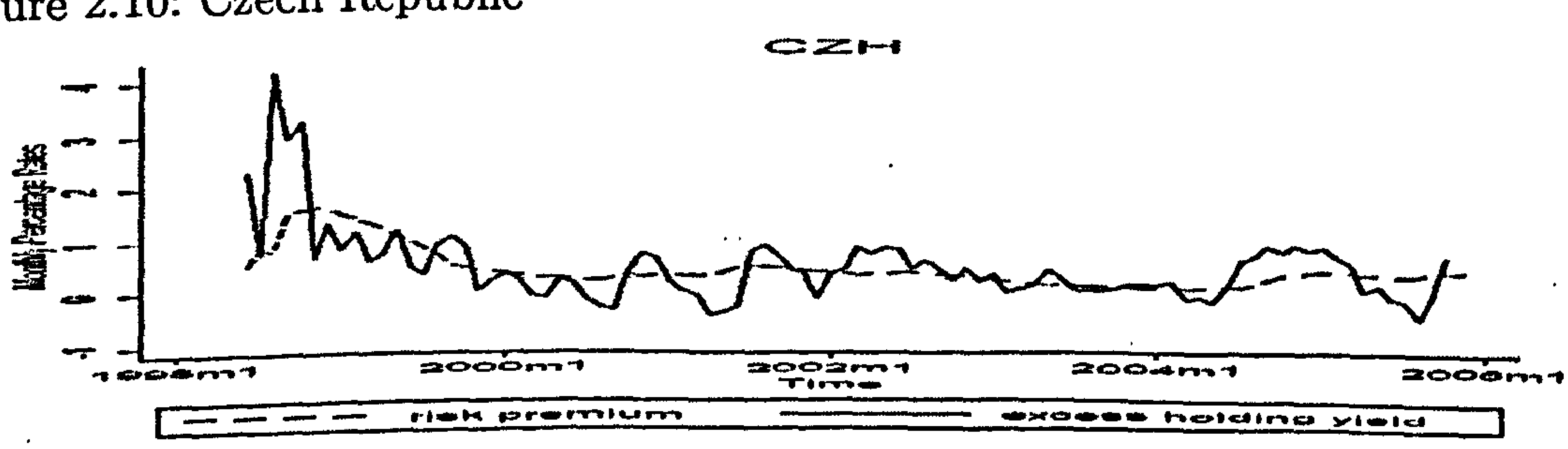


Figure 2.11: Germany

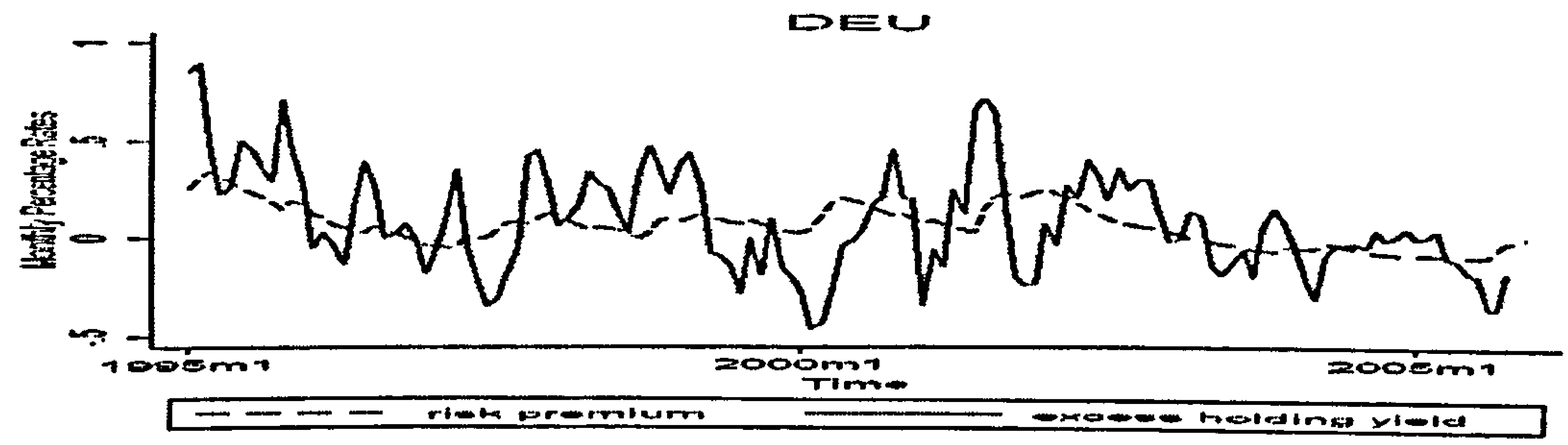


Figure 2.12: Denmark

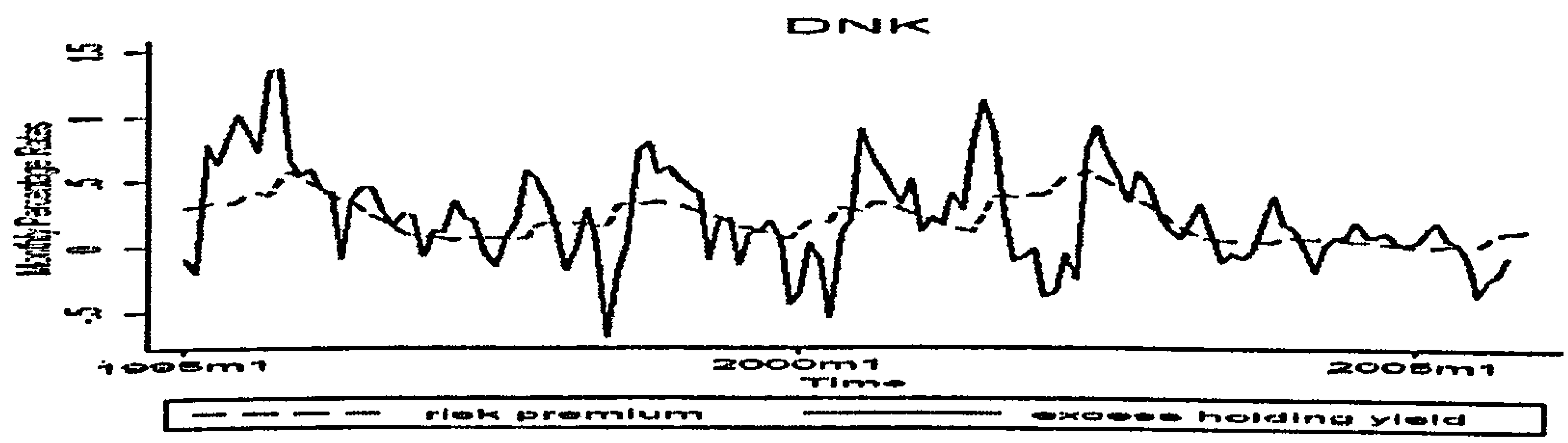


Figure 2.13: Spain

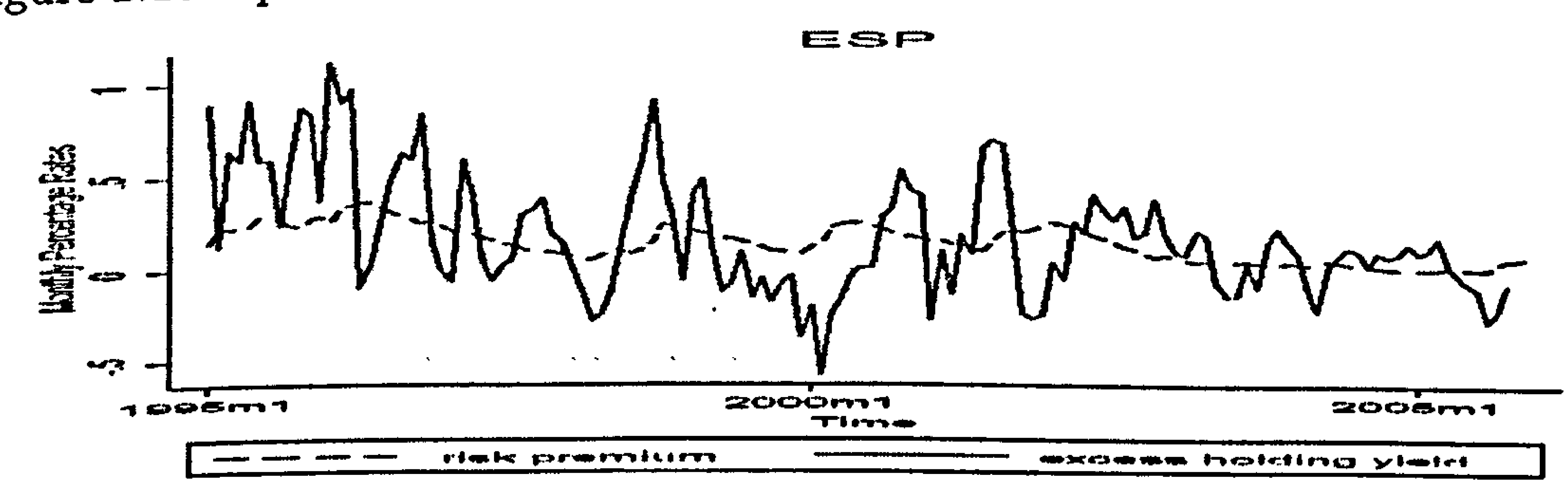


Figure 2.14: Finland

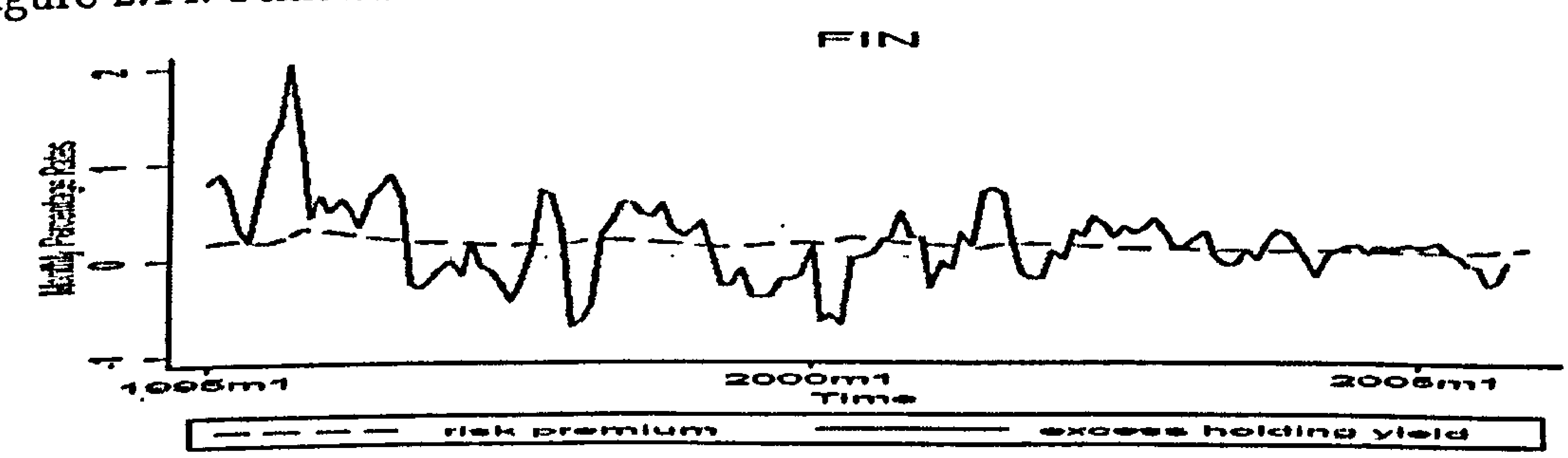


Figure 2.15: France

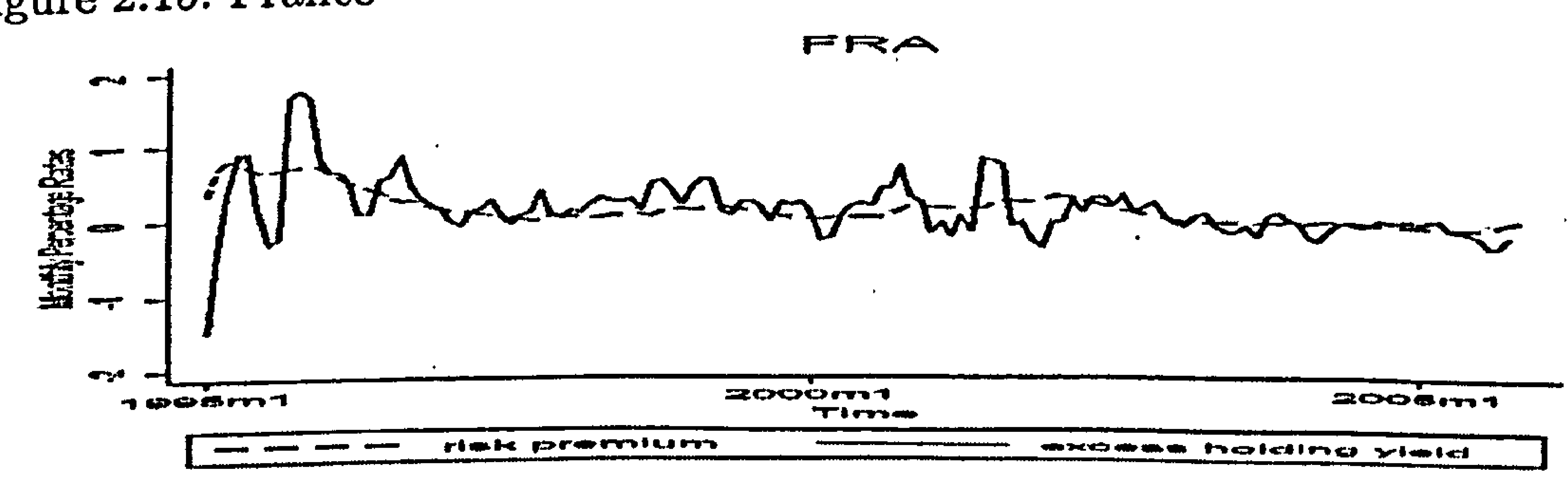


Figure 2.16: Greece

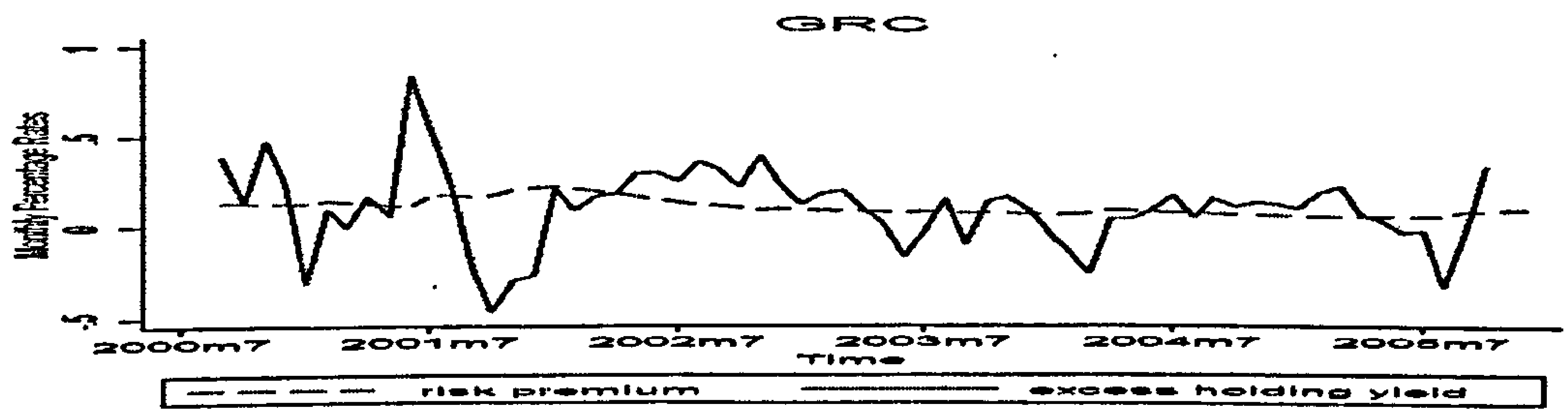


Figure 2.17: Hong Kong

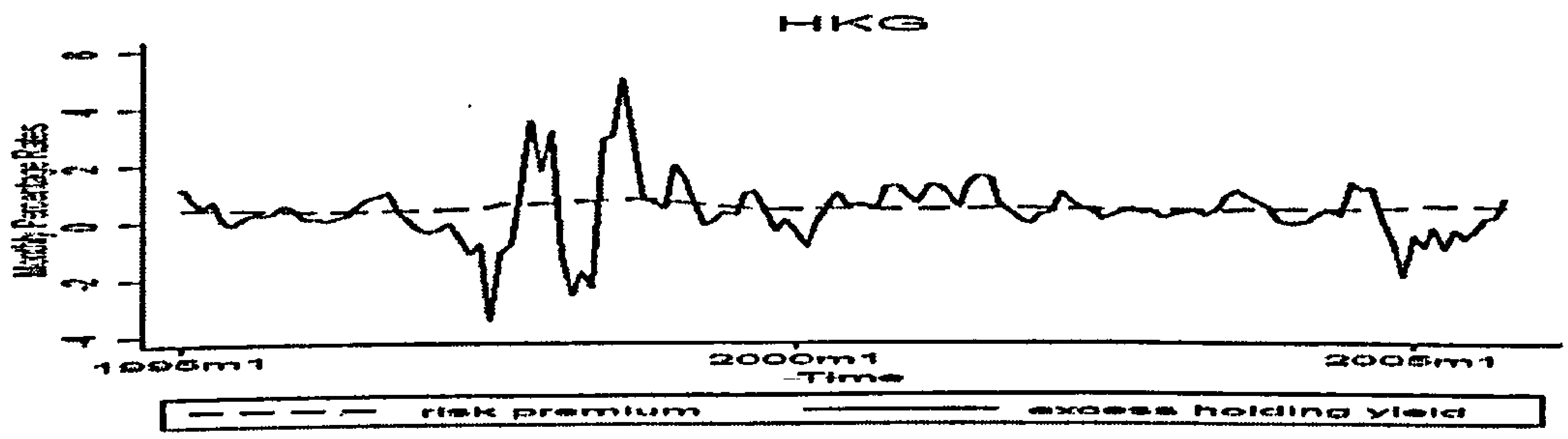


Figure 2.18: Hungary

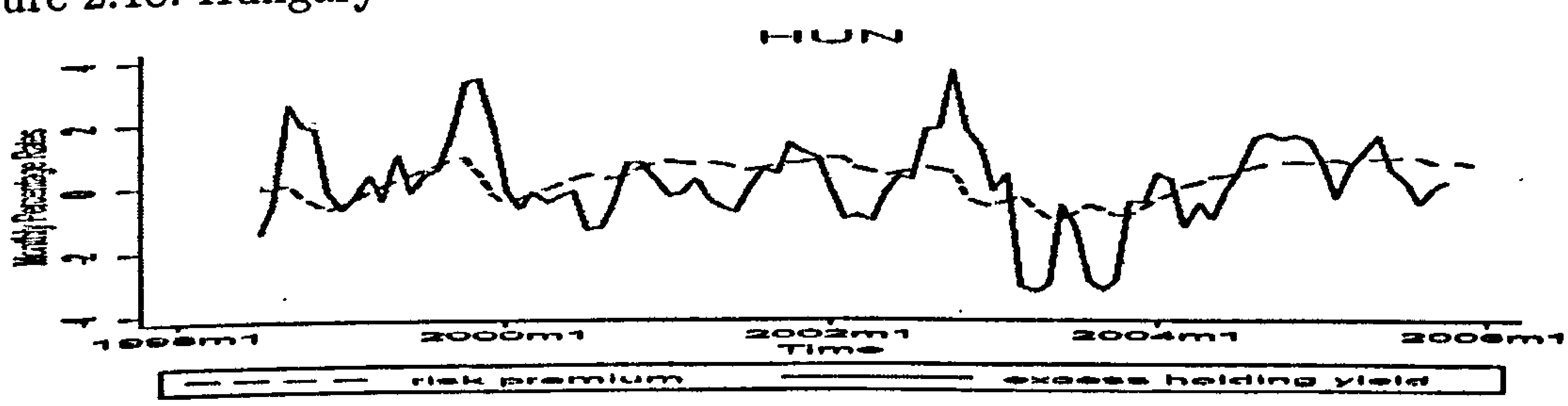


Figure 2.19: Indonesia

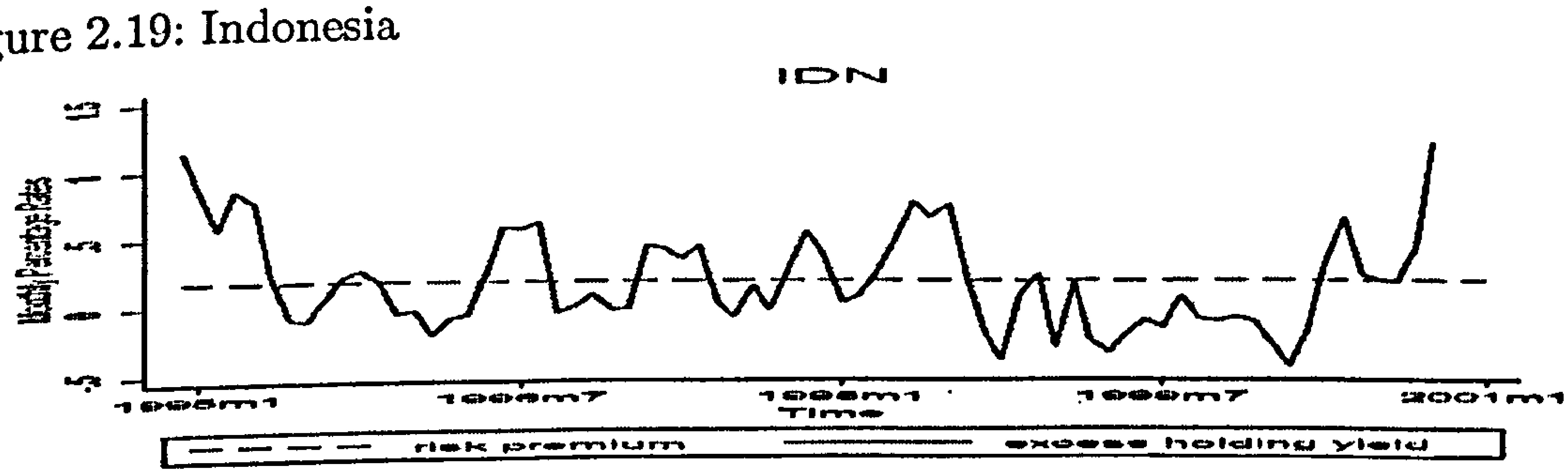


Figure 2.20: India

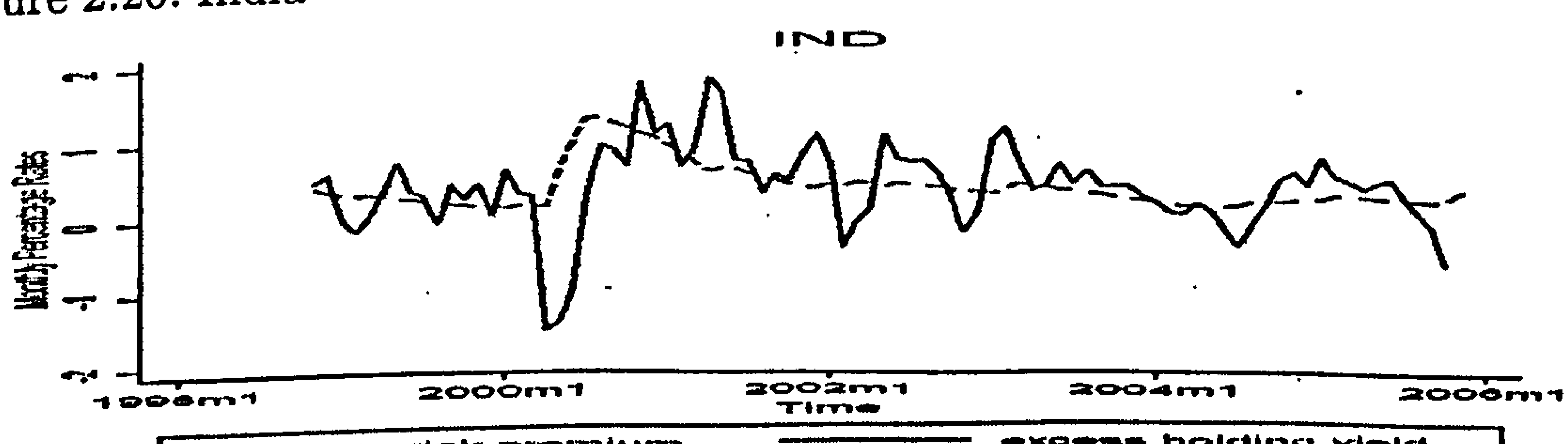


Figure 2.21: Ireland

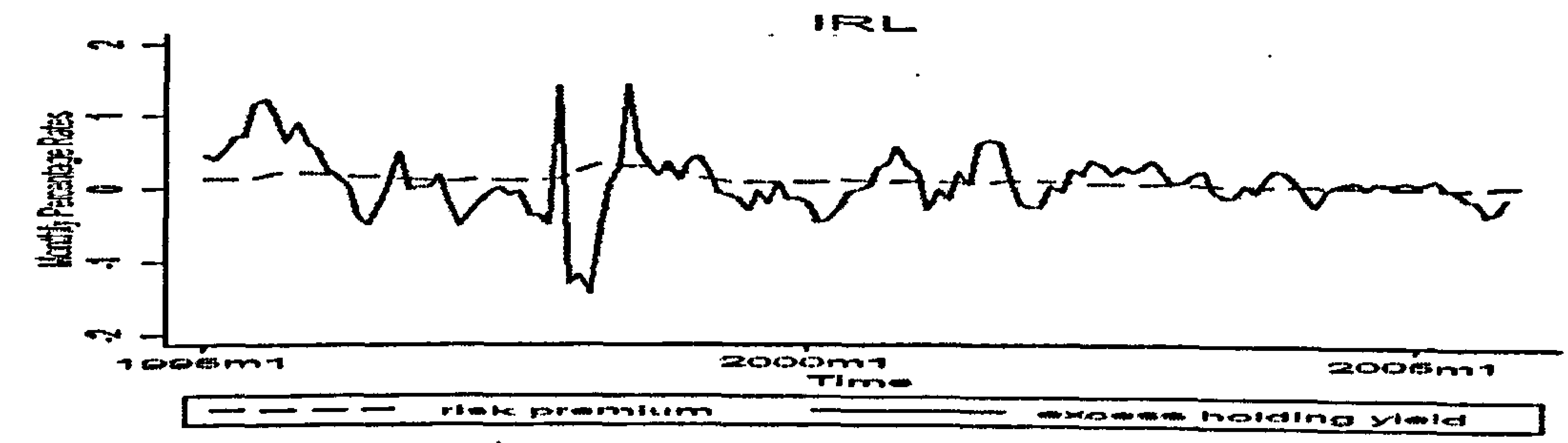


Figure 2.22: Israel

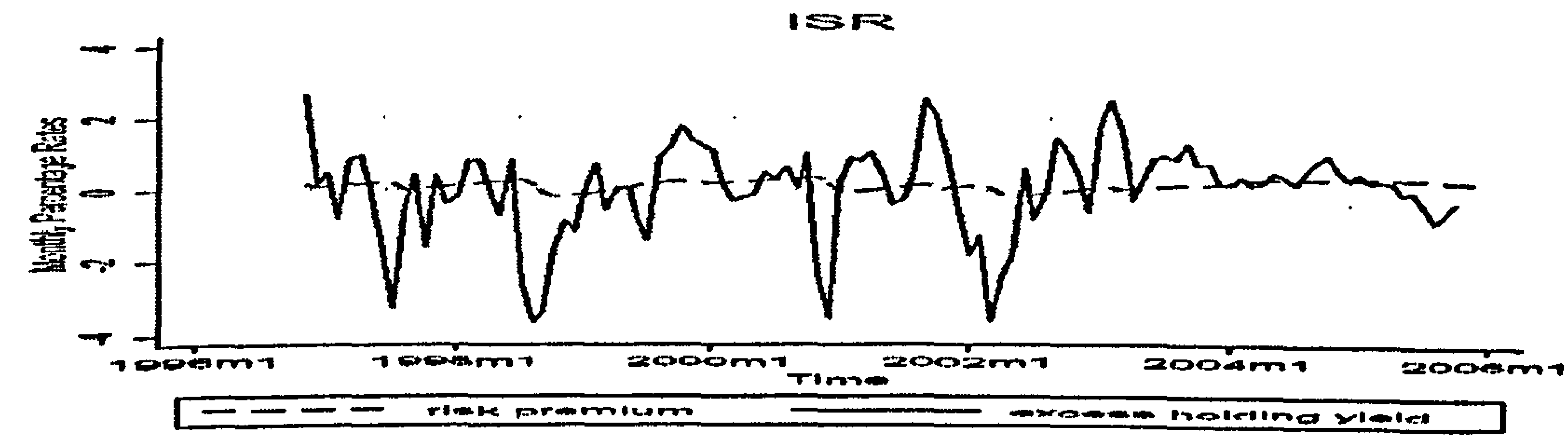


Figure 2.23: Italy

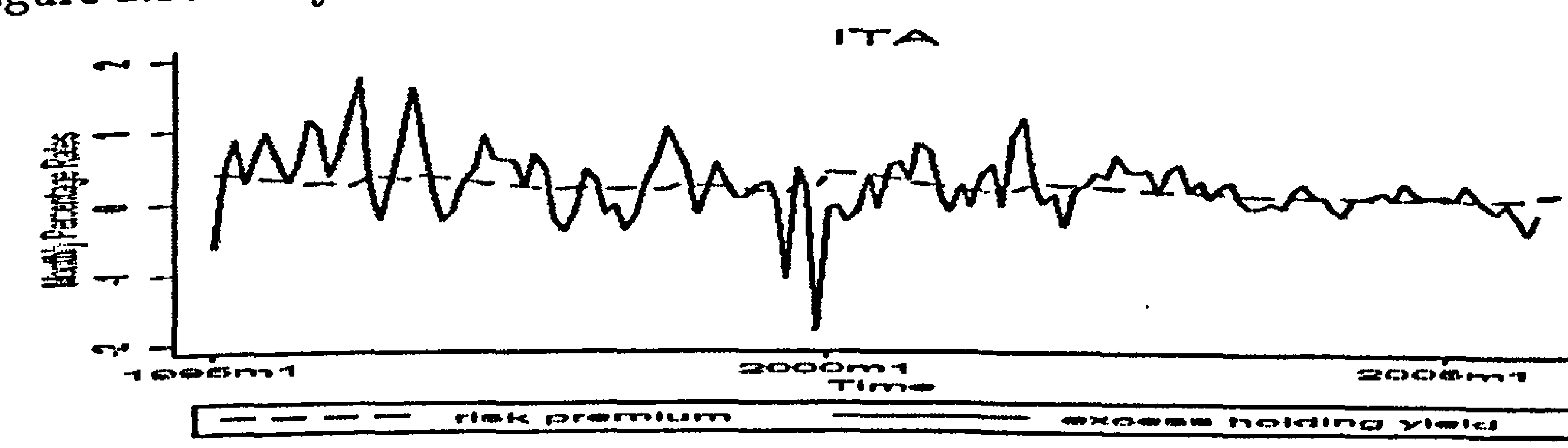


Figure 2.24: Japan

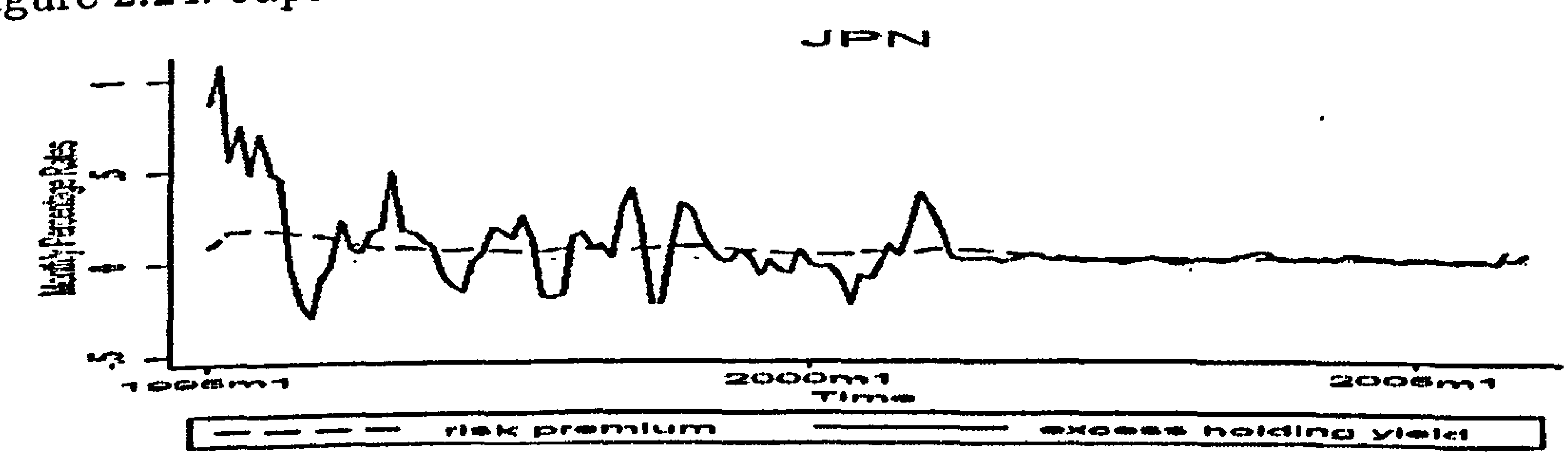


Figure 2.25: Korea

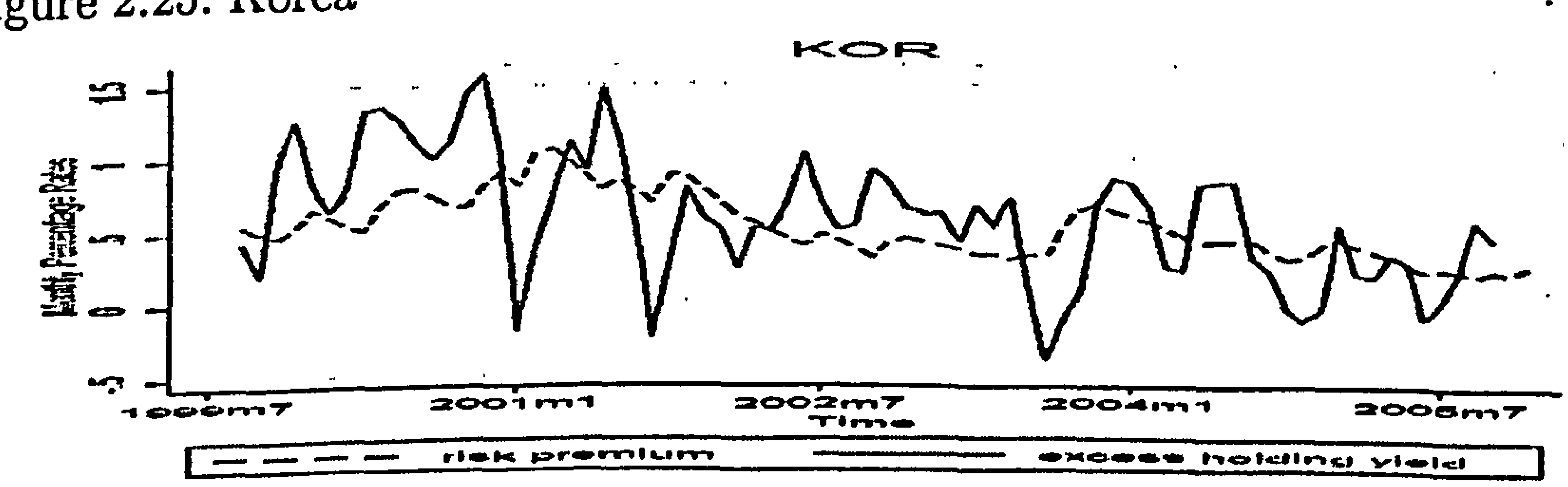


Figure 2.26: Sri Lanka

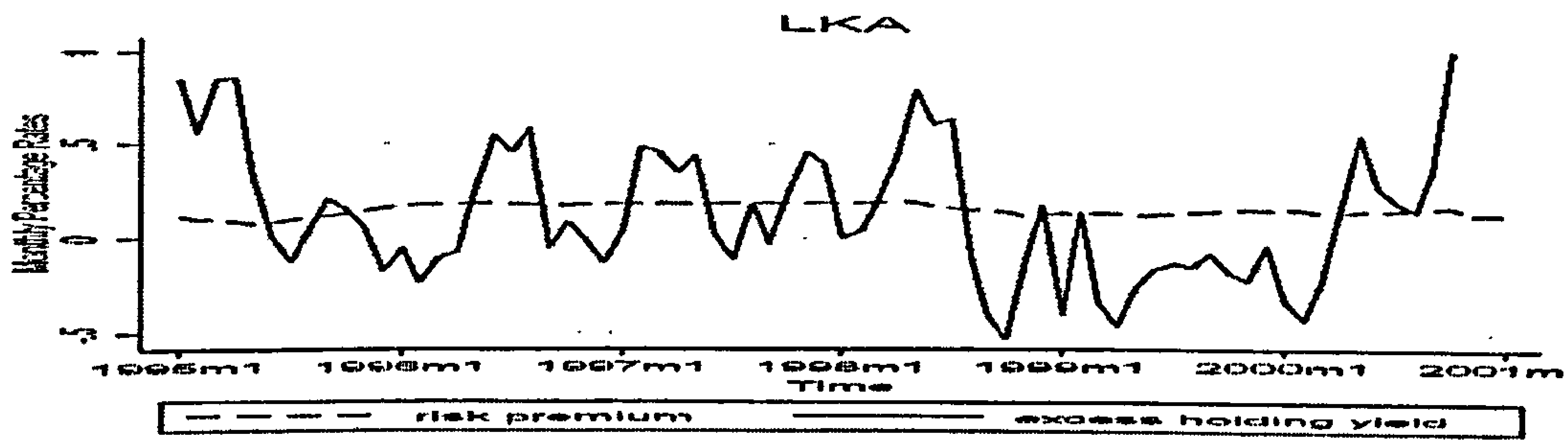


Figure 2.27: Mexico

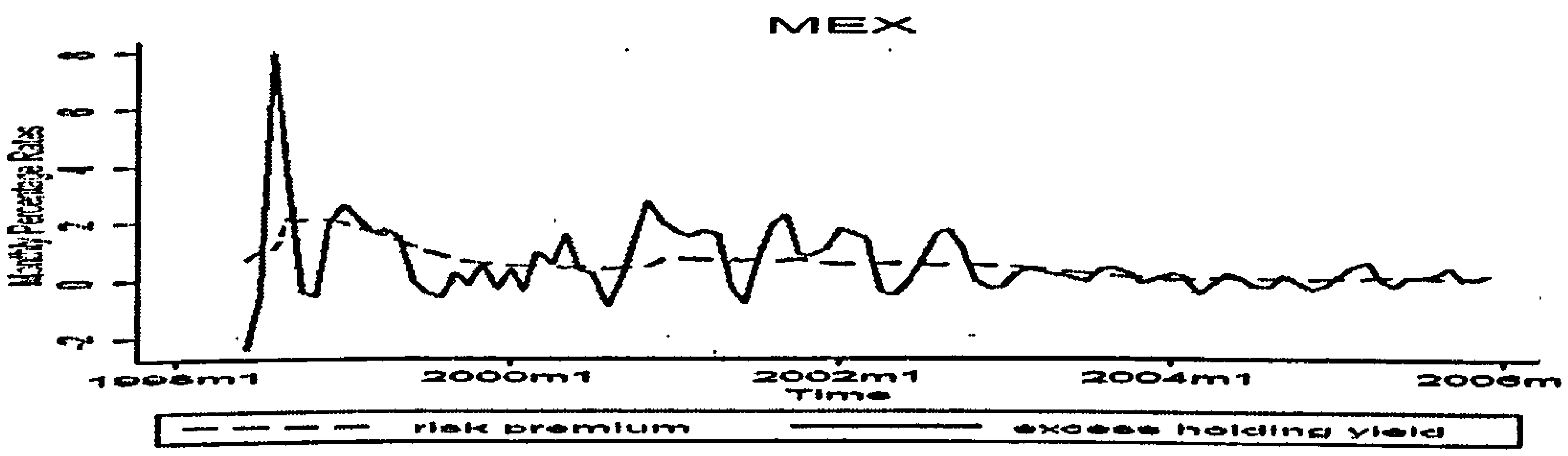


Figure 2.28: Malaysia

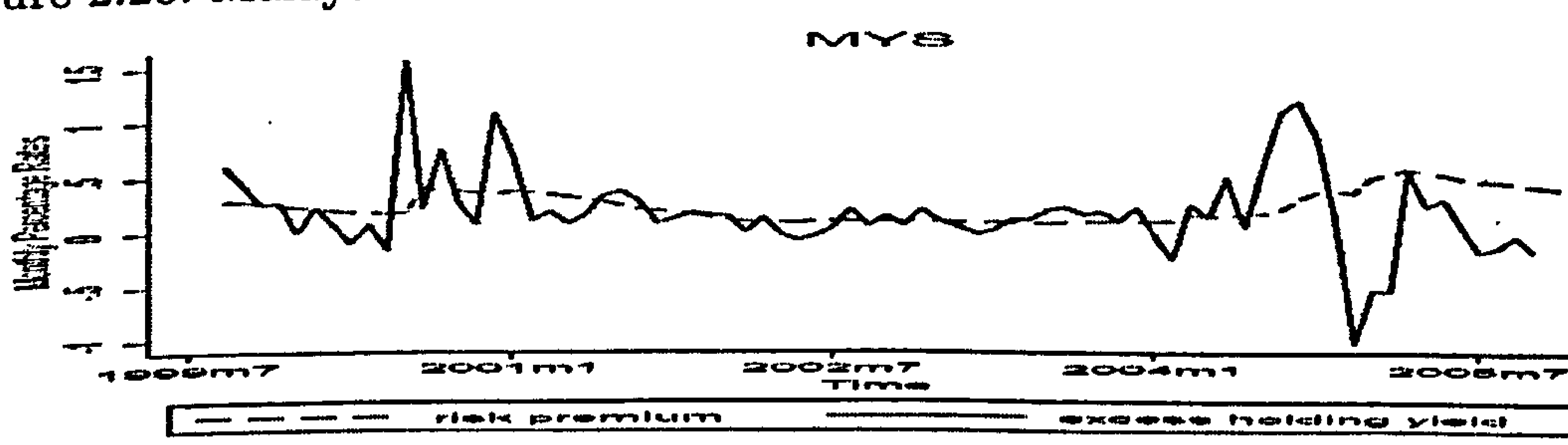


Figure 2.29: Netherlands

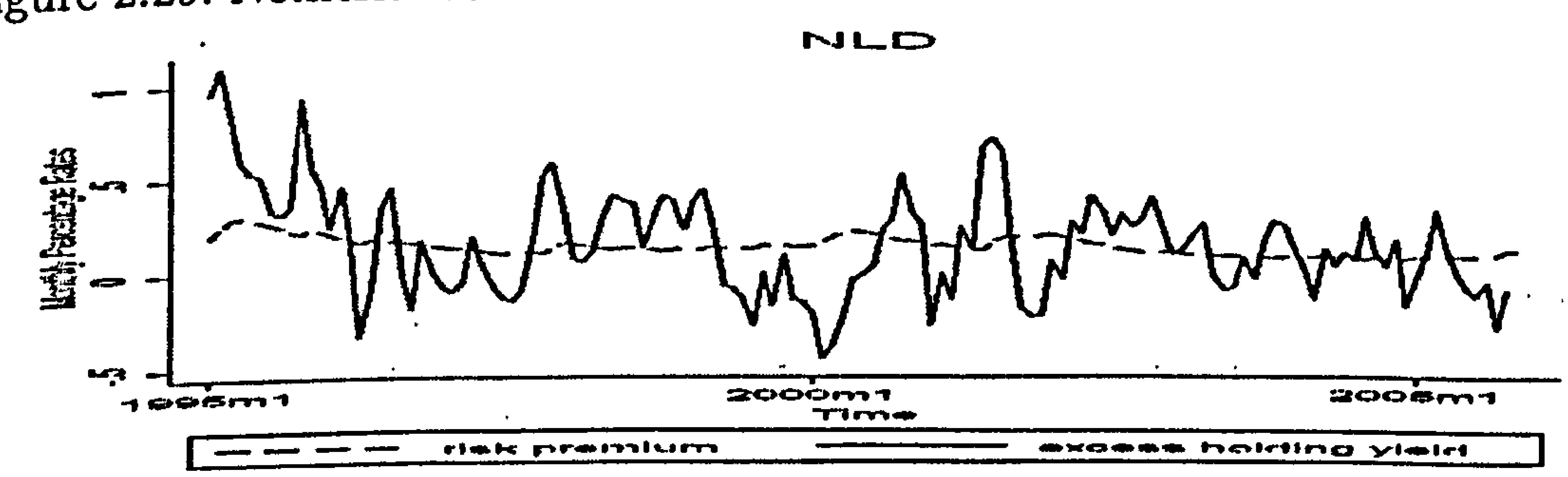


Figure 2.30: Norway

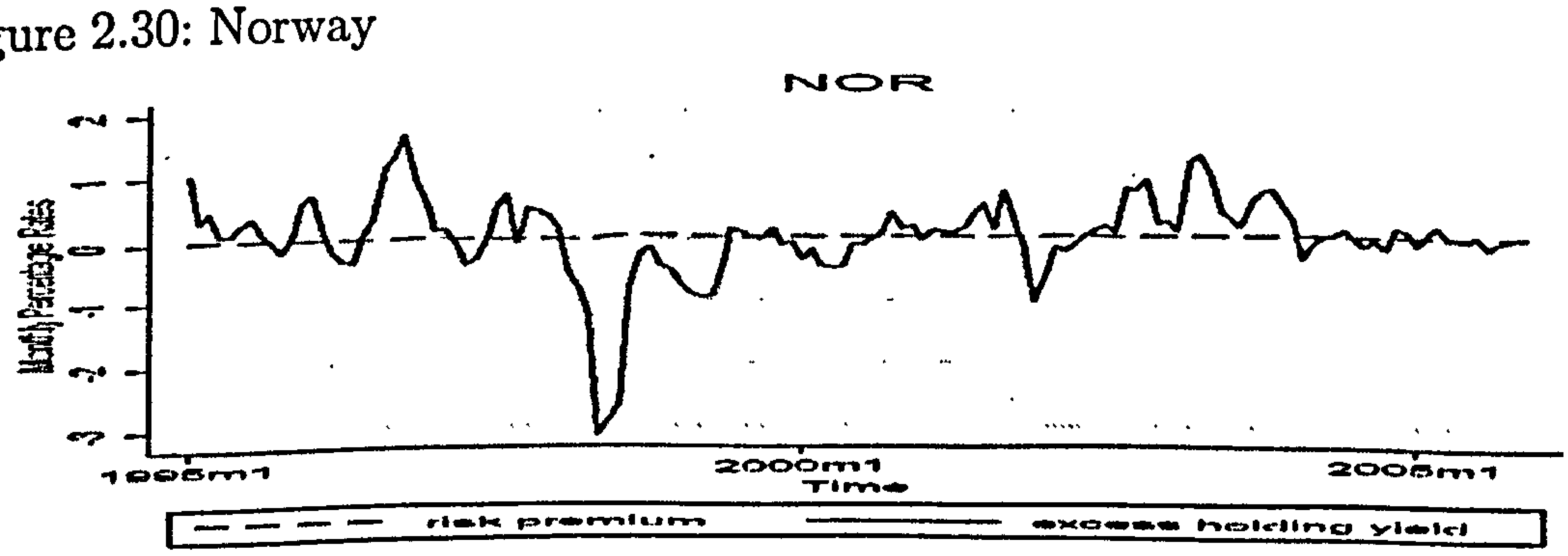


Figure 2.31: New Zealand

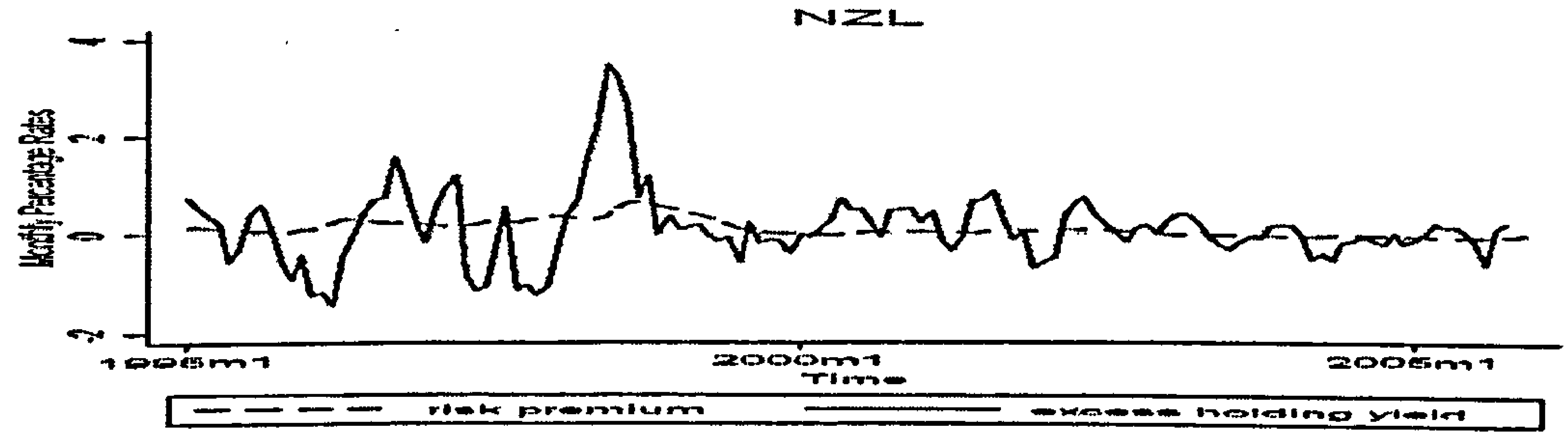


Figure 2.32: Philippines

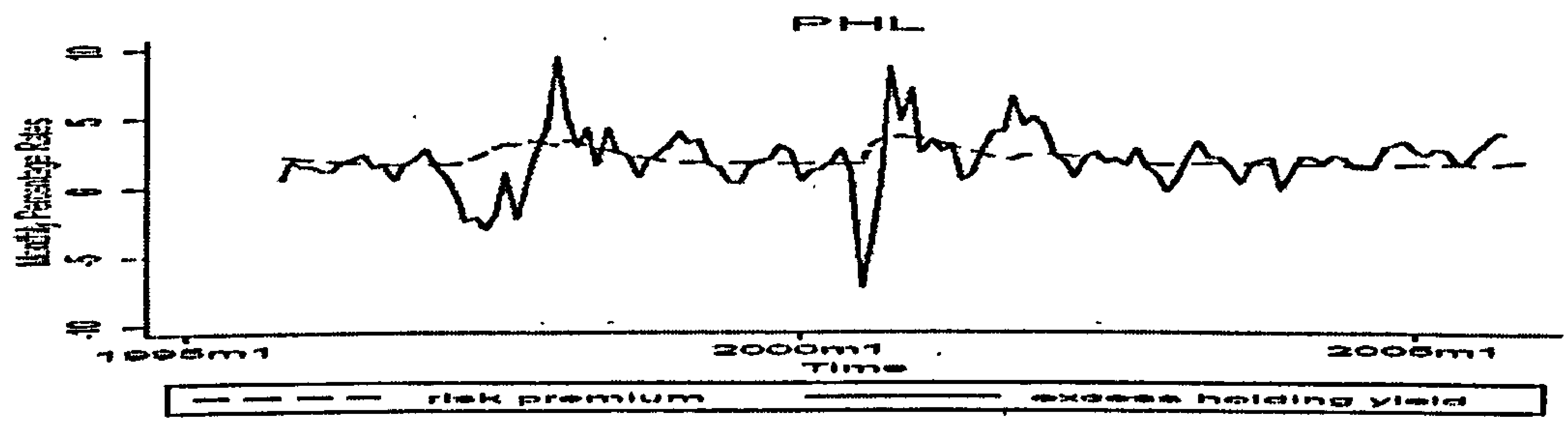


Figure 2.33: Poland

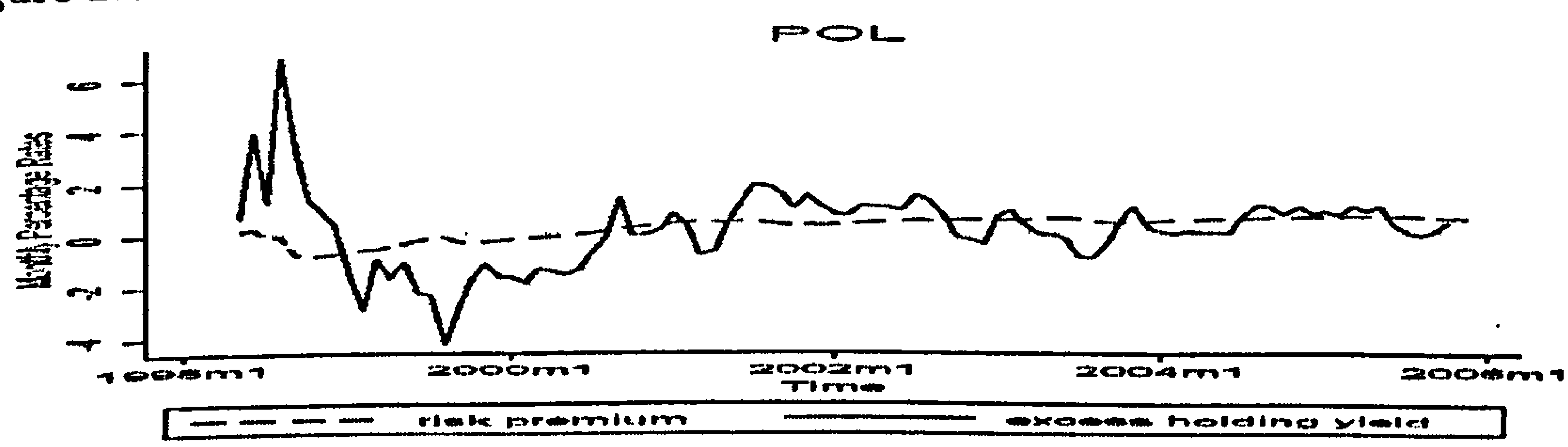


Figure 2.34: Portugal

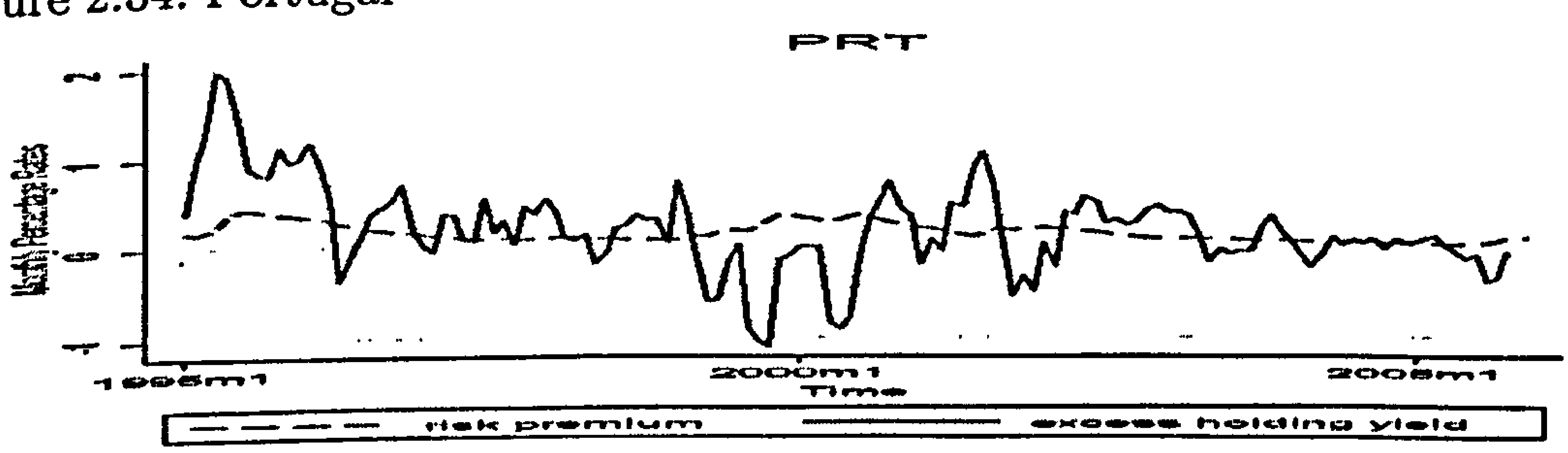


Figure 2.35: Singapore

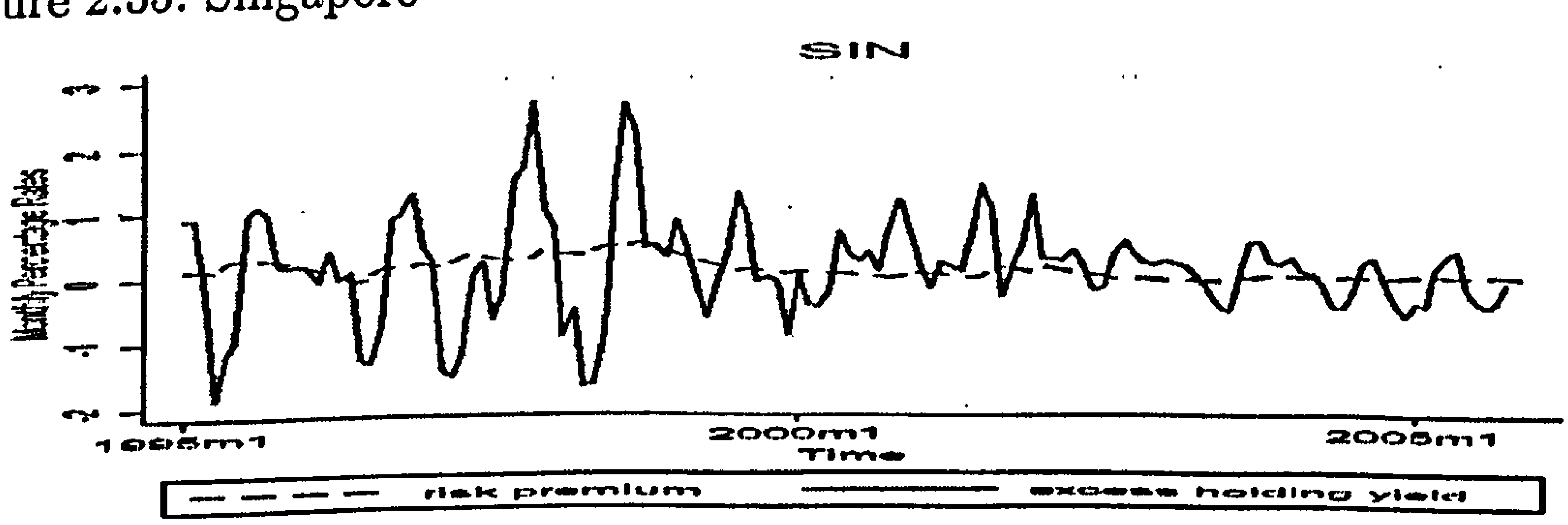


Figure 2.36: Slovak Republic

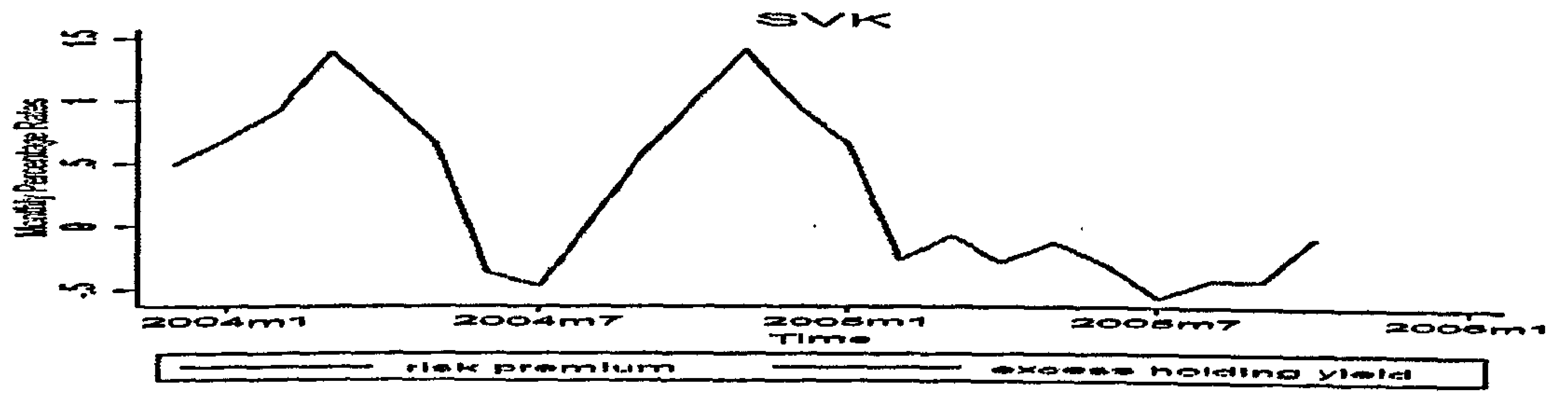


Figure 2.37: Sweden

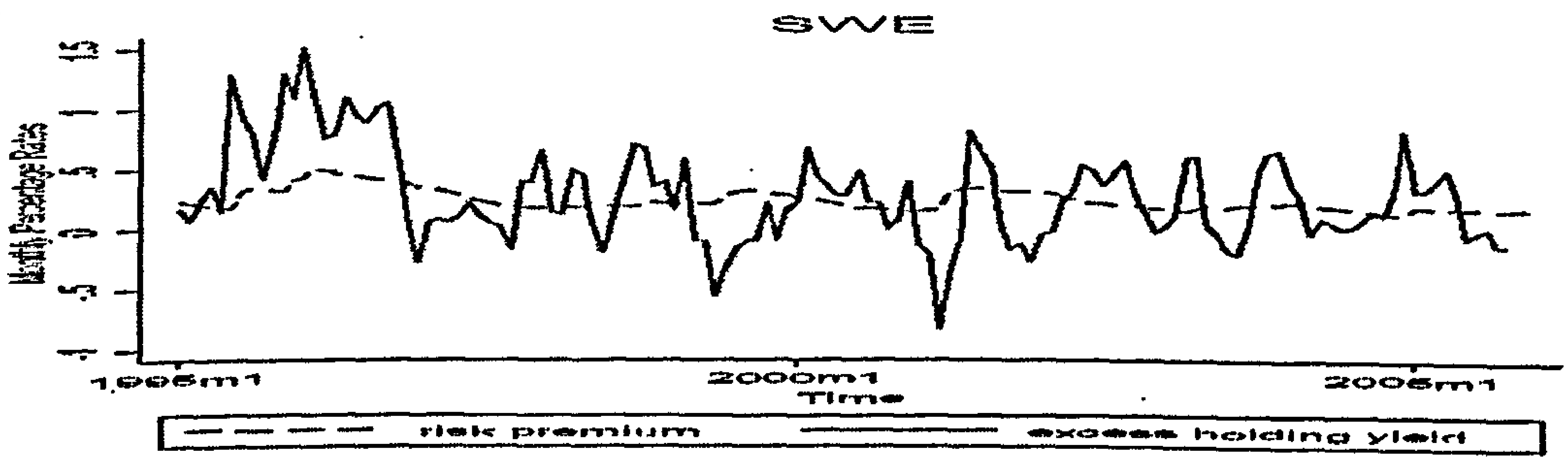


Figure 2.38: Thailand

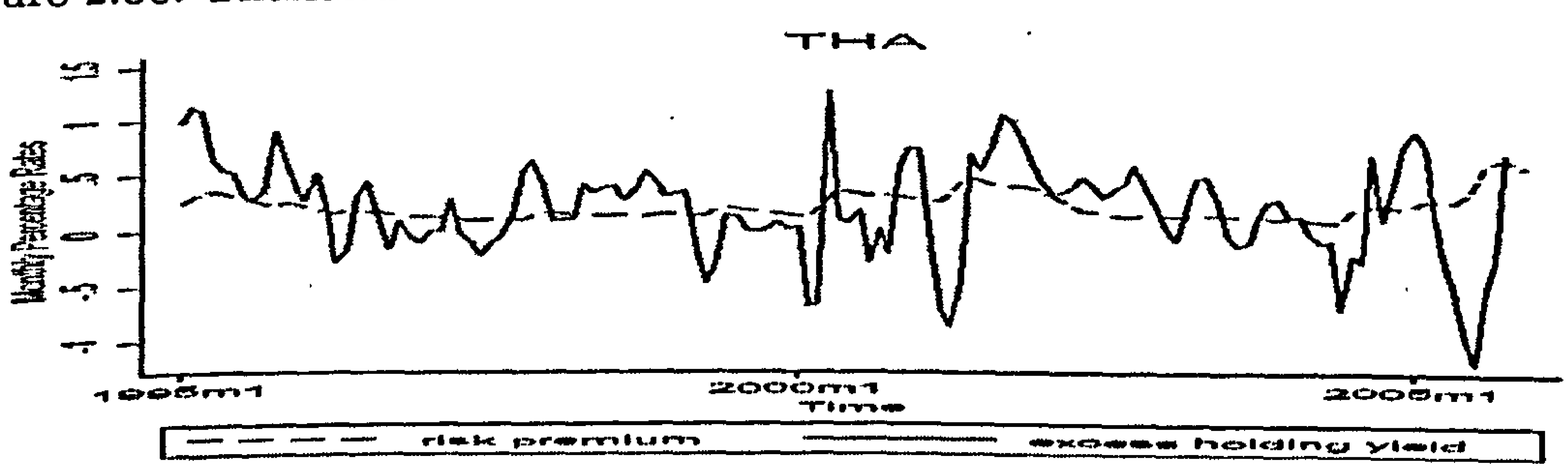


Figure 2.39: Turkey

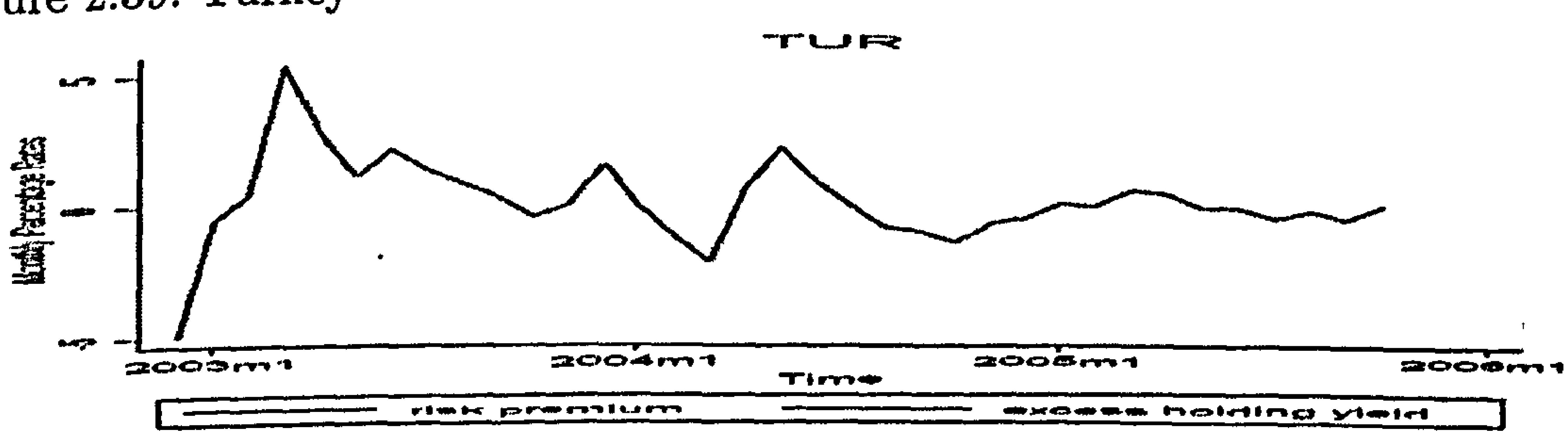


Figure 2.40: United Kingdom

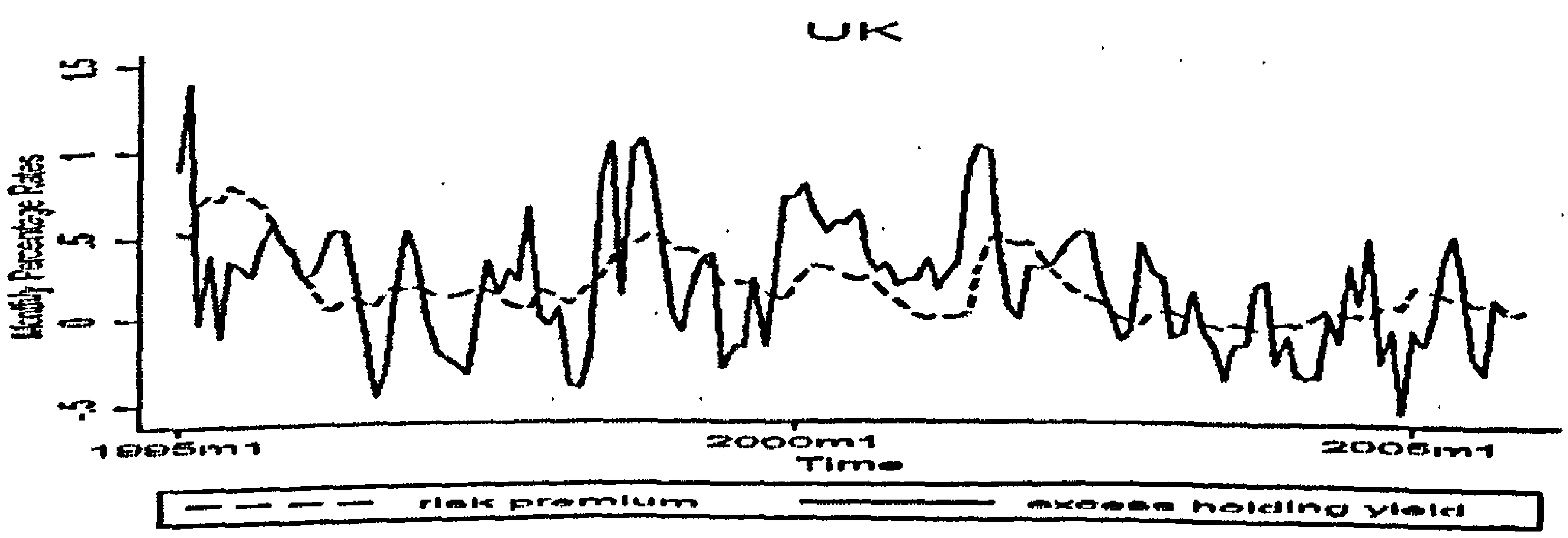


Figure 2.41: Uruguay

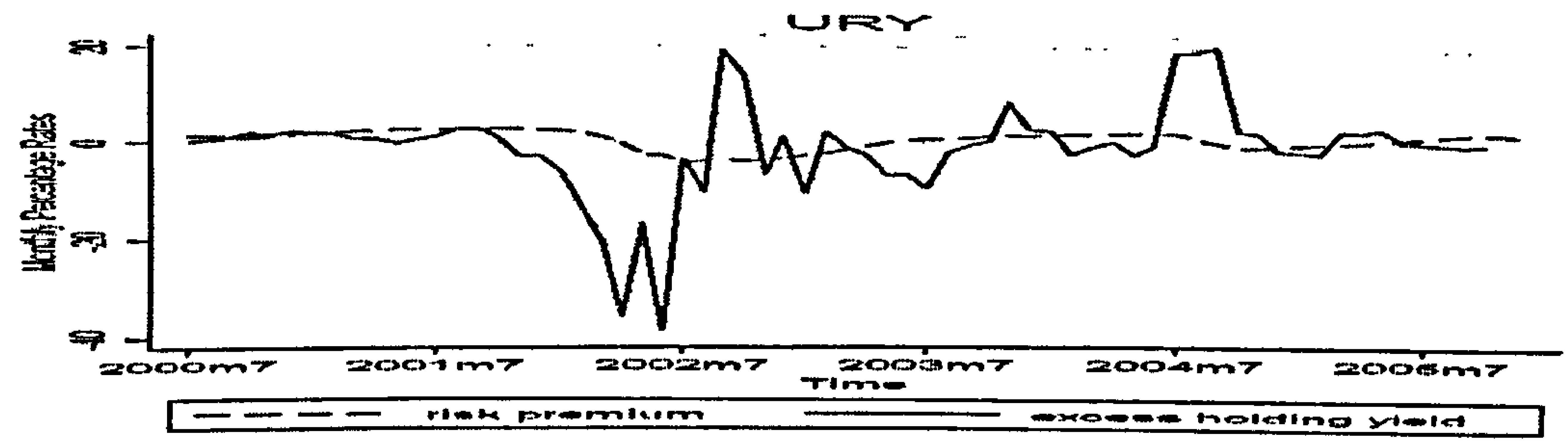


Figure 2.42: United States of America

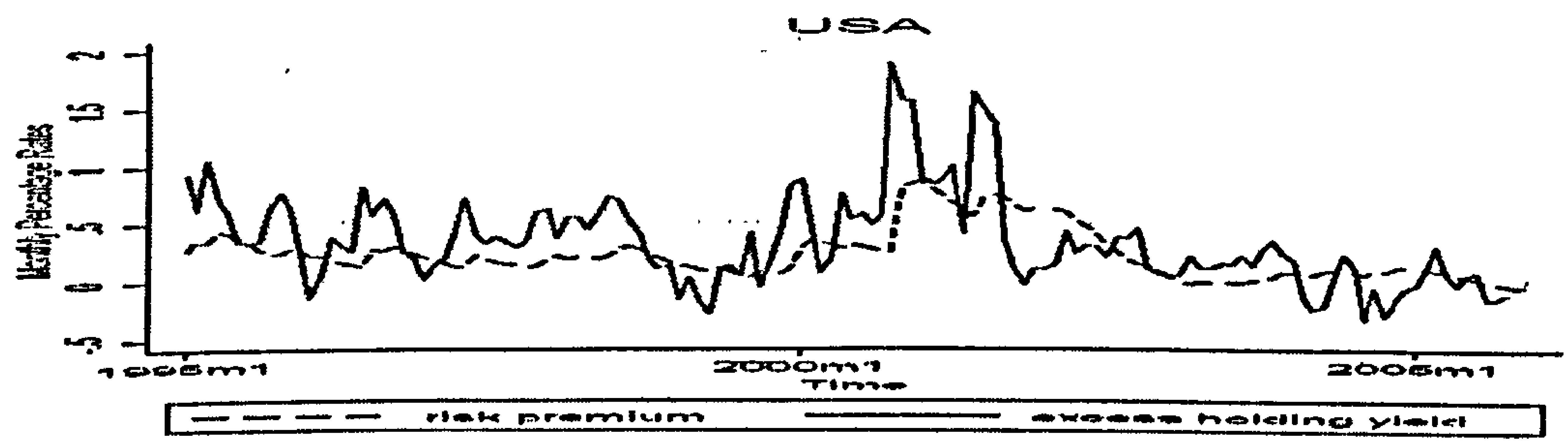


Figure 2.43: South Africa

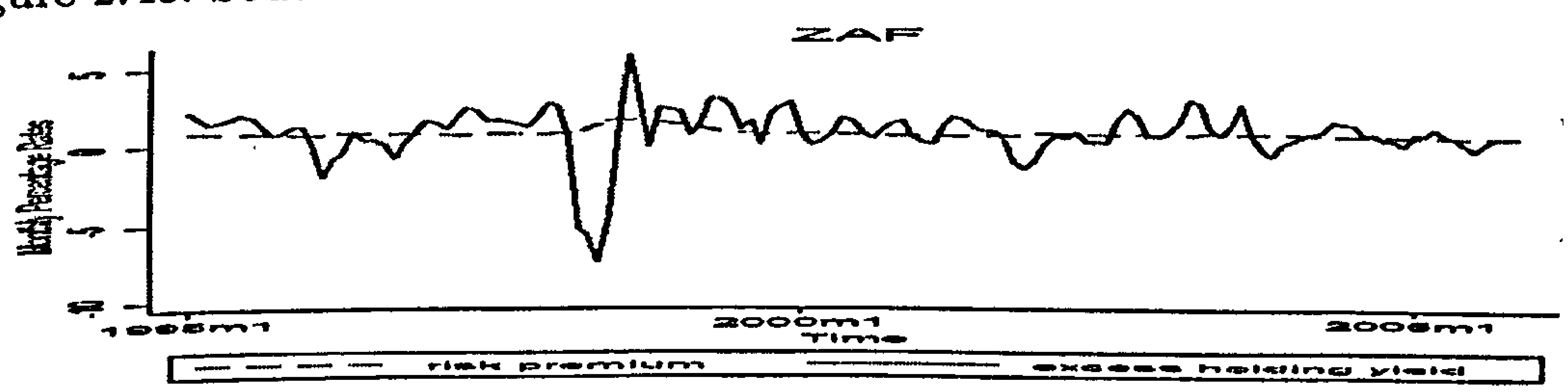


Figure 2.44: Average risk premium for 3 months versus 6 months treasury bills (1994-2006).

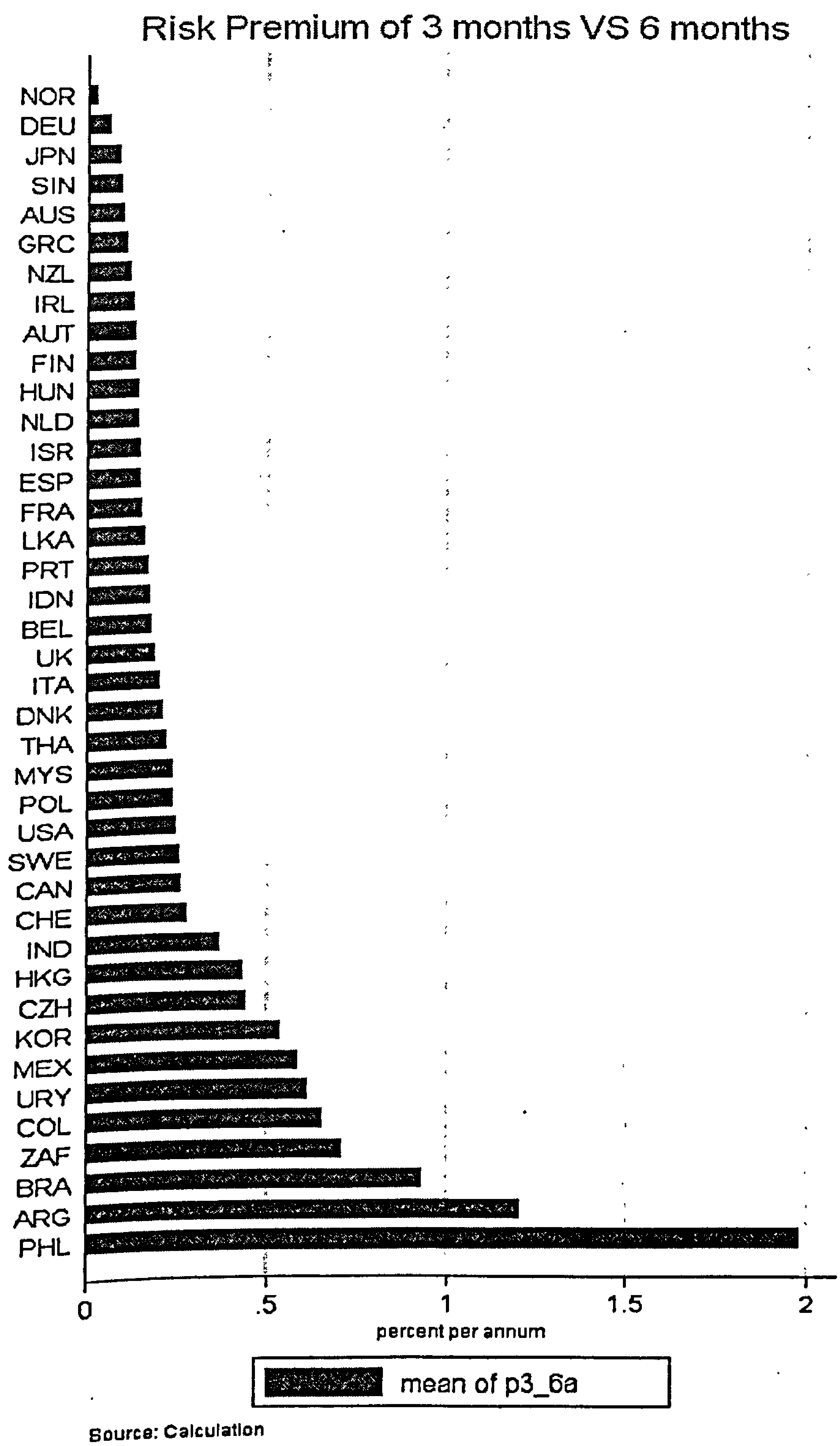


Figure 2.45: Average risk premium for 3 months versus 12 months treasury bills.

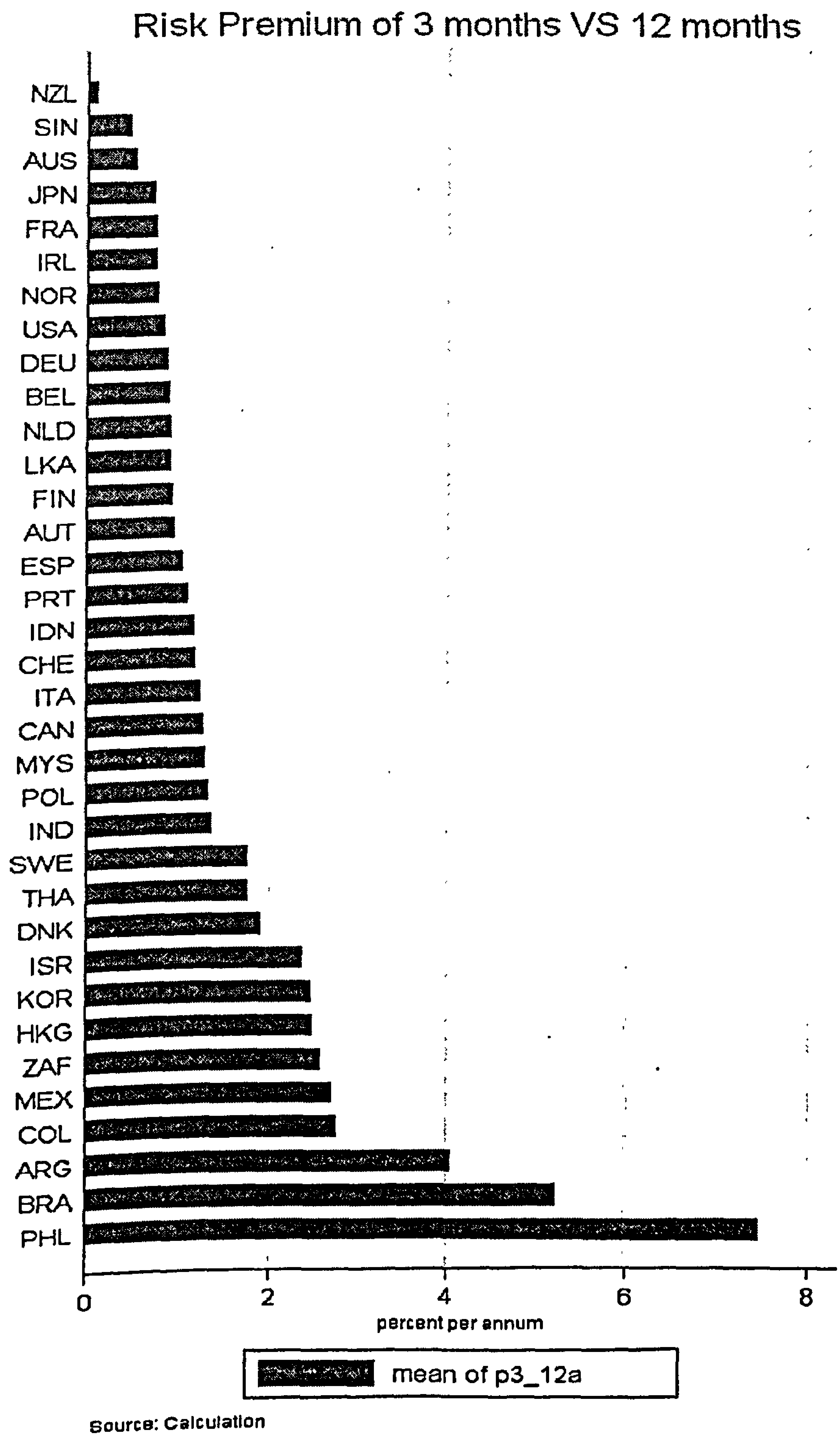
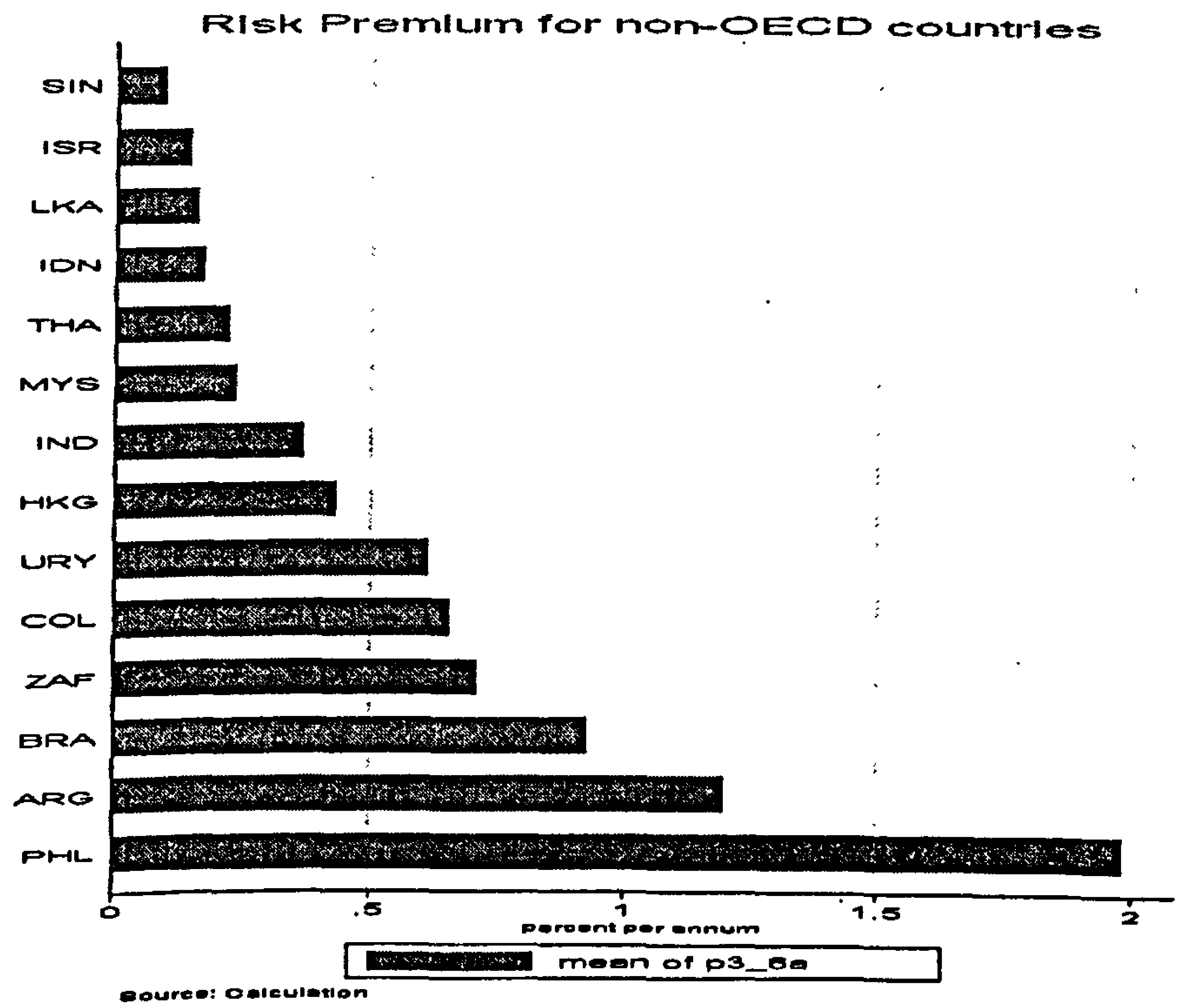


Figure 2.46: Risk premia for 3 months versus 6 months treasury bills: comparisons by country groups

A: Risk premia for non-OECD countries



B: Risk premia for OECD countries

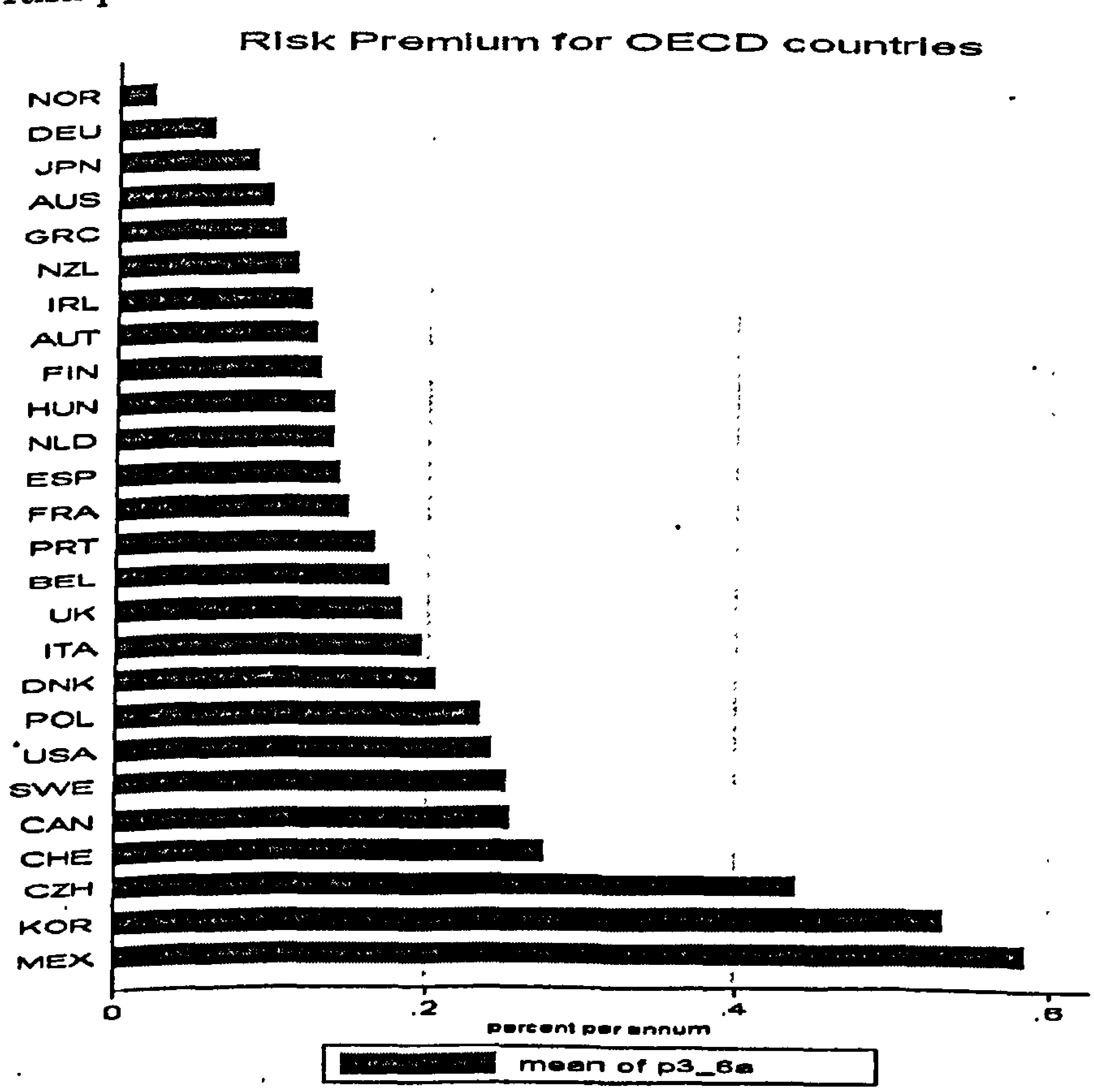
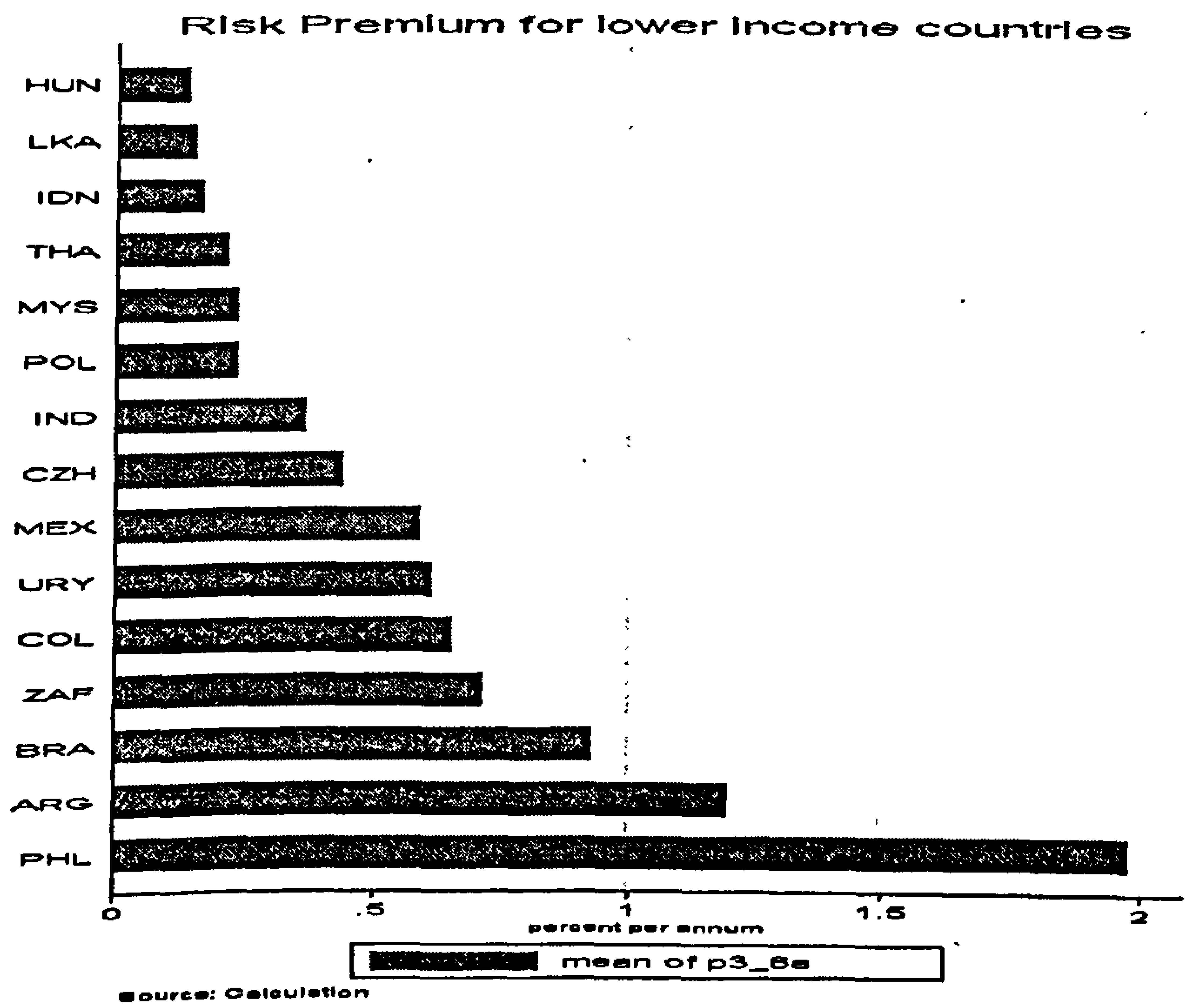


Figure 2.47: Risk premia for 3 months versus 6 months treasury bills: comparisons by income

A: Risk premia for lower income countries



B: Risk premia for high income countries

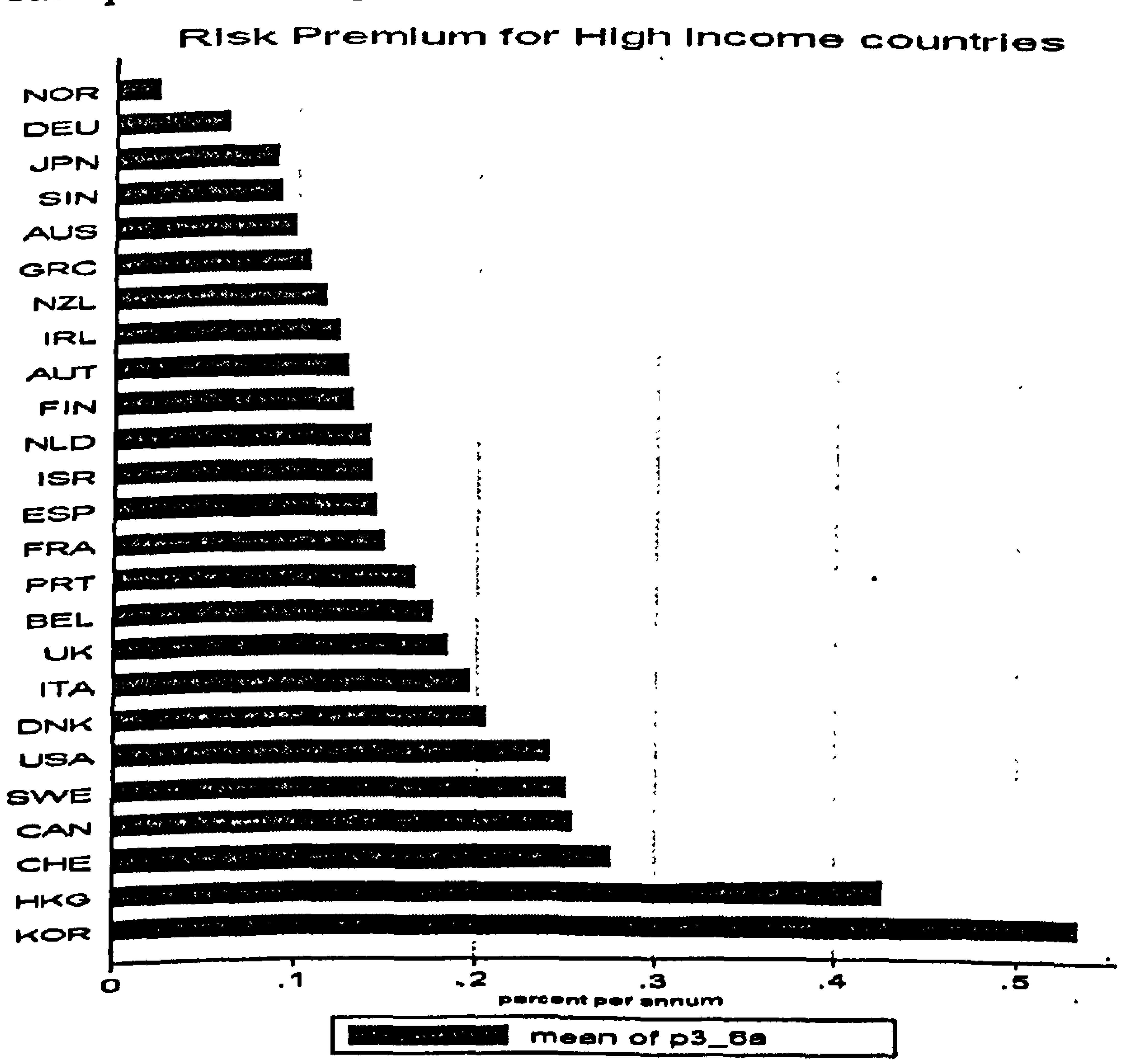


Figure 2.48: Variable plots, average 1994-2006

Figure 2.48A: Initial level of income

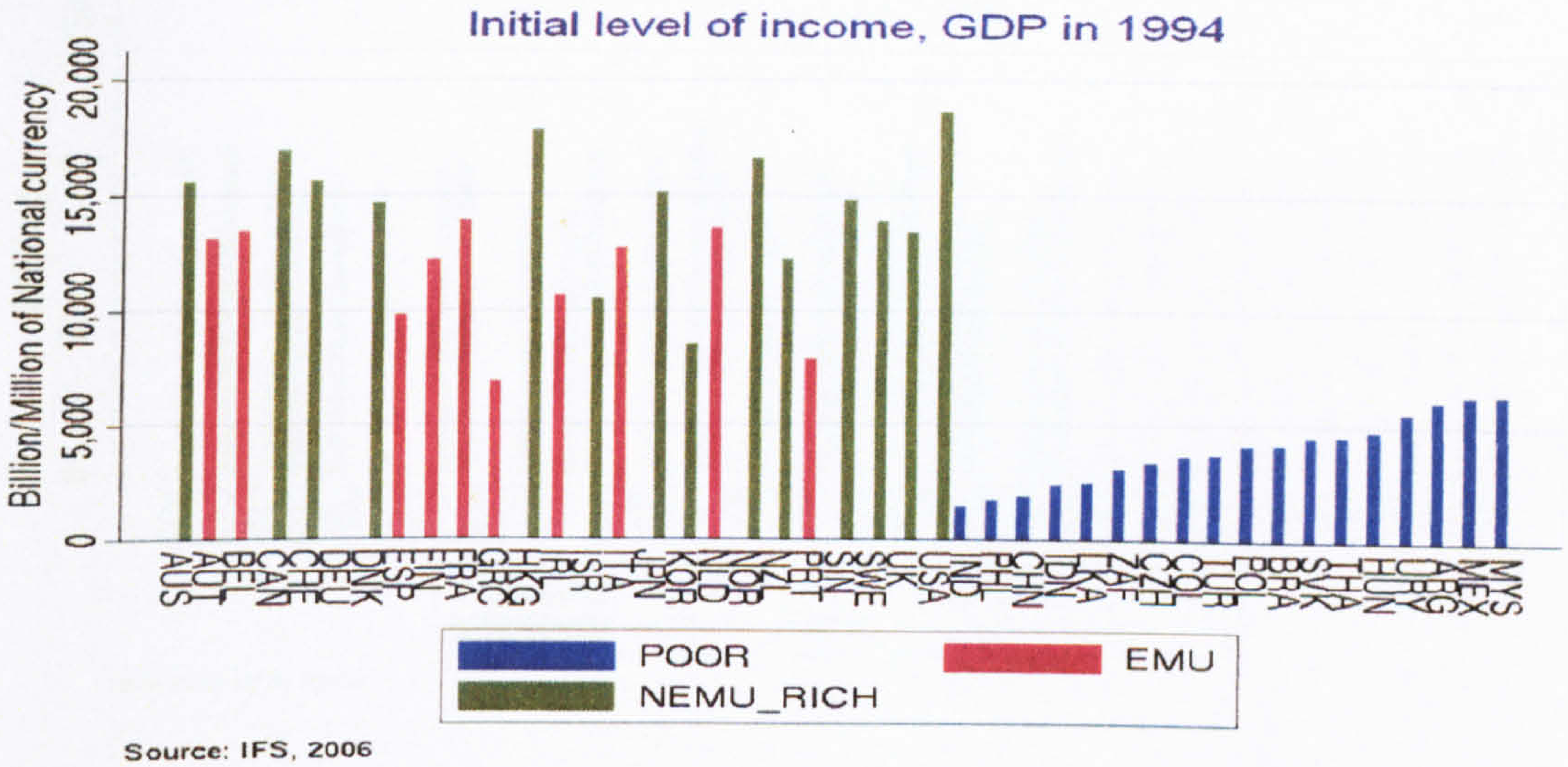


Figure 2.48B: Average economic growth over 1994-2006.

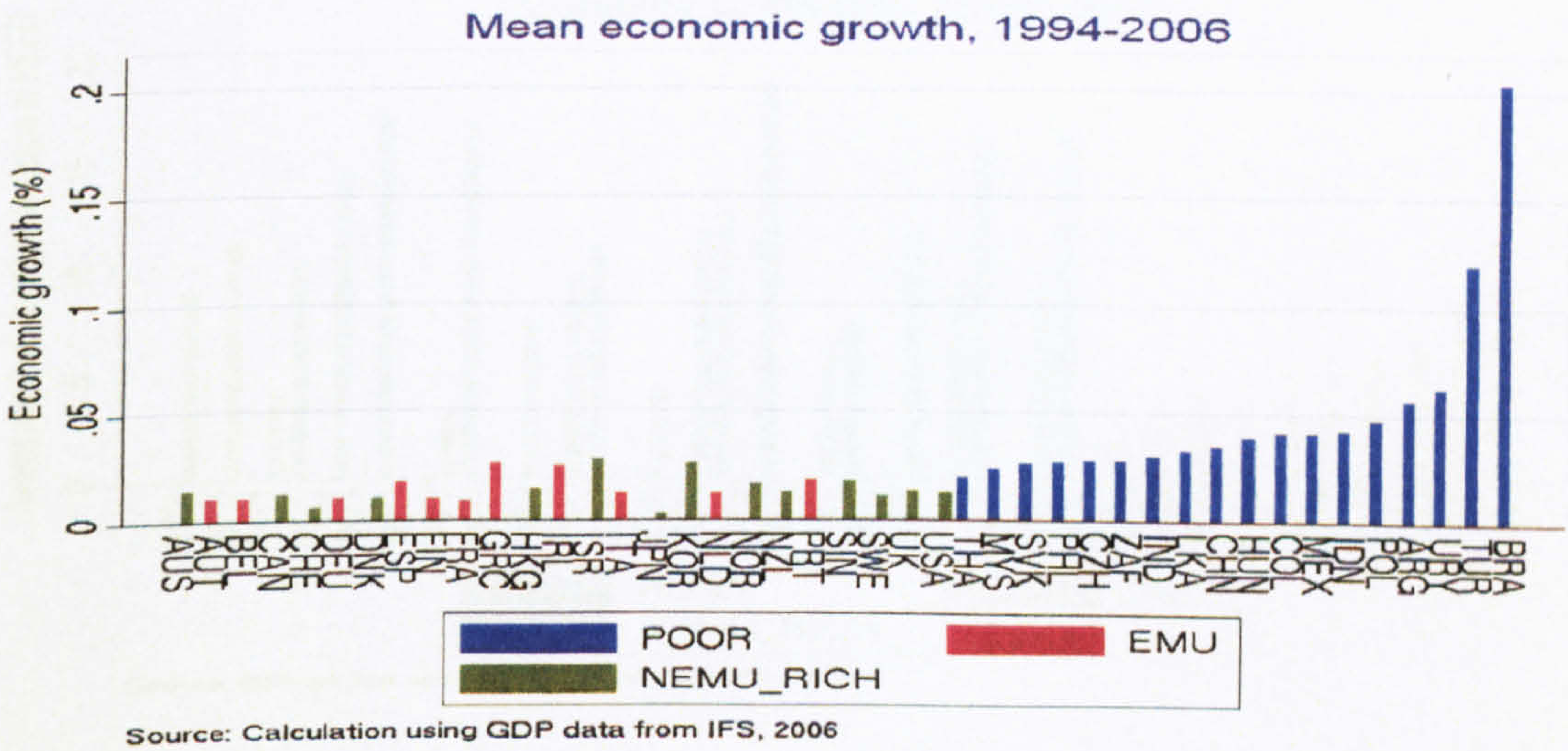


Figure 2.48C: Average inflation over 1994-2006.

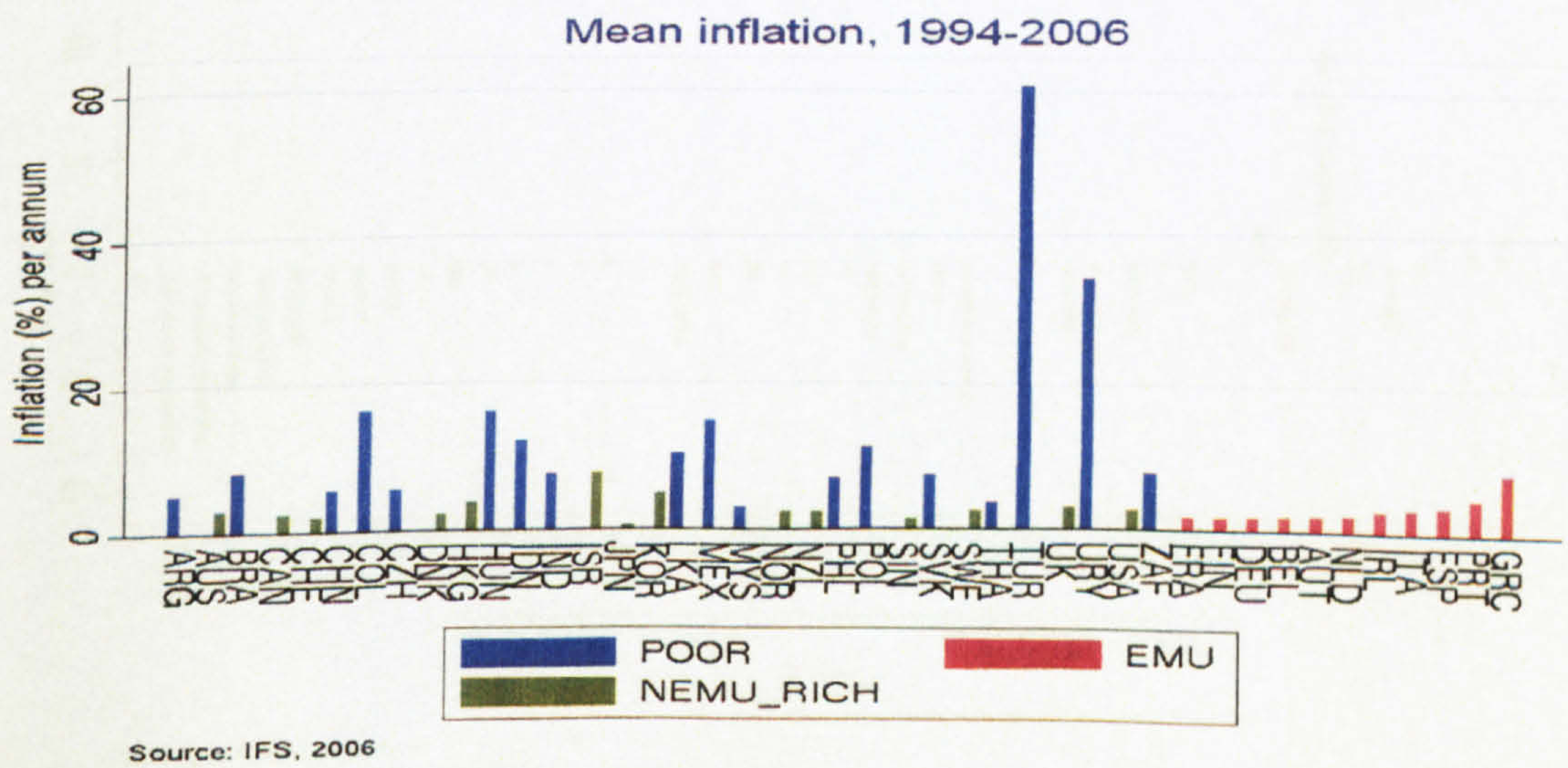


Figure 2.48D: Average real effective exchange rate over 1994-2006.

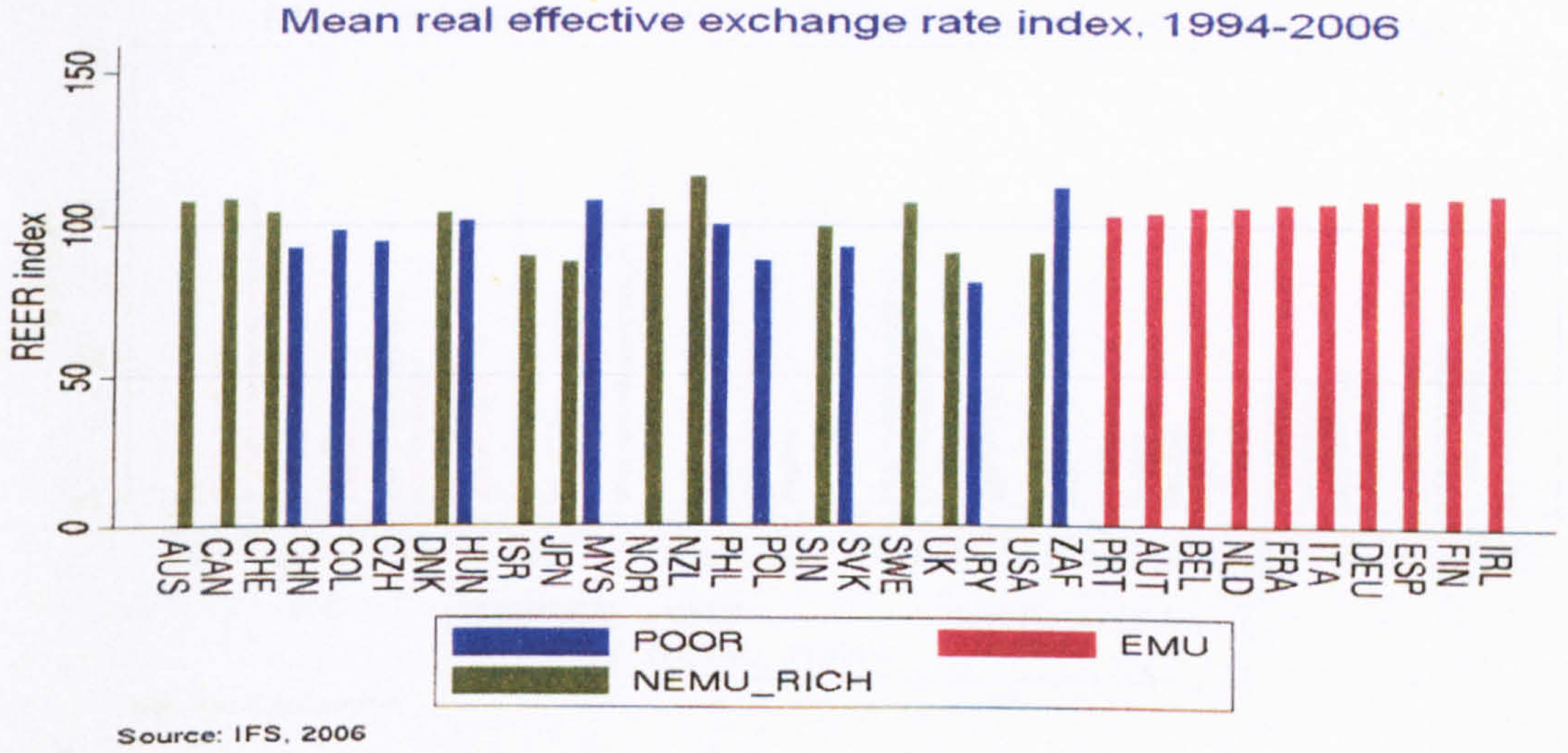


Figure 2.48E: Average standard deviation of real effective exchange rate over 1994-2006.

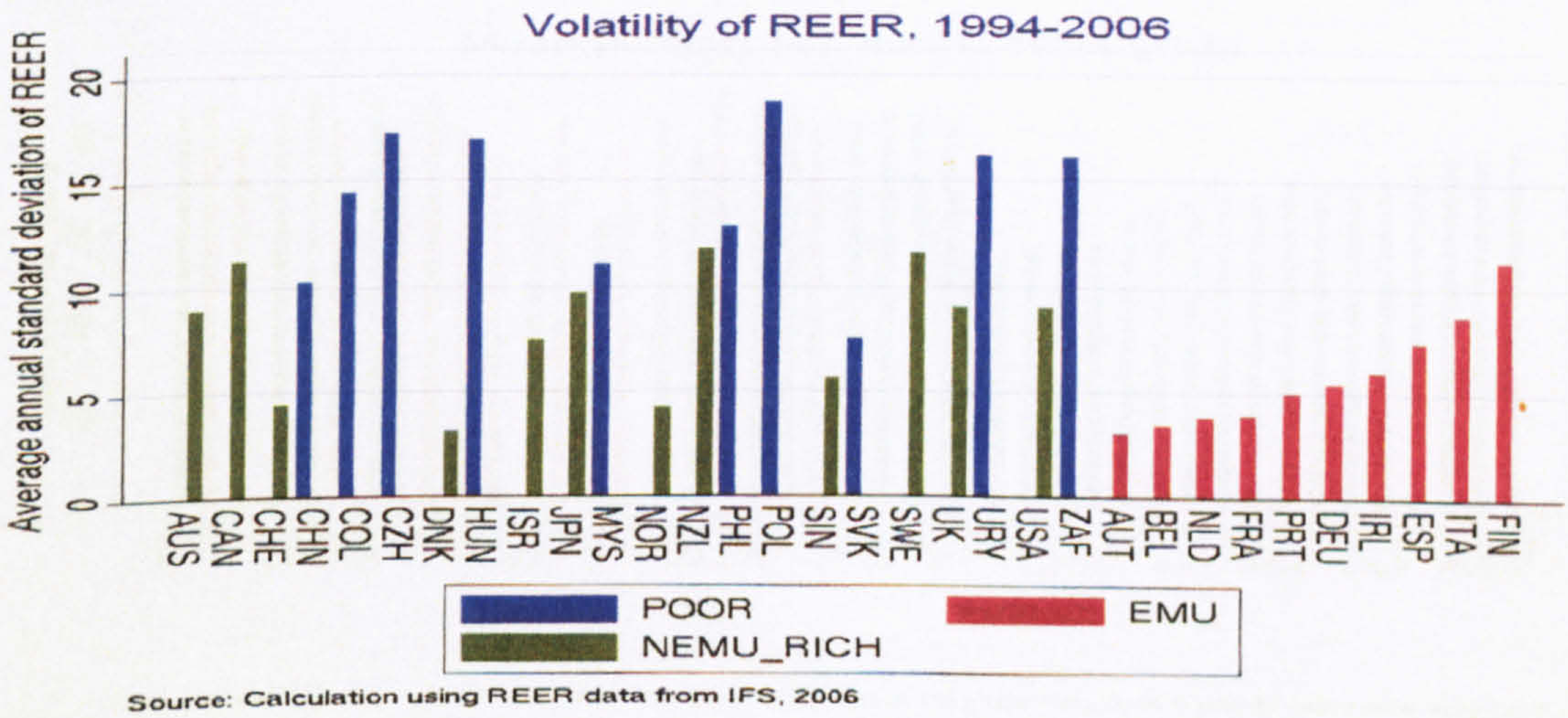


Figure 2.48F: Government budget deficit (% GDP), 1994-2006.

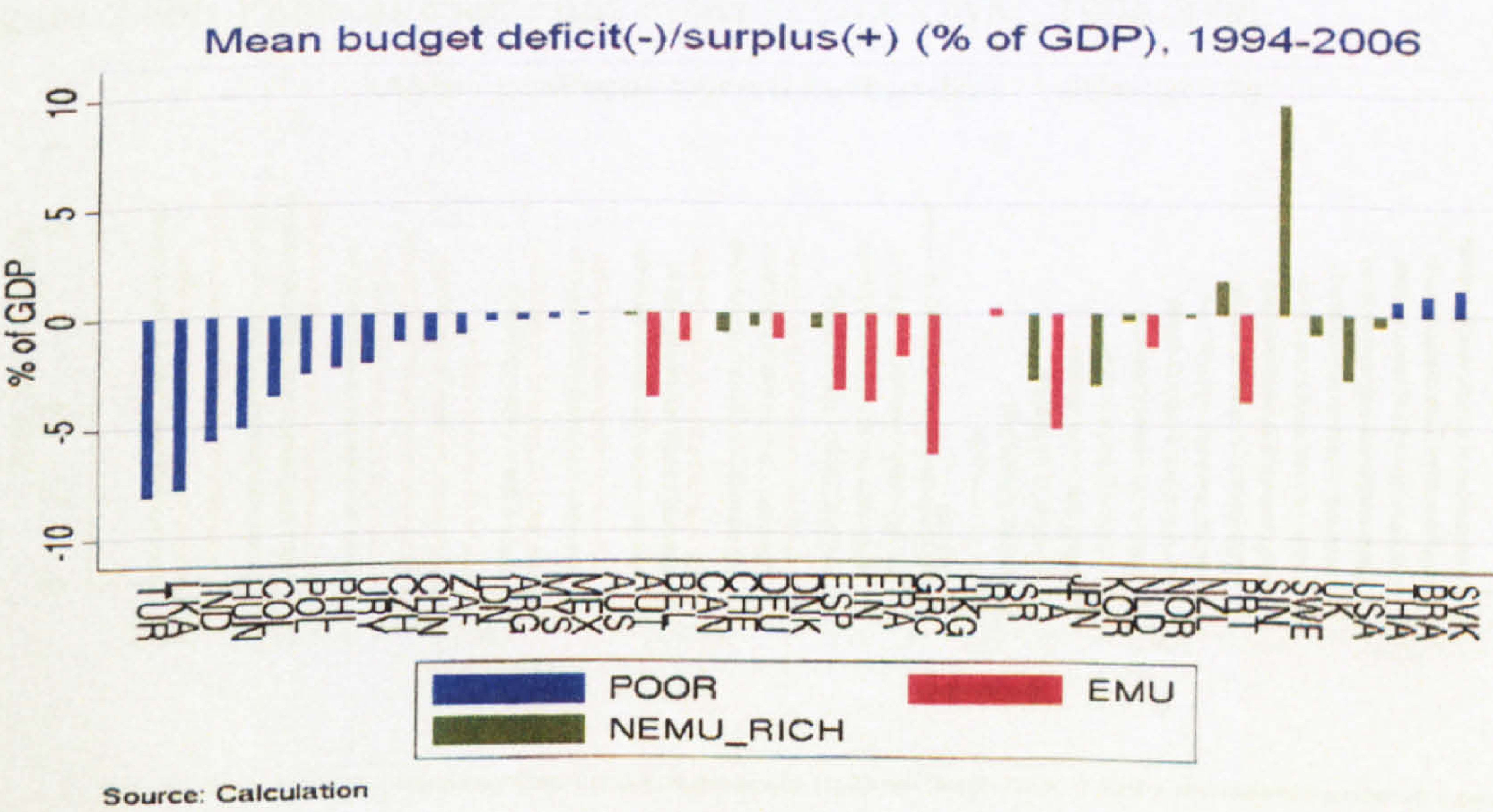


Figure 2.48G: Government debt (% GDP), 1994-2006

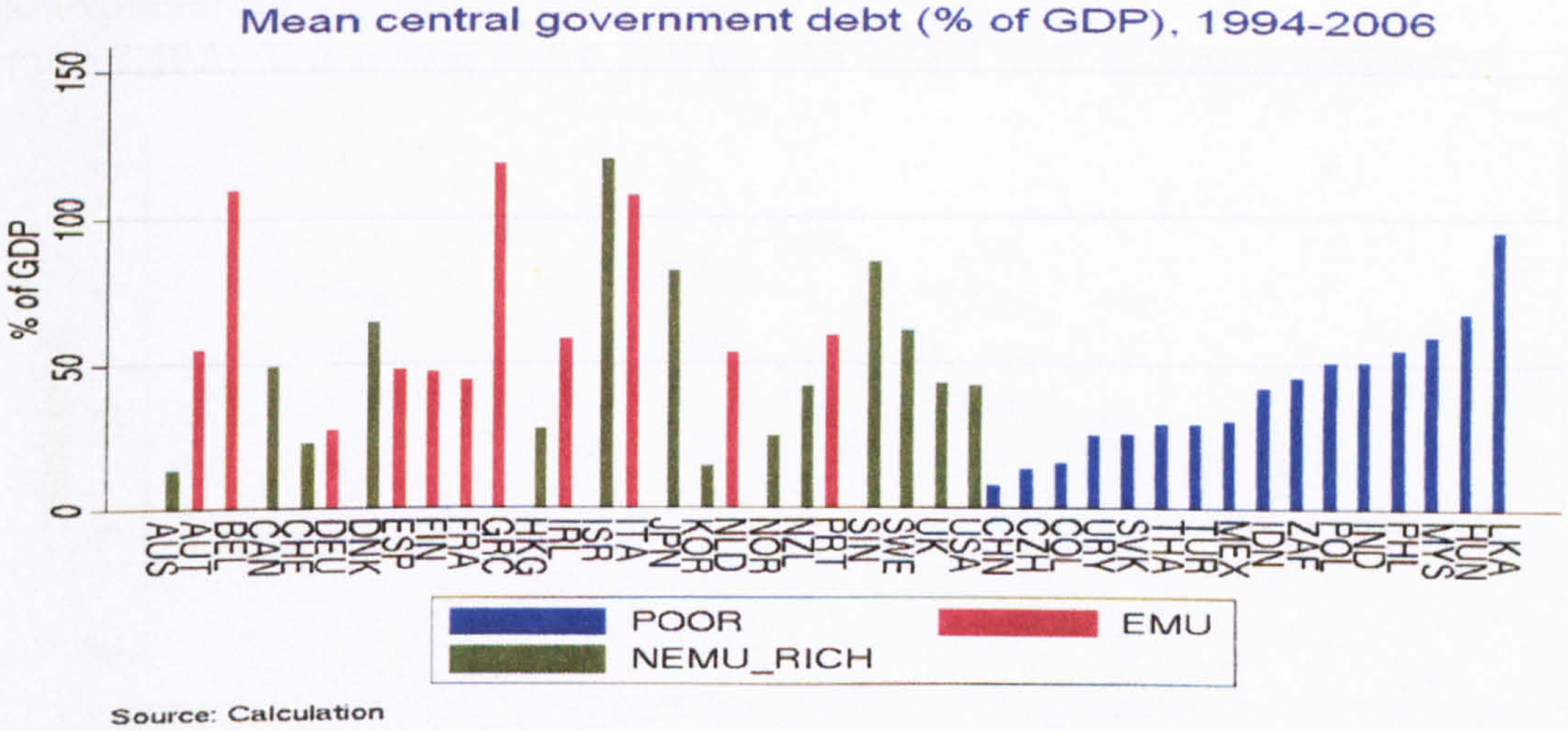


Figure 2.48H: Average political risk rating from International country risk guide (ICRG), 1994-2006.

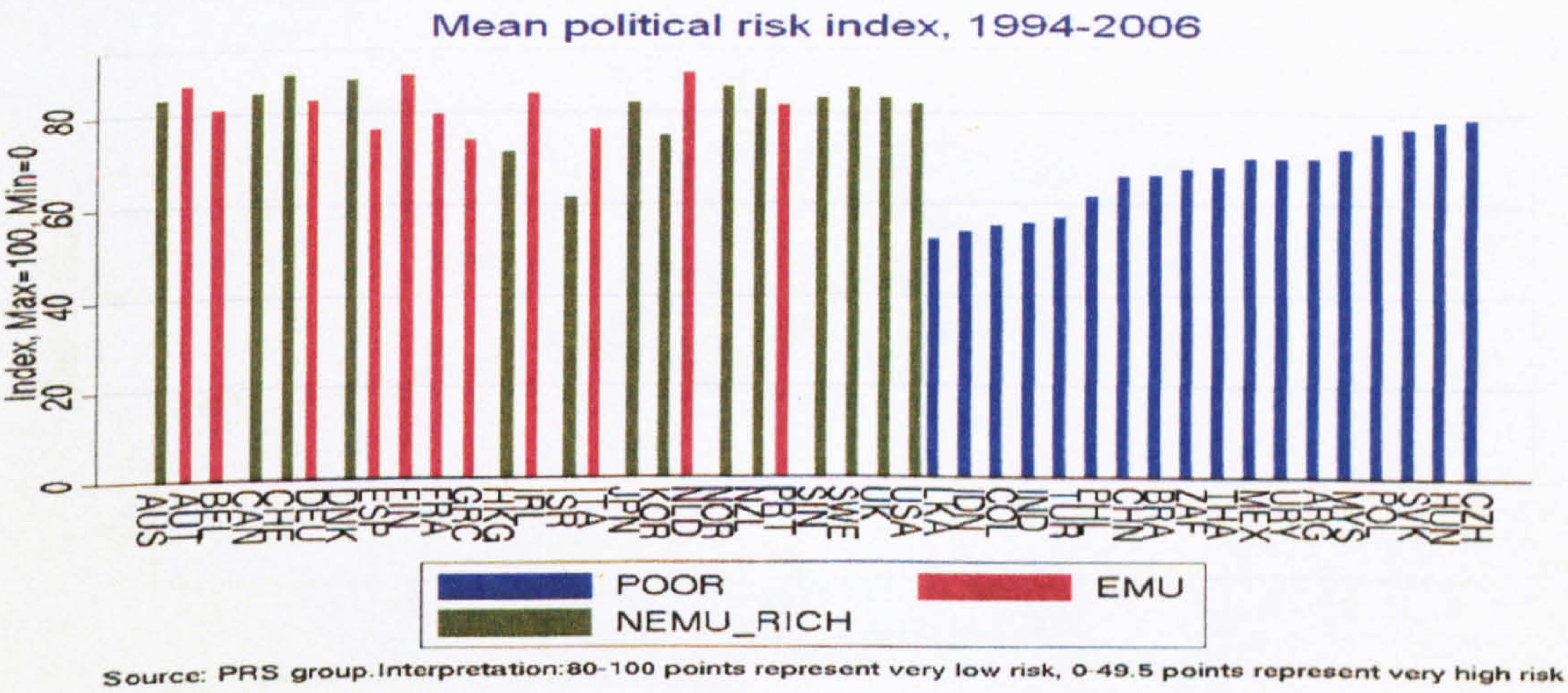


Figure 2.48I: Political constraint index (POLCON5), 1994-2006.

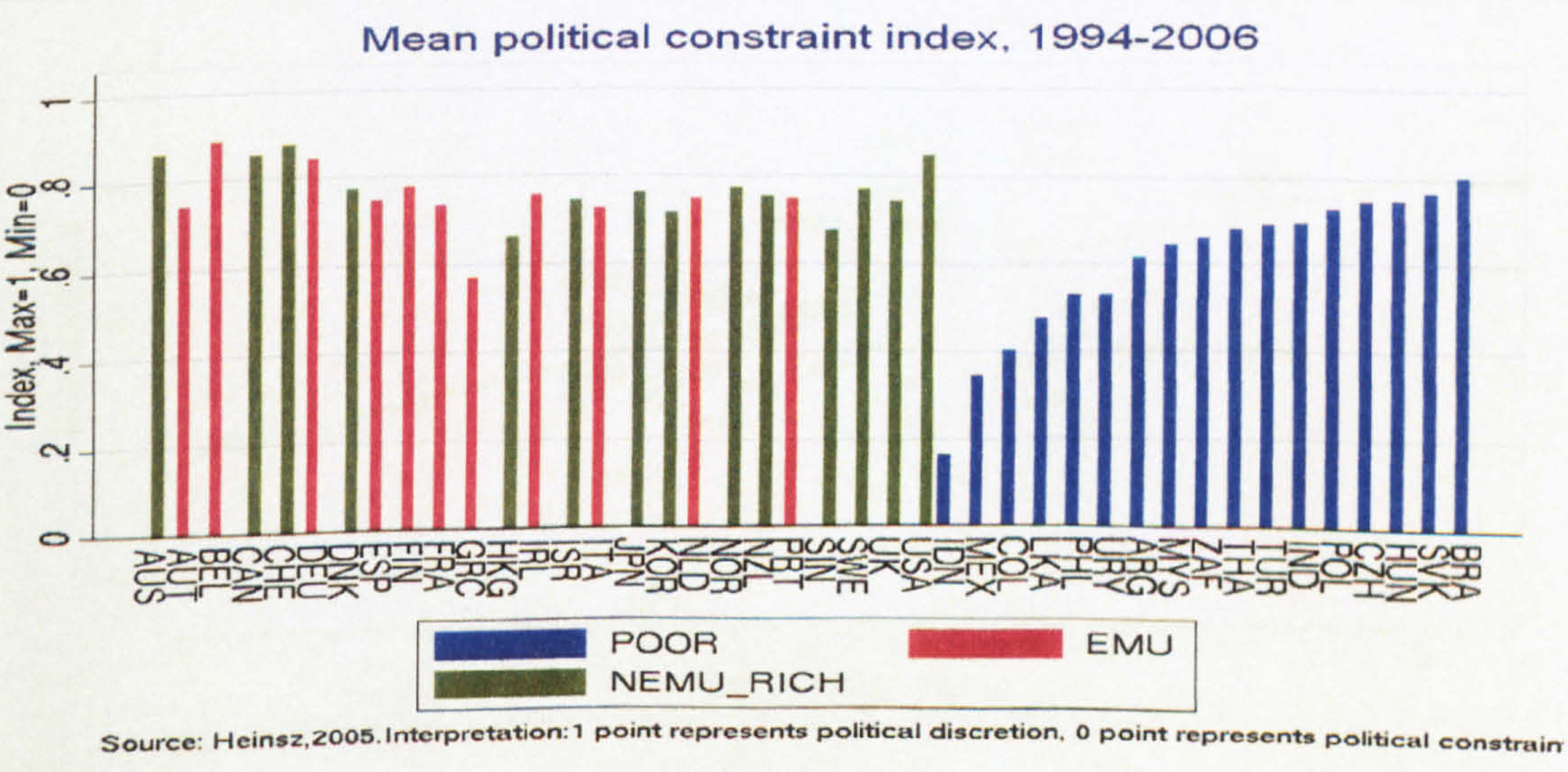


Figure 2.49: Bivariate regression plots of mean value of country's risk premia and explanatory variables (1994-2006).

Figure 2.49A: The average risk premia and initial level of country's income

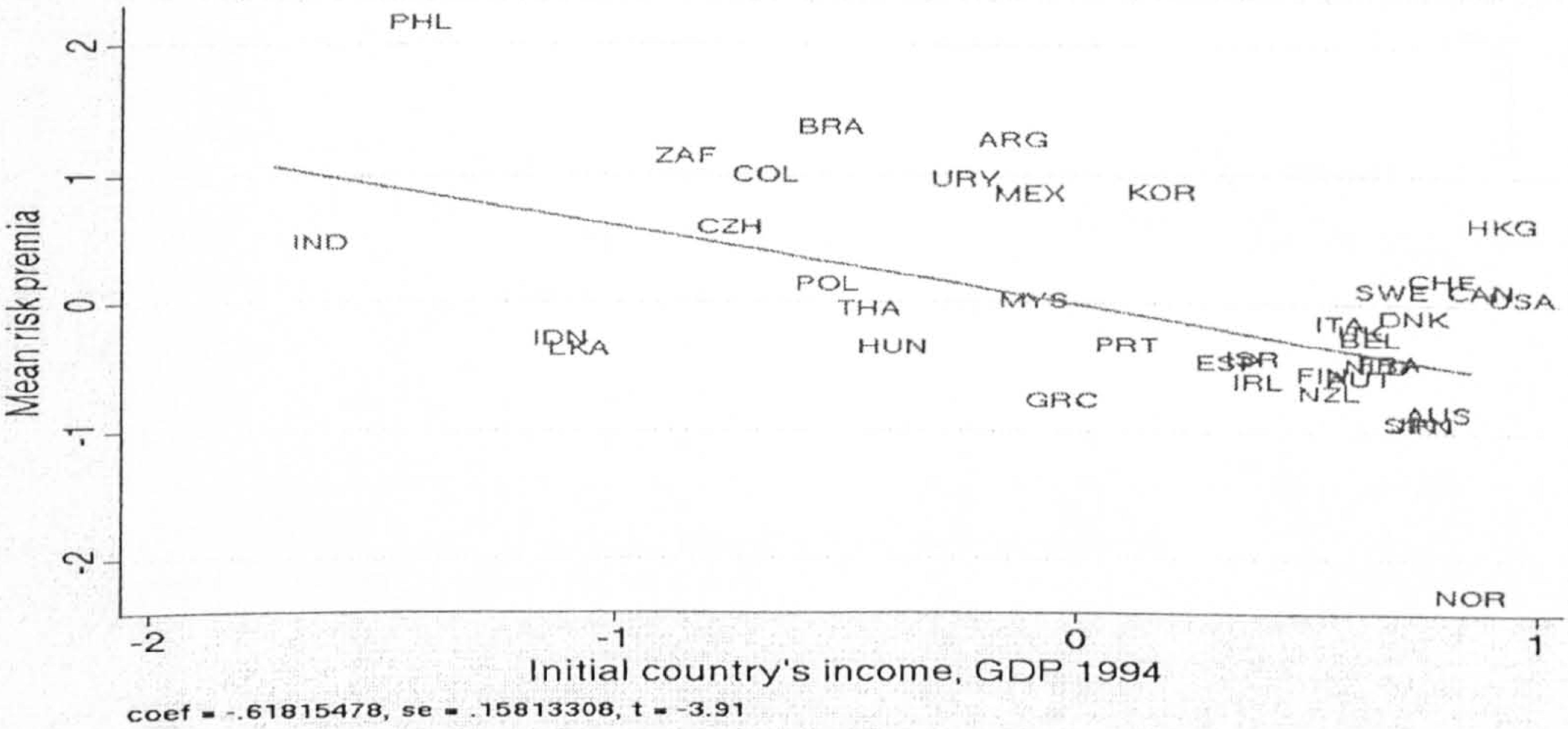


Figure 2.49B: The average risk premia and economic growth.

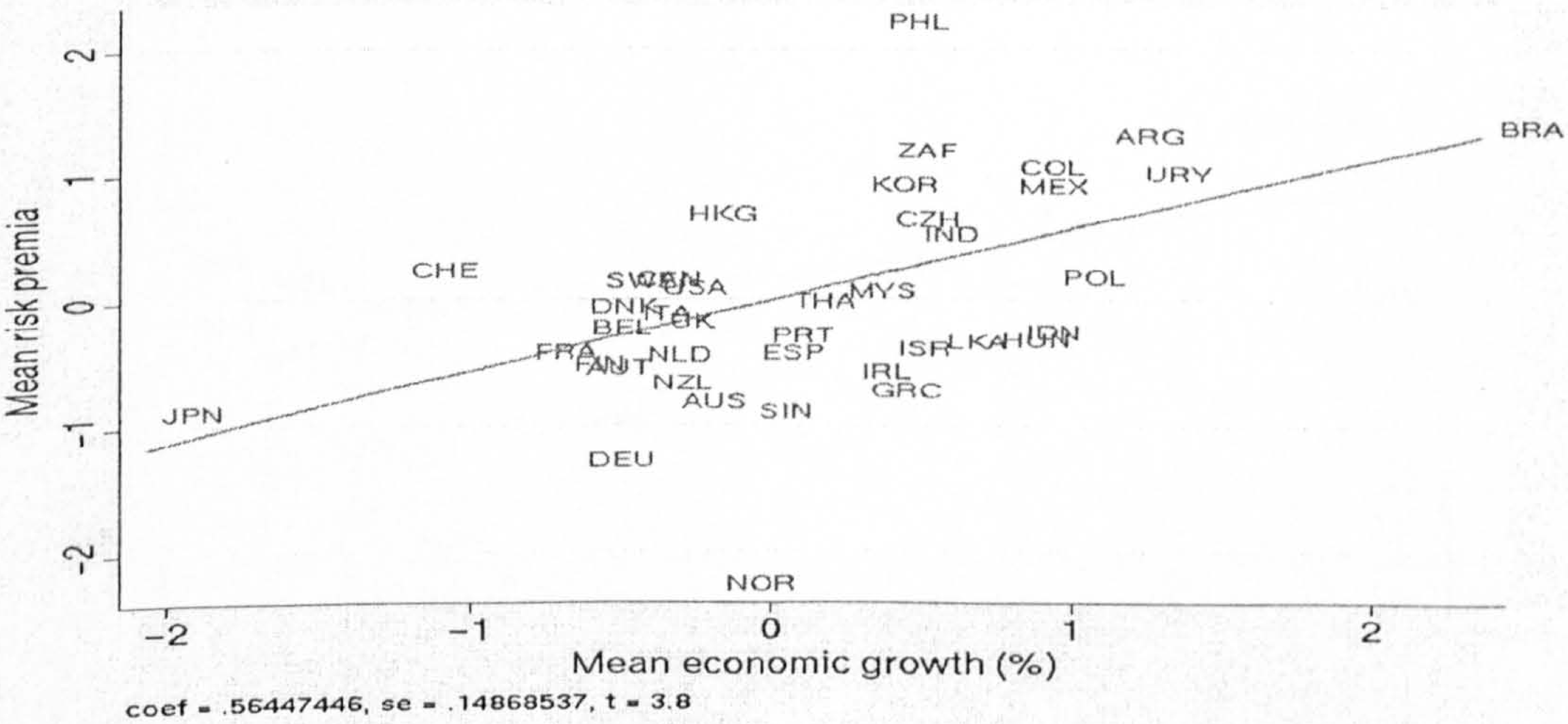


Figure 2.49C: The average risk premia and inflation.

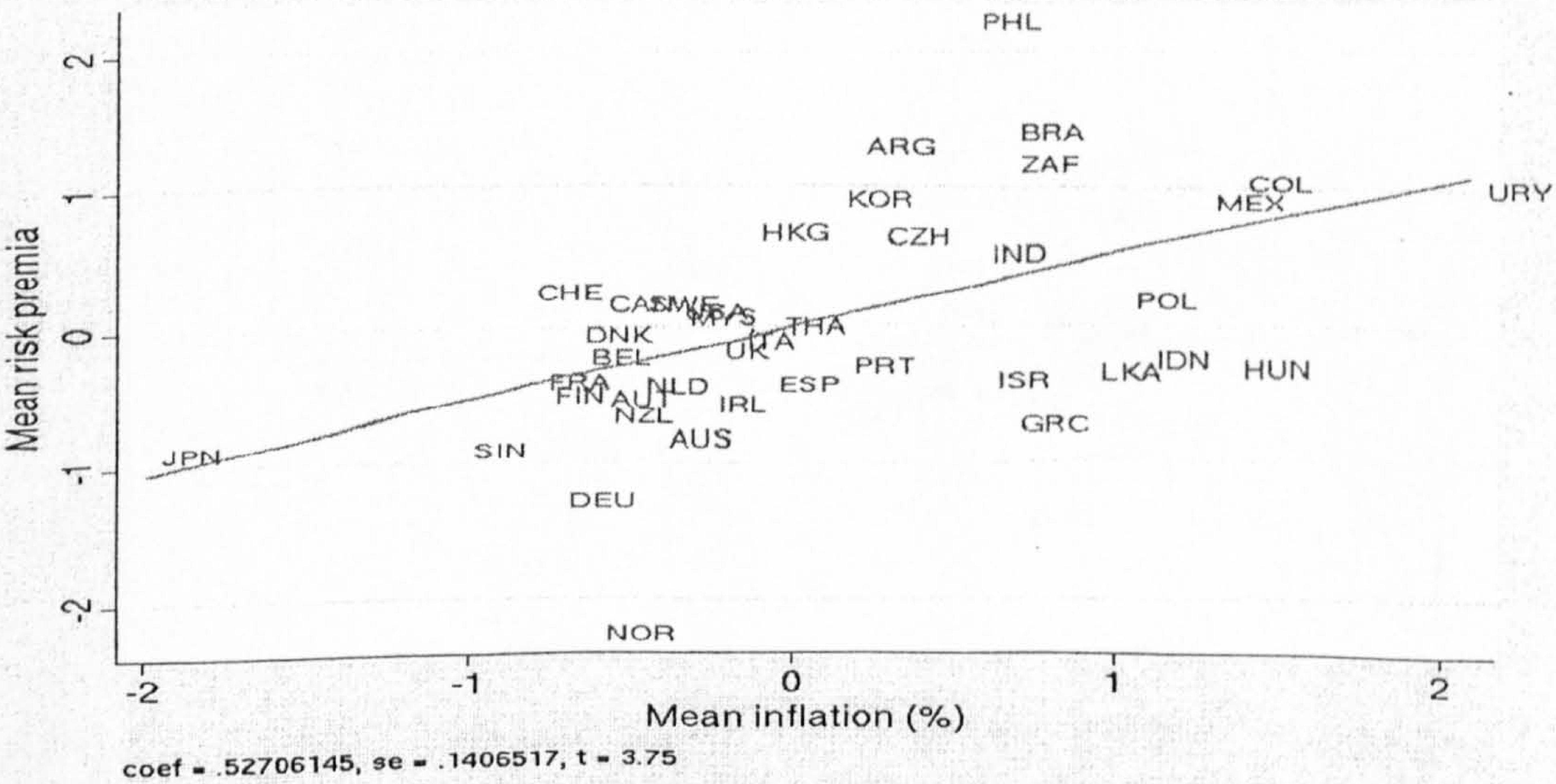


Figure 2.49D: The average risk premia and real effective exchange rate

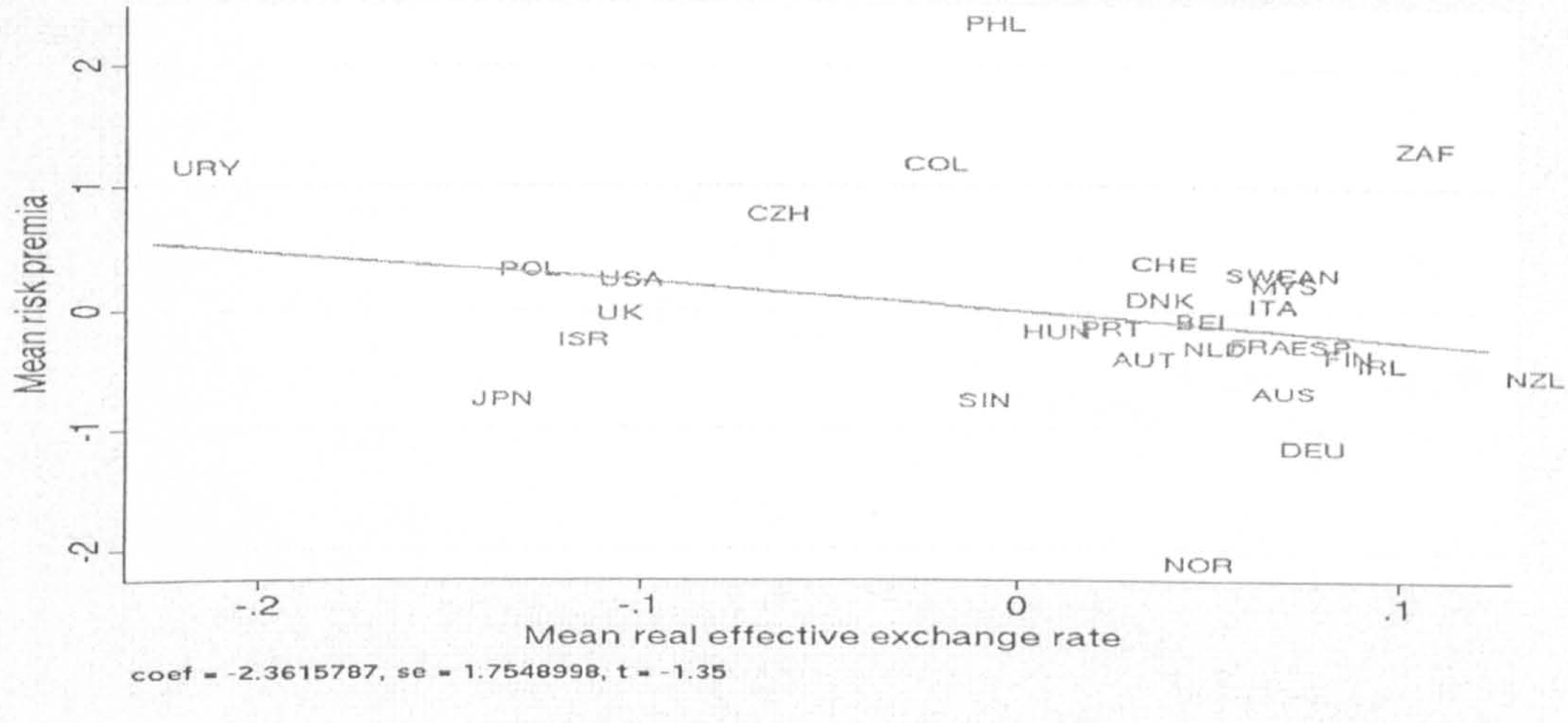


Figure 2.49E: The average risk premia and volatility of the real effective exchange rate.

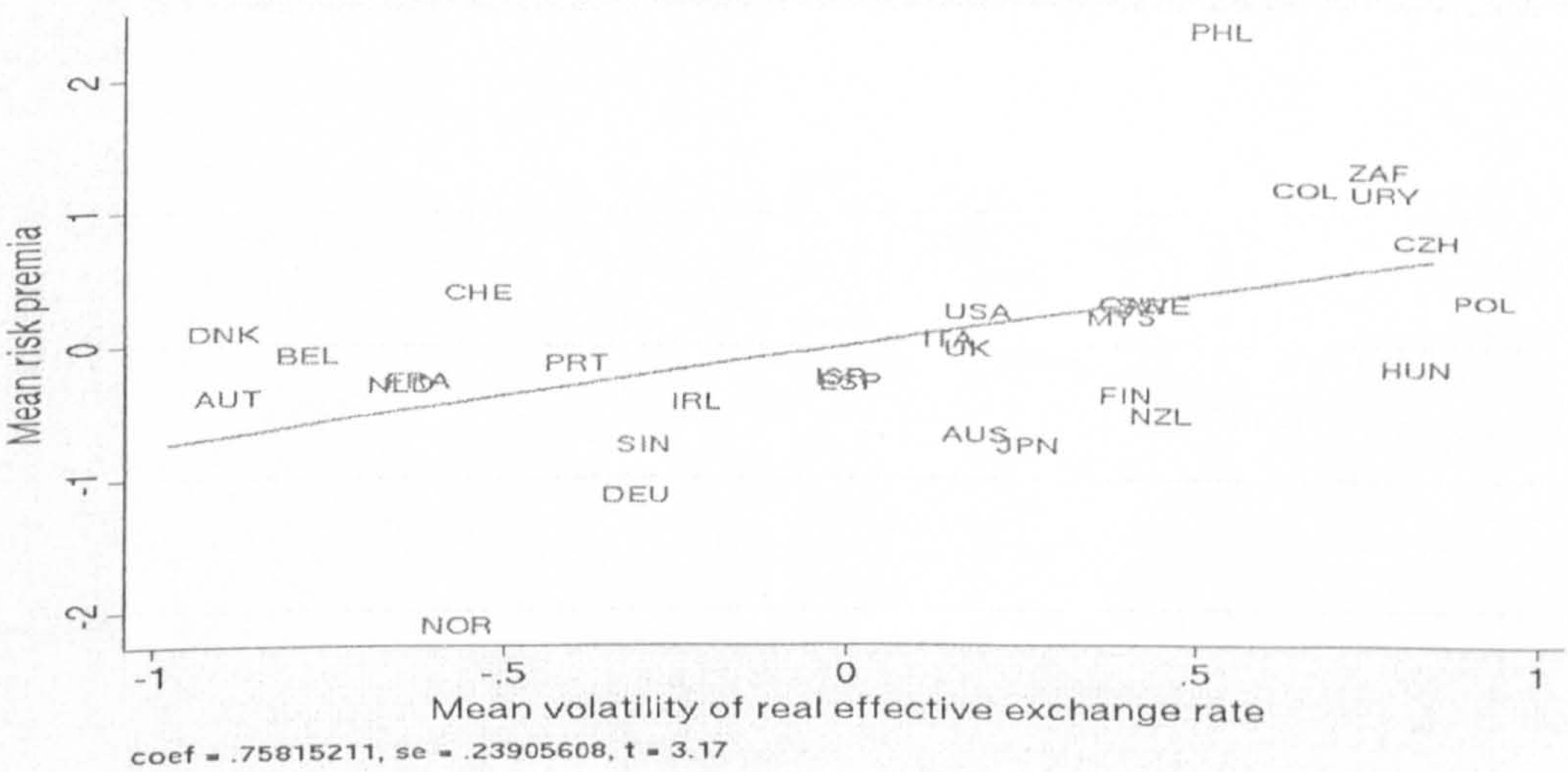


Figure 2.49F: The average risk premia and government budget deficit (% of GDP)

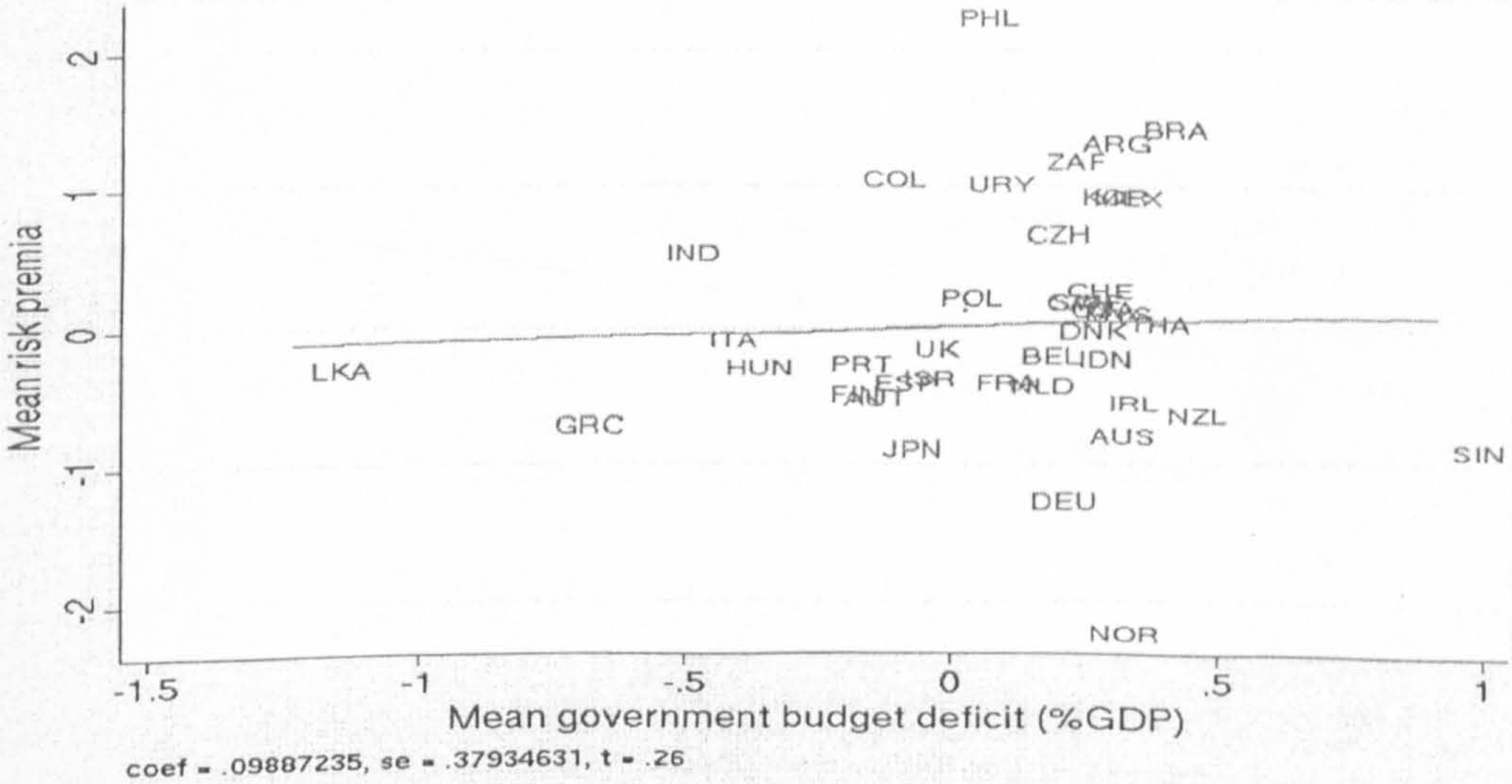


Figure 2.49G: The average risk premia and government debt (% of GDP).

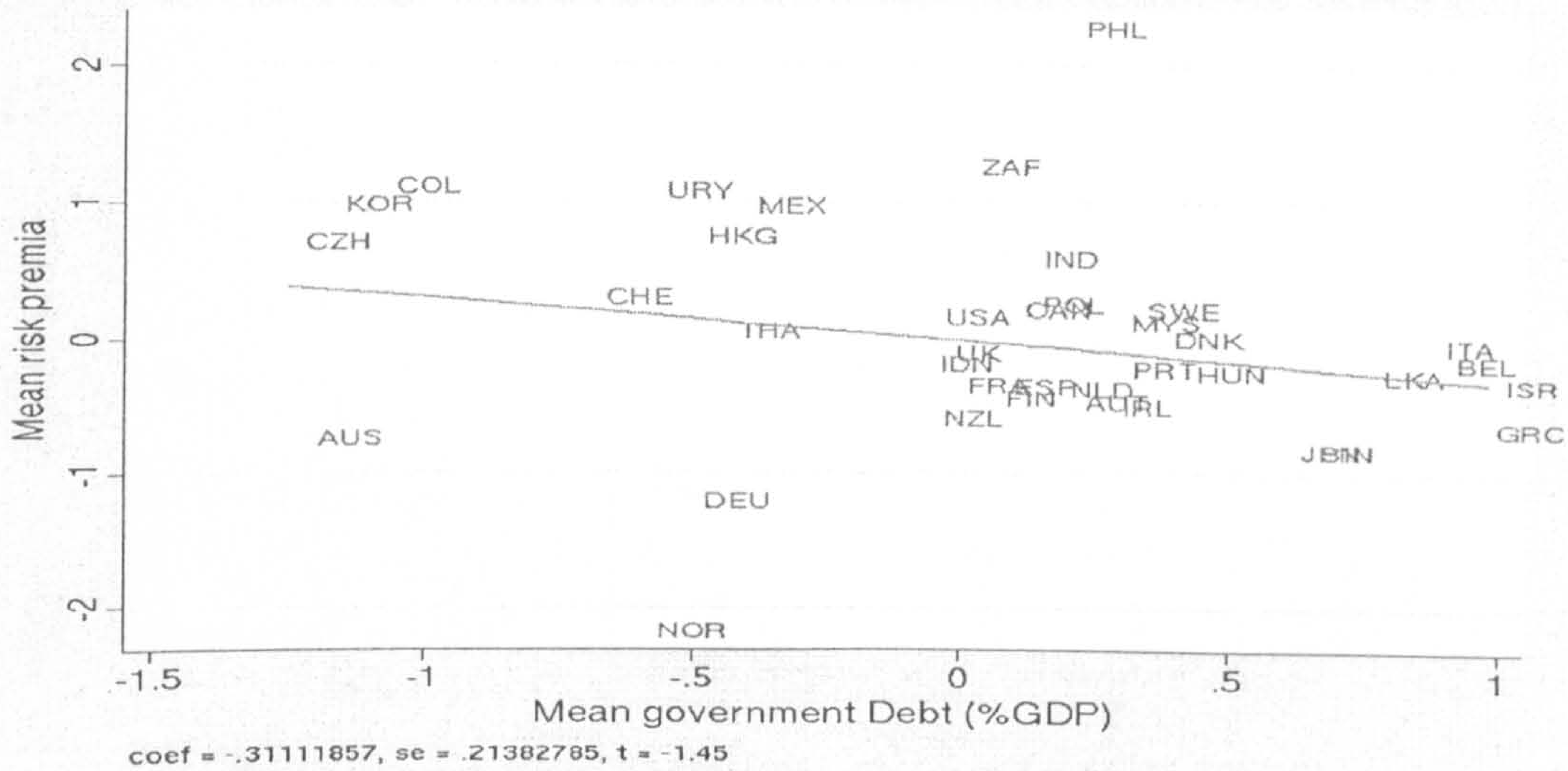


Figure 2.49H: The average risk premia and the political risk index (ICRG).

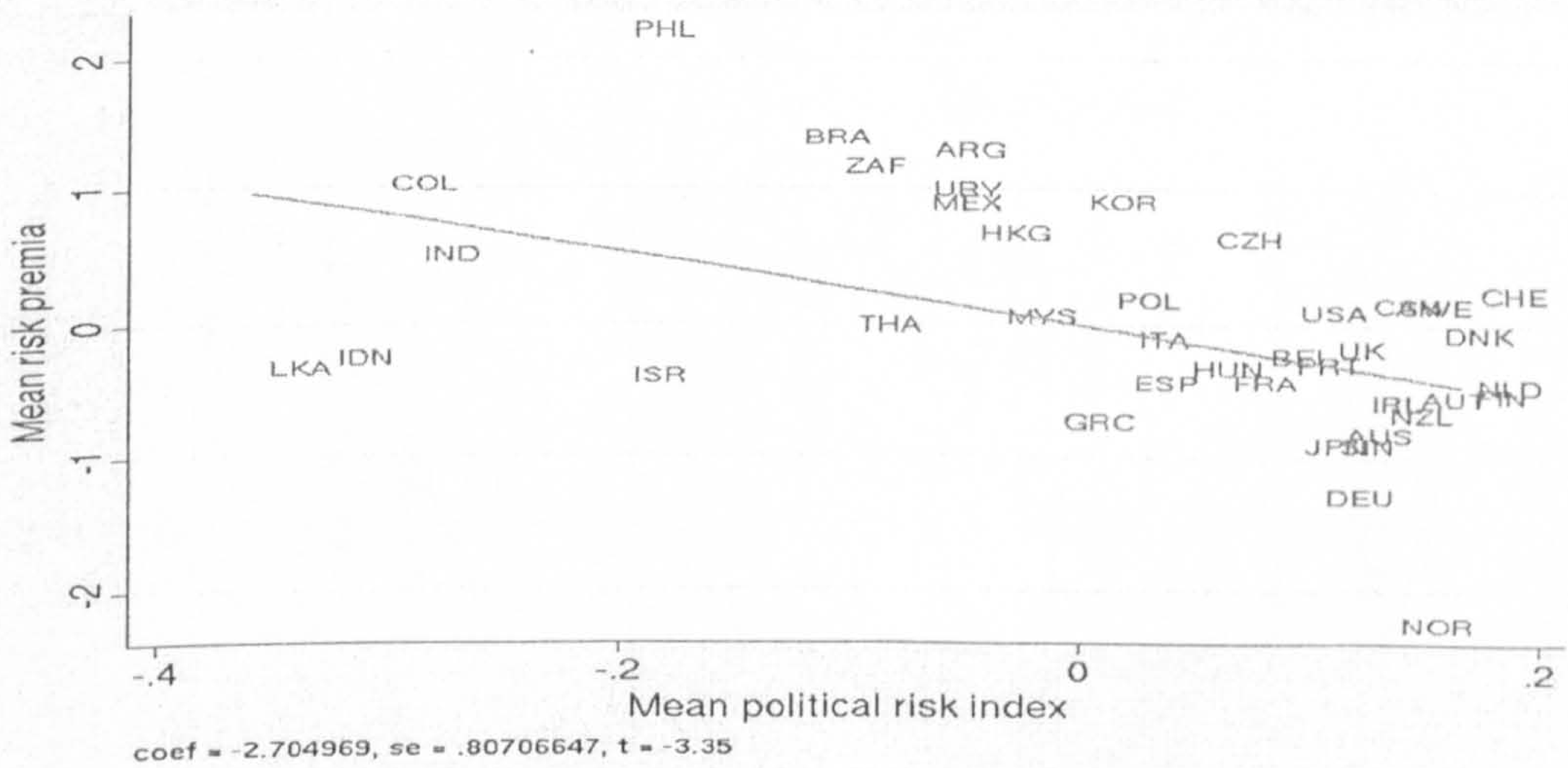


Figure 2.49I: The average risk premia and the political constraint index (POLCON5).

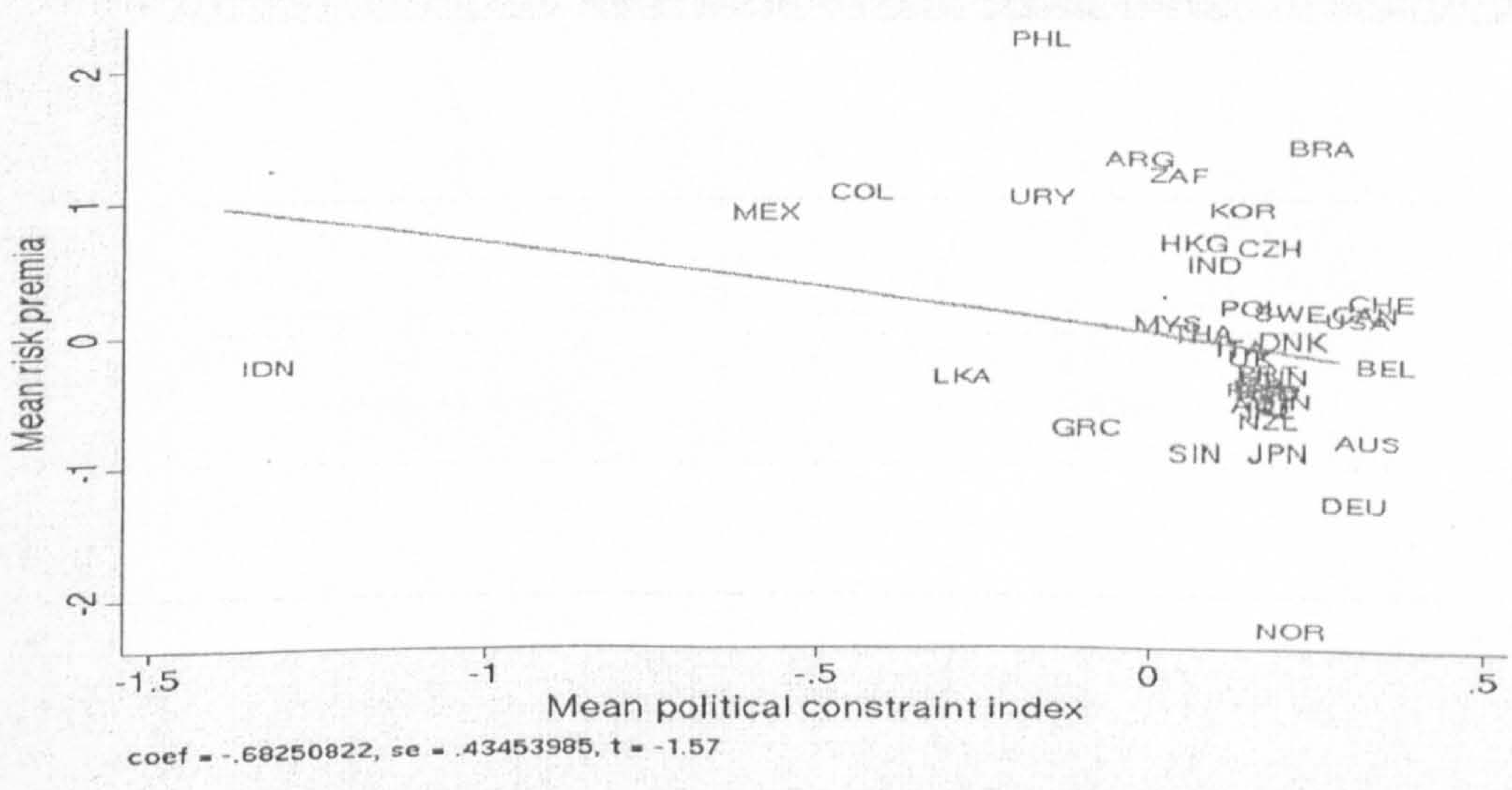
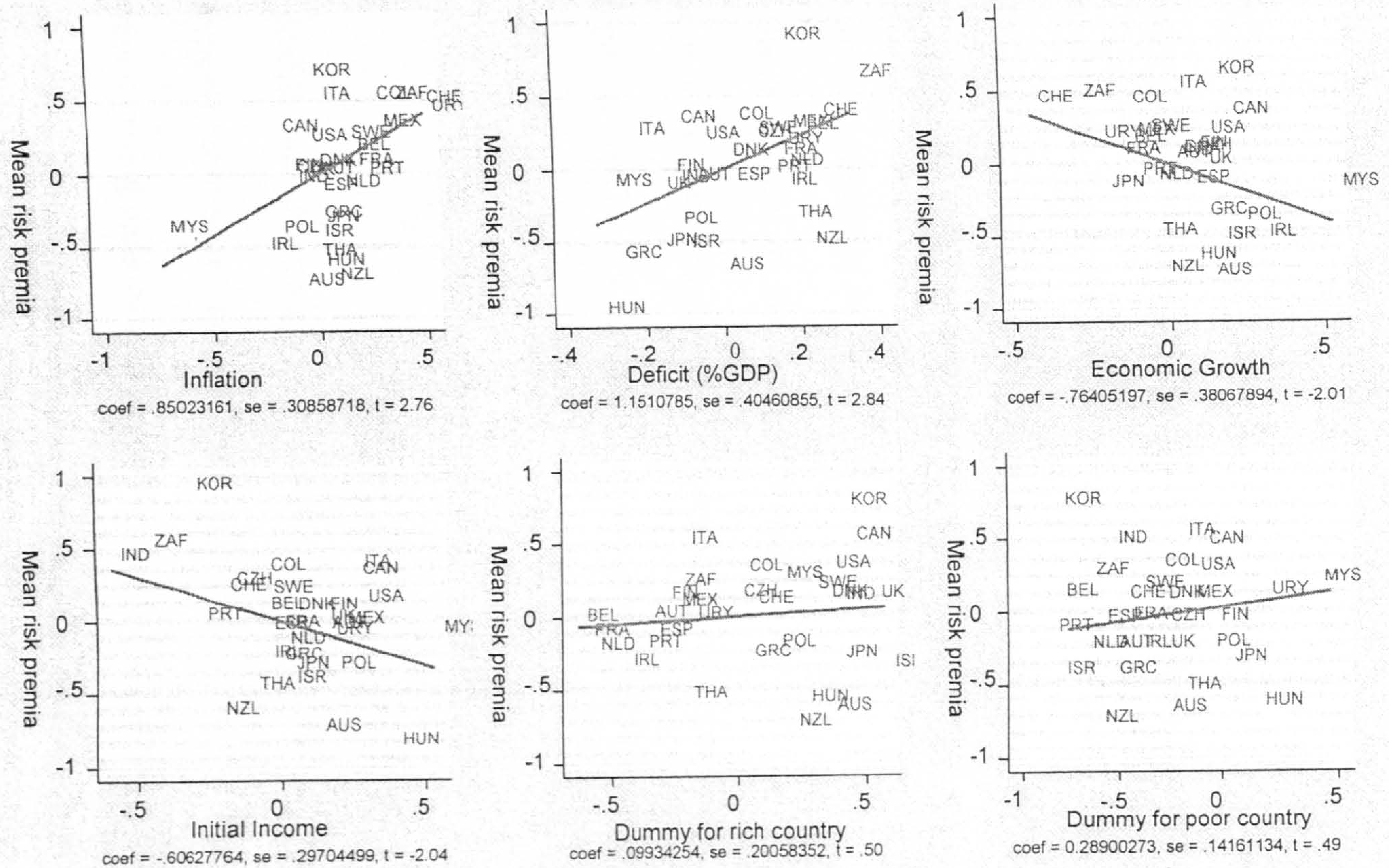


Figure 2.50: Scatter plot of the risk premia and the explanatory variables from model in column (4) of table 2.7.



Chapter 3
Fractional integration and the forward premium puzzle

3.1 Introduction

This chapter revisits the empirical literature testing the efficiency of forward markets for foreign exchange. In previous work, forward rate unbiasedness has been rejected. Generally, the literature has found that the future exchange rate change is negatively related to the forward premium. This chapter assumes rational expectations and attempts to explain the forward premium anomaly by firstly the presence of foreign exchange risk premia and secondly statistical artifacts of the data.

This chapter focuses on the statistical artifacts of the data and finds the conditional mean of the forward premium series to be fractionally integrated while the return on the spot exchange rate is stationary. This yields an unbalanced test regression of the Uncovered Interest Parity, which causes the violation of the hypothesis. This study also pays attention to the relationship of the exchange rate risk premia and the test of the forward market efficiency.

The contribution of this paper to the literature can be divided into two parts. First, this paper examines the time series properties of the forward premium in the 10 most commonly traded currencies using daily data from 1994 to 2007. This paper applies both parametric and semiparametric estimation methods to examine the fractionally integrated behaviour of the forward premium series. The results confirm earlier finding (namely, Baillie and Bollerslev, 1994; Maynard and Phillips, 2001 and Choi and Zivot, 2005) that there is fractionally integrated behaviour in the forward premium series. Moreover, we found that the spot return in all currencies are stationary and follow $I(0)$ process. Thus, it is inappropriate to apply conventional regression analysis to test the hypothesis of the forward rate being an unbiased predictor of the future spot rate. Additionally, we found that even controlling for regime shifts, the fractional integrated behaviour still persists. However, there is evidence that structural breaks cause spurious long memory in the forward premium series of some of the sample currencies analysed in this study.

Secondly, the rejection of the unbiasedness hypothesis is also generally attributed to the presence of a time varying risk premia, which has led to an intensive search for their proper specification. This chapter introduced a new methodology to model the exchange rate risk premia. In contrast to Baillie and Bollerslev (2000) who suggested that the time varying foreign exchange risk premium is extremely small at the monthly level, this chapter argues that the time varying risk premium is significant at the daily level. The risk premia estimation of this chapter is related to that of Baillie and Bollerslev (1990) who attempted to estimate time varying foreign exchange risk premium models in the forward rate forecast error series. They use the multivariate Generalised Autoregressive Conditionally Heteroskedastic (GARCH) model to describe the behaviour of the forward rate forecast error. However, we argue that the forward rate forecast error process appears to have a long memory feature in the conditional variance for all of the currencies in this study. Thus this chapter suggests using the Fractionally Integrated GARCH (FIGARCH) models in estimating the exchange rate risk premia. An implication of the robustness of the FIGARCH model and importance of the long memory volatility parameter strongly suggests that the FIGARCH model is the best model to explain the long memory property in the forward rate forecast error series.

The organization of the paper is as follows. Section 3.2 reviews the concept of the unbiasedness hypothesis in exchange rate markets and surveys several empirical studies that investigate this hypothesis. Section 3.3 reviews some of the important advances in the theoretical modelling of the exchange rate risk premium. Section 3.4 describes the data sources and their time series properties. Section 3.5 provides empirical evidence of the forward premium anomaly by performing basic rolling regression analysis. The causes of the forward premium anomaly are addressed in sections 3.6 to 3.8. Sections 3.6 and 3.7 explain the forward premium anomaly by the evidence of persistence mismatch between the spot return and the forward premium. Empirical estimates of the fractional differencing parameter (correcting for multiple structural breaks) are provided in section 3.6, followed by the stationarity test for the spot rate returns in section 3.7. Section 3.8 explains the forward premium anomaly by the existence of the exchange rate risk premium. Lastly, section 3.9 gives out the summary remarks and conclusions.

3.2 The unbiasedness hypothesis

Interest arbitrage refers to the international flow of short-term liquid capital to earn higher returns abroad (Salvatore, 1998). The test of the market unbiasedness hypothesis generally considers the Covered Interest Parity condition, the Uncovered Interest Parity, and the Forward Rate Unbiasedness hypothesis. These hypotheses will be discussed in detail in the following section.

Covered Interest Arbitrage refers to the spot purchase of foreign currency required for investments in assets market and offsetting simultaneous forward sales (swap[81]) of foreign currency to cover the foreign exchange risk (Salvatore, 1998). The Covered Interest Parity condition (CIP) states that there should be no profitable covered arbitrage opportunity. The CIP condition compares two assets which are identical in every relevant respect (such as default and political risk) except currency of denomination in the market and assumes no barriers to arbitrage across international financial markets. The standard log-linear form of the CIP condition is defined as follows:

$$f_{t,k} - s_t = \left(i_{t,k} - i^*_{t,k}\right), \tag{3.10}$$

where $f_{t,k}$ is the natural log of the forward rate for a contract expiring k periods in the future, s_t is the exchange rate - defined as the natural log of the price of foreign currency in units of domestic currency at time t, $i_{t,k}$ is the k-period yield on the domestic financial instrument, and $i^*_{t,k}$ is the corresponding yield on the foreign instrument. Equation (3.10) is an equilibrium condition that holds regardless of investor preferences toward risk since it is based on riskless arbitrage.

Early studies usually tested the CIP by examining the arbitrage condition among countries that have comparable economic and political risk aspects such as EU member countries (Aliber, 1973; Dooley and Isard, 1980; Frankel and MacArthur, 1988; Sarno and

81 A currency swaps (or cross currency swap) is a foreign exchange agreement between two parties to exchange a given amount of one currency for another and, after a specified period of time, to give back the original amounts swapped. A swap is considered to be a foreign exchange transaction (short leg) plus an obligation to close the swap (far leg) being a forward contract.

Taylor, 2002a; Taylor, 1987 and Taylor, 1989). These studies use OLS regression analysis of the equation $(f_{t,k} - s_t) = a + b\left(i_{t,k} - i^*_{t,k}\right) + \varepsilon_t$. The null of CIP is $H_0 : a = 0, b = 1$, and if there are transactions costs, these may show up as $a \neq 0$. These studies found that the data strongly support the covered interest parity and efficient market hypotheses i.e. there were no profitable covered arbitrage opportunities. The recent work on CIP employ more advanced econometric analysis such as Balke and Wohar (1998) and Peel and Taylor (2002) which examine the dynamic behavior and significance of deviations from CIP using non-linear Threshold Autoregressive (TAR) models. The results of this work support the earlier finding i.e. the arbitrage profits from CIP is very small if it exists.

As mentioned earlier, equation (3.10) is a risk free arbitrage condition that holds regardless of investor preferences. However, if investors are risk averse, the forward rate can differ from the expected future spot rate by a premium that compensates for the perceived riskiness of holding domestic versus foreign assets as follows;

$$f_{t,k} - E_t s_{t,t+k} = \eta_{t,t+k}, \tag{3.11}$$

where $E_t s_{t,t+k}$ is market expectation at time t of exchange rate that will prevail at time $t+k$ and $\eta_{t,t+k}$ is the risk premium associated with investment in foreign assets when the future foreign currency receipts are not covered in the forward market. If $\eta_{t,t+k}$ equals zero, there are two implications. Firstly, agents are risk neutral. Secondly, it gives an econometric implication that the forward rate is an unbiased predictor of the future spot rate. This condition is called Forward Rate Unbiasedness (FRU). Under risk neutrality, if FRU does not hold, there would be profitable opportunities available by speculating in the forward market.

Substituting equation (3.11) into (3.10) yields a relationship between the expected change in the spot exchange rate, the interest rates differential and the risk premium, $E_t s_{t,t+k} - s_t = \left(i_{t,k} - i^*_{t,k}\right) - \eta_{t,t+k}$. If the risk premia term, $\eta_{t,t+k}$ equals zero, the relationship defines the uncovered interest parity condition (UIP),

$$E_t s_{t,t+k} - s_t = \left(i_{t,k} - i^*_{t,k}\right). \tag{3.12}$$

From equation (3.12), investment in foreign assets is risky because the future foreign currency receipts are not covered in the forward market as in (3.10). Moreover, it can be inferred that the UIP is the condition for equilibrium in the capital account under the assumption of risk neutrality[82] when agents only consider the expected return. Equation (3.12) is not directly testable since market expectations of future exchange rate movements are difficult to observe in practice.

According to Chinn and Meredith (2005), the concept of UIP is tested jointly with the assumption of rational expectations in exchange markets. Under rational expectations, future realizations of exchange rates at time $t+k$ can be expressed as the sum of its value expected at time t and a white noise error term $\xi_{t,t+k}$ that is uncorrelated with all information known at time t, including the interest differential and the spot exchange rate:

$$s_{t+k} = E_t s_{t,t+k} + \xi_{t,t+k}. \tag{3.13}$$

Substituting equation (3.13) into equation (3.12) yields $\Delta s_{t,t+k} = \left(i_{t,k} - i^*_{t,k}\right) + \xi_{t,t+k}$. An empirically testable equation for the realized change in the exchange rate from t

82 The UIP condition assumes that the market is dominated by risk-neutral investors and that neither risk averse rational speculators nor noise traders have a powerful influence on market prices (Cuthbertson and Nitzsche, 2004).
See Engel (1996) for the approximations and simplifying assumptions of the risk neutrality.

to $t+k$ can be written as

$$\Delta s_{t,t+k} = a + b\left(i_{t,k} - i^*_{t,k}\right) + \varepsilon_{t,t+k}, \tag{3.14}$$

where $\varepsilon_{t,t+k}$ is a rational expectation forecast error term, $\Delta s_{t,t+k} = s_{t+k} - s_t$ is the percentage depreciation of the currency over k periods and $\left(i_{t,k} - i^*_{t,k}\right)$ is the difference between current k period domestic interest rates and foreign interest rates. From equations (3.13) and (3.14), we can infer that the UIP hypothesis is founded on the joint assumptions of rational expectations and risk neutrality. Thus the UIP is sometimes referred to as the risk neutral efficient market hypothesis (RNEMH). Additional assumptions are that there is free capital mobility and an absence of taxes on capital transfers. The UIP condition holds if $a = 0$, $b = 1$ and $\varepsilon_{t,t+k}$ is serially uncorrelated with information available at time t. The unit slope coefficient implies that high interest rates in the foreign country should not imply high returns to foreign investment but should signal an equal expected depreciation of the foreign currency.

However, empirical studies generally find that UIP is rejected. b in the regression equation (3.14) is frequently found to be *negative* which implies that in periods when the interest differential in favour of the foreign currency is positive, the foreign currency tends to *appreciate*. Cuthbertson and Nitzsche (2004) suggested that the negative slope coefficient does not necessary imply profit on average by holding assets in countries that have interest rates higher than domestic countries. Their explanation is that the intercept may offset the impact of the interest differential term. Normally, financial assets in highly risky countries offer higher interest rates *than normal.* These countries tend to have monetary instability, political unrest, a weak and non-diversified export base, then the high expected return to invest in these countries may be a payment for these systematic risk. Another interesting point is that the R-squared in the regressions is usually quite low. Cuthbertson and Nitzsche (2004) interpret these phenomena that the bets are highly risky, although they do pay off in terms of positive returns over a run of bets.

Note that the specification of the UIP condition in equation (3.14) does *not* require the CIP condition to hold. An alternative specification of the UIP condition can be presented below (in equation (3.15)). It implies that the forward premium should be equal to the market expectation of the exchange rate depreciation, given that the covered interest parity condition holds. Combining equation (3.10) with equation (3.12) and assuming that there is no risk premia in holding foreign assets ($\eta_{t,t+k} = 0$), the UIP condition can be expressed as follows;

$$E_t \Delta s_{t,t+k} = (f_{t,k} - s_t) = \left(i_{t,k} - i^*_{t,k}\right), \tag{3.15}$$

where $E_t \Delta s_{t,t+k} = E_t s_{t,t+k} - s_{t+k}$ and the term $f_{t,k} - s_t$ is defined as the forward premium, or the percentage difference between the current forward and spot exchange rates. By arbitrage, the forward premium must equal the interest differential. If it did not, a strategy of borrowing in the foreign currency, then changing the proceeds into domestic currency, investing the domestic currency and then selling forward would yield a riskless profit (Froot and Thaler, 1990). The market respects this arbitrage condition; for example, banks allow forward rates to be set by interest differentials. Froot and Thaler (1990) further conclude that under risk neutrality and rational expectations, the forward premium should also be an unbiased estimate of the subsequent exchange rate changes. The UIP in this form implies that the forward premium should predict changes in the exchange rate[83]. The test of UIP is then defined as

$$\Delta s_{t,t+k} = a + b\,(f_{t,k} - s_t) + \varepsilon_{t,t+k}. \tag{3.16}$$

Again, the test is under the null hypothesis of $H_0 : a = 0, b = 1$. If the forward rate embodies all information available at time t, then it should also be true that $\varepsilon_{t,t+k}$ is serially uncorrelated. Generally and perhaps surprisingly, empirical work finds that $b \neq 1$ and is usually negative e.g. Frankel, 1980; Fama, 1984; Bekaert and Hodrick, 1993. These works use a variety of exchange rates against the US dollar. Froot (1990) provides an extensive literature review and also found that b is frequently estimated to be less than zero. Froot and Thaler (1990) find that the average coefficient across some 75 published estimates of equation (3.16) is -0.88. In these, a few are positive, but none of them have the coefficient, b statistically greater than or equal to unity. The finding that b is negative and often significantly different from zero may be interpreted that one can make predictable profits by betting against the forward rate (Obstfeld and Rogoff, 1999).

Much research effort has been devoted to explaining the rejection of the UIP condition or the risk neutral efficient market hypothesis (RNEMH). Generally, the explanations can be classified into three groups. The first explanation states that it is due to expectational errors. The second explanation states that it is due to the foreign exchange risk premium. The third explanation is a modern view and focuses on the statistical properties of the data.

83 Equivalently, UIP and CIP together imply FRU which states that the forward rate should equal to the market expectation of the future spot rate,

$$E_t s_{t+k} = f_{t,k}.$$

In the expectational errors explanation, there is a failure of the rational expectations component of the joint hypothesis. The literature identifies at least four possible issues[84]: rational bubbles; learning about regime shifts (Lewis, 1989a,b), the peso problem (Rogoff, 1979; Evans and Lewis, 1995), or inefficient information processing (Bilson, 1981).

The research presented in this chapter focuses on the second and the third views, namely, the foreign exchange risk premium and the statistical properties of the data. A detailed discussion of these two explanations will be presented in the following paragraphs.

The foreign exchange risk premium explanation is theoretically based and states that the forward rates are biased predictors of actual exchange rate movements because there exists a risk premium on one country's currency relative to another's. In order to specify the relationship between the risk premium and the forward premium, we first consider the risk premia equation in (3.11). The risk premium term can be re-specified by subtracting and adding the spot rates on the left hand side of equation (3.11) as follows

$$\eta_{t,t+k} = (f_{t,k} - s_t) - (E_t s_{t,t+k} - s_t). \tag{3.17}$$

The forward premium now consists of two parts which are expected depreciation and the risk premium. The risk premium term is the payment for incurring risk of buying foreign currency.

Previous literature finds that the beta coefficient from ordinary least squares (OLS) estimate of equation (3.16) is negative. Fama (1984) originally attempted to explain this anomaly by arguing that the time varying risk premium term can lead to bias and inconsistency in the OLS estimates of b.

84 A detailed discussion of these issues can be found in Sarno and Taylor (2002b) and Lewis (1995).

Combining equations (3.13) and (3.17), the equation for the forward rate unbiasedness now becomes

$$\Delta s_{t,t+k} = a + b\left(f_{t,k} - s_t\right) + \left(\xi_{t,t+k} - \eta_{t,t+k}\right) + \varepsilon_{t,t+k}. \tag{3.18}$$

If $\eta_{t,t+k}$ and $(f_{t,k} - s_t)$ are correlated, then the OLS estimates are inconsistent.

A small positive or a negative slope coefficient in equation (3.16) might be explained by a rational expectation risk premium that is extremely variable (this is explained below). To better understand Fama's result, the asymptotic ordinary least square (OLS) estimate of the b coefficient of equation (3.16) is expressed as

$$\text{plim}(b^{OLS}) = \frac{Cov(f_{t,t+k} - s_t, \Delta s_{t,t+k})}{Var(f_{t,k} - s_t)}. \tag{3.19}$$

Following Obstfeld and Rogoff (1999)[85], define the risk premium, $\eta_{t,t+k}$ as the bias in the (log) forward premium, as in equation (3.11) and (3.17), with the implication that

$$(f_{t,k} - s_t) = (E_t s_{t,t+k} - s_t) + \eta_{t,t+k}. \tag{3.20}$$

Again, under rational expectations, the difference between the expected and the realised exchange rate or $\xi_{t,t+k}$ in equation (3.13) must be uncorrelated with all variables observable on date t, including that date's forward premium. In particular,

$$E_t\{(f_{t,k} - s_t)(s_{t+k} - E_t s_{t,t+k})\} = 0.$$

Thus, equation (3.19) under rational expectations can be rewritten as

$$\text{plim}(b^{OLS}) = \frac{Cov(f_{t,k} - s_t, E_t s_{t+k} - s_t)}{Var(f_{t,k} - s_t)}. \tag{3.21}$$

85 pages 588-591, Obstfeld and Rogoff (1999).

Fama's argument has 2 aspects. Firstly, if the resulting b from the OLS regression is negative in a large sample, then there must be negative covariance between the risk premium and the expected change in the spot rate. This can be illustrated by replacing $(f_{t,k} - s_t)$ in equation (3.21) with the right hand side of equation (3.20). Therefore, the numerator of equation (3.21) can be rewritten as

$$Cov(f_{t,k} - s_t, E_t s_{t+k} - s_t) = Var(E_t s_{t+k} - s_t) + Cov(E_t s_{t+k} - s_t, \eta_{t,t+k}).$$

Hence, $p\lim(b^{OLS})$ can be negative if $Cov(E_t s_{t+k} - s_t, \eta_{t,t+k}) < 0$.

Secondly, Fama (1984) argues that if the beta coefficient is estimated to be below $\frac{1}{2}$ in large sample, then the risk premium must be more variable than the expected change in the exchange rate. To illustrate this, multiply both sides of equation (3.21) by $Var(f_{t,k} - s_t)$, and again use equation (3.20) to substitute out for $f_{t,k} - s_t$. Noting that $Var(f_{t,k} - s_t) = Var[(E_t s_{t,t+k} - s_t) + \eta_{t,t+k}] = Var(E_t s_{t,t+k} - s_t) + 2Cov(E_t s_{t+k} - s_t, \eta_{t,t+k}) + Var(\eta_{t,t+k})$, equation (3.20) can now be written as

$$\begin{aligned} & \text{plim}(b^{OLS})\{Var(E_t s_{t,t+k} - s_t) + 2Cov(E_t s_{t+k} - s_t, \eta_{t,t+k}) + Var(\eta_{t,t+k})\} \\ = \;& Var(E_t s_{t+k} - s_t) + Cov(E_t s_{t+k} - s_t, \eta_{t,t+k}, \end{aligned}$$

where $\text{plim}(b^{OLS}) < \frac{1}{2}$ implies that

$$\frac{1}{2}[Var(E_t s_{t,t+k} - s_t) + Var(\eta_{t,t+k})] > Var(E_t s_{t,t+k} - s_t),$$

so that

$$Var(\eta_{t,t+k}) > Var(E_t s_{t+k} - s_t). \tag{3.22}$$

Hence, the risk premium must be more variable than expected future exchange rate changes for the estimated beta coefficient to be less than $\frac{1}{2}$. This is a significant challenge to attempts to model the exchange rate risk premium. However, Obstfeld and Rogoff (1999) give warning that the puzzle posed by inequality in equation (3.22) should not be overstated since the expected exchange rate change appears to be empirically small in major currencies. It is not easy to reject the hypothesis that the log exchange rates follow a random walk[86], in which case $Var(E_t s_{t+k} - s_t) = 0$. Thus the surprising fact may be that the expected exchange rate change are typically very small, and not that the variance of the risk premium is so large.

Generally, inference from empirical analysis of equation (3.16) has been conducted under the assumption of a short memory forward premium and exchange rate return. However, this assumption is incorrect if one of these two series are empirically nonstationary. Recent research has cast doubt on the validity of the conventional wisdom that the forward premium is stationary; see e.g. Crowder, 1994; Evans and Lewis, 1995; and Mark, Wu and Hai, 1993. The empirical evidence shows that the forward premium series are in fact highly persistent. This leads us to the last view of the explanation for the rejection in UIP hypothesis.

The last view explains the rejection of the UIP hypothesis by focusing on the statistical artifacts of the data. As the modern literature pays more attention to the time series properties of the returns on the nominal spot exchange rate and the forward premium, a possible explanation for a negative beta coefficient in equation (3.16) is based on 1) the long memory behaviour of the forward premium; and 2) the existence of structural breaks in the forward premium. Baillie and Bollerslev (2000) proposed that the anomaly is caused

86 See Obstfeld and Rogoff (1999) page 591.

by a very persistent autocorrelation in the forward premium. The forward premium series tend to follow fractionally integrated (long memory) process, while the rate of return on the spot exchange rate is a stationary process. This creates an unbalanced test regression in equation (3.16). Maynard and Phillips (2001) also found evidence of fractionally integrated behaviour in the forward premium. They supported Baillie and Bollerslev's (2000) argument that the traditional asymptotic regression for unbiasedness may not be suitable due to the difference in persistence between the two series mentioned above. Recently, Choi and Zivot (2005) pointed out the importance of structural breaks and confirmed that both explanations are important.

To illustrate this further, we refer back to equation (3.18). The return on the spot exchange rate, $\Delta s_{t,t+k}$ and the forecast error, $\xi_{t,t+k}$ are widely accepted to be stationary (Cornell, 1977; Meese and Singleton, 1982; Corbae and Ouliaris, 1986, Baillie and Bollerslev, 1989; Baillie and Bollerslev, 1993; Maynard and Phillips, 2001; Choi and Zivot, 2005). Empirically, the logarithm of the spot rate is found to be a non-stationary $I(1)$ variable, and hence $\Delta s_{t,t+k}$ is $I(0)$. The rational expectation forecast error, $\xi_{t,t+k}$ itself must be stationary $I(0)$, so that it would not be forecastable from the past information.

The study of the persistence of the forward premium has recently gained substantial interest from researchers. Initially, the studies of Crowder, 1994; Evans and Lewis, 1995; and Mark, Wu and Hai, 1993, argued that the forward premium is non-stationary. However, later work rejects the view that the forward premium has a unit root. Instead, the forward premium is found to follow a fractionally integrated process (Choi and Zivot, 2005; Baillie and Bollerslev, 1994a; Maynard and Phillips, 2001).

Among the first group of papers that suggested a unit root process in the forward premium, Crowder (1994) tested for unit root in daily forward premium data from 1980-1985 in British pound, German Deutsche mark and Canadian dollar, all relative to the US dollar. He found that the null of an $I(0)$ process in the forward premium is rejected using the KPSS test of Kwiatowski, Phillips, Schmidt and Shin (1992) while the unit root hypothesis was not rejected in these three forward premium series using augmented Dickey-Fuller (ADF) tests[87].

Baillie and Bollerslev (1994a) argued that Crowder's findings do not necessary guarantee a unit root in the forward discount. This is because the test methodology in Crowder's work forced a choice between $I(0)$ and $I(1)$ processes. They suggested that the degree of persistence has led Crowder's and other studies to erroneously conclude that the forward premium consists of a unit root process. Their study also provided a good literature review that calling for a careful interpretation of the KPSS and ADF tests. To quote Baillie and Bollerslev (1994a, pp. 566) "...although Crowder's KPSS test statistics are providing evidence against $I(0)$ behaviour, this should not automatically be interpreted as being suggestive of an $I(1)$ process. Similarly, as demonstrated by Diebold and Rudebusch (1989), the conventional ADF test for a unit root, or $I(1)$ behaviour, has very low power against fractionally integrated alternatives."

Baillie and Bollerslev (1994a) allow for a fractionally integrated $I(d)$ process in the forward premium, for $0 < d < 1$. They use the same data as Crowder (1994) to compare

87 The other two papers that argue in favour of unit root in the forward premium are Evans and Lewis (1995) and Mark, Wu and Hai (1993). However, they dispense the unit root test and proceed the test for cointegraiting relationship between the spot and forward rates which is the main objective of their papers.

the correlograms of the spot exchange rate, the return on the spot exchange rate, and the forward premium, and then estimate ARFIMA models of the forward premium. Monthly exchange rate data for Canada, Germany and UK are used from January 1994 to December 1991. The spot rate (s_t) itself showed strong evidence of a unit root while the degree of persistence in the forward premium rate's ($f_t - s_t$) autocorrelations is substantially less. Although the effects of innovations in the forward premium series are moderately persistent, these innovations eventually die out at a slow hyperbolic rate of decay. Thus ARFIMA's models are suggested to be better estimators for the forward premium. The models report the point estimate for the order of fractional integration, equal to 0.445, 0.767, and 0.551 for Canada, Germany and UK, respectively (each exchange rate series is in terms of number of US dollars per unit of foreign currency). Interestingly, Baillie and Bollerslev's results imply a prediction for the risk premium series. As the forward premium follows a fractionally integrated process while the spot rates is $I(1)$; therefore, the risk premium itself should be fractionally integrated according to equation (3.18).

Maynard and Phillips (2001) joined Baillie and Bollerslev (1994a) in finding evidence of a fractionally integrated process in the forward premium using ARFIMA models, for the case of higher frequency data. The data in their study are daily from November 1986 to March 1988. The samples are from 6 major currencies, in particular AUS, CAD, FR, DM, YEN, GBP against the US Dollar. The results show that there is evidence of non-stationary long memory behaviour in the forward premium. The model reports that all estimates of the fractionally integrated parameters are between $0.858 \leq d < 1$. Their paper concluded that the principal failure of unbiasedness hypothesis is the mismatch in persistence between the spot return and the forward premium series. Additionally, they argued that traditional asymptotic theory may not be applicable to test forward rate unbiasedness hypothesis. The non-standard nature of the limit theory is primarily attributable to the long memory of the forward premium. They propose new hypothesis test statistics which have nonstandard limiting distributions with long left tails, which may explain the forward premium anomaly as statistical artifacts.

Choi and Zivot (2005) suggested that the long memory property in the data may be due to the presence of structural breaks or regime switches. They test for and estimate the multiple mean break model by Bai and Perron (1998, 2003a) and then adjust for the structural breaks in the forward premium. Allowing for structural breaks drastically reduces the degree of persistency in the forward premium. However, evidence of (stationary) long memory in all of the forward premium series still persists. They use monthly exchange rate data in terms of US dollars for five G7 countries: Germany, France, Italy, Canada and UK over the period January 1976 to January 1999. The corresponding estimated value of the fractional differencing parameters after adjusting for structural breaks are 0.284, 0.332, 0.374, 0.516 and 0.357, respectively, which are lower than previous estimates that don't incorporate structural breaks.

3.3 Risk premia

This section aims to present the literature reviews for the estimation of the foreign exchange risk premium. There are two approaches used within the literature in interpreting the risk premia.

The first approach is a traditional approach which examines various specifications of the fundamental determinants of risk premium. The explanation of the risk premia is through the standard discrete time consumption-based asset pricing (CAPM) model of Lucas (1982), which takes into account the real returns to forward market speculation. Studies using this approach generally proxy the risk premia (in equation (3.11)) by the conditional heteroskedasticity or time dependence between the conditional covariances of the exchange rate and the fundamental variables such as consumption and prices e.g. Engel and Rodrigues (1989), Hodrick (1989), Kaminsky and Peruga (1990), Mark (1988) and Baillie and Bollerslev (1990). First of all, agents optimize their consumption plan (C_t, C_{t+k}) such that the expected real returns of the current and future consumption streams of the representative investor in the forward market must be zero[88]. That is,

$$E_t \left[\left(\frac{F_t - S_{t+k}}{P_{t+k}^D} \right) \left(\frac{U'(C_{t+k})}{U'(C_t)} \right) \right] = 0, \tag{3.23}$$

where P_{t+1}^D is the time $t+k$ domestic price level, $U'(C_{t+k})/U'(C_t)$ equals marginal rate of substitution in terms of utility derived from current and future consumption, and F_t and S_{t+k} are forward exchange rate and future spot exchange rate. If the utility function is characterized by the constant relative risk aversion[89], equation (3.23) can be rewritten as

$$E_t \left[\left(\frac{F_t - S_{t+k}}{P_{t+k}^D} \right) \left(\frac{C_t}{C_{t+k}} \right)^{\varkappa} \right] = 0, \tag{3.24}$$

where $\varkappa$ is the coefficient of relative risk aversion. Assuming that all variables in equation (3.24) are jointly log normally distributed (Sarno and Taylor, 2002), the Taylor series

88 See Sarno and Taylor (2002b) page 22-23 for the full derivation.

89 In case of a constant relative risk aversion, agent's utility function can be written as

$$U(C_t) = (1-\varkappa)^{-1} C_t^{1-\varkappa}.$$

expansion to the second order of this equation is as follows,

$$f_{t,k} - E_t s_{t+k} = 0.5var_t(s_{t+k}) - cov_t(s_{t+k}, p^D_{t+k}) - \varkappa cov_t(s_{t+k}, q^D_{t+k}), \tag{3.25}$$

where p^D_{t+k} is a natural log of the domestic price level, q^D_{t+k} denotes the logarithm of intertemporal marginal rate of substitution $(U'(C_{t+k})/U'(C_t))$ of domestic agents and var_t and cov_t denote conditional variance and covariance based on information available at time t. According to equation (3.11), $f_{t,k} - E_t s_{t+k}$ is the foreign exchange risk premium. Thus, using CAPM methodology, the risk premium can be estimated by the right hand side of equation (3.25).

Note that despite the expected real profit from forward rate speculation being zero (as in equations (3.23) and (3.24)), there is a wedge between the expected spot rate and the forward rate when they are expressed in log form (as in equation (3.25)). The wedge contains 3 conditional second moment terms as expressed on the right hand side of equation (3.25). FRU is violated if any or all of these three terms are significantly different from zero.

The simple asset pricing model assumes that agents are completely risk neutral and care about the real returns, hence $\varkappa = 0$. Equation (3.24) then reduces to

$$E_t\left[\frac{F_{t,t+k} - S_{t+k}}{P^D_{t+k}}\right] = 0. \tag{3.26}$$

Assuming again that prices and exchange rates are jointly lognormally distributed, equation (3.26) can be written as follows,

$$f_{t,k} - E_t s_{t+k} = 0.5var_t(s_{t+k}) - cov_t(s_{t+k}, p^D_{t+k}), \tag{3.27}$$

Note that even under rational expectations and risk neutrality ($\varkappa = 0$), the right hand side of equation (3.27) still contains the two conditional second moment terms. These are called

Jensen Inequality terms (JIT). Including the JIT terms, the FRU does not hold even if agents are risk neutral. These two terms imply that risk neutrality does not imply the UIP (see Frenkel and Razin, 1980 and Engel, 1984 for a more detailed discussion).

Some papers define the term $cov_t(s_{t+k}, q^D_{t+k})$ in equation (3.25) as the time dependent risk premium. However, using this measure when derived from the CAPM model is largely unsuccessful in explaining the deviations from UIP. Kaminsky and Peruga (1990) estimated equation (3.25) using monthly data on consumption and prices and proxying the conditional covariance of s_{t+k} with q^D_{t+k} using an ARCH model. Unfortunately, it is found to be negligible. Other researches also support their view such as that by Baillie and Bollerslev, 1989, 1990; and Bekaert and Hodrick, 1993. They argue that the time paths of consumption and prices are relatively smooth and the foreign exchange rate data show little evidence of conditional heteroskedasticity, thus the influence of the second and third terms on the right hand side of (3.25) should be empirically small. Moreover, Lewis (1995) finds that empirically the covariance between exchange rates and inflation, $cov_t(s_{t+k}, p^D_{t+k})$ is quite small and near zero. Engel (1984) and Cumby (1988) have also found that the behavior of the excess return in real terms (the term on the left hand side of equation (3.26)) is not very much different from the nominal term[90]. Thus it is unlikely that this term can help explaining important fraction of excess return behaviour.

Other works that apply the Lucas model in equation (3.24) in explaining the risk premium are Hansen and Hodrick, 1980; Hodrick and Srivastava, 1984, 1986; Giovannini and Jorion, 1987; Bekaert and Hodrick, 1992 and Bekaert, 1994. This research in general

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90 When taking the Taylor series expansion of equations (3.24) and (3.26) to obtain the risk premium expression, we obtain $cov_t(s_{t+k}, p^D_{t+k})$. Engel (1984) and Cumby (1988) argue that this term is small.

finds that the model has limited success in explaining the risk premium. A key problem is that the risk premium model is not robust across different datasets and time periods (Lewis, 1995). Moreover, as in the equity premium puzzle, it is difficult to rationalize a large absolute value risk premium because the consumption series is not that variable. For the risk premium to explain a significant portion of the forward rate forecast error or excess returns, either there must be a very large coefficient of relative risk aversion $\varkappa$, or consumption must be highly correlated with the exchange rate. Sarno and Taylor (2002b) provide a rationale for this latter possibility by arguing that the less that the forward exchange position provides a hedge against variations in future consumption, the greater the covariation between the consumption and exchange rate. However, they also point out that consumption empirically tends to be fairly smooth in advanced economies, while the nominal exchange rate in these countries is typically a lot more volatile. Thus the covariations between price and consumption, and exchange rate are empirically found to be small.

Moreover, Baillie and Bollerslev (1989, 1990) and Bekaert and Hodrick (1993) estimate equation (3.27) by using alternative forms of ARCH, and GARCH in mean specifications for $var_t(s_{t+k})$. They found that $var_t(s_{t+k})$ is not statistically significant and that omitting it from the UIP regression does not affect the estimates of b in equation (3.18), which remains negative. Cuthbertson and Nitzsche (2004) also find that in general the JIT in the foreign exchange market are small and can be ignored in practice. Thus, we can conclude that estimating the foreign exchange risk premium by CAPM model does not yield satisfactory results in the literature.

Obstfeld and Rogoff (1999) pointed out the problem in understanding foreign exchange risk premium that it changes sign as expected depreciation does (as illustrated in equation (3.16)). Sometimes the risk premium runs against a country's currency and sometimes in favour of it. The extension of the simple CAPM analysis is to relate the stochastic properties of global output to forward premium anomaly. Under assumption that each country consumption growth is proportional to world income growth with CRRA utility, they suggested replacing the former with the growth rate of per capita global output[91] in the Euler equation (3.24) as follows,

$$E_t\left[\left(\frac{F_t - S_{t+k}}{P^D_{t+k}}\right)\left(\frac{Y^w_t}{Y^w_{t+k}}\right)^{\varkappa}\right] = 0, \tag{3.28}$$

where Y^w_t is per capita global output at time t. Approximate equation (3.28) by Taylor series expansion yields

$$f_{t,k} - E_t s_{t+k} = 0.5var_t(s_{t+k}) - cov_t(s_{t+k}, p^D_{t+k}) - \varkappa cov_t(s_{t+k}, y^w_{t+k}), \tag{3.29}$$

where y^w_{t+k} is the natural log of Y^w_{t+k} and the covariance term $cov_t(s_{t+k}, y^w_{t+k})$ can change sign over the course of the world business cycle. However, no study has yet succeed in using the equation (3.29) for regression analysis. This is not surprising. Finding the per capita global output is straight forward but data is only available annually. Thus there is problem of a limited number of observations in the regression.

The second approach specifies "statistical" models of the risk premium. This approach tests for certain patterns in or across excess exchange rate returns. This approach is introduced by Domowitz and Hakkio (1985) who give an explanation of the risk premia

91 See detailed derivation in Obstfeld and Rogoff (1999) page 580 and 592.

through the conditional variance of the market forecast error. Their work proceeds test to the UIP condition using data from the United Kingdom, France, Germany and Switzerland during Jun. 1973- Sept. 1982 using monthly data[92]. They argued that the inter-temporal capital asset pricing model is highly stylised, however, it is quite restrictive in its assumption on preferences and technology, and the direct estimation and testing of the model is difficult. They thus suggested using the econometric model to capture some major aspects of risk in a foreign exchange contract. The risk premia are defined as a function of the conditional variance of the error in forecasting the spot rate using the forward rate,

$$\eta_{t,t+k} = var_t(f_{t,k} - s_{t+k}), \tag{3.30}$$

where the conditional variance of the forecast errors $(f_{t,k} - s_{t+k})$ were assumed to follow the ARCH in mean process introduced by Engel, Lilien and Robins (1987). The test for UIP is to add the risk premium term to equation (3.16) and perform an OLS regression but again it was still rejected for the majority of currencies. For this study, the monthly data in United Kingdom and Japan exhibit a significant risk premium in the regression, while the data in Germany, France and Switzerland did not reject the null hypothesis of no risk premium. Furthermore, the sign of the risk premium varies for different time periods. Finally, the authors suggest two alternative approaches for future research. The first suggestion is to specify the risk premium as a function of the conditional covariance matrix for all currencies. Baillie and Bollerslev (1990) explore this possibility. The second suggestion is to define the risk premium as the conditional variances of the forecast errors of the domestic and foreign money supplies. Up to my knowledge, no work pursues this methodology.

92 This was a floating currency period.

Baillie and Bollerslev (1990) suggest that the reason for the lack of success in using measures of the risk premia to explain the rejection of the UIP is probably because of the assumption that the conditional covariance of future spot rates are time invariant as in Domowitz and Hakkio (1985). Instead of testing for the unbiasedness hypothesis by performing the OLS regression as in Domowitz and Hakkio (1985), the main objective of their research is to find proper specification of the foreign exchange risk premium and explain the possibility that the rejection of the unbiasedness hypothesis is due to the existence of a time varying exchange rate risk premia. They examine the conditional variance and covariance of the forward rate forecast error using a multivariate GARCH model. The forward rate forecast error is defined as the difference between the future spot rate and the forward rate, $s_{t+k} - f_{t,k}$. The equation for forward rate forecast error is

$$s_{t+k} - f_{t,k} = \eta_{t,t+k} + u_{t+k}, \tag{3.31}$$

when $\eta_{t,t+k}$ is the foreign exchange risk premia and u_{t+k} is a random innovations in the period between the market setting the forward exchange rate and the actual realisation of the spot rate k periods later. The risk premia is proxied by the conditional variance of the multivariate GARCH(1,1) model of the forward rate forecast error[93]. They consider the time series modelling of u_{t+k} as the $[k]$th-order moving average process such that

$$u_{t+k} = \sum_{j=1}^{[k]} \tilde{\delta}_j \varepsilon_{t+k-j} + \varepsilon_{t+k}, \tag{3.32}$$

93 Baillie and Bollerslev (1990) proxy the risk premium by the first term of equation (3.25) i.e. the conditional variance of the future spot exchange rate ($Var_t(s_{t+k})$) is equivalent to the conditional variance of the forward rate forecast error ($Var_t(s_{t+k} - f_{t,t+k})$).

where the innovations, ε_{t+k} are serially uncorrelated with mean zero and finite unconditional variance. Data used in their study are weekly spot and one-month maturity forward rates (approximately 4 weeks maturity). They thus set the autocorrelation structure for u_{t+k} as the 4th order moving average ($k = 4$).

They suggested that the weekly exchange rate data should possess more ARCH type behavior than the monthly data used for example by Domowitz and Hakkio (1985)[94] and examines 4 currencies: Pound sterling, Swiss Franc, German Mark, and French Franc, all of them are against the US dollar from Mar. 1980 to Feb. 1989. Thus, $\eth_j$ is generalised to be a matrices of four sample currencies in the multivariate framework.

In deriving the implied model for the forward rate forecast error using MA(4), three assumptions must hold, namely, risk neutrality $(\eta_{t,t+k})$, rational expectation $[E_t\,(u_{t+k}) = 0]$, and that the first differences in nominal exchange rates follow an uncorrelated process. The purpose for parametrising the forward rate forecast errors as the moving average process (MA(4)) is to use it to provide some bounds on the magnitude of the risk premium. They then perform the check for the presence of the time varying risk premia by comparing sample estimates of the residual variances with those implied by the MA(4) process.

In testing for the significance (of the proxy) of the time varying risk premia, they use the Lagrange Multiplier (LM) test for the inclusion of the own conditional variance in explaining any deviation of the forward rate forecast error from the fixed MA(4) process. Unfortunately, the test for the presence of the time varying risk premium gains little support

94 The frequency matters in characterising the ARCH effect. Bailie and Bollerslev (1989) note a sharp decline in ARCH effects, as the sampling frequency declines from weekly, fortnightly, and monthly. Other studies such as Booth and Glassman (1987), Domowitz and Hakkio (1985), and Hodrick (1989) studied the monthly spot exchange rate data and observed only minimal ARCH effects and approximate normality.

which confirms the finding in Domowitz and Hakkio (1985). Although the ARCH effects appear to be much more pronounced with weekly data, the conditional variance remains an insignificant determinant of the forward rate forecast error for three of the four currencies. Using the conditional covariance between the currencies as proxies for the other components of the time varying risk premia also shows no much support to the idea the risk premium is a simple linear function of the corresponding covariances. Only the UK data shows some evidence that the conditional covariances give explanation to the data in addition to the own conditional variances.

From the literature, we can make some concluding remarks and provide research agenda as follows.

The estimation of the risk premium by the CAPM model does not gain much successfulness due to the limitation of the data, e.g., a small covariation between consumption and exchange rate since the data of the former is fairly smooth (Kaminsky and Peruga, 1990; Sarno and Taylor, 2002; Baillie and Bollerslev 1989, 1990; and Bekaert and Hodrick, 1993)), and a small covariation between exchange rate and inflation (Lewis, 1995; Engel, 1984; and Cumby, 1988) and the foreign exchange rate data show little evidence of conditional heteroskedasticity (Baillie and Bollerslev 1989, 1990; and Bekaert and Hodrick, 1993).

The second approach which specifies "statistical" models of the risk premium is not yet very much successful; however, there is a room for new research. Although the test for the presence of time varying risk premium gains little support when using monthly data from 1973-1982 in Domowitz and Hakkio (1985). We argue that there is a possibility that using the higher frequency data and longer time span would give more information to the model. Moreover, in estimating the conditional variance of the forecast error, we can find a model that fits the data better than the ARCH-in-mean model and the multivariate GARCH model. Thus, this chapter employ the second approach in estimating the risk premium (in equation (3.30)) to the test of UIP (in equation (3.18)) hypothesis as follow

$$\Delta s_{t,t+k} = a + b\left(f_{t,k} - s_t\right) + \varkappa var\left(s_{t+k} - f_{t,k}\right) + \varepsilon_{t,t+k}. \tag{3.33}$$

The next section aims to study the time series properties of the three variables in equation (3.33).

3.4 The Variables and their time series properties

This section aims to describe the time series properties of three variables in equation (3.33) which are the return on the spot rate, the forward premium and the forward rate forecast error and their source of data.

The first two are variables in the regression for the uncovered interest parity. $\Delta s_{t,t+k} = s_{t+k} - s_t$ is defined as the realised exchange rate depreciation. It is the difference between natural log of the spot rate at date t and natural log of the spot rate at date $t+k$, where k equals to 1 month maturity. $(f_{t,k} - s_t)$ is the forward premium which is the difference between natural log of the time t forward rate with maturity k and natural log of the time t spot rate. We multiply the natural log of all rates by 100, so that the forward premium and the spot rate return are in percentages[95].

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95 $100 * (\ln f_{t,k} - \ln s_t)$ and $100 * (\ln s_{t+k} - \ln s_t)$, respectively.

The third variable is the forward rate forecast error which is defined as $s_{t+k} - f_{t,k}$. Again, we multiply it by 100 so that the forward rate forecast error is the percentage deviation of the realized spot exchange rate k period later from the forward contract rate. The conditional variance of the forward rate forecast error is the proxy for the time varying exchange rate risk premia ($\eta_{t,t+k}$ in equations (3.17) and (3.18)). This is in line with Domowitz and Hakkio (1985) and Baillie and Bollerslev (1990). The maturity of the forward rate forecast error, k in this study is 1 month.

The foreign exchange rates and the forward rates data were provided by Bloomberg. These data are market rates provided by the central bank of each country. We consider the exchange rate data on the eight most traded currencies in the foreign exchange market which are the Eurozone Euro (EUR), the British Pound (GBP), the Japanese Yen (JPY), the Swiss Franc (CHF), the Australian Dollar (AUD), the Canadian Dollar (CAD), the Swedish Krona (SEK), and the Norwegian Krone (NOK). Two additional sample countries are the New Zealand Dollar (NZD) and the Danish Krone (DKK). The base currency is the United States Dollar (USD). The sampling interval is from January 1994 to June 2007, except for the EUR, NOK and SEK, which starts in January 1999. The sampling frequency of the exchange rate data is daily. The spot and forward exchange rate data are the average value of the day. However, these daily data are available on trading days only i.e. excluding weekends and holidays. The length or maturity of the forward rate contract k, is defined as 1 month or approximately 20 working days.

Considering that the sampling interval of the forward rate forecast error series ($s_{t+k} - f_{t,k}$) is daily which is finer than the forecast interval (monthly), the forward rate forecast errors will be serially correlated (see Hansen and Hodrick, 1980; Hakkio, 1981; Baillie 1989; and Baillie and Bollerslev, 1990). In order to deal with this problem, this chapter models the conditional mean of the forward rate forecast error as the k-th order moving average process. This is in line with Baillie and Bollerslev (1990) as stated in equation (3.32). This chapter analysed one month forward contracts and daily data, the maturity of the contract equals 20 working days, we thus apply MA(20) to filter out the serial correlation from the forward rate forecast error series. This case also applies to the spot return, $\Delta s_{t,t+k}$ series which will also be filtered by MA(20) process.

Table 3.1 provides a brief description of the exchange rate data and the summary statistics for the calculated 1 month return on the spot exchange rates or the realised exchange rate depreciation $\Delta s_{t,t+k}$, the calculated 1 month forward premium $(f_{t,k} - s_t)$, and the calculated forward rate forecast error $(s_{t+k} - f_{t,k})$. Before proceeding to the empirical results, it is also useful to discuss the time series properties of these three series by observing their time series graphically and the sample autocorrelation functions.

Plots of the forward premium $(f_{t,k} - s_t)$ series in the 10 sample countries are presented in figures 3.1A-3.1J. Most of series show structural breaks and persistence in the mean. The plots of the autocorrelation function (ACF) of the forward premium series of these countries are presented in figures 3.2A-3.2J. The ACF of the forward premium series all countries have the similar behaviour. There is no evidence of seasonality or any cyclical patterns and the series are not white noise since at least one autocorrelation coefficient in each case is different from zero. However, the series tend to be nonstationary as the autocorrelation coefficients decays less than exponentially as the lag length increases. Instead, the autocorrelation coefficients at various lags are very high even up to a lag of 200-300 working days. Considering the Box-Pierce Q statistics, one can reject the null hypothesis that all the true autocorrelation coefficients up to lag 200-300 are equal to zero. In facts, all the lags are highly significant. This indicates that the forward premium series may display long memory. It is best modelled by a fractionally integrated ARMA process.

Plots of the spot return ($\Delta s_{t,t+k} = s_{t+k} - s_t$) filtered by the MA(20) process are presented in figures 3.3A-3.3J. The series tend to be stationary. Plots of the autocorrelation coefficients for the spot return in figures 3.4A-3.4J also show that they do not have long memory behaviour. To further confirm this finding, the stationarity of the spot return series will be examined in the next section using ADF and KPSS tests.

Figures 3.5A to 3.5J present time series plots of the forward rate forecast error ($s_{t+k} - f_{t,k}$) on 10 currencies (in term of US Dollar) filtered by MA(20) process. The volatility of the forward rate forecast error appears to be serially correlated. The individual series vary wildly, but they do so within a range which itself changes slowly over time. In general, the forecast error lies in the range between -2 percentage points to 2 percentage points, excluding CAD where the range of the series variability is much narrower. For CAD, the series ranges from around -1 percentage points to 1 percentage points. This is not surprising, since the US and Canadian dollar follow each other very closely.

The volatility of the series in AUD, CAD, and NZD during 1998-2007 appears to be higher than the previous periods. Considering figures 3.6A-3.6J which are time series plots of the absolute value of the forward rate forecast error, the fat periods for these 3 currencies become more clearly visible. For JPY, during the period around 1997-1998 when the Asian economic crisis took place, volatility is higher than in the other periods. However, all the currencies show clear evidence of changing volatility through time. Moreover, the series are characterized by volatility clustering where high volatility tend to be surrounded by other high volatility days (and vice versa). This is a clear sign of presence of ARCH effects in the series.

Figures 3.7A-3.7J illustrate plots of the autocorrelation coefficients for the forward rate forecast error after filtering for the MA(20) process. Generally the series of the forward rate forecast error shows no trend behaviour or seasonality. The decay of the sample autocorrelation coefficient does not indicate nonstationary in the series. Thus, there is no evidence of positive predictability from past values of the forward rate forecast error to its current value (a short memory process in the level). However, considering the autocorrelation coefficients of the absolute value and the squared value, the errors are strongly time dependent, the autocorrelations tend to be fairly large even for large values of lags which confirms the volatility clustering in the series.

3.5 Empirical evidence of the forward premium anomaly: the rolling regression

In this section, we perform an empirical analysis of the forward premium anomaly during 19904-2007 using traditional regression methodology. We perform the forward market efficiency test and ask whether one can reject the null hypothesis that $a = 0, b = 1$ in equation (3.16). If the forward rate contains all available information at time t, then $\varepsilon_{t,t+k}$

should be serially uncorrelated. Hence,

$$\Delta s_{t,t+k} = a + b\left(f_{t,k} - s_t\right) + \varepsilon_{t,t+k}. \tag{3.34}$$

As described earlier, the sampling frequency of the exchange rate data is daily. The length or maturity of the forward rate contract is specified by k, which is defined as 1 month or approximately 20 working days. Thus, $\Delta s_{t,t+k} = s_{t+k} - s_t$ is defined as the realised exchange rate depreciation or the difference between the spot rate at date t and the spot rate at date $t + 20$. $(f_{t,k} - s_t)$ is the forward premium or the difference between the time t forward rate and the spot rate. The null hypothesis tested is that $a = 0$, $b = 1$ and $\varepsilon_{t,t+k}$ has a conditional mean of zero.

When using daily data in the UIP regression, one needs to be careful with the data overlapping issue. If the sampling frequency is equal to the maturity time of the forward contract, so that $k = 1$, then $\varepsilon_{t,t+k}$ will be serially uncorrelated. If $k > 1$, this gives rise to the overlapping data problem and $(k-1)th$ order serial correlation (autocorrelation) in $\varepsilon_{t,t+k}$. When $\varepsilon_{t,t+k}$ is serially correlated, the usual OLS standard errors estimates are biased (Greene, 2000). Consequently, one needs to adjust the standard errors of the regressors. We use the Newey-West (1987) Heteroskedasticity and Autocorrelation Consistent estimator (HAC)[96]. This is a robust, consistent estimator for autocorrelated disturbances.

In this case the sampling frequency is daily but the maturity of the forward contract is 1 month, hence the data overlap. Thus, the HAC is employed to correct the heteroskedas-

96 See Greene (2000), pages 464 and 537 for further details of the Newey-West consistent estimator. Using HAC, it must be determined in advance how large is the lag order (k) of the series. The maximum lag, k must be large enough that the autocorrelations at lags longer than k are small enough to ignore.

ticity and autocorrelation issue. The organisation this section is that: we first perform the UIP regression using the HAC methodology, and then we perform the rolling regression.

Table 3.2 presents the estimated slope coefficient $\widehat{b}$ coefficients of the UIP regression (in equation (4.64)) based on the daily data for 10 currencies and their standard error adjusted by Newey and West (1987). The last column of the table presents the test statistics for the null hypothesis of estimated beta equal to untiy. The results base on 95 percent confidence interval. The results show that the estimate for $\widehat{b}$ significantly different from the UIP hypothesis of unity in all currencies except in the case of GBP. Moreover, the estimated beta coefficients show negative sign in all sample currencies. The implication for the negative beta coefficient is that a positive forward premium was, over the period of studies, associated with their currency depreciation (in unit of US dollar) for all currencies. This finding is consistent with previous literature such as Frankel, 1980; Fama, 1984; Bekaert and Hodrick, 1993; and Baillie and Bollerslev (2000). At the same time, the R^2 is very low in all currencies, and it is in the range of (0.001, 0.043). This indicates that the forward premium actually explains very little of the spot rate return in the UIP regression (equation (4.64)). This finding is the same as other literatures mentioned above using other currencies and sampling frequencies. Thus, the regression results in all sample currencies confirm the earlier finding of the forward premium anomalies.

We next perform the rolling regression to test for the UIP hypothesis. Sarno and Taylor (2002b) point out that the regression based UIP test used in the previous literature assumes that the beta coefficient is constant over some interval of time and therefore that it shows the parity condition over the period of time. The rolling regression in this study partially corrects this imperfection. With the rolling regression, the beta coefficient measured over the subsample moves forward through time.

In every country, the estimates of β are from 5-year rolling regression[97] (or 1,300 trading days). This approach first estimates equations (3.16) using data from day 1 through day 1,300 of the dataset, then using data from day 2 through day 1,301, day 3 through day 1,302, and so on[98]. In other words, in the rolling regression, the number of observations is held constant and the starting and ending points are shifted.

The resulting beta coefficients from the rolling regressions and the 95 percent confidence interval based on the HAC standard errors are plotted in figures 3.8A-3.8J. The beta equal to unity line is drawn which represents the point where the uncovered parity condition holds. Using the 95 percent confidence interval, the beta coefficients for the regressions with one month maturity are significantly different from unity in most of the period of study in all currencies, except GBP[99]. Additionally, the beta coefficients from the rolling regression are also less than zero in most cases. This is consistent with the empirical research discussed in the previous section.

Other research that performs the rolling regression to test for the unbiasedness hypothesis is in Baillie and Bollerslev (2000). They test for the parity condition in German DM/ US dollar using monthly data from Mar. 1973 to Nov. 1995. The results show that

97 Alternatively, one could also perform recursive and reverse recursive analyses, in which the starting or ending time period is held fixed and the sample size grows. I did not select these 2 analyses. Recursive analyses are often used in forecasting situations as information become more available with time (see Kmenta; 1997, page 423-424) and this is a different application.

98 The beta coefficient obtained from the rolling regression is based on the trading dates not the calendar dates.

99 Note that initially, I also study the case for maturity of 1 week, 3 months, 6 months and 12 months. The results show that the longer the maturity, beta becomes less closer to unity. In most cases, beta becomes even more negative.

the estimated slope coefficient in equation (4.64) significantly different form unity and it is also significantly less than zero using 5 year rolling regression.

The concluding remarks for this section is that the estimated slope coefficients in equation (4.64) is likely to be uninformative about the true value of b, resulting in a forward premium anomaly.

In the next section, we consider an empirical evidence of "mismatch in persistence", which is the main source of the forward premium anomaly. The purpose of the following sections is to argue that the principal failure of the unbiasedness hypothesis (as in equation (4.64)) is the difference in persistence between the spot return series and the forward premium series. The spot return ($\Delta s_{t,t+k}$) series is stationary while the forward premium ($f_{t,k} - s_t$) series has long memory. Thus, the UIP regression is invalid and the slope coefficients from OLS regression are highly biased. Additionally, the presence of structural breaks in the forward premium series is found to increase the persistence of the series.

3.6 Empirical evidence of the fractional integration in the forward premium

Empirical evidence (Baillie and Bollerslev, 1994a; Maynard and Phillips, 2001; and Choi and Zivot, 2005) supports the view that the forward premium series is fractionally integrated with a differencing parameter that is significantly different from zero and unity. In this section, we extend previous analysis by using the daily forward premium data for ten mostly traded currencies during Jan. 1994 - Jun. 2007. We find that a fractionally integrated model appears to fit the data quite well. The framework for ARFIMA model, method of estimation and the model identification, and the results will be discussed in detail as follows.

3.6.1 The ARFIMA model

This section closely follows Baillie and Bollerslev (1994a). For any covariance stationary time series process y_t, the impulse response weights or moving average representation is given by,

$$y_t = \sum_{j=0}^{\infty} \psi_j \varepsilon_{t-j},$$

where $t = 1, ..., T$. ε_t is Gaussian white noise with mean zero, finite variance σ^2, and serially uncorrelated, and $\sum_{j=0}^{\infty} \psi_j < \infty$.

The basic ARMA(p, q) model of the y_t process is

$$y_t = \phi_1 y_{t-1} + ... + \phi_p y_{t-p} + \varepsilon_t + \theta_1 \varepsilon_{t-1} + ... + \theta_1 \varepsilon_{t-q},$$

the $I(0)$ property associated with stationary and invertible ARMA models imposes exponentially decaying impulse response weights on the y_t process and corresponding exponentially decaying autocorrelation coefficients. On the other hand, non-stationary $I(1)$ process imply *complete* persistence on both the impulse response weights and the autocorrelation coefficients.

A more flexible class of process is the ARFIMA(p, d, q) model, introduced by Granger and Joyeux (1980), Granger (1980, 1981) and Hosking (1981). This model is useful for series that exhibit significant autocorrelation between observations widely separated in time, and for example is suitable for financial data such as inflation rates, exchange rates and interest rates. In such a case, y_t is said to display long memory (or long-term dependence) and may be best modelled by a fractionally integrated ARMA process. The ARFIMA(p, d, q) model for y_t is written as

$$\Phi(L)(1-L)^d(y_t - \mu) = \theta(L)\varepsilon_t, \tag{3.35}$$

where μ is mean of y_t. $E(\varepsilon_t) = 0, E(\varepsilon_t^2) = \sigma^2$, and $E(\varepsilon_t \varepsilon_s) = 0$ for $s \neq t$. L is a lag operator ($Ly_t = y_{t-1}$). $\Phi(L) = (1 - \Phi_1 L - ... - \Phi_p L^p)$ is the autoregressive polynomial and $\theta(L) = (1 + \theta_1 L + ... + \theta_q L^q)$ is the moving average polynomial in the lag operator L. All the roots of $\Phi(L)$ and $\theta(L)$ lie outside the unit circle. d denotes the fractional differencing parameter[100], which determines the degree of long range persistence in y_t. The y_t process defined by equation (3.35) is said to be integrated on order d (or $I(d)$) for $d \neq 0$ if by differencing d times, it may be expressed as a stable and invertible ARMA process (Abadir and Taylor, 1999). A value of $d = 0$ implies short memory stationarity, $d = 1$ corresponds to a unit root, and for $d < 1$, the impulse response weights are finite, which implies that shocks to level of the series are eventually mean reverting. An ARFIMA process is said to be covariance stationary when $-0.5 < d < 0.5$, and it is invertible (with mean μ) and has innovations that eventually disappear hyperbolically. There are 2 cases for the intermediate values of d. First, for $0 < d < 0.5$, the process is called stationary long memory. Second, for $-0.5 < d < 0$, it is called anti-persistent memory. For $0.5 < d < 1$, the y_t process does not have a finite variance, thus, it is nonstationary (with initialization parameter μ) and is

100 The term $(1-L)^d$ is the fractional difference operator defined by the following binomial expansion: $(1-L)^d = \sum_{j=0}^{\infty} \delta_j L^j = \sum_{j=0}^{\infty} \binom{d}{j} (-L)^j$.

recurrent but shocks are non-permanent and the unconditional variance grows at a slower rate than in the case of a unit root.

3.6.2 Methodology and model identifications

This section discusses the methodology in estimating the fractional differencing parameters, d. We first present the framework for the parametric and semiparametric estimators for d.

Then, we consider correcting for the multiple structural breaks, using Bai and Perron's (1998, 2003a) method, in the forward premium data in order to get better estimates of the long memory parameter, d. Their methodology in testing for and estimate multiple structural breaks will be presented as follows.

Estimating the fractional differencing parameters

We follow Maynard and Phillips (2001) in applying both the semiparametric and parametric frequency domain approach to estimate d. The main estimators in this study are the semiparametric estimators, whereas the parametric estimators are presented as a robustness check.

The semiparametric estimators are the modified log periodogram regression (MLP) estimators by Kim and Phillips (1999, 2006), and the log-periodogram regression (LP) method by Geweke and Potter-Hudak (1983). The parametric estimator is the Exact maximum likelihood estimator. These estimators are applied in Maynard and Phillips (2001), and Choi and Zivot (2005). A detailed discussion of each estimator is presented as follows.

Comparing the parametric and semiparametric estimators, Maynard and Phillips (2001) suggested that the advantages of the semi parametric estimator are its robustness to non-stationarity and its capacity to work well over the region $\frac{1}{2} < d < 2$. However, the limitation of the semiparametric method is the possibility of finite sample bias in the estimation that may arise from the strongly autoregressive short memory (Agiakloglou, Newbold and Wohar, 1993). The parametric ARFIMA estimates are less robust in the large sample but on the other hand they are less prone to finite sample bias. However, Maynard and Phillips (2001) found that the two sets of estimates match fairly closely.

The log-periodogram regression (LP) estimators of the fractional differencing parameter, d are widely used in economic applications since it involves a straight forward modification to the periodogram ordinates and it is easy to apply in empirical work. Diebold and Rudebusch (1989) studied the long memory process of quarterly post World WarII US real GNP data using the LP estimator. Cheung and Lai (1993) studied the purchasing parity condition and applied the LP estimator to test for cointegration of the series. Barkoulas, Baum, and Oguz (1997) studied the persistency of five long term interest rates using the LP estimator. This semiparametric estimator is based on a regression of the ordinates of the log spectral density on a trigonometric function. Following the notation from Baillie (1996), we can estimate d by examining

$$(1-L)^d y_t = u_t,$$

where $u_t \sim I(0)$. The spectral density of y_t is given by

$$f(\omega)_y = \left|1-e^{-i\omega}\right|^{-2d} f(\omega)_u = [4\sin^2(\omega_j/2)]^{-d} f(\omega)_u, \tag{3.36}$$

where $f(\omega)_y$ and $f(\omega)_u$ are spectral densities of y_t and u_t at frequency ω respectively. Rearrange equation (3.36) by taking logs, and adding and subtracting $\log[f_u(0)]$ to both sides of this equation to yield

$$\log[f_y(\omega_j)] = \log[f_u(0)] - d\log[4\sin^2(\omega_j/2)] + \log[f_u(\omega_j)/f_u(0)]. \tag{3.37}$$

Geweke and Potter-Hudak (1983), henceforth GPH, suggested estimating d from a regression based on (3.37) using the first m spectral ordinates $\omega_1, \omega_2, ..., \omega_m$, from the periodogram of y_t, that is $I_y(\omega_j)$. Hence, for $j = 1, 2, ..., m$,

$$\log[I_y(\omega_j)] = a + b\log[4\sin^2(\omega_j/2)] + v_j, \tag{3.38}$$

where $v_j = \log[f_u(\omega_j)/f_u(0)]$ and v_j is assumed to be i.i.d. with zero mean and variance $\pi^2/6$. From equation (3.36) and (3.37), the GPH estimator, d is the slope of the least squares regression, where the dependent and the explanatory variables are $\log[I_y(\omega_j)]$ and $\log[4\sin^2(\omega_j/2)]$, respectively in the sample $j = 1, 2, ..., m$. The estimated long memory parameter from LP is $\hat{d}_{LP} = -\hat{b}$. If the number of ordinates m is chosen such that $m = g(T)$, where $g(T)$ is such that $\lim_{T\to\infty} g(T) = \infty$, $\lim_{T\to g}[g(T)/T] = 0$, $\lim_{T\to g}[(\log(T)^2)/g(T)] = 0$, then the OLS estimator of d in equation (3.38) will have the limiting distribution

$$\left(\hat{d}_{LP} - d\right) / \left[var\left(\hat{d}_{LP}\right)\right]^{1/2} \Longrightarrow N(0,1).$$

where $var\left(\hat{d}_{LP}\right)$ is obtained from the OLS regression formula, either using the regression residual variance or alternatively setting it as $\pi^2/6$. GPH proved consistency and asymptotic normality of $\hat{d}_{LP}$ only for $d < 0$. Later on, Robinson (1990) proved the consistency for $0 < d < 0.5$. The disadvantage of the LP estimator is that $\hat{d}_{LP}$ posses serious bias and is very inefficient when u_t is $AR(1)$ or $MA(1)$ and the AR or MA parameter is quite large (Agiakloglou, Newbold, and Wohar, 1993).

The modified log periodogram (MLP) estimator[101] is the nonlinear version of the log-periodogram regression (LP) method, where the non-stationarity range ($d \geqslant 0.5$) is allowed in the estimation and is also robust for $AR(1)$ and $MA(1)$ errors. Thus the best estimator for the fractional differencing parameters in this work is the MLP estimator. In using this estimator, the dependent variable is modified to reflect the distribution of d under the null hypothesis that $d = 1$. In case of MLP estimator, one can write the parallel equation to (3.38) as follow. From Kim and Phillips (2006), the MLP estimation[102] involves testing the presence of long memory when $0.5 < d < 1$ which is described as follows.

$$\log I_{vx}(\omega_j) = c - 2d \log\left|1 - e^{i\omega_j}\right| + b(\omega_j), \tag{3.39}$$

where

$$b(\omega_j) = a(\omega_j) + \log \frac{I_{vx}(\omega_j; d)}{I_{vx}(\omega_j)} + O_p\left(\frac{1}{j^{1-d}}\right)$$

$$a(\omega_j) = \log\left[I_u(\omega_j)/f_u(0)\right], c = \log f_u(0).$$

The MLP regression estimator of d is obtained by regressing $\log I_{vx}(\omega_j)$ on $\log\left|1 - e^{i\omega_j}\right|$ over frequencies $\{\omega_j\}, j = 1, 2, ..., m$. The estimator is defined as

$$\hat{d}_{MLP} = -\frac{1}{2}\left[\sum_{j=1}^{m} x_j^2\right]^{-1}\left[\sum_{j=1}^{m} x_j \log I_{vx}\right],$$

101 The MLP shows a significant superiority over LP. The limit theory for the estimated d from MLP is the same as that of the LP estimator in the stationary case (Robinson, 1995; Hurvich, Deo and Brodsky, 1998). In contrast, the LP estimator has a mixed normal limit theory when $d = 1$ (Phillips, 1999), and is inconsistent when $d > 1$ (Kim and Phillips, 1999). Moreover, the simulations study by Kim and Phillips (1999) prove that the MLP is superior to LP in the nonstationary case where $0.5 < d < 1$.

102 The detailed derivation can be found in Kim and Phillips (2006).

where $x_j = \log|1-e^{i\omega_j}| - \overline{\log|1-e^{i\omega_j}|}$ and $\overline{\log|1-e^{i\omega_j}|} = \frac{1}{m}\sum_{j=1}^{m}\log|1-e^{i\omega_j}|$. As in the case of the LP estimator, the distribution of $\hat{d}_{MLP} \sim N(d, \pi^2/24m)$.

The practical problem is the choice of the number of periodogram ordinates to be used in the regression, m where it is written as

$$m = T^{\alpha}, \tag{3.40}$$

where T is the sample size. There is no optimal bandwidth rule over the full range $0 < d < 1$. Based on the simulation experiment, Kim and Phillips (2001) suggested that the optimal m is at $0.7 < \alpha < 0.8$ for the MLP estimation method, and applied $m = T^{0.7}$, $m = T^{0.75}$ and $m = T^{0.8}$ to analyse the extended Nelson-Plosser data in their work. Following this, Maynard and Phillips (2001) chose $\alpha = 0.75$ with 3,000 observations. Choi and Zivot (2005) used three different values of α which are 0.7, 0.75 and 0.8. This chapter follows Kim and Phillips (2001) and Choi and Zivot (2005) to apply a bandwidth[103] of $m = T^{0.7}, T^{0.75}, T^{0.8}$.

The computation of the ARFIMA model using the log periodogram regression (LP) is implemented using the ARFIMA1.04 package of Doornik and Ooms (2006) within the Ox programming suite (see Doornik, 2006). The code for the modified log periodogram

103 In the LP estimation, we apply the same bandwidth choice as in the case of MLP estimator. Geweke and Porter-Hudak (1983) suggested that the optimal bandwidth choice in the LP estimator of $T^{1/2}$ be used, and this choice has been widely adopted in the applied literature such as Diebold and Rudebusch (1983), and Chueng and Lai (1993). However, it is now known that this choice is not optimal in general. (see Henry and Robinson, 1996; Hurvich, Deo, and Brodsky, 1998; and Hurvich and Deo, 1999). Many researches attempted to find the optimal bandwidth rule such as Henry and Robinson (1996), Delgado and Robinson (1996) and Hurvich and Deo (1999). However, these works have limitations. The optimal m depends on d and their plug-in procedure is iterative. Moreover, there is distributional restriction. So far, there is still no rigorous justification of the optimality of the choice of ordinates for the LP estimator.
The main estimator for the fractional integration parameter in this chapter is the MLP estimator, where the results for LP estimator are for the robust check. Thus, the choice of harmonic ordinates for the LP estimator follows those of MLP estimator.

regression (MLP) is provided by Chang Sik Kim written in Gauss (see Kim and Phillips, 1999, 2006). Later on, Baum and Wiggins (2000) made the code of Kim and Phillips available in the environment of STATA. They also made an improvement on the original code by removing the deterministic trends from the series before application of the estimator[104]. The estimation results from the Gauss code of Chang Sik Kim and the STATA code of Baum and Wiggins (2000) yield very similar results. However, this study reports estimation results from Baum and Wiggins (2000).

The parametric ARFIMA estimates allows for autoregressive and moving average terms p and q. The parameters of the ARFIMA model are estimated by the exact maximum likelihood estimators (ML) with the assumption of conditional normality in ε_t. The criticism of the exact maximum likelihood estimation is that d can be severely biased in the presence of unknown nuisance parameters for regressor variables, even if there is only a constant to measure the unknown mean (see Doornik and Ooms, 2004). However, the presentation of the results from ML in this work is used as a robustness check. We utilise ARFIMA$(1, d, 0)$ and $(0, d, 1)$ models in estimating the value of d which is in line with Maynard and Phillips's (2001) work.

The computation of the ARFIMA model using exact maximum likelihood is implemented using the AFRIMA1.04 package of Doornik and Ooms (2006) within the programming environment of Ox (see Doornik, 2006). For the purpose of illustration, we write the ARFIMA$(1, d, 1)$ model as,

104 Phillips (1999) suggested that deterministic trends should be removed from the series before application of the estimator. Using Baum and Wiggins (2000)'s program, the linear trend is extracted from the series by default.

$$(1 - \Phi L)(1 - L)^d(f_t - s_t - \mu) = (1 + \theta L)\varepsilon_t,$$

where the orders of p and q can easily be derived. Here, the generic dependent variable y_t is replaced by the forward premium, $f_t - s_t$.

In the previous literature, the choice of p, q varies case by case. The majority of previous research considers a pure autoregressive process case where $q = 0$ such as Hol and Koopman, 2002; Baillie and Bollerslev, 1994a; and Baillie, Han and Kwon, 2002. Hol and Koopman (2002) consider ARFIMA$(1, d, 0)$ for the conditional second moment of the intra-day returns on the Standard and Poor 100 stock index. In their work, not all parameters of an ARFIMA$(1, d, 1)$ can be identified from the data. The original test for long memory of the forward premium by Baillie and Bollerslev (1994a) applied an ARFIMA$(2, d, 0)$ model on the 1 month forward premium in term of number of US dollars per unit of foreign currency. The fractional differencing parameters were estimated by approximate maximum likelihood. The resulting values of d were 0.445, 0.767 and 0.551 in Canada, Germany and UK. Baillie, Han and Kwon (2002) applied seasonal ARFIMA$(1, d, 0)$ on the conditional mean specification of the monthly CPI inflation series on eight developed countries, and found that the series follow stationary long memory process. They also found similar long memory properties in the second moment of inflation from applying ARFIMA-FIGARCH model.

Multiple mean break model

The purpose of this subsection is to briefly review the methodology used by Bai and Perron (1998,2003a) to estimate the unknown structural break dates. Their model considers a structural change in mean model that allows the errors to be serially correlated and heteroskedastic.

The structural change model we consider is defined by the multiple linear regressions with ϑ breaks ($\vartheta + 1$ regimes) as follows;

$$y_t = z_t^{'} \varrho_j + u_t, \tag{3.41}$$

where $t = T_{j-1} + 1, ..., T_j$ for $j = 1, ..., \vartheta + 1$ and $T_0 = 0$ and $T_{\vartheta+1} = T$. In this model, y_t is the observed dependent variable at time t, or the forward premium $(f_{t,k} - s_t)$ in this case. $z_t(q \times 1)$ is vector of covariates and ϱ_j is the corresponding vectors of coefficients. $z_t^{'} \varrho_j$ is treated as the mean of the forward premium for each regime. u_t is the disturbance at time t. The structural break points are represented by $T_1, ..., T_\vartheta$ and they are explicitly treated as unknown.

The model in equation (3.41) is called a pure structural change model since all the coefficients are subject to change. The variance of u_t need not be constant. Breaks in the variance are also permitted provided they occur at the same dates as the break in the parameters of the regression.

For each ϑ partition $(T_1, ..., T_\vartheta)$, the method of estimation for ϱ_j is based on the least-squares principles by minimizing the sum of squared residuals,

$$S_T(T_1, ..., T_\vartheta) = \Sigma_{i=1}^{\vartheta+1} \Sigma_{t=T_{i-1}+1}^{T_i} [y_t - z_t^{'} \varrho_j]^2. \tag{3.42}$$

Let $\hat{\varrho}(\{T_j\})$ denotes the resulting estimates. Substituting them in the objective function, the estimated break points $(T_1, ..., T_\vartheta)$ are written as

$$(\hat{T}_1, ..., \hat{T}_\vartheta) = \arg \min_{T_1, ..., T_\vartheta} S_T(T_1, ..., T_\vartheta), \tag{3.43}$$

where the minimization is taken over all partitions $(T_1, ..., T_\vartheta)$ such that $T_i - T_{i-1} \geq q$. The regression parameter estimates are the associate least squares estimates at the estimated ϑ–partition $\{T_j\}$, i.e. $\hat{\varrho} = \hat{\varrho}(\{T_j\})$. Next, we consider the test statistics for multiple breaks as follows.

The first test is a test of no break versus a fixed number of breaks, the $\sup F_T(l)$ which is defined as the F statistic of no structural breaks ($\vartheta = 0$) versus a fixed number of breaks ($\vartheta = k$). The F statistic is obtained by maximizing the difference between the restricted (without z_t) and unrestricted sums of squared residuals over all potential break dates. If a significant break is found, the full sample is divided into subsample at the break point, and the test is then performed on each of the subsample.

The other two test statistics test the null hypothesis of no structural break against an alternative of unknown number of breaks given some upper bound L (the maximum number of breaks allowed). These two tests are the double maximum statistics ($UD\max$) and the weighted double maximum statistics ($WD\max$). They are defined as follow

$$\begin{aligned} UD\max &= \max_{1 \leq l \leq L} \sup F_T(l), \\ WD\max &= \max_{1 \leq l \leq L} w_l \cdot \sup F_T(l), \end{aligned}$$

where w_l are the weights to the individual $\sup F_T(l)$ tests such that marginal p-values are equal across values of l.

The critical values of $\sup F_T(l)$, $UD\max$, and $WD\max$ tests statistics are presented in Bai and Perron (1998).

The last test is a test of l versus $l+1$ breaks, denoted $\sup F_T(l+1|l)$. This is a test of the null hypothesis of l breaks against the alternative that an additional break exists. The test is based on the difference between the sum of squared residuals obtained with l breaks and that obtained with $l+1$ breaks. One reject the null in favour of a model with $l+1$ breaks if the overall minimal value of the sum of squared residuals (over all segments where an additional breaks is included) is sufficiently smaller than the sum of squared residuals from the l breaks model. The critical values for this test can be found in Bai and Perron (1998, 2003a, 2003b).

There are two alternative ways in selecting the number of breaks, namely the "sequential procedures" and the information criterion. The "sequential procedure" is a criteria based on the sequential application of the $\sup F_T(l+1|l)$ test using the sequential estimates of the breaks. The other procedures to estimate the number of breaks are to minimize a Bayesian Information Criterion (BIC) suggested by Yao (1988) or a modified Schwarz's Criterion (LWZ) suggested by Liu, Wu and Zidec (1997).

In the Monte Carlo experiments of Bai and Perron (2000), the "sequential procedure" is found to be more reliable than the model selection criteria. They suggested that the BIC works well when breaks are present but less so under the null hypothesis of no break, especially if serial correlation is present. On the other hand, the LWZ criterion works better under the null hypothesis (even with serial correlation) by imposing a higher penalty. However, Bai and Perron (2000) suggested that this higher penalty translates into a very bad performance when breaks are present. A more detailed discussion can be found in Bai and Perron (1998, 2003a). This chapter adopt the sequential procedure in estimating structural breaks.

The practical recommendation by Bai and Perron (2003a) is to apply the sequential procedure in the presence of multiple breaks and check the $UD\max$ or $WD\max$ tests to see if at least one break is present. If the tests indicate the presence of at least one break (reject the null hypothesis of no structural breaks), then the number of breaks can be decided based on a sequential examination of the sup $F_T(l+1|l)$ statistics constructed using global minimizers for the break dates.

The computation of the multiple mean breaks model is implemented using the GAUSS code of Bai and Perron (1998, 2003a).

3.6.3 The empirical results

This section presents the resulting estimates of the fractional differencing parameters in the forward premium series by using the daily data for ten mostly traded currencies during Jan. 1994 - Jun. 2007. The approach is summarized into 3 steps as follows. First, the methods to estimate log memory parameter, d will be discussed. For our purpose, we have applied the modified log periodogram regression approach (MLP) of Kim and Phillips (1999, 2006) as the main estimator which is widely used for the case of nonstationary d.

Second, we consider the argument of Choi and Zivot, 2005; Diebold and Inoue, 2001; Granger, 1999; Granger and Hyung, 2004; and Lamourex and Lastrapes, 1990, that the presence of structural breaks may also artificially produce high persistence in volatility.

We test for and estimate the multiple mean break model in the forward premium data using Bai and Perron's (1998, 2003a) method. Third, we adjust for the structural breaks in the forward premium and reestimate the long memory parameter using the MLP regression on the mean-break adjusted data.

Step 1: Estimating d without allowing for structural breaks

Estimates of the fractional differencing parameter (d) and its standard error are shown in Tables 3.3 and 3.4. Tables 3.3 presents the estimated parameters using the semiparametric estimation techniques whereas table 3.4 contains estimated parameters from the parametric method by exact maximum likelihood[105] ($\hat{d}_{ML}$) of the ARFIMA$(1, d, 0)$ and $(0, d, 1)$ models[106].

Panel on the left hand side of table 3.3 contains the estimated long memory parameters $\hat{d}_{MLP}$, and its corresponding standard error and test statistics estimated by the modified log periodogram (MLP) method of Kim and Phillips (1999,2006). Panel of the right hand side of table 3.3 contains estimated long memory parameters, $\hat{d}_{LP}$ and its corresponding test statistics estimated by the log periodogram method of Geweke and Potter-Hudak (1983). The first column of table 3.3 presents the power or the alpha value in equation

105 Using the ARFIMA package on Ox (see Doornik and Ooms, 2006) the ML estimator was applied to the first differenced series and the resulting estimate of d was then increased by one.

106 There has been attempted in testing of higher order AR and MA polynomials in the ARFIMA estimation, i.e., ARFIMA(1,d,1) and ARFIMA(0,d,2) models. However, the results are not reported in the thesis. There is problem of flat log-likelihood in some sample currencies using ARFIMA(1,d,1) model. In case of ARFIMA(0,d,2), some sample currencies show that MA(2) is not significant. Lastly, the resulting log-likelihood from the estimation show that ARFIMA(0,d,1) and ARFIMA(1,d,0) are superior to ARFIMA(1,d,1) and ARFIMA(0,d,2) models. There is also a parsimony argument in favour of lower-order processes. Using lower order is tested to be preferred.

Lastly, these specifications correspond to Maynard and Phillips (2001) who selected the ARFIMA(1,d,0) and ARFIMA(0,d,1) models in estimating the fractional integration parameters of the forward premium data.

(3.40), which determines the number of harmonic ordinates to be included in the spectral regression, m of the MLP and LP estimations. The results can be discussed as follows.

The Modified Log Periodogram (MLP) estimation results in table 3.3 show that the estimated value for the fractional integration parameter, $\hat{d}_{MLP}$ are positive in all cases. The estimated standard errors are found to be small for all currencies[107]. We reject the null of $\hat{d}_{MLP} = 0$ in all currencies and for all values of α. On the other hand, the results show mixed evidence when testing for the null of $\hat{d}_{MLP} = 1$. For example, one fails to reject the null of unit root process in the case of CAD, CHF, DKK, GBP and NOK for every value of α, while the forward premium series in AUD show no evidence of a unit root for every value of ordinates, m. The stationarity test for the rest of the sample currencies such as EUR, JPY, NZD and SEK show mixed evidence with respect to different values of α. The resulting estimated values of $\hat{d}_{MLP}$ can be summarised as follows.

In case of AUD, one can conclude that the forward premium series have mean reverting properties with infinite variance, but finite cumulative impulse response weights. This is represented by the value of $\hat{d}_{MLP}$ being statistically less than unity using 5 percent confidence interval and the forward premium series fall into the nonstationary range where $0.5 < \hat{d}_{MLP} < 1$. In case of CAD, CHF, DKK, GBP and NOK, the forward premium series are found to follow the unit root process where the estimated $\hat{d}_{MLP}$ are statistically not different from unity. Using 5 percent confidence interval, the test statistics fail to reject the null of unit root in the forward premium series.

107 Note that the standard errors generally increase when α gets smaller. Moreover, the point estimate falls as α increases.

Panel on the right hand side of table 3.3 presents the estimated $\hat{d}_{LP}$ from LP estimator. All the $\hat{d}_{LP}$ are positive and are highly significantly different from zero. This corresponds to the results from MLP estimator. The estimated values of d from LP estimator are very close to those from MLP estimator and all the series fall into nonstationary range of $\hat{d} > 0.5$.

Table 3.4 presents the parametric estimation results using exact maximum likelihood estimation. All of the forward premium series are found to be nonstationary. Considering the mean reversion of the series, the ARFIMA$(1, d, 0)$ and the ARFIMA$(0, d, 1)$ give contradictory results. All estimated d are less than 1 when using the ARFIMA$(1, d, 0)$ model (column 1 of table 3.4). In the ARFIMA$(0, d, 1)$ model (column 2 of table 3.4), the impulse response weights of the forward premium for AUD, CAD, EUR, GBP, NOK and SEK are infinite ($d > 1$), which implies that shocks to level of the series are not mean reverting. However, the ARFIMA$(1, d, 0)$ model is more preferred according to Akaike information criteria. Moreover, the estimated d from the ARFIMA$(1, d, 0)$ model is more reliable since its estimated standard errors are lower than in the ARFIMA$(0, d, 1)$ model.

Overall, the semiparametric and parametric estimators show mixed evidences of stationarity and mean revertability. This is may be because of the structural changes in the forward premium series. In the next subsection, we identify and estimate for potential structural breaks. Finally, we estimate the fractional differencing parameters using the mean break adjusted data.

Step 2: Estimating the mean break model

Figures 3.1A-3.1J provide visual evidence of structural changes in the mean of the series. Following Bai and Perron (2003a) and Choi and Zivot (2005), we apply our procedure with only a constant as a regressor (i.e. $z_t = \{1\}$) and impose 15 percent trimming on each end of the sample and between the break dates and allow a maximum of 5 breaks[108].

Table 3.5 reports the significant break dates (at 5 percent level) from the Bai and Perron's tests for multiple structural breaks. The mean forward premium before adjusting for the breaks is presented in the second column. The third column is $\mathrm{TB}_{i=1,..,5}$ the estimated mean after each subsequent breaks, which corresponds to $z_t' \varrho_j$ in equation (3.41), followed by asymptotic standard error. The last column is the estimated break dates with 95 percent confidence interval.

Table 3.6 represents the multiple structural change tests results. Note that the final number of breaks is determined by the sequential procedure. As mentioned in the previous subsection the information criteria are biased downward and the estimated number of breaks chosen by BIC and LWZ is only for illustration purpose.

Figures 3.9A-3.9J plot the forward premium, along with their changing means, for all ten countries.

108 Hence each segment has at least 510 days (15% of around 3,400 observations in our data). The choice of value of trimming (ε) is specified by the program, it varies by the maximum number of break allows as follows. For each option, the maximal value of breaks is: 10 for $\varepsilon = 5\%$; 8 for $\varepsilon = 10\%$, 5 for $\varepsilon = 15\%$, 3 for $\varepsilon = 20\%$ and 2 for $\varepsilon = 25\%$. See details in Bai and Perron (2003).

Considering the time span of our data (from Jan 1994 to Jun 2007), allowing 5 breaks is sufficient. Bai and Perron (2003a) also suggest that the maximum break of 5 is sufficient for most empirical applications and applied this to the study of the quarterly US real interest rates from 1961:1 to 1986:3.

Anyway, this chapter found that the results are not sensitive to an alternative trimming values of $\varepsilon = 5\%$ (when allowing for maximum of 10 breaks).

In estimating the number of break points during 1994-2007 for 10 sample currencies, one can summarise the results from tables 3.5 and 3.6 as follows.

First, the data suggest the presence of at least one structural change in all of the forward premium series. From table 3.6, the $UD\max$, $WD\max$ and $\sup F_T(1)$ through $\sup F_T(5)$ tests of all series are significant at 1 percent level.

Second, the lag orders (number of structural changes) selected by sequential procedure and the $\sup F_T(l+1|l)$ statistics varies when using different series. The detailed description of the results is as follow.

The forward premium series in NOK exhibit one structural break. The sequential $\sup F_T(l+1|l)$ test statistics (in table 3.6) suggest one break for NOK on 20 July 2004 presumably due to the Norges bank stopped the loosening monetary policy[109]. The results can be seen in table 3.5 and figure 3.9H. The break date has a rather small confidence interval (between 16 Jul. 2004 and 30 Aug. 2004). Before the break date the mean of the forward premium is estimated to be 0.206 percent while the estimated intercept after break is -0.119 percent. Interestingly, the mean of the forward premium over the period of observation before adjusting for the structural break is 0.093 percent. Thus, taking structural break into account, we drastically obtain more accuracy in explaining behaviour of the forward premium before and after the Norges bank's change in the monetary policy. We expect

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109 On 22 January 2003, Norges Bank cut both overnight rate and the deposit rate by 50 basis points each, bringing rates down to 6% and 8% respectively. This is to defend its currency against the strong Euro. The Norges bank cut interest rates 8 more times following the January rate cut, finally ending the loosening of its monetary policy in March 2004 with its key interest rate at 1.75%.

The Norwegian rates stayed until around Mid-2005 when the oil price soared. Norway was the 10th largest oil producing, and the third largest oil exporting country in the world in 2006. Thus, oil price soar made the Norwegian Krone grew even stronger.

that the value of the estimated long memory parameters should be reduced after taking into account the structural breaks as in Choi and Zivot (2005).

Considering the tests results in table 3.6 for the case of NOK in details, sup $F_T(1)$ through sup $F_T(5)$ are all significant at 1 percent level. The $UD\max$ and $WD\max$ tests are also significant at 1 percent level which suggest the presence of structural change in these three countries, but the sup $F_T(l+1|l)$ statistics are not significant when $l \geq 1$ which indicates only one break. This confirms the number of break dates selected by the sequential procedure. We therefore conclude that there is only one structural break in case of NOK.

The forward premium series of CHF, EUR, and GBP are estimated to have 2 breaks. In the case of EUR at least one structural change occurred at around 4 Jan. 2001. We also find evidence of another structural break around 20 Jan. 2005. These two breaks significantly affect both the level and the slope coefficient of the forward premium[110]. The two break dates are precisely estimated since the estimated 95 percent confidence intervals cover only a few days before and after. The differences in the estimated means over each segment are significant and point to a decrease of 0.274 percent in Jan. 2001 and an increase of 0.225 percent in Jan. 2005. In case of CHF, the first and second breaks occurred on 8 Jan. 2001 and 4 Apr. 2005, respectively. For GBP, the first and second breaks occurred on 17 Sept. 2001 and 14 Jun. 2005, respectively.

110 In case of EURO, the empirical evidence is that after the introduction of the Euro in 1999, its exchange rate against other currencies fell heavily, especially against the U.S. dollar: at October, 26, 2000, it had fallen to an all-time low. This results in the first break.

After the first break, the Euro then began steadily appreciating, and reached parity with the U.S. dollar on July 15, 2002. The second break occur when the Euro reached a peak at the end of 2004 since the U.S. dollar fell against all major currencies, fuelled by the double deficit in the US accounts. The dollar recovered in 2005 and was stable throughout the second half of 2005 and so on.

The forward premium series of AUD, DKK, JPY, and SEK are found to have more frequent structural changes as the sequential procedure estimates the number of break at 3. Lastly, the forward premium series of CAD and NZD are estimated to have 4 breaks.

Step 3: Estimating d using the mean break adjusted data

Tables 3.7 and 3.8 present the estimated long memory parameters for the break adjusted forward premium data using both the semiparametric and parametric estimators. The break adjusted data is the residual series, $u_t = y_t - z_t'\varrho_j$ in equation (3.41).

The panel on the left hand side of table 3.7 shows Kim and Phillips (1999, 2006) Modified Log Periodogram (MLP) regression estimate of the fractional integration parameter ($\hat{d}_{MLP}$) and its standard error, using the same range of bandwidth as in the first step. The panel on the right hand side of the same table presents the estimation results from the traditional Log Periodogram (LP) regression estimate. Table 3.8 presents estimation results from the parametric estimator of the ARFIMA$(1, d, 0)$ and ARFIMA$(0, d, 1)$ models, respectively.

First and overall, the results show that allowing for structural breaks reduces the persistence of the daily forward premium data in ten mostly traded currencies. The results are robust across all currencies and model specifications. Considering the MLP estimator, there is a meaningful reduction in the magnitude of $\hat{d}_{MLP}$in all currencies and the order of ordinates $m = T^{\alpha}$. For example, the point estimate of the forward premium in AUD falls from $\hat{d}_{MLP} \in (0.811, 0.897)$ to $\hat{d}_{MLP} \in (0.739, 0.889)$. The point estimate for CAD also decreases from $\hat{d}_{MLP} \in (0.979, 1.027)$ to $\hat{d}_{MLP} \in (0.889, 0.924)$ and the point estimate for SEK declines from $\overset{\curlywedge}{d}_{MLP} \in (0.851, 0.976)$ to $\overset{\curlywedge}{d}_{MLP} \in (0.790, 0.911)$. This shows that the estimation of the long memory parameter is sensitive to the multiple structural breaks.

In addition, after removing the break, the previous failure to reject the unit root in the forward premium series no longer obtains in all currencies; using 5 percent confidence interval, the null of $\overset{\curlywedge}{d}_{MLP} = 1$ is rejected in every value of α. Considering these two points, we thus support the argument of Choi and Zivot, 2005; Diebold and Inoue, 2001; Granger, 1999; Granger and Hyung, 2004; and Lamourex and Lastrapes, 1990, that the presence of structural breaks can artificially produce high persistence in volatility. This is called "spurious long memory process".

In comparison to Choi and Zivot (2005), the reduction in magnitude of d after allowing for structural breaks is less drastic in my work. In Choi and Zivot (2005), after the adjustment for multiple structural breaks, four out of five forward premium series appear to change from nonstationary to stationary long memory process. However, after removing the breaks, the evidence of nonstationary long memory in the forward premium still persists in all cases in my work. The forward premium displays less persistence than a unit root process ($d < 1$), but still too much persistence to satisfy stationarity. This is possibly due to different sample currencies and sample period used. They use monthly exchange rate data in terms of US dollars for five G7 countries: Germany, France, Italy, Canada and UK over the period January 1976 to January 1999, whereas we employ the daily data of ten mostly traded currencies[111] from around Jan. 1994 - Jun. 2007. Thus we can conclude the finding that the forward premium series of all sample countries is covariance nonstationary

111 The list of sample currencies can be found in table 3.1.

because its variance is not finite (Hosking, 1981). However, the process is mean reverting, since an innovation has no permanent effect on the values of the forward premium series.

Second, the other benefit of the structural break adjustment is that there is less inconsistency in the estimation results across different model specifications, i.e. MLP, LP, ARFIMA$(0, d, 1)$ and ARFIMA$(1, d, 0)$. The four sets of estimates match fairly closely. This is consistent with Maynard and Phillips (2001) who find that estimated values for d are robust to change in model specifications.

Lastly, it is useful to compare the estimation results to the previous literature. Maynard and Phillips (2001) applied a bandwidth of $m = T^{0.75}$ in estimating the parameter d using the MLP estimator. They examined daily data from November 1986 to March 1988 and the corresponding estimated d values in their work are 0.957, 0.937, 0.882 and 0.993 for AUD, CAD, JPY and GBP, respectively. Considering the same bandwidth as Maynard and Phillips (2001), the estimated values of the long memory parameter are fairly close to their work. The corresponding values of $\hat{d}_{MLP}$ using the break adjusted data in this work are 0.845, 0.914, 0.882, and 0.906 in AUD, CAD, JPY and GBP, respectively. However, it is useful to note that the different value of $\hat{d}_{MLP}$ is likely to be because of difference in periods of observations and the time span of the data; this chapter employs a longer datasets than their work.

The results from this chapter obtain somewhat larger estimates of d than those first reported by Baillie and Bollerslev (1994a) and in the later work by Choi and Zivot (2005). Baillie and Bollerslev (1994a) used monthly data and found point estimates for the order of fractional integration using an $ARFIMA(2, d, 0)$ model of 0.445 and 0.551 for

CAD and GBP, respectively. The corresponding values of d using the $ARFIMA(1,d,0)$ and the $ARFIMA(0,d,1)$ on the break adjusted data in this work are $(0.935, 0.954)$ and $(0.849, 0.968)$ for CAD and GBP, respectively. However, the reason that we obtain higher value of $\hat{d}_{MLP}$ is clearly because of different model estimation method, period of estimations and frequency of the data. As mentioned earlier, this chapter employ the MLP estimator to study the daily forward rate data from Jan. 1994 to Jun. 2007.

In comparison with the results from Choic and Zivot (2005), the resulting estimated long memory parameters in this chapter are more than in their work. They estimated the long memory parameter using the bandwidth of $m = \{T^{0.70}, T^{0.75}, T^{0.80}\}$ with the MLP estimator. Considering the break adjusted data, the estimated $\hat{d}_{MLP}$ in their work are $(0.450, 0.517, 0.582)$ in Canada and $(0.303, 0.369, 0.455)$ in the UK, respectively. Using break adjusted data, the estimated parameter from MLP, $\hat{d}_{MLP}$ using the same bandwidth as their work, are found to be $(0.924, 0.914, 0.889)$ and $(0.940, 0.906, 0.863)$, in Canada and UK, respectively. However, the different finding is due to the different time period of studies and frequency of the datasets since they used monthly data from Jan. 1976 to Jan. 1999.

Conclusion

Using the daily data for ten mostly traded currencies during Jan. 1994 - Jun. 2007, this chapter confirms the finding of previous literatures that the forward premium data follow (nonstationary) fractionally integrated process, which attributes to the forward premium anomalies. The presence of structural breaks also proves to play an important role in explaining the long memory of the forward premium data. Allowing for structural breaks reduces the persistence of the forward premium across all currencies and model specifications. Moreover, the criticism that the presence of the structural breaks tend to cause the spurious long memory process in the data as in Granger, 1999; Granger and Hyung, 2004; Diebold and Inoue, 2001 is supported when applying the forward premium data in this study.

This study also confirms the argument of Baillie and Bollerslev (2000) that the statistical artifacts of the data are important. There is evidence of fractional integration in the forward premium series where the fractional differencing parameters of all currencies are significantly different from zero though on the whole less than unity. The very persistent autocorrelation in the forward premium contributes to the forward premium anomaly. The estimated values for d are quite similar to those obtained by Maynard and Phillips (2001). Considering the break adjusted data, the majority of the forward premium series are characterised by nonstationary long memory process although they are also mean reverting ($0.5 < d < 1$).

Lastly, upon finding evidence of fractional integration in this study, Maynard and Phillips (2001) suggested that traditional statistical theory may not be applicable to many of the regressions commonly used to test forward rate unbiasedness. Thus, it is necessary to develop the new limit theories and new methods to test the unbiasedness hypothesis.

The next section further confirms the finding of an unbalanced order of integration in the unbiasedness regression by testing for stationarity of the return on the spot rate.

3.7 Evidence of stationarity in the spot return

This subsection aims to study the time series properties of the return on the spot rates ($\Delta s_{t,t+k}$) series in equation (3.33). We first report the unit root tests of this series for each currency in tables 3.9A and 3.9B.

We perform the Elliott-Rothenberg-Stock (ADF-GLS, 1996) efficient test for an autoregressive unit root. The results are presented in table 3.9A. The null hypothesis is that the returns on the spot rate contain a unit root, and the alternative is that the series were generated by a stationary process. Table 3.9B presents Kwiatkowski, Phillips, Schmidt, Shin (KPSS, 1992) test statistics for the null of level stationary versus an alternative of a unit root of the spot return series at different values of truncation lag. The maximum lag order for these two tests is by default calculated from the sample size using a rule provided by Schwert (1989). For the ADF-GLS t-test, the optimal lag order is calculated by the Ng-Perron (1995) sequential t test on the highest order lag coefficient, stopping when that coefficient's p-value is less than 0.10.

The first column of table 3.9A presents the t-test statistics from the ADF-GLS test, followed by the optimal lag length. The ADF-GLS test results show that one can overwhelmingly reject the null hypothesis of a unit root at 5 percent significance levels in all currencies. We next consider the KPSS test which can be used in conjunction with the ADF test to investigate the possibility that the returns on the spot rate series are fractionally integrated.

Further analysis using the KPSS test reveals that the spot returns of all the currencies are indeed stationary. The KPSS test results in table 3.9B show that the spot returns in all currencies are level stationary. Using a 5 percent level of significance, one cannot reject the null hypothesis that the spot returns are level stationary in all countries for the truncation lags of up to their maximum.

The finding that the spot returns are stationary in all currencies is not surprising. There is overwhelming evidence that the logarithm of the spot exchange rate is nonstationary and follows $I(1)$ process, while the return on the spot rate is stationary (Cornell, 1977; Meese and Singleton, 1982; Corbae and Ouliaris 1986; Baillie and Bollerslev, 1989; Baillie and Bollerslev, 1993; Baillie and Bollerslev, 1994a).

These findings establish that there is unbalanced order of integration between the spot returns and the forward premium. Thus it is insufficient to test the hypothesis of the forward rates being an unbiased predictor of the future spot rates by using traditional regression analysis.

3.8 Estimating the exchange rate risk premia

This section aims to estimate the foreign exchange risk premia following Domowitz and Hakkio (1985) and Baillie and Bollerslev (1990). The risk premium is proxied by the conditional variance of the forward rate forecast error as represented by equation (3.30),

$$\eta_{t,t+k} = var_t(f_{t,k} - s_{t+k}),$$

where $(f_{t,k} - s_{t+k})$ is the forward rate forecast error, defined as the difference between the realised future spot rate and the current forward rate, $s_{t+k} - f_{t,k}$. s_{t+k} is a natural log of the spot rate at time $t + k$ and $f_{t,k}$ is the natural log forward rate at time t with maturity k.

In the previous section we found that the forward premium followed a fractionally integrated (ARFIMA) process, but its conditional variance was assumed to be constant over time. This section aims to model the long memory in the conditional variance for the forward rate forecast error series. The Fractionally Integrated Generalised Autoregressive Conditional Heteroskedasticity (FIGARCH) model explains the fractional integration (the long memory) behaviour in the conditional variance of the series.

The analysis begins with a description of the conditional variance properties of the forward rate forecast error and the rationale for applying this model, followed by the theoretical modelling of the FIGARCH model. Lastly, it presents the estimation results.

3.8.1 Fractional Integration behaviour in the conditional variance of the forward rate forecast error

Referring back section 3.4, this chapter apply the MA(20) process to filter out the serial correlation from the 1 month forward rate forecast error series. We primarily observe the autocorrelation plots of the residuals from the MA(20) process (figures 3.7A-3.7J). These plots show that there are no significant autocorrelations in all lags, thus the filtered forward rate forecast error series is the short memory process in the level.

We then plot the autocorrelation functions of the squared and the absolute value of the residuals from the MA(20) process (figures 3.7A-3.7H) and found that they display persistence pattern. This suggests long memory behaviour in the second moment of the series. Moreover, the autocorrelation coefficients display a slow hyperbolic rate of decay. The autocorrelation coefficients at various lags are very high even up to a lag of 100 working days. Thus, the fractional integration behaviour is likely to exist in the second moment of the filtered series.

Lastly, we proxy the foreign exchange risk premium by applying the parametric model of long memory in the conditional variance of the filtered forward rate forecast error series.

3.8.2 The Fractionally Integrated GARCH model

This section aims to estimate the degree of persistence in the variation in the forward rate forecast error using the FIGARCH process. Originally, the Fractionally Integrated GARCH (FIGARCH) model is introduced by Baillie, Bollerslev and Mikkelsen (1996), henceforth BBM. So far, none of the literature applies it to the foreign exchange risk premium context.

The model that is postulated to describe the filtered forward rate forecast error is

$$(s_{t+k} - f_{t,k}) = \sum_{j=1}^{[k]} \theta_j \varepsilon_{t+k-j} + \varepsilon_{t+k}, \tag{3.44}$$

where $k = 20$. Equation (3.44) represents the conditional mean specifications of the model where $(s_{t+k} - f_{t,k})$ is the forward rate forecast error series. We assume that ε_t is an innovation in the forward rate forecast error series, that is, it has mean zero conditional on time $t-1$ information.

The FIGARCH(P, δ, Q) model of BBM is given by

$$\varepsilon_t = v_t h_t, \tag{3.45}$$

$$[1 - \beta(L)]h_t^2 = \alpha_0 + [1 - \beta(L) - \phi(L)(1 - L)^\delta]\varepsilon_t^2, \tag{3.46}$$

where $\beta(L) = \beta_1 L + \beta_2 L^2 + ... + \beta_P L^P)$, $\phi(L) = [1 - \alpha(L) - \beta(L)](1 - L)^{-1}$ and $\alpha(L) = (\alpha_1 L + ... + \alpha_Q L^Q)$ and all have their roots outside the unit circle. $(1 - L)^{\delta}$ accounts for the long memory of the process and α_0 is a constant term.

Equation (3.45) gives a decomposition of the innovations, ε_t. v_t is an independently and identically distributed (*i.i.d.*) random variable with $E_{t-1}(v_t) = 0$ and $Var_{t-1}(v_t) = 1$. h_t is a positive time varying and measurable function with respect to the information set available at time $t - 1$. $E_{t-1}(.)$ and $VAR_{t-1}(.)$ refer to the conditional expectation and variance with respect to this same information set. The $\{\varepsilon_t\}$ process is serially uncorrelated with mean zero, but the conditional variance of the process, h_t^2 is changing over time (Baillie, Bollerslev, Mikkelson (1996).

To capture persistence in the series volatility, equation (3.46) is the conditional variance specification of the innovations. h_t^2 is defined as the time t conditional variance[112] of ε_t. We model the changing volatility of a time series ε_t by using the Fractionally Integrated Generalised Autoregressive Conditional Heteroskadasticity (FIGARCH) methodology. Conditional on time $t - 1$ information, the innovation is assumed to be normally distributed: $\varepsilon_t \sim N(0, h_t^2)$. In particular, the FIGARCH(P, δ, Q) model implies a slow hyperbolic rate of decay for the lagged squared innovations in the conditional variance function. The FIGARCH class of the process is covariance stationary for $0 < \delta < 1$, shocks to conditional variance will ultimately die out. FIGARCH model implies a finite persistence

112 With the time varying h_t^2 the unconditional distribution of ε_t has fatter tails than a normal distribution $[K(\varepsilon_t) \geqslant 3]$. This is because the variability of the conditional variance affects higher moments of the unconditional distribution of ε_t. [see Campbell, Lo and MacKinlay (1997, page 480)]

of volatility shocks, i.e., there is a long memory behaviour and slow rate of decay after a volatility shock.

If we introduce the conditional variance (or the conditional standard error) as an explanatory variables in equation (3.44), we get ARCH-in-Mean model (ARCH-M) of Engle, Lilien and Robins (1987), i.e.

$$\mu_t = \mu + \nu h_t^b, \tag{3.47}$$

with $b = 1$ to include the conditional standard deviation and $b = 2$ for the conditional variance. The ARCH-M is usually applied to financial time series to represent the relation ship between the expected return of the series and its expected risk. Domowitz and Hakkio (1985) model the forward rate forecast error using the ARCH-in-mean relationship. The estimated coefficient of the expected risk ν is interpreted as a measure of the risk-return trade-off. Integrating equation (3.47) to the FIGARCH framework yields the FIGARCH-in-mean model. In the estimation part, we examine the possibility of this specification.

Chung (1999) introduced a slight modification in the conditional variance function of BBM. He argues that BBM's parameterisation of the FIGARCH model may have a specification problem that causes difficulty in estimation and interpretation of the resulting estimates. Chung (1999) proposed a slightly different process in the conditional variance function of equation (3.46) as follows;

$$[1-\beta(L)]h_t^2 = [1-\beta(L)]h^2 + \{1-\beta(L)-\phi(L)(1-L)^\delta\}\left(\varepsilon_t^2 - h^2\right),$$

where h^2 is the unconditional variance of ε_t and $v_t \equiv (\varepsilon_t^2 - h_t^2)$.

The computation of the FIGARCH model is implemented using the G@RCH4.2 package of Laurent and Peters (2004) within the programming environment of Ox. The estimation of the FIGARCH model assumes conditional normality of the process using the Maximum likelihood estimation for parameters of the process. The results are presented in the following subsection.

3.8.3 The empirical results

Tables 3.10 and 3.12 contain estimation results using the FIGARCH$(1, \delta, 1)$ models for currencies' forward rate forecast error for both BBM's and Chung's specifications, respectively. We also found that there is no FIGARCH-in-mean relationship in the forward rate forecast error series. The estimated coefficient of the expected risk, ν in equation (3.47) is found to be statistically insignificant for both $b = 1$ and $b = 2$ cases.

The estimated parameters from FIGARCH$(1, \delta, 1)$ using BBM specification are presented in table 3.10. There is strong evidence of long memory in the conditional variance. The long memory parameter in the conditional variance specification (δ) is significantly different from zero, and lies in the range between 0.279 and 0.528. This means that the conditional variance matters in determining deviation of the forward rate from the expected future spot rate. Since $0 < \delta < 1$, thus the propagation of shocks to the mean and variance of forward rate forecast error is proved to occur at a slow hyperbolic rate of decay. The value of the skewness and the kurtosis of the standardised residuals of the estimated FIGARCH model and their p-values are reported. Moreover, the Jarque-Bera normality test (Jarque and Bera, 1987) is also reported.

In modelling the foreign exchange risk premium, this paper proposed that the FIGARCH model fits the data better than the multivariate GARCH model as used by Baillie and Bollerslev (1990) and the ARCH-in-mean model as used by Domowitz and Hakkio (1985)). Reliable estimates and inference of the risk premium depend on well-specified conditional heteroscedasticity models. Thus, checking the adequacy of a fitted model becomes an important issue for model selection. Tse (2002) argued that misspecification in the mean and variance results in inconsistency and loss of efficiency in the estimated parameters. It is useful to discuss the misspecification test results for the model in table 3.11 as follows.

The misspecification tests of the FIGARCH model by BBM are presented in table 3.11. The first panel reports the Box-Pierce statistics at lags 10, 15 and 20 for the standardised residuals and the squared standardised residuals under the null hypothesis of no autocorrelations. The second panel is the residual based diagnostic (RBD) for conditional heteroskedasticity of Tse (2002). The residual-based diagnostics are constructed to test for certain residual patterns implied by the deviation of the fitted model from its underlying assumptions. This test diagnoses model misspecifications concerning on conditional heteroscedasticity in time series models. This is the test of the null of model adequacy against alternatives of model misspecification. The third panel is the sign bias t test (SBT). This is a diagnostic test of Engle and Ng (1993) that investigates possible misspecification of the conditional variance equation and test for the presence of leverage effect. This is the test of the null of no sign bias against the alternative of the presence of leverage effect. The next panel is the Adjusted Pearson Chi-square Goodness-of-fit test under the null of a correct distribution of the innovations. Considering the test results in table 3.11, one can argue that the model captures the dynamic of the forward rate forecast error data in all currencies. The Q-statistics on standardised residuals and the squared standardised residuals, and RBD test with various lag values as well as the adjusted Pearson Chi-square goodness of fit test with different cells number accept the null hypothesis of correct specifications in all currencies. Additionally, there is no sign bias in the conditional variance function in all currencies. The last panel is Engle's LM ARCH test (Engle, 1982), it tests the presence of ARCH effects in a residuals. For each specified order, the squared series is regressed on p of its own lags. The test statistics is distributed $\chi^2(p)$ under the null hypothesis of no ARCH effects.

The overall conclusion is that the FIGARCH model appears to be a good specification for the filtered forward rate forecast error series. The post estimation test results for the FIGARCH model using BBM's specification show that there is no misspecification. The model well captures the dynamic of the forward rate forecast error in all currencies.

It is useful to examine the results from using an alternative specification of the FIGARCH model as suggested by Chung (1999). Table 3.12 contains the estimated FIGARCH models for currencies' forward rate forecast error using Chung's specification. There is no FIGARCH-in-mean relationship. This is the same as the estimation results using BBM's specification. This proves the robustness of the model. The estimated long memory parameter in the conditional variance specification (δ) is significantly different from zero, and lies in the range of 0.283 to 0.451. This confirm the evidence that the volatility of the forward rate forecast error tends to change quite slowly over time, i.e. the effect of shocks takes a considerable time to decay. The estimated parameter values from Chung's specification are quite close to those estimated by BBM's specification. The misspecification test results are presented in table 3.13. Again, Chung's specification appears to fit the data quite well.

Lastly, Chung (1999) derived the sufficient condition for non-negative conditional variances for the case of FIGARCH$(1, \delta, 1)$ specification[113] of BBM as

$$\begin{aligned} \beta_1 - \delta &\leqslant \phi_1 \leqslant \frac{2-\delta}{3} \\ \text{and } \delta(\phi_1 - \frac{1-\delta}{2}) &\leqslant \beta_1(\phi_1 - \beta_1 + \delta). \end{aligned} \tag{3.48}$$

The estimation results show that the sufficient condition to ensure that the conditional variance is positive almost surely for all t is fulfilled in all sample currencies using BBM specification. In contrast, the positivity constraint for the FIGARCH$(1, \delta, 1)$ is observed in 1 out of 10 currencies using Chung's specification (in case of CHF). Generally, the standard errors of the estimated parameters in the conditional variance function from Chung's specification are relatively higher than in BBM specification.

From the estimation results and misspecification test results, the BBM's specification proves to fit the data better. Hence, the exchange rate risk premium will be generated from the conditional variance of the forward rate forecast error using BBM's specification.

Additionally, this paper argue that the fractionally integrated GARCH model fits the forward rate forecast error data adequately. Comparing to Domowitz and Hakkio (1985), we find that the estimated coefficient of the expected risk ν in equation (3.47) is statistically insignificant, hence there is no ARCH (or FIGARCH)-in-Mean relationship in our data.

113 See prove in appendix A of Chung, 1999.

Comparing to Baillie and Bollerslev (1990) which applied the multivariate GARCH process to the forward rate forecast error series, the FIGARCH model in this work allows us to develop more flexible class of the process for conditional variance that are more capable of explaining and representing the observed temporal dependencies in the series volatility. The GARCH framework assumes no persistence in the volatility of shocks. However, the autocorrelation plots of the absolute and squared residuals (in figures 3.7A-3.7J) prove that there is persistence of shocks. Thus, the foreign exchange risk premium is proved have a long memory behaviour and slow rate of decay after a volatility shock.

Domowitz and Hakkio (1985) add the estimated risk premium in regression for the Uncovered Interest Parity as in equation (3.33)) and perform the OLS regression. However, this chapter argues that it is inappropriate to use the OLS regression since one need to eliminate the nonstandard nature of the limit theory which attributes to the long memory of the forward premium and the foreign exchange risk premium. Additionally, Maynard and Phillips (2001) suggest that the finding of short memory in the spot return and a long memory in the forward premium already causes for rejection in the unbiasedness hypothesis. The contribution of this section is to propose a better estimator for the foreign exchange risk premia. Additionally, finding that the estimated value of the foreign exchange risk premia is non-trivial and that the forward premium is nonstationary proves that the Uncovered Interest Parity does not hold.

3.9 Conclusion

This work corroborates earlier findings that a principal reason for rejection of the forward rate unbiasedness hypothesis is differences in persistence between the forward premium series and the spot rate return and the existence of the time varying exchange rate risk premium. In the risk premia estimation part, this work suggests a new method in estimating the foreign exchange rate risk premium.

This chapter extends the research of Baillie and Bollerslev (1994a) by studying 10 mostly traded currency using daily data. In accordance with their work, the forward premium in the foreign exchange market, $s_t - f_{t,k}$ is best characterised by a (nonstationary) fractionally integrated process ($I(d)$ where $0 < d < 1$), while the spot return series are found to be stationary which follow an $I(0)$ process in all sample currencies. This implies imbalance in the traditional regression of the return on the spot rate on the forward premium.

The presence of structural breaks also proves to play an important role in explaining the long memory of the forward premium data. We correct for structural breaks in the forward premium data using Bai and Perron (1998,2003a). Allowing for structural breaks reduces the persistence of the forward premium across all currencies and model specifications. Nevertheless, the forward premium still follows the (nonstationary) fractionally integrated process after correcting for multiple structural breaks.

In sum, the finding that there is an unbalance in the order of integrations proves that the statistical artefacts of the data contribute to the rejection of the hypothesis. This chapter thus argued that it is inappropriate to use OLS regression since one need to eliminate the non-standard nature of the limit theory which attributes to the long memory of the forward premium already causes for rejection in the unbiasedness hypothesis.

This chapter suggests a new methodology in estimating the exchange rate risk premium by modelling the conditional variance of the forward rate forecast error $(s_{t+k} - f_{t,k})$ using the fractionally integrated GARCH model. In estimating the foreign exchange risk premia, the FIGARCH model is found to be econometrically superior to regular stable GARCH model as in Baillie and Bollerslev (1990) and the ARCH-in-mean model as used by Domowitz and Hakkio (1985). Finding that the risk premia exists can explain why the Uncovered Interest Parity does not hold.

Lastly, to further study the high frequency dynamics of the forward premium and attempt to impose the balance in the UIP regression by adding the risk premia to the UIP regression is not plausible here. To find fractional cointegration between the risk premium and the forward premium which will lead to a reasonable estimate of beta coefficient requires that these two variables have the same order of integration, and the results in the chapter find that they are not.

Table 3.1: Summary statistics

Panel 1: Daily one month return on spot exchange rates (with US dollars), $(s_{t+k} - s_t)*100$

Data	Obs	Mean	Std. Dev.	Skewness	Kurtosis	JB test
Australian Dollar (AUD)	3390	0.104	2.716	-0.125	3.023	8.88
British Pound (GBP)	3477	0.162	2.109	-0.040	3.188	237.00
Canadian Dollar (CAD)	3468	-0.120	1.663	-0.240	3.296	45.86
Danish Krone (DKK)	3430	-0.115	2.623	-0.176	2.959	17.88
EURO (EUR)	2184	0.132	2.668	0.246	2.912	22.72
Japanese Yen (JPY)	3475	0.051	3.126	-0.716	5.163	78.30
New Zealand Dollar (NZD)	3280	0.164	2.972	-0.282	3.192	74.77
Norwegian Krone (NOK)	2177	-0.194	2.812	-0.190	2.882	14.40
Swedish Krone (SEK)	2181	-0.108	2.831	-0.160	2.740	15.42
Swiss Franc (CHF)	3476	-0.103	2.839	-0.360	2.940	33.29

Panel 2: Forward premium (%), $(f_{t,k} - s_t)*100$

Data	Obs	Mean	Std. Dev.	Skewness	Kurtosis	JB test
Australian Dollar (AUD)	3410	-0.113	0.129	-0.357	2.058	198.70
British Pound (GBP)	3497	-0.090	0.094	-0.199	1.977	186.40
Canadian Dollar (CAD)	3488	-0.012	0.099	0.029	2.459	42.98
Danish Krone (DKK)	3450	-0.042	0.129	0.413	1.835	293.10
EURO (EUR)	2204	0.048	0.134	-0.091	1.375	245.40
Japanese Yen (JPY)	3495	-0.332	0.143	0.438	1.785	333.30
New Zealand Dollar (NZD)	3300	-0.202	0.144	0.561	2.306	146.50
Norwegian Krone (NOK)	2197	0.093	0.207	0.345	1.926	149.10
Swedish Krone (SEK)	2201	-0.045	0.172	0.309	1.484	245.70
Swiss Franc (CHF)	3496	-0.216	0.127	0.437	1.842	457.00

Panel 3: Forward rate forecast error (%), $(f_{t,k} - s_{t+k})*100$

Data	Obs	Mean	Std. Dev.	Skewness	Kurtosis	JB test
Australian Dollar (AUD)	3390	0.217	2.745	-0.128	2.981	9.28
British Pound (GBP)	3477	0.252	2.115	-0.030	3.189	123.01
Canadian Dollar (CAD)	3468	-0.108	1.679	-0.280	3.265	55.46
Danish Krone (DKK)	3430	-0.074	2.651	-0.173	2.928	17.94
EURO (EUR)	2184	0.084	2.699	0.239	2.881	22.12
Japanese Yen (JPY)	3475	0.383	3.148	-0.678	4.997	184.20
New Zealand Dollar (NZD)	3280	0.371	2.999	-0.289	3.139	73.19
Norwegian Krone (NOK)	2177	-0.289	2.841	-0.199	2.846	16.50
Swedish Krone (SEK)	2181	-0.064	2.870	-0.167	2.727	16.89
Swiss Franc (CHF)	3476	0.113	2.867	-0.350	2.899	16.94

Note: The exchange rate data are observed during 3rd January 1994 to 16th June 2007, except, EUR, NOK and SEK. The exchange rate data for these 3 countries are observed from 4th January 1999 to 16th June 2007.

Table 3.2: The Unbiaseness Regression
(Dependent variable: the spot return)

	beta	cons	$F(H_0\text{:beta}=1)$	R^2
GBP	-1.129	0.001	2.70	0.001
	(1.295)	(0.002)	[0.100]	
AUD	-4.202	-0.004	16.31	0.040
	(1.288)***	(0.002)*	[0.000]	
CAD	-2.158	-0.002	11.05	0.016
	(0.950)**	(0.001)	[0.001]	
CHF	-1.283	-0.003	8.38	0.008
	(0.789)	(0.002)	[0.004]	
DKK	-3.967	-0.003	23.53	0.038
	(1.024)***	(0.002)*	[0.000]	
EUR	-4.111	0.003	15.22	0.043
	(1.310)***	(0.002)*	[0.000]	
JPY	-2.008	-0.006	18.25	0.017
	(0.704)***	(0.002)***	[0.000]	
NZD	-2.373	-0.004	10.60	0.014
	(1.036)**	(0.003)	[0.001]	
NOK	-1.353	-0.001	6.34	0.010
	(0.935)	(0.002)	[0.012]	
SEK	-3.229	-0.003	15.36	0.039
	(1.079)***	(0.002)	[0.000]	

Notes:
[1]The first column represents the estimated beta coefficients of the UIP regression. The second column is the estimated constant. The numbers in parentheses are the Newey-West (1987) standard errors of the corresponding parameter estimates.
*, **,*** indicate 10%, 5% and 1% significance respectively.
[2]The third column represent the F-test statistics for the null of estimated beta equal to unity. The figures in the squared bracket are the corresponding p-values [Prob>F].
[3]The last column represents the R-squared of the OLS regression.

Table 3.3: Semi-parametric estimates for d before adjusting for structural breaks in the forward premium data

Country	Power	ordinates	MLP: *d*	MLP: t(H0:d=0) P>|t|	MLP: z(H0:d=1) P>|z|	LP: *d*	LP: t(H0:d=0) P>|t|
1. AUD	0.7	297	0.897 (0.038)	23.787 [0.000]	-2.772 [0.006]	0.892 (0.039)	22.606 [0.000]
	0.75	446	0.876 (0.032)	27.221 [0.000]	-4.097 [0.000]	0.848 (0.032)	26.299 [0.000]
	0.8	670	0.811 (0.027)	30.470 [0.000]	-7.641 [0.000]	0.740 (0.026)	28.206 [0.000]
2. CAD	0.7	301	1.022 (0.035)	29.452 [0.000]	0.592 [0.554]	1.032 (0.038)	27.135 [0.000]
	0.75	453	1.027 (0.031)	33.407 [0.000]	0.894 [0.371]	1.023 (0.030)	33.569 [0.000]
	0.8	682	0.979 (0.025)	39.214 [0.000]	-0.876 [0.381]	0.990 (0.025)	39.496 [0.000]
3. CHF	0.7	302	0.983 (0.032)	30.736 [0.000]	-0.453 [0.651]	0.986 (0.032)	30.931 [0.000]
	0.75	454	0.969 (0.037)	26.544 [0.000]	-0.832 [0.405]	0.936 (0.027)	35.245 [0.000]
	0.8	683	0.970 (0.033)	28.993 [0.000]	-0.821 [0.412]	0.874 (0.023)	37.661 [0.000]
4. DKK	0.7	299	1.018 (0.033)	31.126 [0.000]	0.493 [0.622]	1.016 (0.031)	33.320 [0.000]
	0.75	450	0.995 (0.027)	36.993 [0.000]	-0.181 [0.856]	1.002 (0.026)	38.738 [0.000]
	0.8	676	0.963 (0.022)	42.815 [0.000]	-1.491 [0.136]	0.982 (0.022)	44.954 [0.000]
5. EUR	0.7	218	0.931 (0.046)	20.036 [0.000]	-1.580 [0.114]	0.938 (0.050)	18.818 [0.000]
	0.75	321	0.913 (0.039)	23.720 [0.000]	-2.424 [0.015]	0.906 (0.042)	21.440 [0.000]
	0.8	472	0.756 (0.034)	22.406 [0.000]	-8.266 [0.000]	0.736 (0.036)	20.458 [0.000]
6. GBP	0.7	302	1.005 (0.037)	27.325 [0.000]	0.148 [0.883]	1.009 (0.038)	26.833 [0.000]
	0.75	454	0.992 (0.031)	31.953 [0.000]	-0.279 [0.781]	1.000 (0.030)	33.433 [0.000]
	0.8	683	0.978 (0.029)	33.516 [0.000]	-0.747 [0.455]	0.909 (0.025)	35.676 [0.000]
7. JPY	0.7	302	0.971 (0.037)	26.158 [0.000]	-0.773 [0.440]	0.984 (0.038)	25.693 [0.000]
	0.75	454	0.953 (0.032)	29.882 [0.000]	-1.565 [0.118]	0.899 (0.031)	29.009 [0.000]
	0.8	683	0.813 (0.027)	30.640 [0.000]	-7.632 [0.000]	0.811 (0.026)	30.995 [0.000]
8. NOK	0.7	218	1.056 (0.040)	26.101 [0.000]	1.296 [0.195]	1.051 (0.033)	31.423 [0.000]
	0.75	320	1.003 (0.034)	29.666 [0.000]	0.079 [0.937]	1.016 (0.029)	34.894 [0.000]
	0.8	471	0.956 (0.043)	22.288 [0.000]	-1.214 [0.225]	0.919 (0.027)	33.433 [0.000]
9. NZD	0.7	290	0.946 (0.037)	25.468 [0.000]	-1.425 [0.154]	0.937 (0.040)	23.204 [0.000]
	0.75	435	0.893 (0.030)	29.585 [0.000]	-3.495 [0.000]	0.892 (0.034)	26.339 [0.000]
	0.8	652	0.816 (0.026)	31.316 [0.000]	-7.337 [0.000]	0.824 (0.029)	28.625 [0.000]
10. SEK	0.7	218	0.976 (0.043)	22.516 [0.000]	-0.561 [0.575]	1.036 (0.043)	24.040 [0.000]
	0.75	321	0.930 (0.042)	22.060 [0.000]	-1.614 [0.106]	0.926 (0.037)	25.322 [0.000]
	0.8	472	0.851 (0.029)	29.752 [0.000]	-5.059 [0.000]	0.780 (0.033)	23.706 [0.000]

Notes: [1]Panel on the left hand side is estimated by the Modified Log Periodogram Regression estimator (MLP) by Kim and Phillips (1999, 2006). Panel on the right hand side is estimated by the Log Periodogram Regression estimator (LP) by Geweke Porter-Hudak (1983).
[2]Column 2 is the choice of the root (T) or the power value (alpha) in equation (3.40) in the main text. A range of power values (0.7, 0.75 and 0.8) is calculated. Column 3 is the number of harmonic ordinates to be included in the spectral regression. The term *d* represents the long memory (fractional integration) parameter of the forward premium data, followed by conventional standard errors. Column 5 is the t-test statistics for the test of d=0 and the corresponding p-values. Column 6 is the test statistics (zd) for the test of d=1 and its p-value.

Table 3.4: Parametric estimates for d before adjusting for structural breaks in the forward premium data

		Estimated d_{ML}			
Country		ARFIMA(1,d,0)		ARFIMA(0,d,1)	
1	AUD	0.714	(0.013)	1.150	(0.071)
2	CAD	0.948	(0.017)	1.010	(0.028)
3	CHF	0.787	(0.013)	0.812	(0.019)
4	DKK	0.909	(0.016)	0.921	(0.020)
5	EUR	0.857	(0.015)	1.332	(0.061)
6	GBP	0.860	(0.014)	1.167	(0.063)
7	JPY	0.767	(0.013)	0.774	(0.022)
8	NOK	0.976	(0.016)	1.048	(0.019)
9	NZD	0.806	(0.015)	0.966	(0.044)
10	SEK	0.831	(0.015)	1.026	(0.046)

Notes:
[1] All models were estimated by exact maximum likelihood (ML).
[2] The numbers in parentheses are asymptotic standard errors of the corresponding parameter estimates.

Table 3.5: Multiple Structural Change Model Estimates for forward premium with 1 month maturity (/US Dollar)

Country	Mean	Estimated intercept			Estimated break date			
AUD	-0.113	TB1	-0.127	(0.002)	*10-Dec-96*	[05-Nov-96	:	01-Jan-97]
		TB2	0.017	(0.001)	*18-Sep-01*	[07-Sep-01	:	26-Sep-01]
		TB3	-0.279	(0.002)	*01-Jul-05*	[27-Jun-05	:	25-Jul-05]
		TB4	-0.088	(0.002)				
CAD	-0.012	TB1	0.063	(0.002)	*21-Mar-96*	[01-Feb-96	:	21-Jan97]
		TB2	-0.148	(0.002)	*25-Mar-98*	[21-Jan-98	:	18-Feb99]
		TB3	-0.049	(0.002)	*24-Apr-01*	[22-Dec-99	:	10-Jun04]
		TB4	0.111	(0.002)	*19-Jan-05*	[11-Jan-05	:	18-Apr-05]
		TB5	-0.076	(0.002)				
CHF	-0.216	TB1	-0.277	(0.002)	*08-Jan-01*	[04-Dec-00	:	17-Oct-01]
		TB2	-0.079	(0.003)	*04-Apr-05*	[29-Mar-05	:	26-May-05]
		TB3	-0.283	(0.004)				
DKK	-0.042	TB1	0.065	(0.002)	*26-Jan-96*	[23-Jan-96	:	14-Oct-96]
		TB2	-0.140	(0.002)	*22-Feb-01*	[19-Feb-01	:	13-Jul-01]
		TB3	0.122	(0.002)	*01-Dec-04*	[22-Jan-04	:	24-Feb05]
		TB4	-0.142	(0.002)				
EUR	0.048	TB1	0.193	(0.007)	*04-Jan-01*	[08-Dec-00	:	09-Jan-01]
		TB2	-0.081	(0.010)	*20-Jan-05*	[05-Jan-05	:	23-Feb-05]
		TB3	0.144	(0.010)				
GBP	-0.090	TB1	-0.055	(0.002)	*17-Sep-01*	[30-Aug-01	:	07-Nov-03]
		TB2	-0.171	(0.002)	*14-Jun-05*	[26-May-05	:	19-Jul-05]
		TB3	-0.003	(0.002)				
JPY	-0.332	TB1	-0.302	(0.003)	*19-Jan-96*	[16-Jan-96	:	24-Feb-98]
		TB2	-0.447	(0.002)	*31-Jul-01*	[04-Jul-01	:	06-Aug-01
		TB3	-0.151	(0.002)	*06-Jun-05*	[26-May-05	:	15-Jul-05]
		TB4	-0.390	(0.003)				
NZD	-0.202	TB1	-0.215	(0.002)	*03-Aug-98*	[16-Apr-98	:	15-Sep-98]
		TB2	-0.000	(0.002)	*14-Sep-01*	[12-Sep-01	:	02-Nov-01]
		TB3	-0.324	(0.003)	*22-Aug-03*	[27-Jun-03	:	28-Sep-04]
		TB4	-0.361	(0.003)	*25-Jul-05*	[07-Jul-05	:	29-Jul05]
		TB5	-0.222	(0.003)				
NOK	0.093	TB1	0.206	(0.004)	*20-Jul-04*	[16-Jul-04	:	30-Aug-04]
		TB2	-0.119	(0.005)				
SEK	-0.045	TB1	-0.186	(0.002)	*20-Apr-01*	[02-Apr-01	:	27-Apr-01]
		TB2	0.160	(0.002)	*04-Mar-04*	[06-Feb-04	:	16-Jun-04]
		TB3	0.010	(0.002)	*10-Jun-05*	[31-May-05	:	27-Jun-05]
		TB4	-0.209	(0.002)				

Notes:
[1]The estimated parameters are based on 95% confidence interval. Asymptotic standard errors are reported in parentheses.
[2] The second column represents the mean of forward premium data before adjusting for the structural breaks.
[3]In the third column, TB1,..,TB5 are the estimated mean of the forward premium after adjusting for multiple structural breaks, with the break date report in italics.
[4]The dates in the square bracket [.] represent 95 percent confident interval for the break dates.

Table 3.6: Multiple Structural Change Tests Results (/US Dollar) for the forward premium series.

Statistics/Country	AUD	CAD	CHF	DKK	EUR	GBP	JPY	NZD	NOK	SEK
					Tests					
SupFT(1)	64.393***	44.302***	11.232**	99.301***	15.862***	33.247***	17.057***	48.307***	16.130***	22.383***
SupFT(2)	279.018***	53.804***	124.953***	70.577***	219.395***	160.657***	215.098***	264.852***	35.659***	245.165***
SupFT(3)	350.124***	43.035***	146.548***	74.352***	268.450***	90.710***	276.756***	224.548***	154.689***	206.441***
SupFT(4)	316.056***	35.544***	163.302***	91.911***	391.070***	78.205***	143.945***	434.673***	80.670***	203.138***
SupFT(5)	317.253***	29.121***	127.992***	118.947***	285.647***	69.434***	56.635***	358.411***	35.541***	204.763***
UDmax	350.124***	53.804***	163.302***	118.947***	391.070***	160.657***	276.756***	434.673***	154.689***	264.763***
Wdmax	696.172***	63.939***	280.862***	261.014***	672.421***	190.919***	398.422***	786.488***	222.689***	580.989***
SupFT(2/1)	469.934***	16.802***	266.176***	124.345***	277.171***	219.924***	257.426***	239.958***	5.371	162.599***
SupFT(3/2)	229.523***	12.012***	4.751	25.024***	1.741	2.141	11.781**	124.666***	1.351	14.199**
SupFT(4/3)	2.472	24.098***	5.068	0.442	3.630	9.518	2.518	22.293***	0.391	1.163
SupFT(5/4)	0.704	4.301	0.598	0.412		9.819		0.255		
					Numbers of Changes Selected					
Sequential	3	4	2	3	2	2	3	4	1	3
LWZ	4	5	4	5	5	5	4	4	4	4
BIC	4	5	4	4	5	5	4	4	4	4

Notes:
[1]The SupFT(l) tests is the F statistics of no structural breaks versus a fixed number of breaks. The number in the parenthesis is the number of breaks.
[2]The UDmax and the WD max are the test statistics used to determine if structural change has occurred against unknown number of breaks.
[3]The SupFT(l+1/l) test is test of the null of l breaks versus the alternative of l+1 breaks.
*, **,*** indicate 10%, 5% and 1% significance respectively.

Table 3.7: Semi parametric estimates for d after adjusting for structural breaks in the forward premium data

			MLP			LP	
Country	Power	Ordinates	*d*	t(H0:d=0) P>|t|	z(H0:d=1) P>|z|	*d*	t(H0:d=0) P>|t|
1. AUD	0.7	297	0.889 (0.041)	21.913 [0.000]	-2.987 [0.003]	0.898 (0.038)	23.706 [0.000]
	0.75	446	0.845 (0.033)	25.594 [0.000]	-5.100 [0.000]	0.876 (0.032)	27.369 [0.000]
	0.8	670	0.739 (0.027)	27.577 [0.000]	-10.530 [0.000]	0.810 (0.026)	30.677 [0.000]
2. CAD	0.7	301	0.924 (0.038)	24.256 [0.000]	-2.053 [0.040]	0.922 (0.038)	24.554 [0.000]
	0.75	453	0.914 (0.030)	30.850 [0.000]	-2.847 [0.004]	0.951 (0.031)	30.492 [0.000]
	0.8	682	0.889 (0.031)	28.573 [0.000]	-3.677 [0.000]	0.938 (0.026)	36.653 [0.000]
3. CHF	0.7	302	0.932 (0.027)	34.794 [0.000]	-2.257 [0.024]	0.970 (0.038)	25.700 [0.000]
	0.75	454	0.877 (0.026)	34.111 [0.000]	-5.007 [0.000]	0.910 (0.030)	30.068 [0.000]
	0.8	683	0.859 (0.025)	34.690 [0.000]	-5.766 [0.000]	0.852 (0.026)	33.206 [0.000]
4. DKK	0.7	299	0.915 (0.031)	29.343 [0.000]	-2.815 [0.005]	0.909 (0.040)	22.927 [0.000]
	0.75	450	0.912 (0.038)	23.735 [0.000]	-2.370 [0.018]	0.906 (0.031)	28.917 [0.000]
	0.8	676	0.902 (0.030)	30.446 [0.000]	-3.313 [0.001]	0.896 (0.025)	35.563 [0.000]
5. EUR	0.7	218	0.930 (0.044)	21.194 [0.000]	-1.603 [0.109]	0.918 (0.046)	19.917 [0.000]
	0.75	321	0.925 (0.038)	24.447 [0.000]	-2.092 [0.036]	0.899 (0.038)	23.869 [0.000]
	0.8	472	0.855 (0.031)	27.512 [0.000]	-4.900 [0.000]	0.836 (0.031)	26.894 [0.000]
6. GBP	0.7	302	0.940 (0.026)	35.612 [0.000]	-2.428 [0.015]	0.973 (0.033)	29.532 [0.000]
	0.75	454	0.906 (0.025)	36.819 [0.000]	-3.837 [0.000]	0.980 (0.029)	33.679 [0.000]
	0.8	683	0.863 (0.023)	37.625 [0.000]	-5.565 [0.000]	0.906 (0.025)	36.853 [0.000]
7. JPY	0.7	302	0.913 (0.039)	23.358 [0.000]	-2.364 [0.018]	0.906 (0.039)	23.191 [0.000]
	0.75	454	0.882 (0.032)	27.345 [0.000]	-3.913 [0.000]	0.878 (0.032)	27.253 [0.000]
	0.8	683	0.788 (0.025)	31.095 [0.000]	-8.422 [0.000]	0.813 (0.027)	30.652 [0.000]
8. NOK	0.7	218	1.094 (0.050)	21.814 [0.000]	2.155 [0.031]	1.038 (0.041)	25.055 [0.000]
	0.75	320	0.911 (0.025)	36.194 [0.000]	-3.611 [0.000]	0.996 (0.036)	27.675 [0.000]
	0.8	471	0.801 (0.026)	30.250 [0.000]	-8.092 [0.000]	0.891 (0.031)	29.150 [0.000]
9. NZD	0.7	290	0.859 (0.040)	21.473 [0.000]	-3.736 [0.000]	0.867 (0.040)	21.849 [0.000]
	0.75	435	0.840 (0.031)	26.820 [0.000]	-5.204 [0.000]	0.846 (0.031)	27.228 [0.000]
	0.8	652	0.766 (0.038)	20.339 [0.000]	-7.916 [0.000]	0.795 (0.025)	31.390 [0.000]
10. SEK	0.7	218	0.911 (0.037)	24.929 [0.000]	-2.495 [0.013]	0.927 (0.043)	21.420 [0.000]
	0.75	321	0.909 (0.034)	27.094 [0.000]	-2.549 [0.011]	0.904 (0.034)	26.394 [0.000]
	0.8	472	0.790 (0.032)	24.548 [0.000]	-7.117 [0.000]	0.851 (0.030)	28.417 [0.000]

Notes: [1]Panel on the left hand side is estimated by the Modified Log Periodogram Regression estimator (MLP) by Kim and Phillips (1999, 2006). Panel on the right hand side is estimated by the Log Periodogram Regression estimator (LP) by Geweke Porter-Hudak (1983).
[2]Column 2 is the choice of the root (T) or the power value (alpha) in equation (3.40) in the main text. A range of power values (0.7, 0.75 and 0.8) is calculated. Column 3 is the number of harmonic ordinates to be included in the spectral regression. The term *d* represents the long memory (fractional integration) parameter of the forward premium data, followed by conventional standard errors. Column 5 is the t-test statistics for the test of d=0 and the corresponding p-values. Column 6 is the test statistics (zd) for the test of d=1 and its p-value.

Table 3.8: Parametric estimates for d after adjusting for structural breaks in the forward premium data

		Estimated d_{ML}			
Country		ARFIMA(1,d,0)		ARFIMA(0,d,1)	
1	AUD	0.739	(0.017)	0.855	(0.033)
2	CAD	0.935	(0.020)	0.954	(0.025)
3	CHF	0.764	(0.016)	0.767	(0.018)
4	DKK	0.907	(0.021)	0.907	(0.021)
5	EUR	0.855	(0.023)	0.893	(0.034)
6	GBP	0.849	(0.016)	0.968	(0.038)
7	JPY	0.754	(0.017)	0.758	(0.020)
8	NOK	0.899	(0.020)	0.921	(0.027)
9	NZD	0.799	(0.018)	0.863	(0.031)
10	SEK	0.825	(0.025)	0.819	(0.027)

Notes:
[1]All models were estimated by exact maximum likelihood (ML).
[2]The numbers in parentheses are asymptotic standard errors of the corresponding parameter estimates.

Table 3.9: Unit root tests for the spot return series

Table 3.9A: The ADF-GLS test statistics for null of unit root

		ADF-GLS Test			Critical Values		
	Country	Test Statistics	Lag Length	RMSE	1%	5%	10%
1	AUD	-3.493***	28	0.651	-3.480	-2.833	-2.546
2	CAD	-4.315***	29	0.395	-3.480	-2.832	-2.546
3	CHF	-4.310***	28	0.683	-3.480	-2.833	-2.546
4	DKK	-7.145***	23	0.616	-3.480	-2.835	-2.548
5	EUR	-3.480***	25	0.657	-3.480	-2.831	-2.545
6	GBP	-4.209***	25	0.527	-3.480	-2.834	-2.547
7	JPY	-3.923***	28	0.693	-3.480	-2.833	-2.546
8	NOK	-4.989***	25	0.658	-3.480	-2.831	-2.545
9	NZD	-5.149***	28	0.745	-3.480	-2.832	-2.546
10	SEK	-4.180***	25	0.717	-3.480	-2.831	-2.545

Notes:
*, **,*** indicate 10%, 5% and 1% significance respectively.

Table 3.9B: The KPSS statistics for null of level stationary. (The 1%, 5% and 10% critical values are 0.739, 0.463 and 0.347 respectively)

	Country									
Lag	AUD	CAD	CHF	DKK	EUR	GBP	JPY	NOK	NZD	SEK
0	0.304	0.394	0.122	0.203	0.331	0.203	0.054	0.191	0.375	0.287
1	0.304	0.395	0.122	0.218	0.344	0.203	0.054	0.191	0.375	0.290
2	0.307	0.390	0.122	0.223	0.352	0.202	0.055	0.191	0.376	0.299
3	0.313	0.384	0.122	0.224	0.360	0.203	0.055	0.192	0.377	0.306
4	0.315	0.386	0.122	0.226	0.367	0.205	0.055	0.194	0.378	0.313
5	0.315	0.392	0.122	0.225	0.370	0.207	0.055	0.194	0.375	0.316
6	0.315	0.400	0.122	0.226	0.370	0.208	0.055	0.193	0.372	0.317
7	0.315	0.406	0.121	0.225	0.370	0.207	0.055	0.193	0.369	0.318
8	0.316	0.407	0.120	0.225	0.369	0.206	0.055	0.193	0.365	0.318
9	0.316	0.409	0.121	0.225	0.372	0.206	0.055	0.192	0.362	0.316
10	0.316	0.409	0.121	0.225	0.376	0.207	0.054	0.193	0.359	0.315
11	0.318	0.409	0.121	0.226	0.385	0.209	0.054	0.194	0.358	0.317
12	0.319	0.409	0.122	0.226	0.390	0.211	0.054	0.196	0.358	0.318
13	0.320	0.407	0.122	0.226	0.392	0.212	0.054	0.198	0.358	0.317
14	0.322	0.407	0.122	0.225	0.394	0.213	0.054	0.199	0.359	0.317
15	0.323	0.408	0.122	0.225	0.395	0.215	0.054	0.199	0.360	0.317
16	0.325	0.408	0.122	0.224	0.397	0.217	0.054	0.199	0.362	0.317
17	0.325	0.408	0.122	0.224	0.399	0.219	0.054	0.198	0.362	0.316
18	0.326	0.407	0.122	0.223	0.399	0.221	0.054	0.198	0.363	0.315
19	0.325	0.406	0.122	0.222	0.399	0.223	0.053	0.198	0.363	0.314
20	0.324	0.403	0.122	0.221	0.395	0.224	0.053	0.198	0.361	0.313
21	0.324	0.402	0.121	0.220	0.391	0.225	0.053	0.198	0.360	0.312
22	0.323	0.400	0.121	0.218	0.387	0.226	0.053	0.198	0.358	0.310
23	0.323	0.399	0.120	0.217	0.383	0.227	0.053	0.198	0.356	0.309
24	0.323	0.399	0.120	0.215	0.378	0.228	0.053	0.197	0.354	0.307
25	0.323	0.398	0.119	0.214	0.375	0.228	0.053	0.197	0.351	0.305
26	0.322	0.397	0.118	0.212	-	0.228	0.052	-	0.349	-
27	0.322	0.395	0.118	0.211	-	0.229	0.052	-	0.347	-
28	0.322	0.393	0.118	0.210	-	0.230	0.052	-	0.346	-
29	-	0.391	0.118	0.210	-	-	0.052	-	-	-

Table 3.10: Estimated FIGARCH Models for currencies daily forward rate forecast error (Using BBM's methodology).

	AUD	CAD	CHF	DKK	EUR	GBP	JPY	NOK	NZD	SEK
α_0	0.004	0.008	0.003	0.002	0.002	0.006	0.006	-0.001	0.008	-0.004
	(0.010)	(0.006)	(0.010)	(0.010)	(0.012)	(0.008)	(0.009)	(0.014)	(0.009)	(0.015)
β_0	0.011	0.002	0.018	0.009	0.007	0.01	0.03	0.004	0.021	0.034
	(0.005)	(0.001)	(0.007)	(0.005)	(0.006)	(0.004)	(0.012)	(0.003)	(0.007)	(0.017)
δ	0.394	0.528	0.372	0.430	0.432	0.403	0.283	0.279	0.346	0.281
	(0.067)	(0.118)	(0.056)	(0.087)	(0.125)	(0.065)	(0.049)	(0.046)	(0.040)	(0.064)
α_1	0.361	0.279	0.376	0.294	0.249	0.343	0.357	0.490	0.270	0.380
	(0.046)	(0.062)	(0.042)	(0.039)	(0.044)	(0.043)	(0.069)	(0.065)	(0.053)	(0.062)
β_1	0.723	0.783	0.716	0.734	0.708	0.706	0.592	0.61	0.574	0.642
	(0.052)	(0.067)	(0.045)	(0.066)	(0.114)	(0.051)	(0.084)	(0.063)	(0.059)	(0.076)
ln(L)	-3153.455	-1422.223	-4665.816	-3045.928	-1976.315	-3862.468	-4294.191	-2126.091	-3955.755	-2315.052
Skew	-0.079	-0.102	-0.128	-0.068	-0.021	-0.075	-0.455	0.086	-0.291	0.130
	[0.058]	[0.014]	[0.000]	[0.101]	[0.691]	[0.035]	[0.000]	[0.103]	[0.000]	[0.013]
E.Kurt	1.732	0.797	1.062	1.002	0.489	1.300	2.536	0.310	1.759	0.620
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.003]	[0.000]	[0.000]
JB	427.150	97.881	222.590	146.280	22.071	335.400	1349.000	11.347	584.910	41.053
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.003]	[0.000]	[0.000]

Notes: The first 10 columns refer to the daily forward rate forecast error series. ln(L) is a the value of the maximised Gaussian log likelihood; and standard errors are presented in parentheses below corresponding parameter estimates. The value of sample skewness and kurtosis and their p-values are also based on the standardised residuals. The values in the square brackets are the p-values of the test statistics.

Table 3.11: test statistics (BBM's specifications)

	AUD	CAD	CHF	DKK	EUR	GBP	JPY	NOK	NZD	SEK
Box Pierce Q statistics (the lag order is in the parenthesis) :										
Standardised residuals										
Q(10)	3.183	10.243	2.812	4.094	4.194	6.895	14.535	4.648	9.776	7.054
	[0.957]	[0.419]	[0.986]	[0.943]	[0.938]	[0.753]	[0.150]	[0.913]	[0.460]	[0.720]
Q(15)	6.277	13.405	7.967	6.345	11.692	15.758	16.673	11.901	11.177	11.140
	[0.959]	[0.571]	[0.925]	[0.973]	[0.702]	[0.398]	[0.339]	[0.687]	[0.740]	[0.743]
Q(20)	13.997	17.141	9.51	12.606	13.244	18.823	22.777	13.638	17.226	13.772
	[0.784]	[0.644]	[0.976]	[0.894]	[0.867]	[0.533]	[0.300]	[0.848]	[0.638]	[0.842]
Squared standardised residuals										
$Q^2(10)$	12.837	3.201	7.944	7.897	17.110	6.728	8.292	8.649	3.640	6.744
	[0.118]	[0.921]	[0.439]	[0.444]	[0.029]	[0.566]	[0.405]	[0.373]	[0.888]	[0.564]
$Q^2(15)$	15.085	4.843	9.358	12.159	19.062	15.798	11.376	8.917	7.677	12.371
	[0.302]	[0.978]	[0.745]	[0.515]	[0.121]	[0.260]	[0.579]	[0.779]	[0.864]	[0.498]
$Q^2(20)$	20.069	9.872	16.082	17.577	21.964	18.463	17.788	13.879	11.062	16.136
	[0.329]	[0.936]	[0.567]	[0.484]	[0.234]	[0.426]	[0.470]	[0.737]	[0.892]	[0.583]
Residual-Based diagnostic for conditional heteroskedasticity of Tse (2001)										
RBD(10)	12.892	3.254	8.058	7.371	14.281	6.802	8.349	8.848	3.732	6.668
	[0.230]	[0.975]	[0.623]	[0.690]	[0.161]	[0.744]	[0.595]	[0.647]	[0.959]	[0.756]
RBD(15)	15.047	5.115	9.388	10.856	17.491	15.724	11.305	9.076	7.841	11.804
	[0.448]	[0.991]	[0.856]	[0.763]	[0.290]	[0.401]	[0.731]	[0.874]	[9.930]	[0.694]
RBD(20)	19.986	9.253	15.879	15.525	20.923	18.181	17.539	14.034	10.866	16.274
	[0.459]	[0.980]	[0.724]	[0.746]	[0.402]	[0.576]	[0.618]	[0.829]	[0.950]	[0.700]
Sign bias t test (Engle and Ng, 1993)										
t	1.417	0.510	0.448	0.361	0.908	0.860	1.003	0.399	0.197	0.490
	[0.157]	[0.610]	[0.654]	[0.718]	[0.364]	[0.390]	[0.316]	[0.690]	[0.844]	[0.624]
Adjusted Pearson Chi-square Goodness-of-fit test										
χ^2, 40 cells	32.773	33.546	38.804	39.271	31.014	42.988	39.265	51.186	34.925	34.112
	[0.749]	[0.717]	[0.479]	[0.458]	[0.815]	[0.304]	[0.458]	[0.092]	[0.656]	[0.692]
χ^2,50 cells	45.988	60.287	52.612	55.948	37.403	52.596	54.350	63.282	46.876	41.673
	[0.596]	[0.129]	[0.336]	[0.230]	[0.887]	[0.337]	[0.278]	[0.083]	[0.560]	[0.762]
χ^2,60 cells	62.142	62.865	64.464	61.265	53.415	59.898	53.211	58.875	42.870	51.545
	[0.365]	[0.341]	[0.291]	[0.395]	[0.681]	[0.443]	[0.688]	[0.480]	[0.943]	[0.744]
ARCH-LM test (the lag order is in the parenthesis) of Engle (1982)										
ARCH(2)	1.287	0.512	0.706	0.452	1.458	1.765	0.483	1.238	0.008	0.153
	[0.231]	[0.599]	[0.494]	[0.636]	[0.233]	[0.171]	[0.617]	[0.290]	[0.992]	[0.858]
ARCH(5)	0.997	0.442	0.477	1.053	1.651	0.738	1.079	0.554	0.285	0.794
	[0.455]	[0.820]	[0.794]	[0.385]	[0.143]	[0.595]	[0.370]	[0.736]	[0.922]	[0.554]
ARCH(10)	1.003	0.325	0.797	0.776	2.956	0.656	0.836	0.890	0.356	0.669
	[0.455]	[0.975]	[0.632]	[0.653]	[0.399]	[0.766]	[0.594]	[0.542]	[0.965]	[0.754]

Notes:
[1]This table presents the misspecification test statistics of the estimates in table 3.10.
[2]The values in the square brackets are the p-values of the corresponding test statistics.

Table 3.12: Estimated FIGARCH Models for currencies daily forward rate forecast error (Using Chung's methodology).

	AUD	CAD	CHF	DKK	EUR	GBP	JPY	NOK	NZD	SEK
α_0	0.005	0.008	0.003	0.002	0.002	0.006	0.006	-0.001	0.008	-0.004
	(0.010)	(0.006)	(0.010)	(0.010)	(0.013)	(0.008)	(0.009)	(0.014)	(0.009)	(0.015)
β_0	0.378	0.108	0.709	0.361	0.063	0.439	0.445	0.402	0.388	0.499
	(0.083)	(0.029)	(0.176)	(0.098)	(0.008)	(0.154)	(0.095)	(0.058)	(0.096)	(0.076)
δ	0.389	0.451	0.404	0.426	0.425	0.426	0.298	0.299	0.338	0.283
	(0.062)	(0.064)	(0.057)	(0.071)	(0.136)	(0.060)	(0.052)	(0.075)	(0.037)	(0.066)
α_1	0.364	0.319	0.360	0.296	0.257	0.336	0.353	0.470	0.276	0.378
	(0.044)	(0.040)	(0.043)	(0.034)	(0.075)	(0.040)	(0.067)	(0.064)	(0.053)	(0.062)
β_1	0.720	0.746	0.732	0.732	0.726	0.719	0.601	0.718	0.572	0.643
	(0.051)	(0.045)	(0.045)	(0.058)	(0.093)	(0.048)	(0.084)	(0.064)	(0.058)	(0.077)
ln(L)	-3155.078	-1421.957	-4664.862	-3046.047	-1971.231	-3862.755	-4294.458	-2125.900	-3955.993	-2315.071
Skew	-0.079	-0.101	-0.128	-0.069	-0.004	-0.073	-0.458	0.102	-0.290	0.130
	[0.061]	[0.015]	[0.000]	[0.099]	[0.936]	[0.042]	[0.000]	[0.051]	[0.000]	[0.013]
E.Kurt	1.734	0.815	1.068	1.007	0.442	1.288	2.540	0.305	1.773	0.621
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.004]	[0.000]	[0.000]
JB	428.260	101.850	225.330	147.690	17.913	328.820	1354.400	12.226	593.170	41.135
	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.000]	[0.002]	[0.000]	[0.000]

Notes: The first 10 columns refer to the daily forward rate forecast error series. ln(L) is a the value of the maximised Gaussian log likelihood; and standard errors are presented in parentheses below corresponding parameter estimates. The value of sample skewness and kurtosis and their p-values are also based on the standardised residuals. The values in the square brackets are the p-values of the test statistics.

Table 3.13: test statistics (Chung's specification)

	AUD	CAD	CHF	DKK	EUR	GBP	JPY	NOK	NZD	SEK
Box Pierce Q statistics (the lag order is in the parenthesis) :										
Standardised residuals										
Q(10)	2.735	10.009	2.910	4.107	4.422	6.809	14.736	4.724	9.875	7.064
	[0.987]	[0.440]	[0.983]	[0.942]	[0.926	[0.743]	[0.142]	[0.909]	[0.452]	[0.719]
Q(15)	5.823	13.181	8.060	6.361	12.736	15.754	16.850	11.390	11.318	11.144
	[0.983]	[0.588]	[0.921]	[0.973]	[0.623]	[0.399]	[0.328]	[0.725]	[0.730]	[0.742]
Q(20)	13.436	16.927	9.562	12.629	14.612	18.791	23.059	13.201	17.416	13.783
	[0.858]	[0.658]	[0.975]	[0.893]	[0.798]	[0.535]	[0.286]	[0.869]	[0.626]	[0.841]
Squared standardised residuals										
$Q^2(10)$	12.806	3.108	8.328	7.849	11.659	7.159	8.327	5.584	3.708	6.730
	[0.119]	[0.927]	[0.402]	[0.448]	[0.167]	[0.520]	[0.402]	[0.694]	[0.882]	[0.566]
$Q^2(15)$	15.039	4.455	9.938	12.066	13.859	16.339	11.379	6.282	7.760	12.372
	[0.305]	[0.985]	[0.699]	[0.522]	[0.384]	[0.231]	[0.579]	[0.935]	[8.860]	[0.497]
$Q^2(20)$	19.983	9.286	17.086	17.436	16.444	18.787	17.629	12.199	11.080	16.121
	[0.334]	[0.953]	[0.517]	[0.493]	[0.562]	[0.405]	[0.480]	[0.837]	[0.891]	[0.584]
Residual-Based diagnostic for conditional heteroskedasticity of Tse (2001)										
RBD(10)	13.063	5.283	10.108	7.251	11.941	7.727	8.410	5.729	4.298	6.682
	[0.220]	[0.871]	[0.431]	[0.702]	[0.289]	[0.656]	[0.589]	[0.838]	[0.933]	[0.755]
RBD(15)	15.126	6.522	11.843	10.663	14.632	16.835	11.344	6.365	8.409	11.860
	[0.442]	[0.970]	[0.691]	[0.776]	[0.478]	[0.329]	[0.728]	[0.973]	[0.907]	[0.690]
RBD(20)	19.949	10.367	18.941	15.292	17.347	19.006	17.281	12.379	11.350	16.276
	[0.461]	[0.961]	[0.526]	[0.759]	[0.630]	[0.521]	[0.635]	[0.902]	[0.937]	[0.699]
Sign bias t test (Engle and Ng, 1993)										
t	1.359	0.560	0.507	0.365	1.144	0.869	1.030	0.485	0.563	0.486
	[0.174]	[0.575]	[0.612	[0.715]	[0.253]	[0.385]	[0.303]	[0.628]	[0.574]	[0.627]
Adjusted Pearson Chi-square Goodness-of-fit test										
χ^2, 40 cells	40.655	36.821	44.554	41.114	34.797	43.328	45.507	49.054	26.947	34.846
	[0.397]	[0.570]	[0.250]	[0.378]	[0.662]	[0.292]	[0.219]	[0.130]	[0.928]	[0.660]
χ^2,50 cells	42.035	50.599	57.723	58.192	38.267	49.51	54.193	62.619	41.669	44.332
	[0.749]	[0.410]	[0.184]	[0.173]	[0.866]	[0.453]	[0.283]	[0.091]	[0.762]	[0.663]
χ^2,60 cells	69.292	48.471	70.277	60.811	56.689	64.341	52.242	60.308	39.320	50.829
	[0.169]	[0.834]	[0.149]	[0.410]	[0.561]	[0.295]	[0.721]	[0.428]	[0.977]	[0.767]
ARCH-LM test (the lag order is in the parenthesis) of Engle (1982)										
ARCH(2)	1.288	0.473	0.709	0.454	0.598	1.875	0.448	0.112	0.009	0.154
	[0.231]	[0.623]	[0.492]	[0.635]	[0.550]	[0.154]	[0.639]	[0.894]	[0.991]	[0.857]
ARCH(5)	0.995	0.402	0.517	1.049	0.979	0.785	1.061	0.150	0.288	0.796
	[0.457]	[0.848]	[0.764]	[0.387]	[0.429]	[0.560]	[0.380]	[0.980]	[0.920]	[0.553]
ARCH(10)	0.996	0.319	0.831	0.771	1.196	0.698	0.837	0.570	0.359	0.668
	[0.464]	[0.977]	[0.599]	[0.658]	[0.288]	[0.727]	[0.593]	[0.840]	[0.964]	[0.755]

Notes:
[1]This table presents the misspecification test statistics of the estimates in table 3.12.
[2]The values in the square brackets are the p-values of the corresponding test statistics.

Figure 3.1: Forward premium series

Figure 3.1A: AUD1M

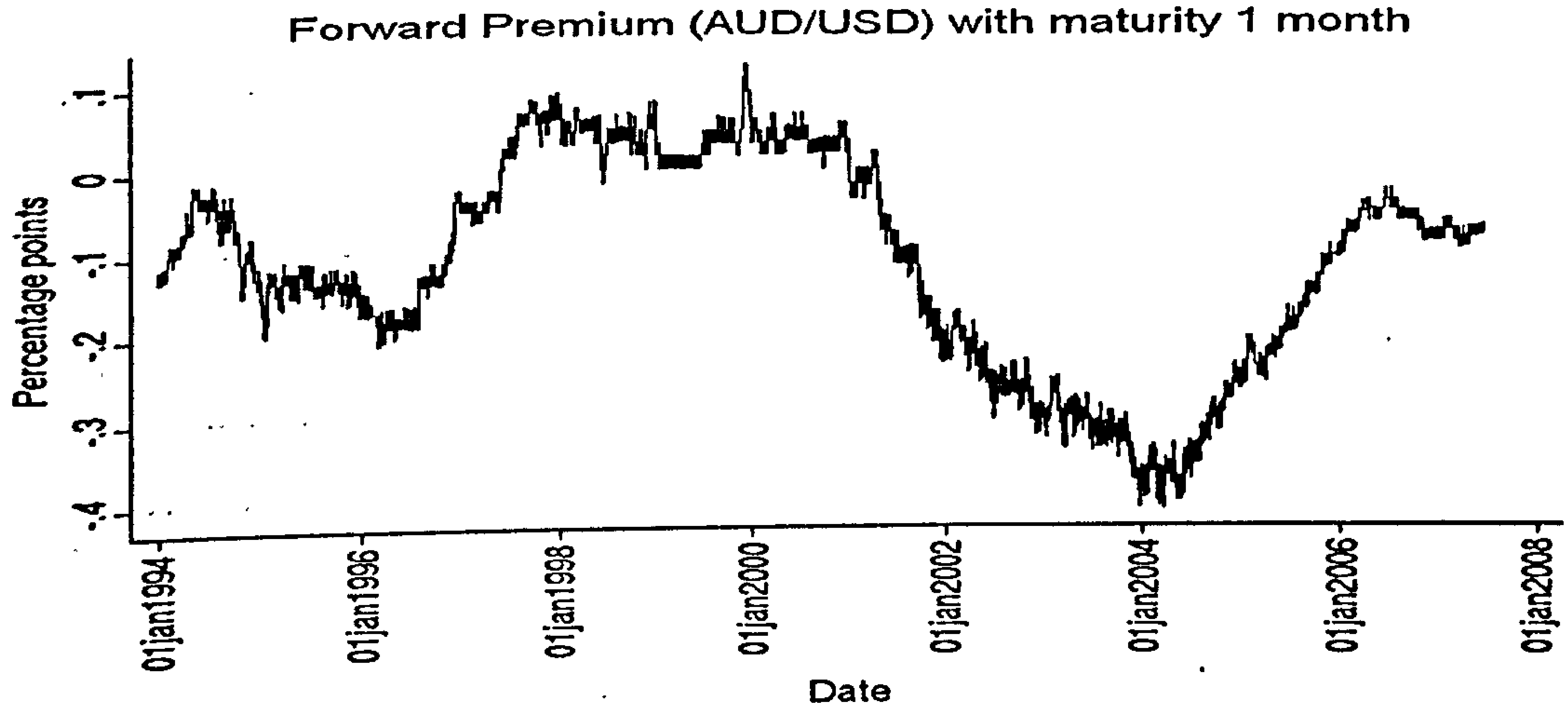


Figure 3.1B: CAD1M

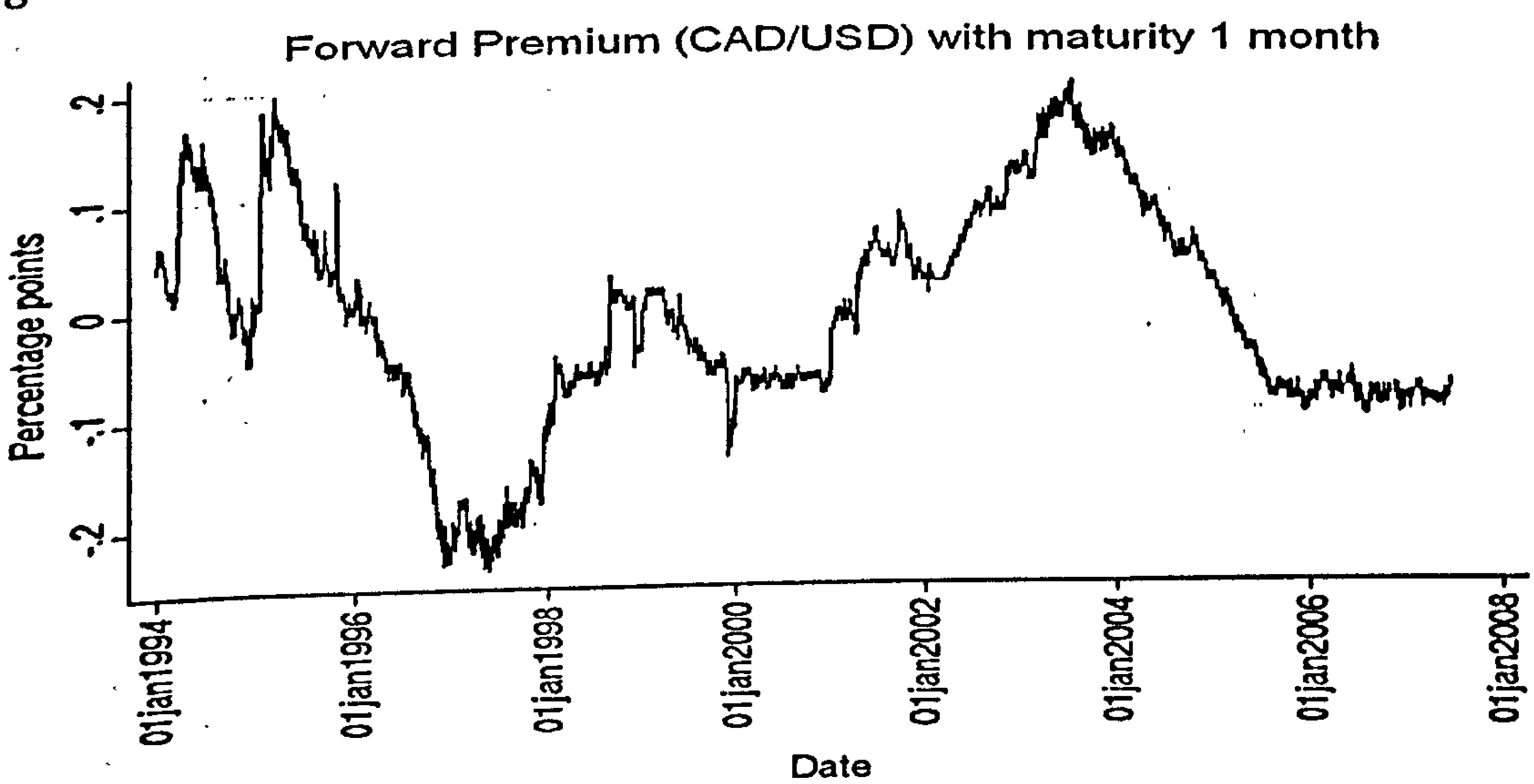


Figure 3.1C: CHF1M

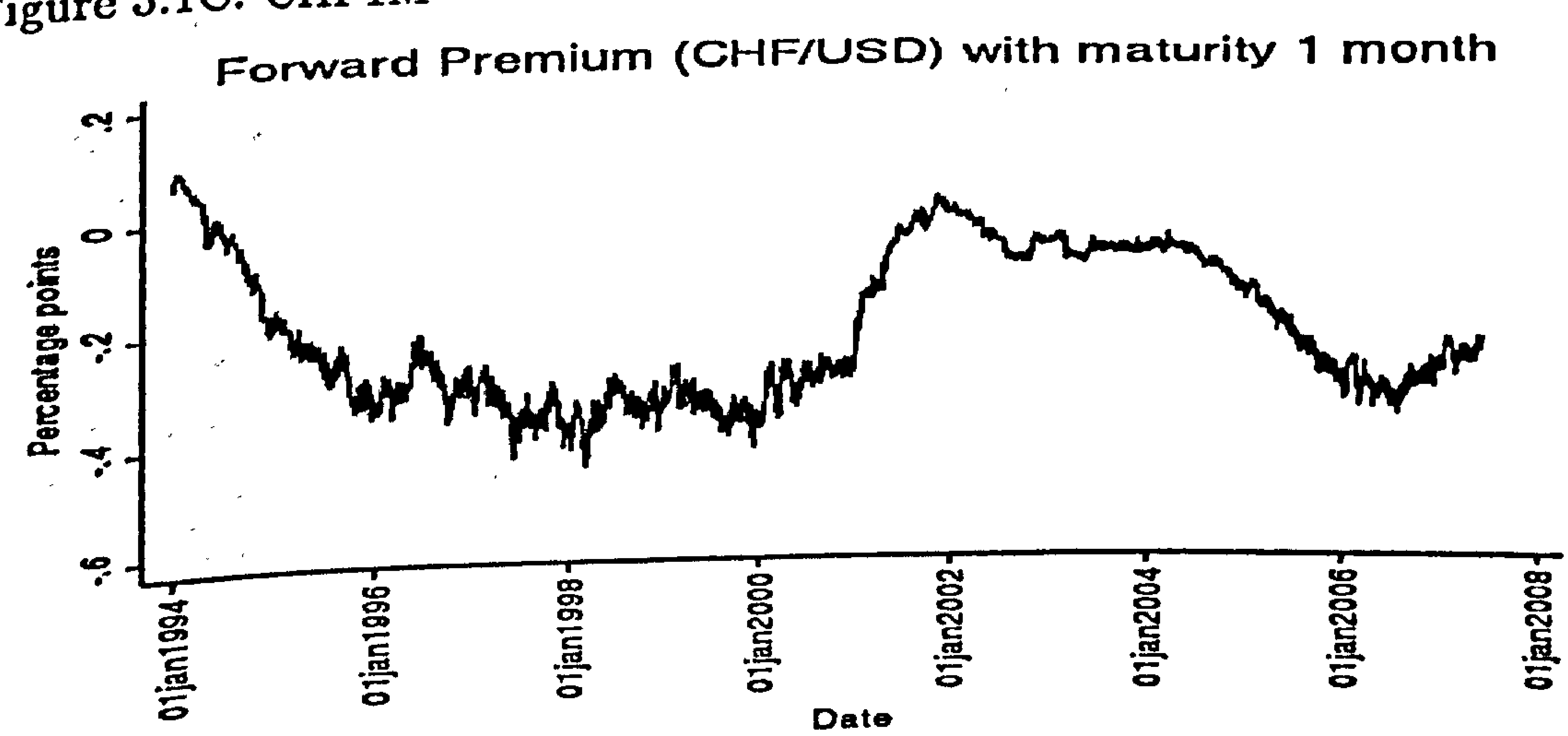


Figure 3.1D: DKK1M

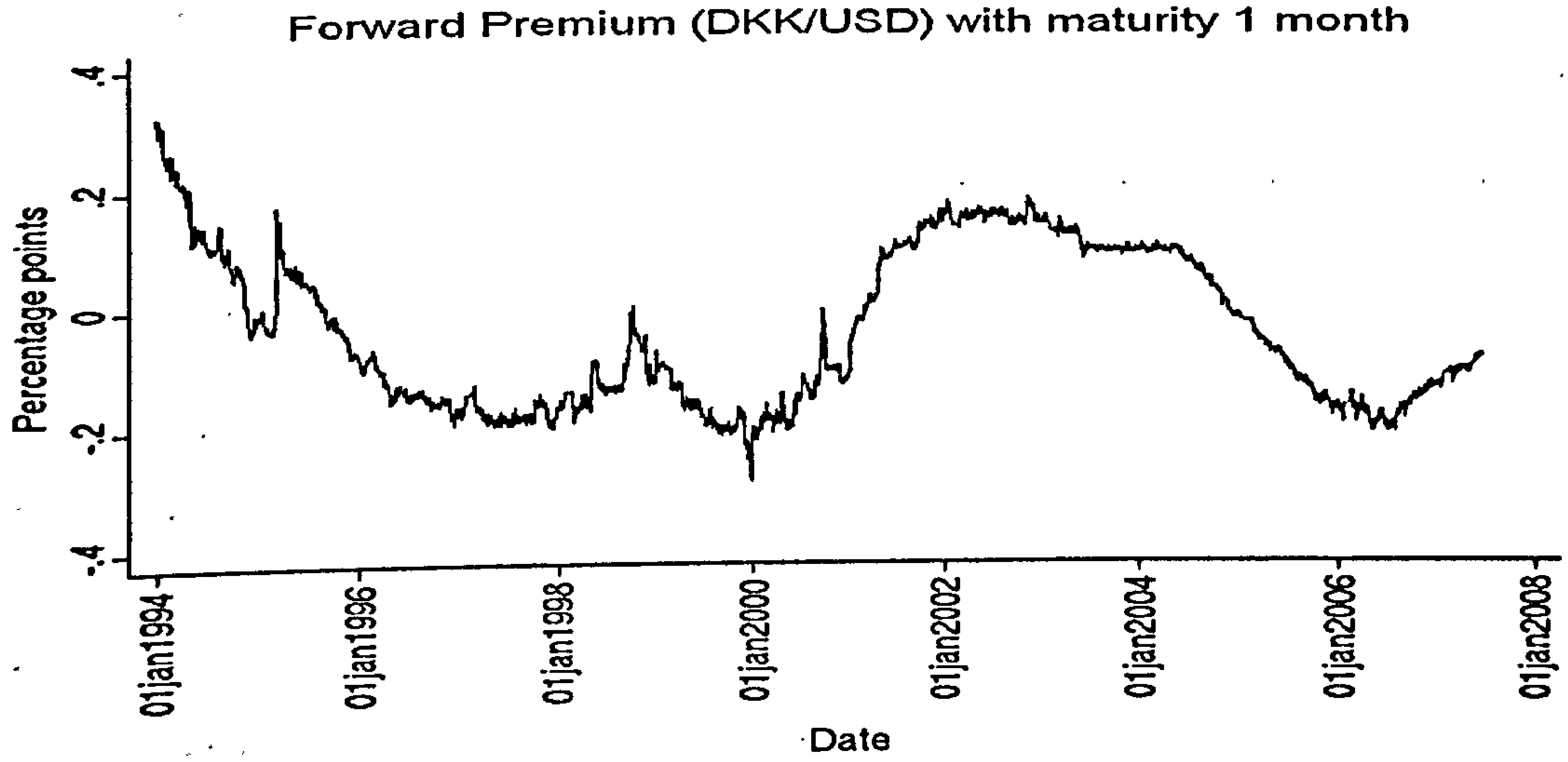


Figure 3.1E: EUR1M

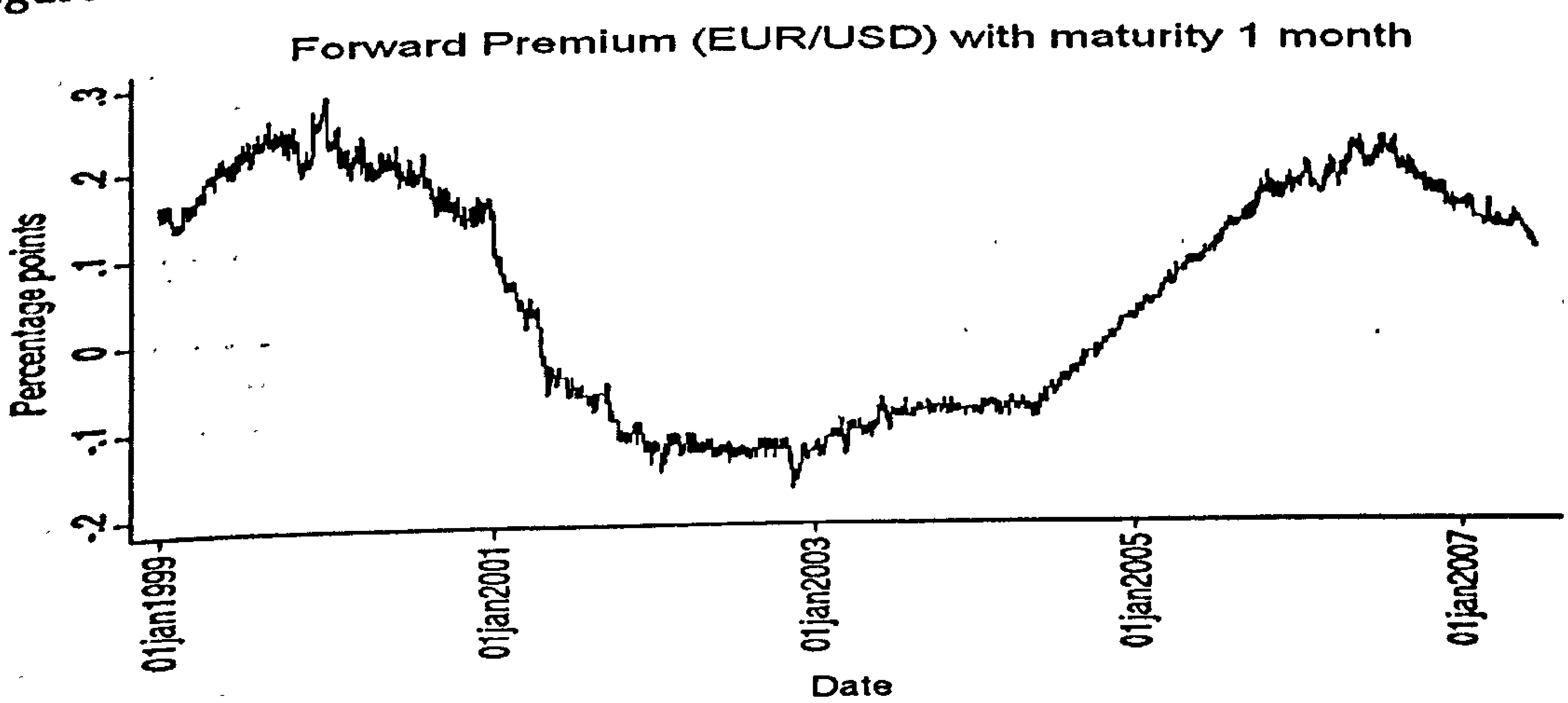


Figure 3.1F: GBP1M

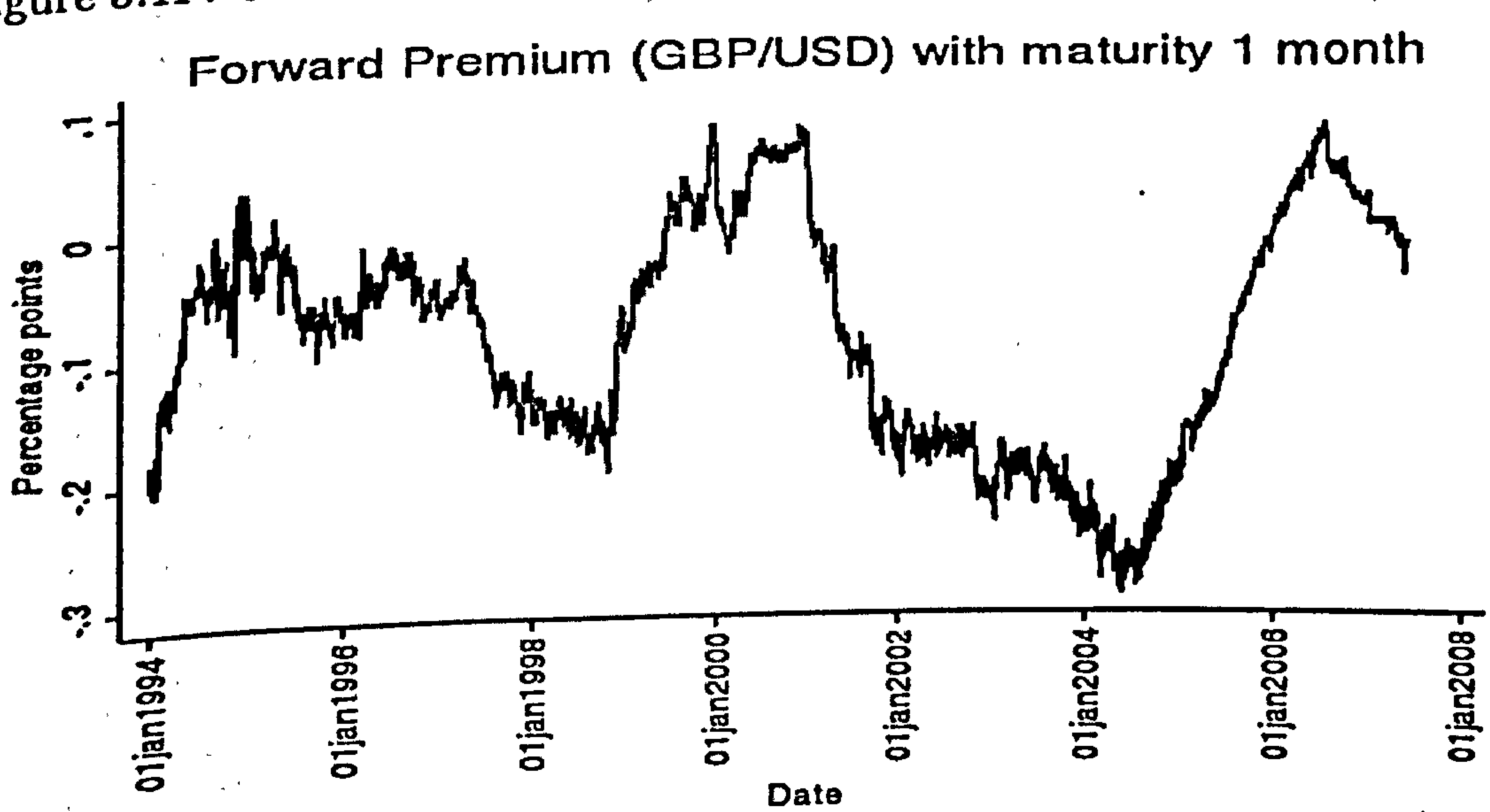


Figure 3.1G: JPY1M

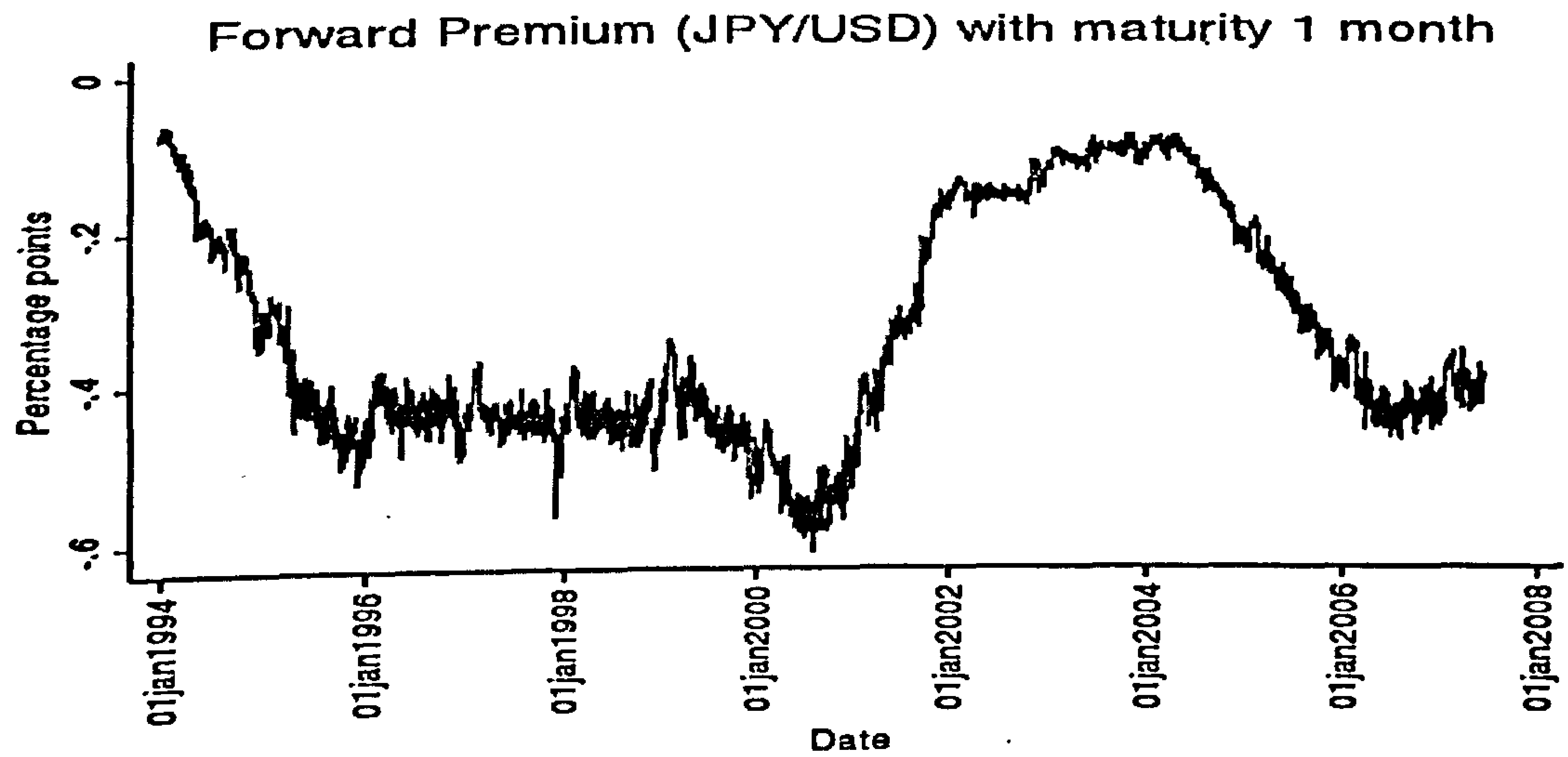


Figure 3.1H: NOK1M

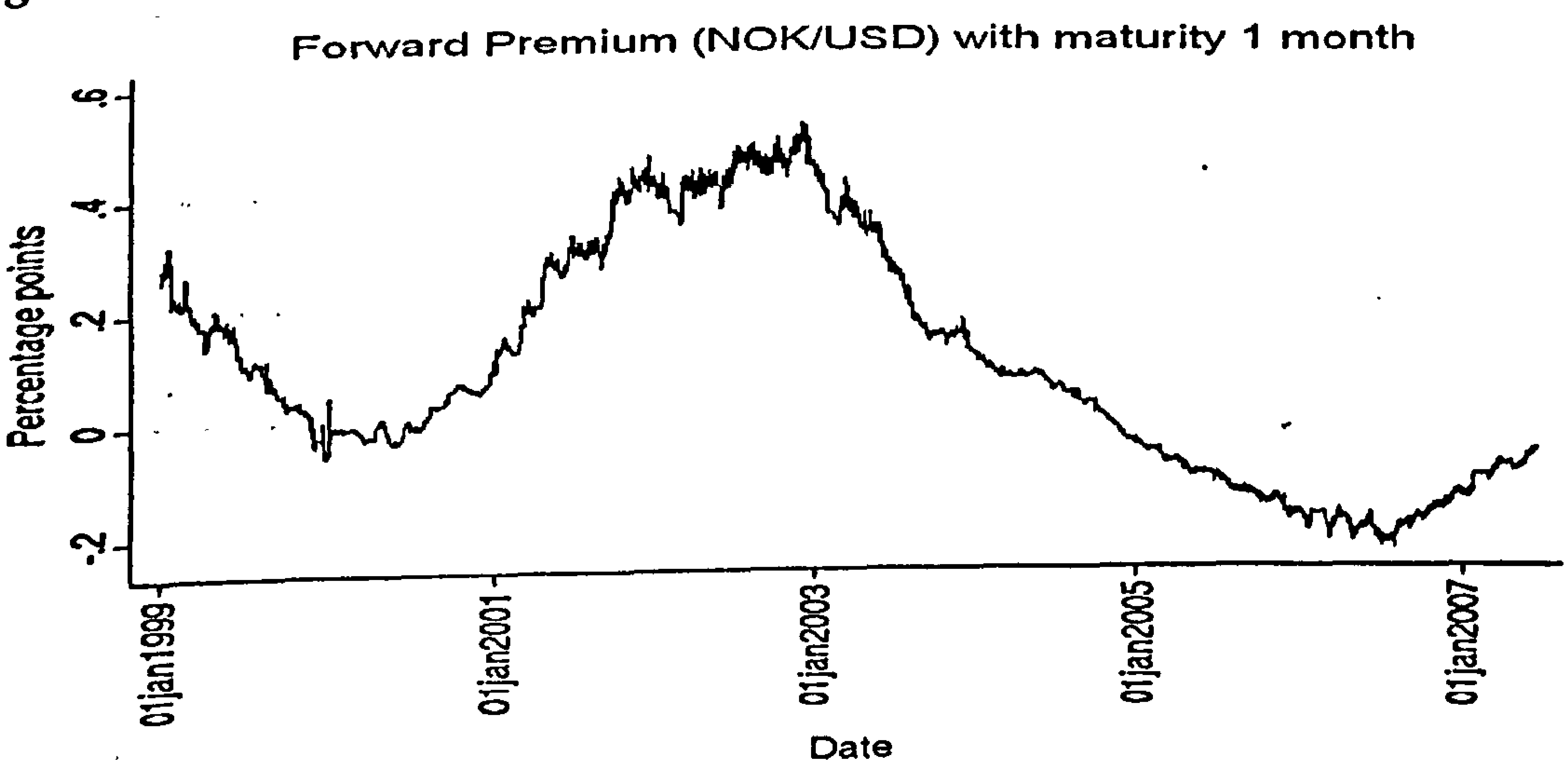


Figure 3.1I: NZD1M

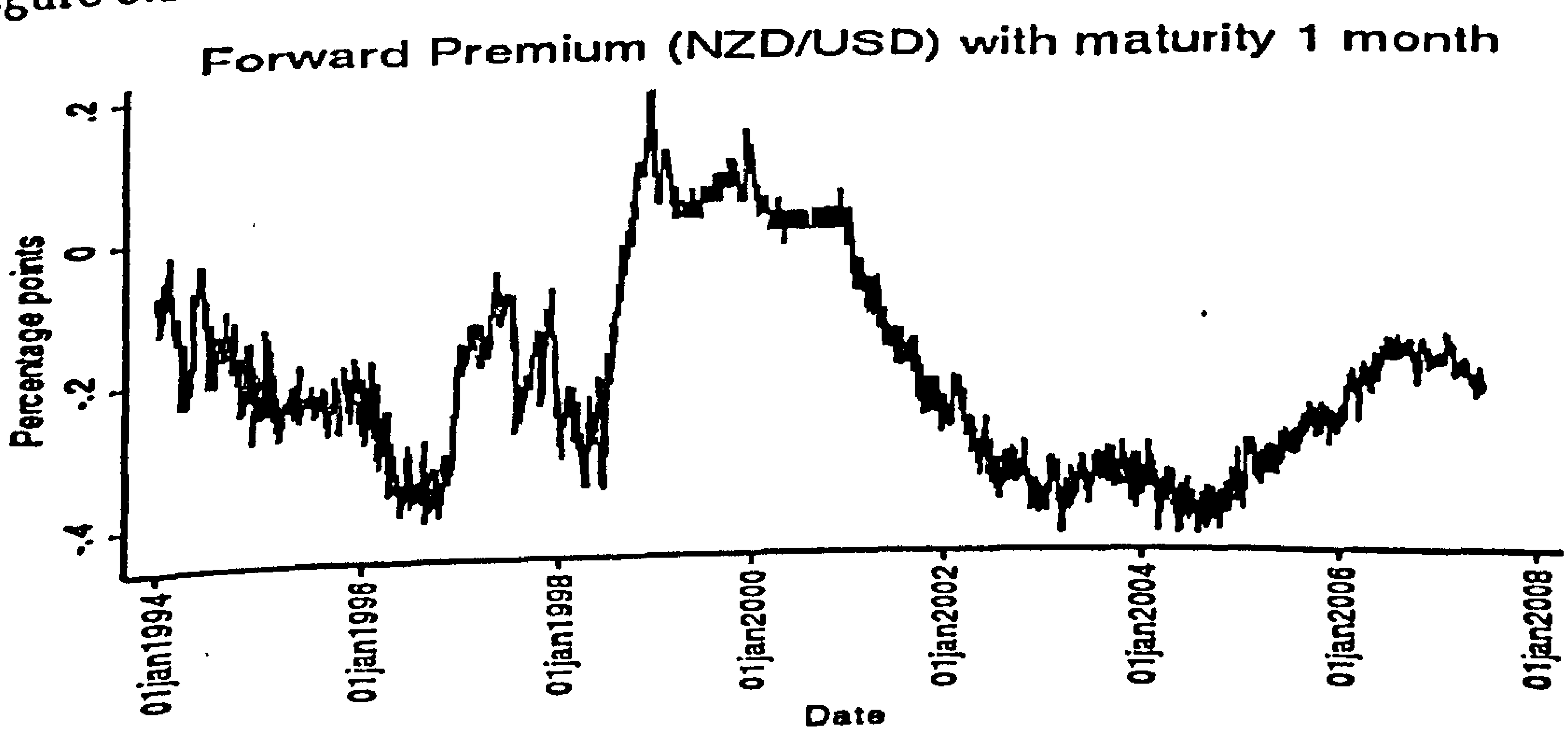


Figure 3.1J: SEK1M

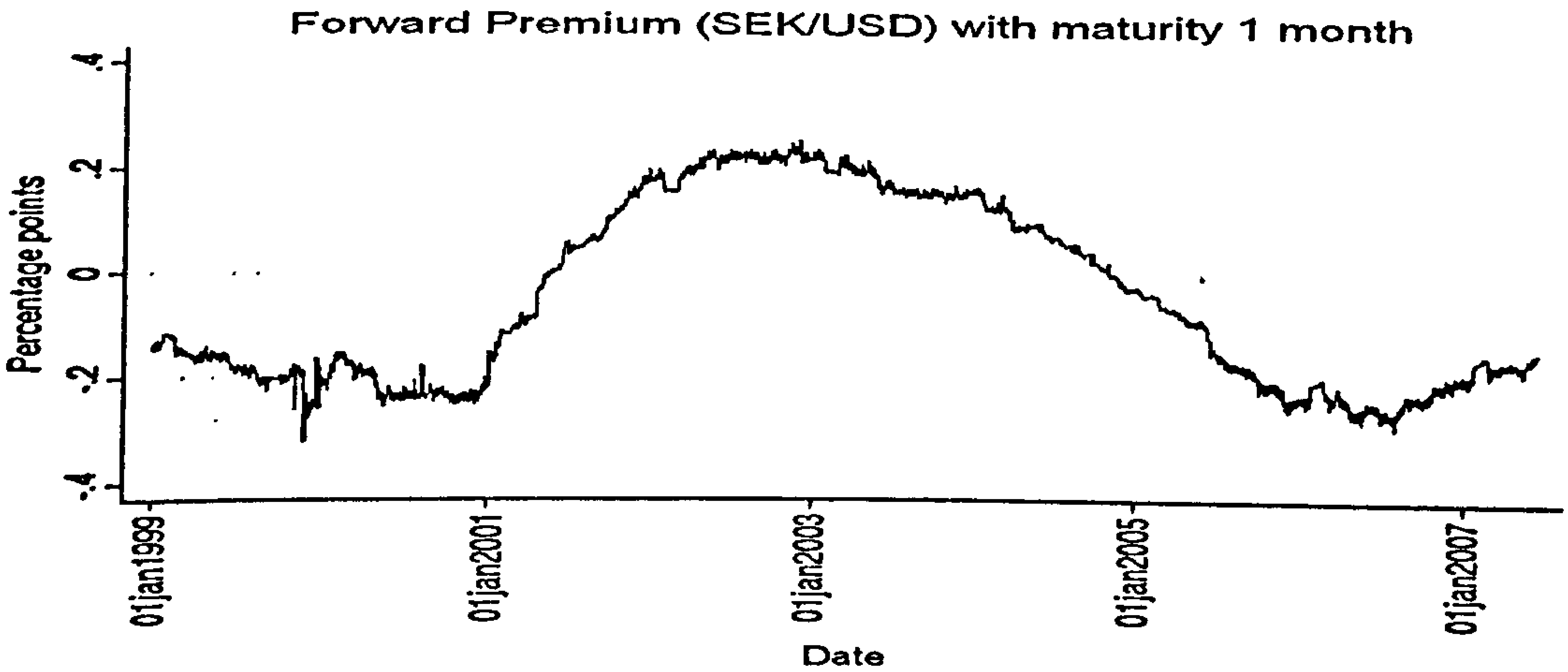


Figure 3.2: Autocorrelations function for the forward premium

Figure 3.2A: AUD

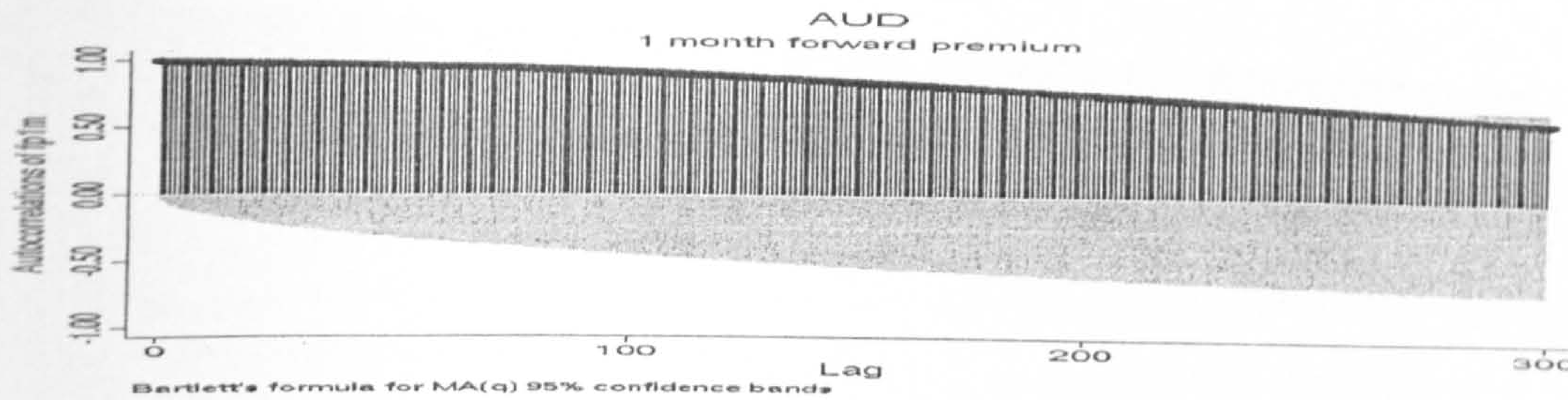


Figure 3.2B: CAD

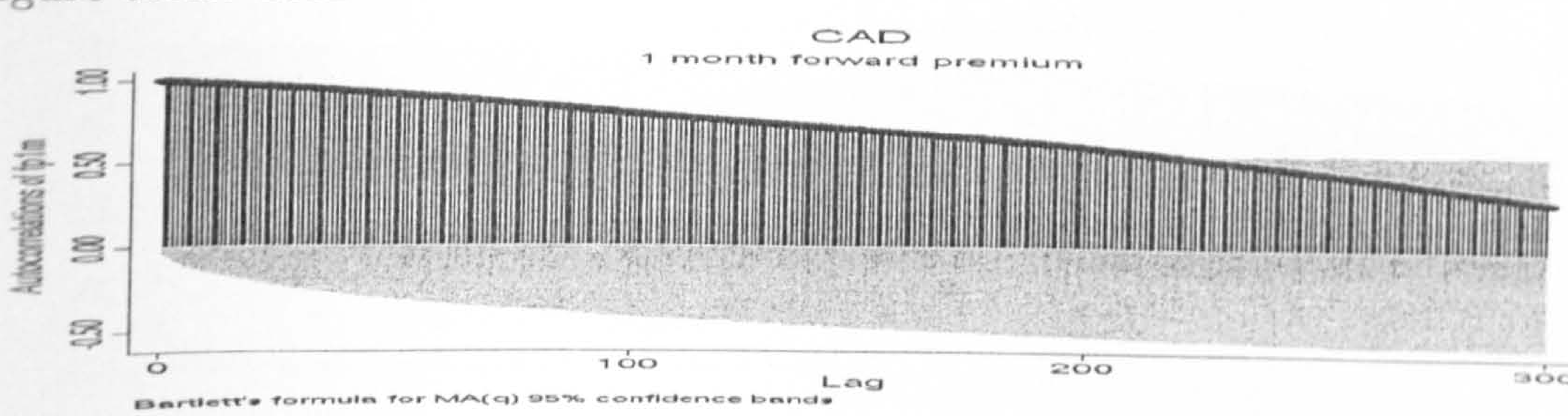


Figure 3.2C: CHF

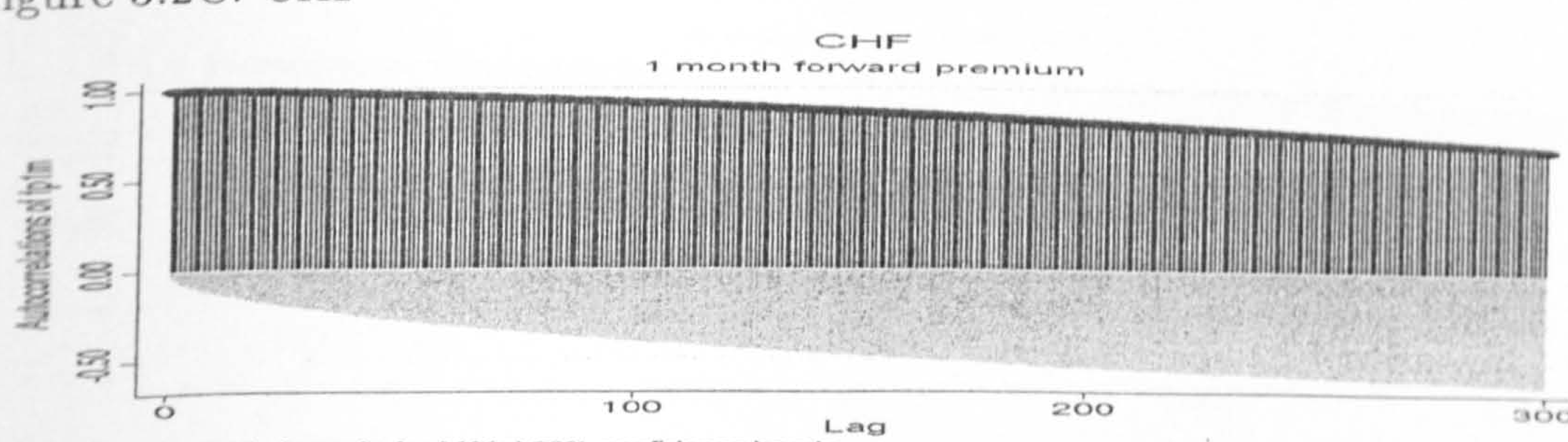


Figure 3.2D: DKK

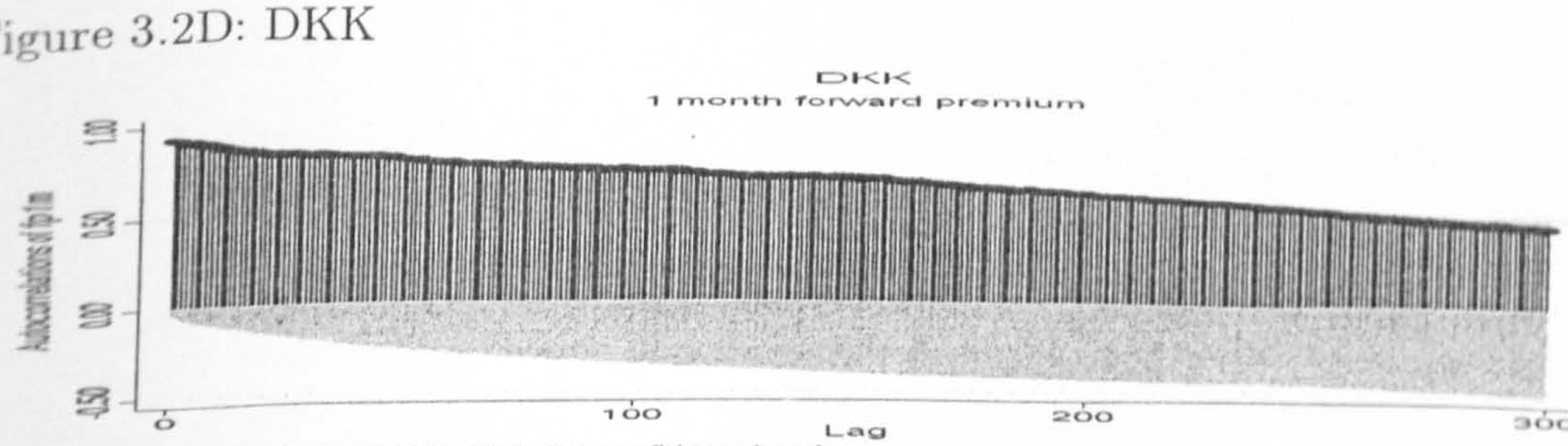


Figure 3.2E: EUR

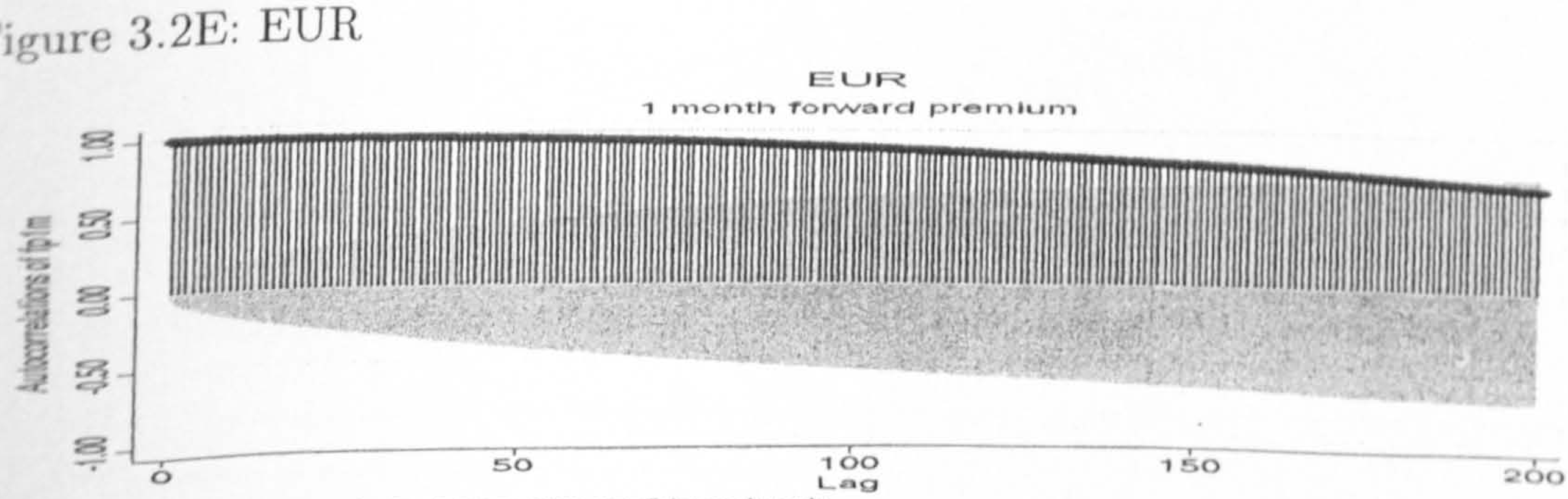


Figure 3.2F: GBP

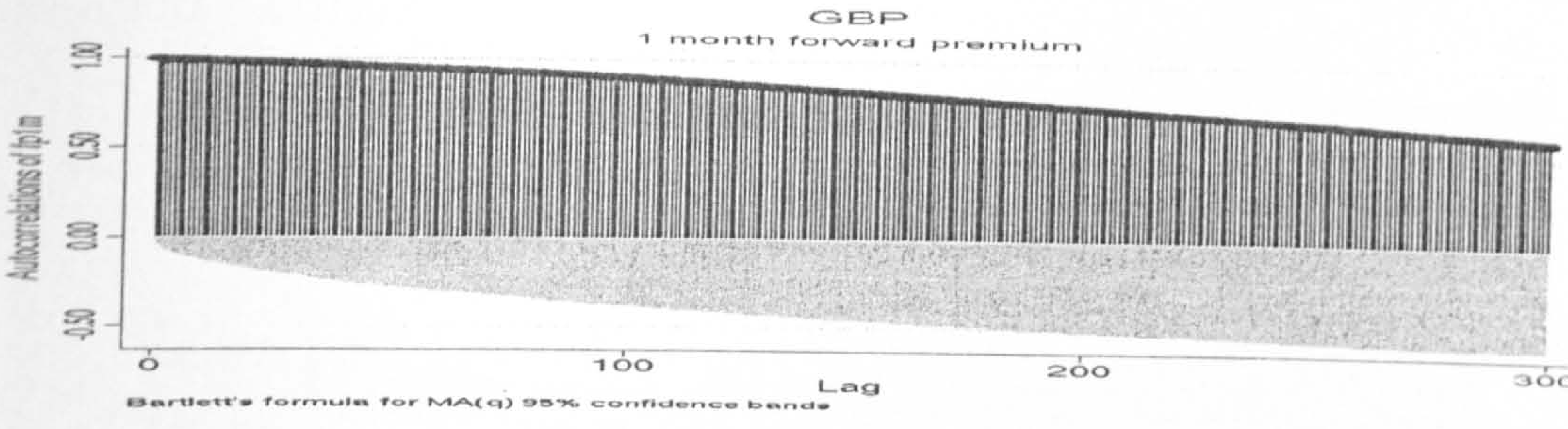


Figure 3.2G: JPY

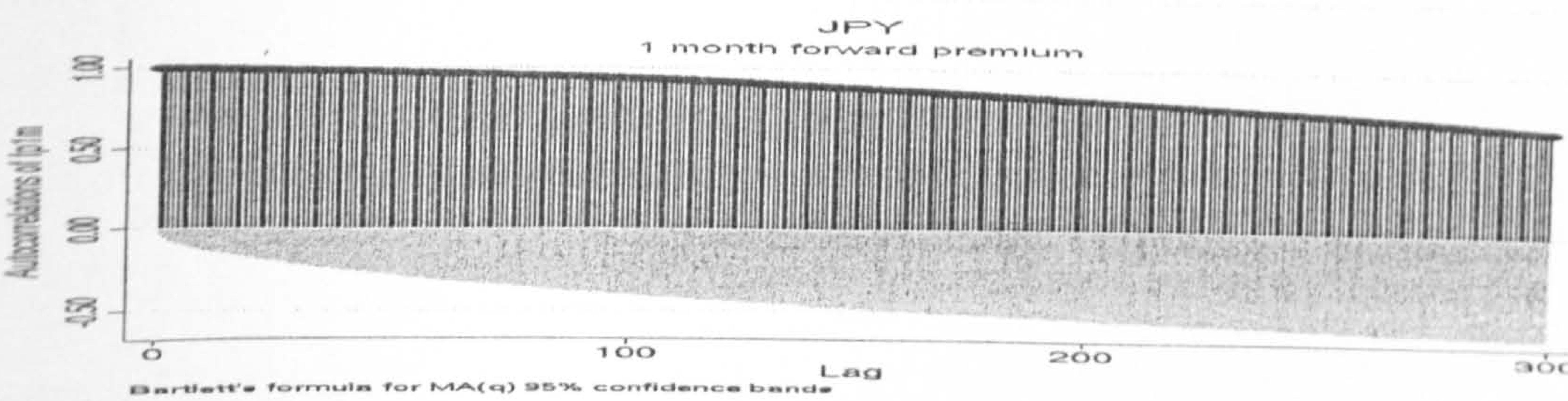


Figure 3.2H: NOK

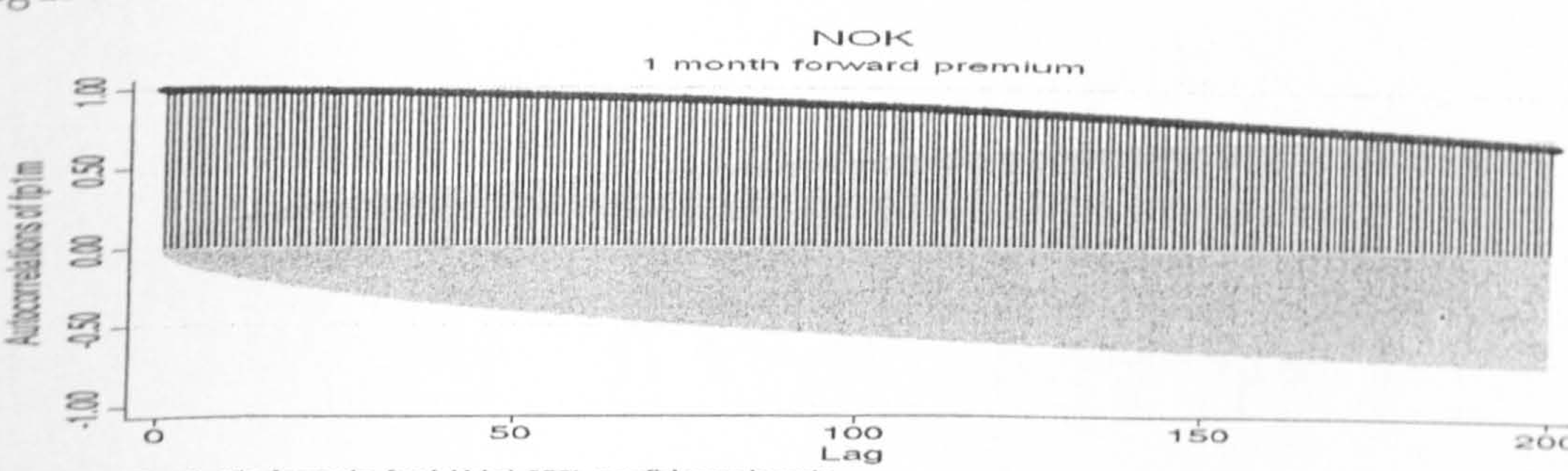


Figure 3.2I: NZD

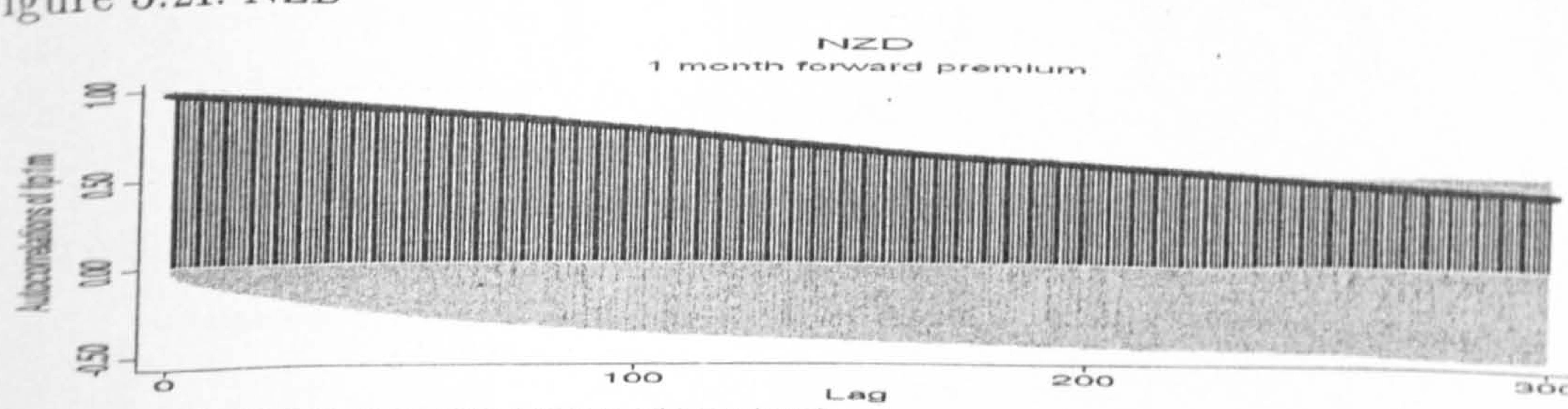


Figure 3.2J: SEK

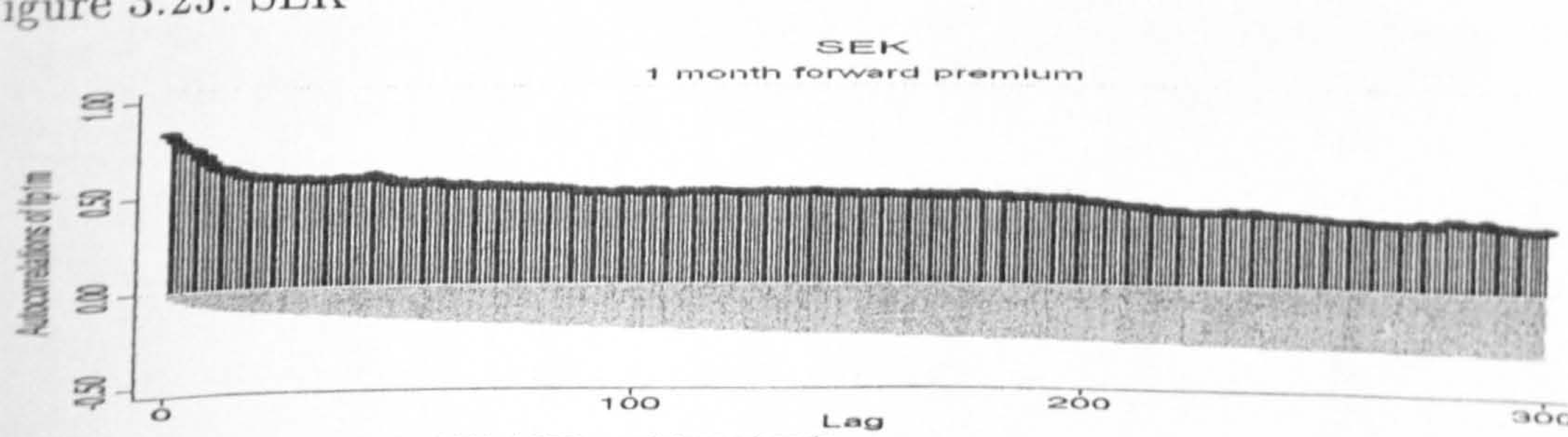


Figure 3.3: The return on the spot rate series filtered by the MA(20) process.
Figure 3.3A: AUD

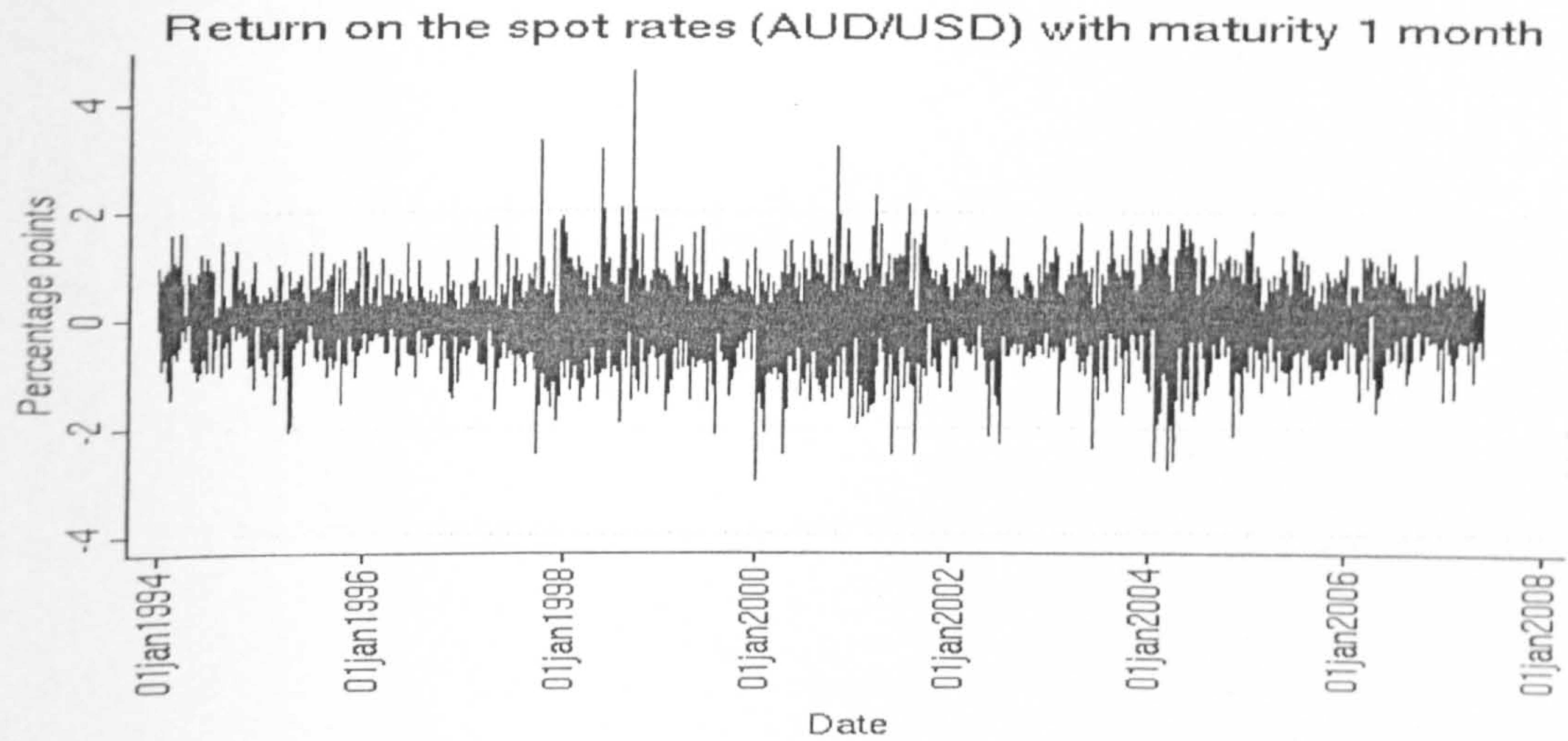


Figure 3.3B: CAD

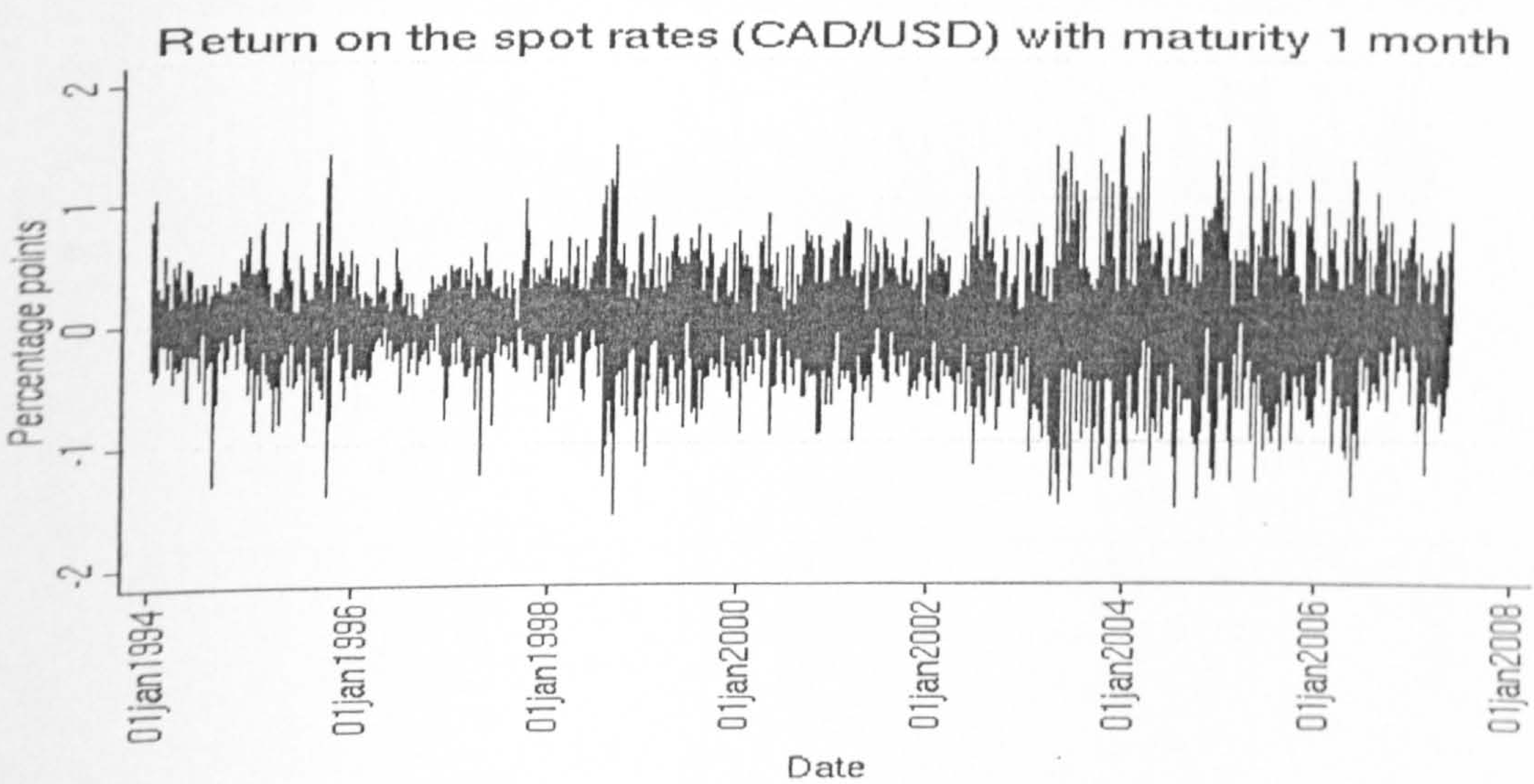


Figure 3.3C: CHF

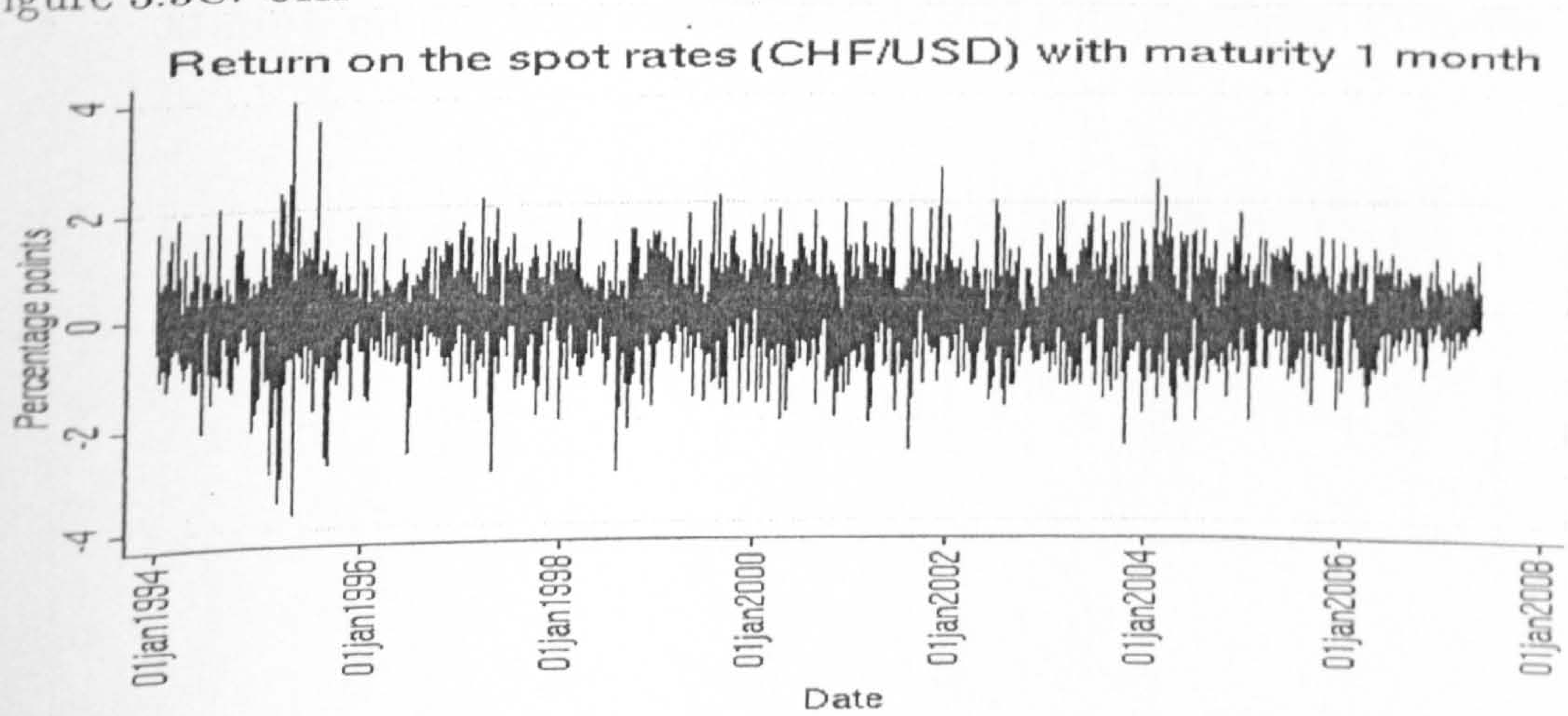


Figure 3.3D: DKK

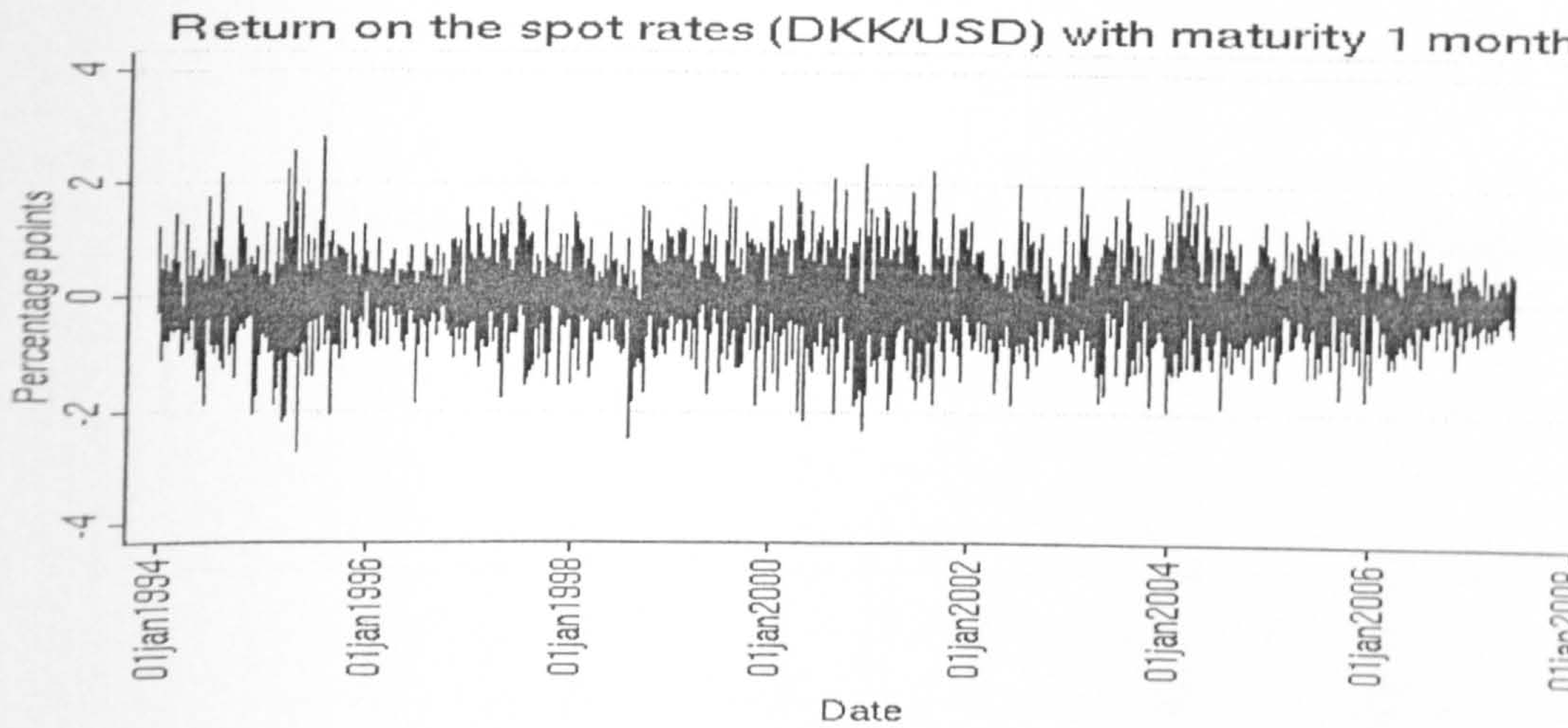


Figure 3.3E: EUR

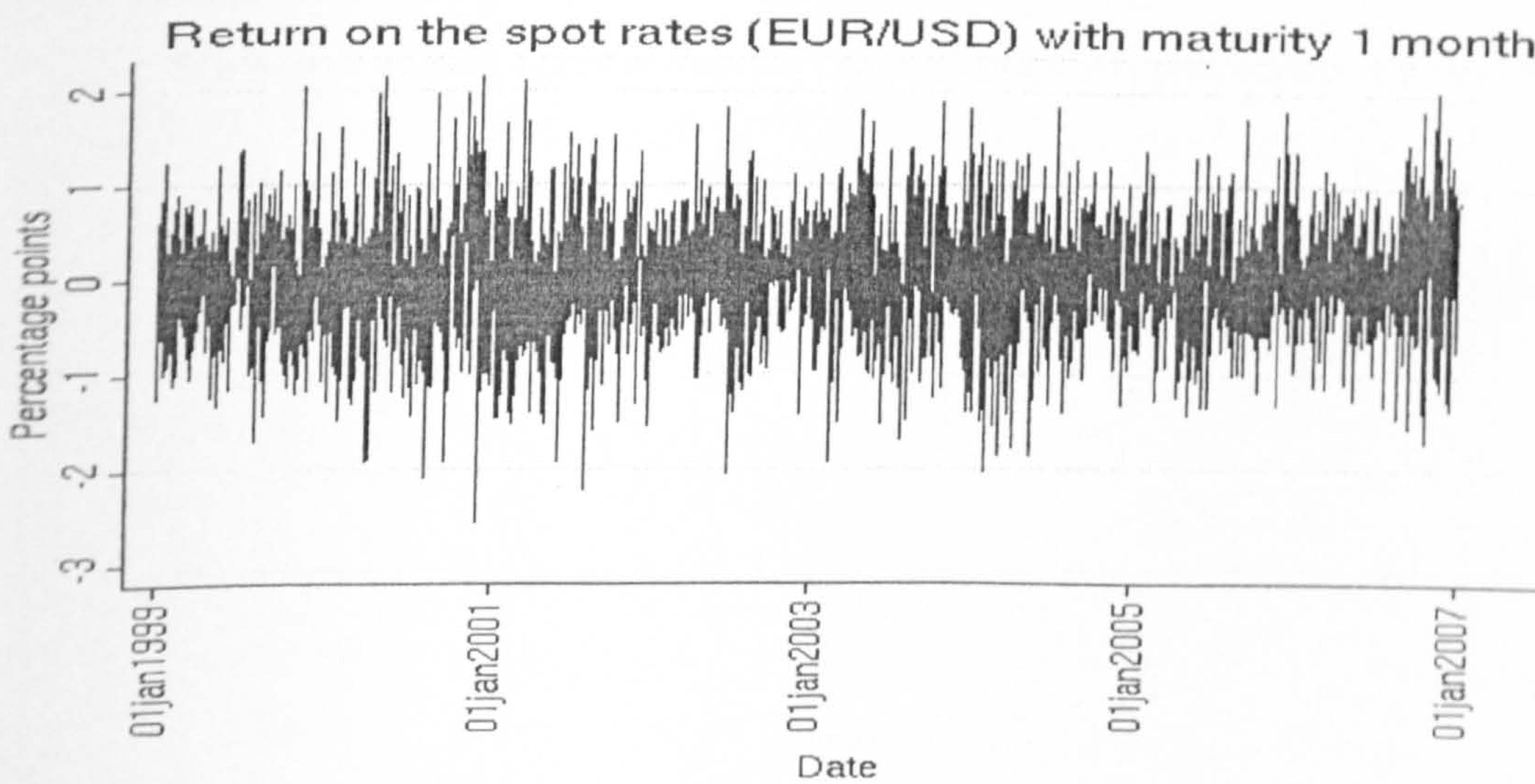


Figure 3.3F: GBP

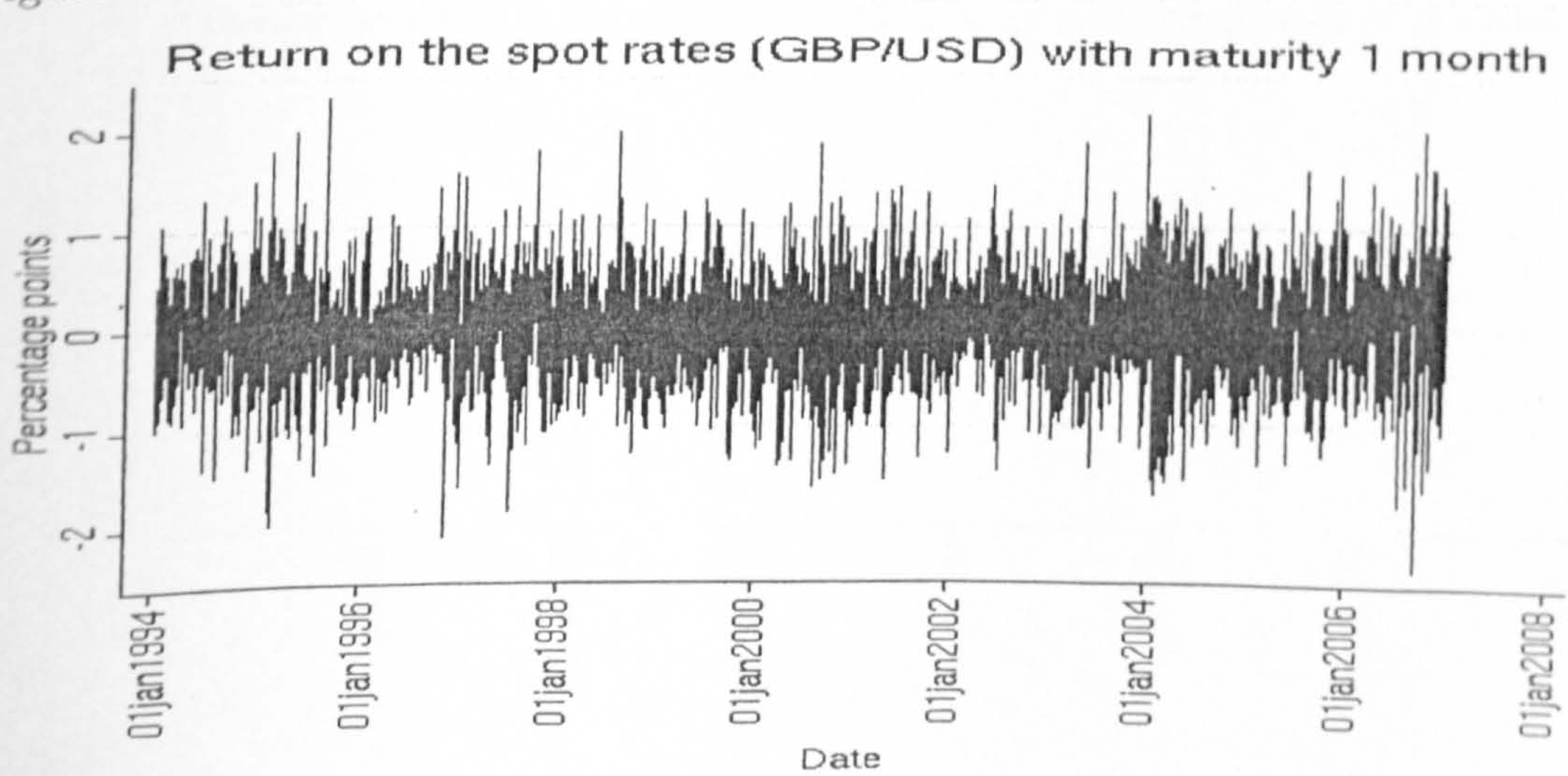


Figure 3.3G: JPY

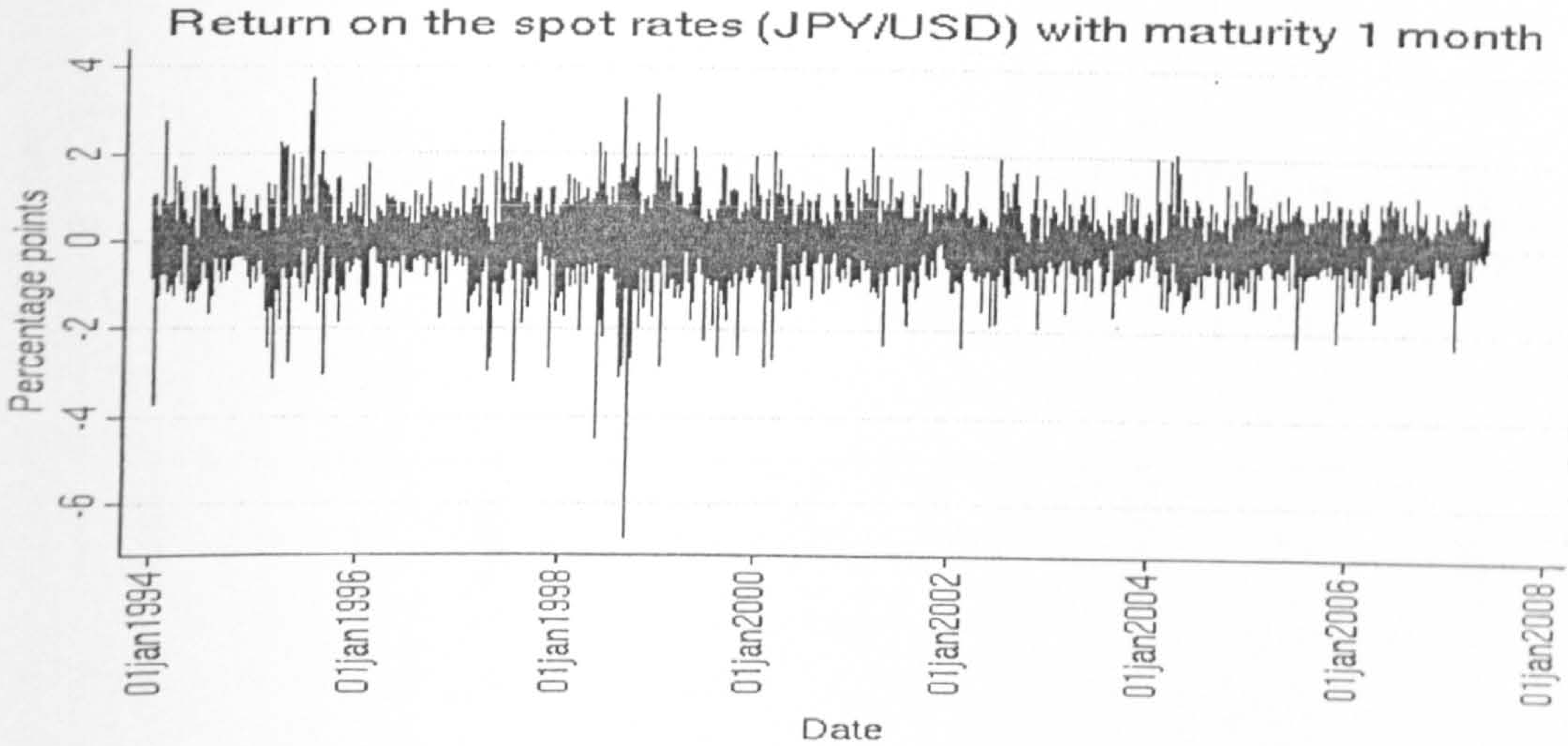


Figure 3.3H: NOK

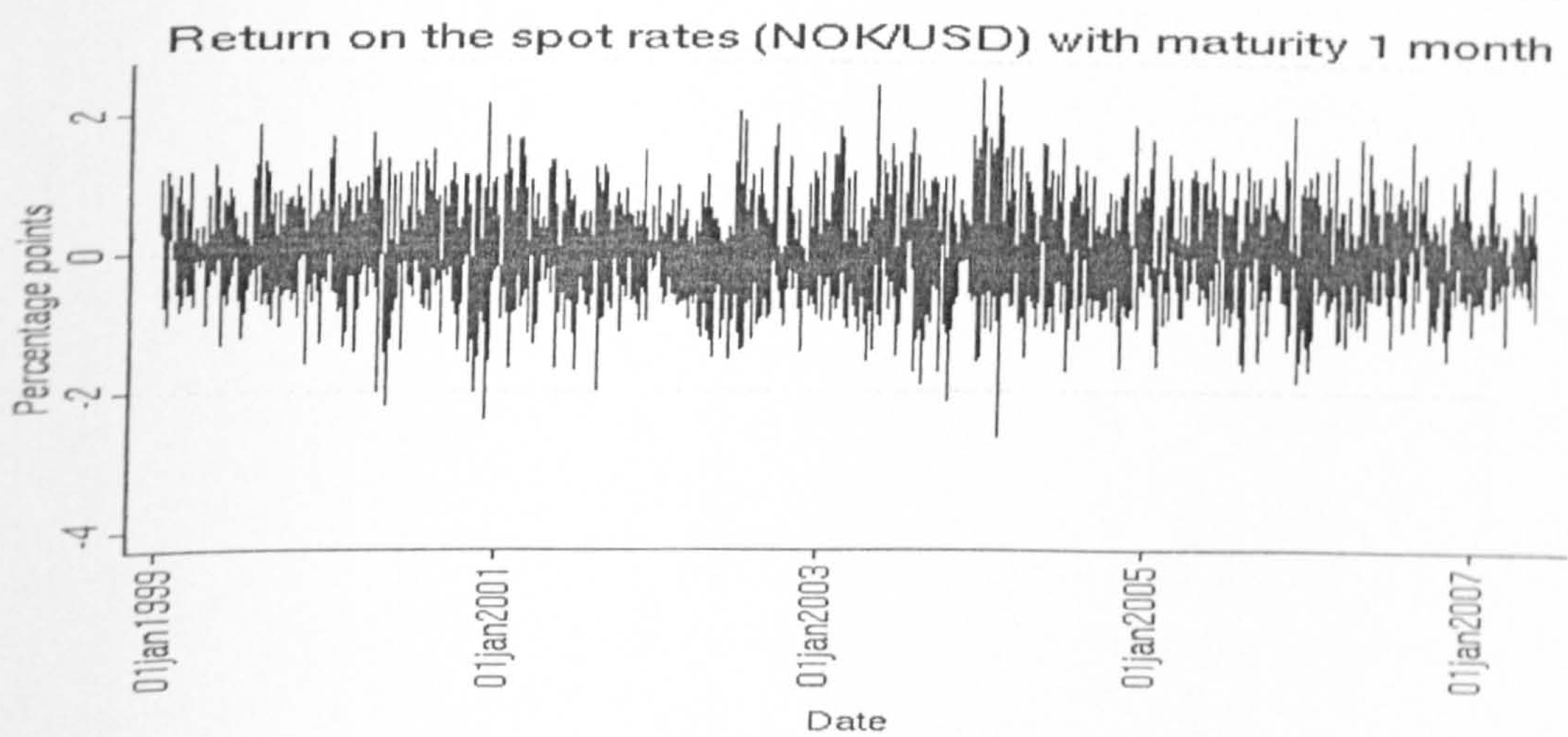


Figure 3.3I: NZD

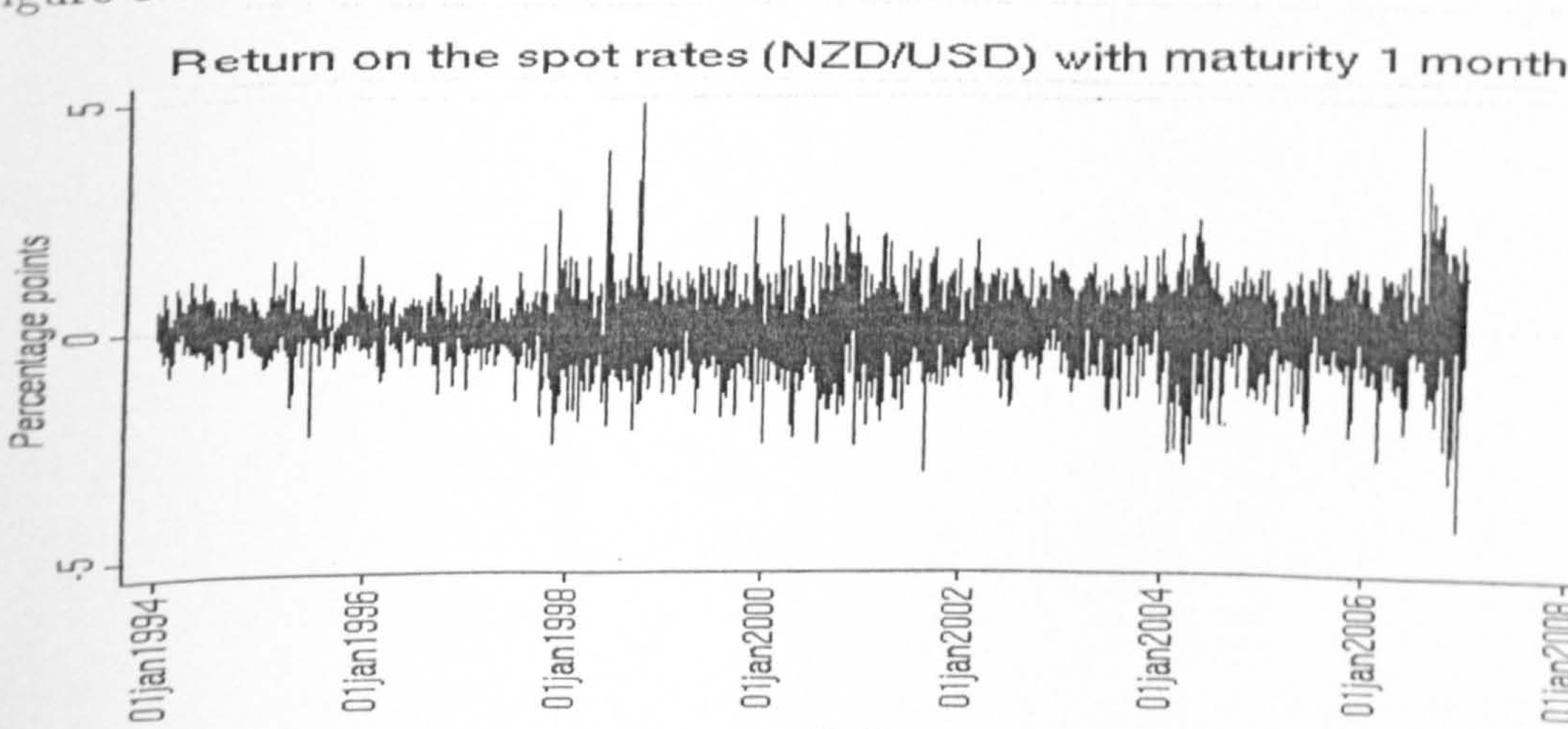


Figure 3.3J: SEK

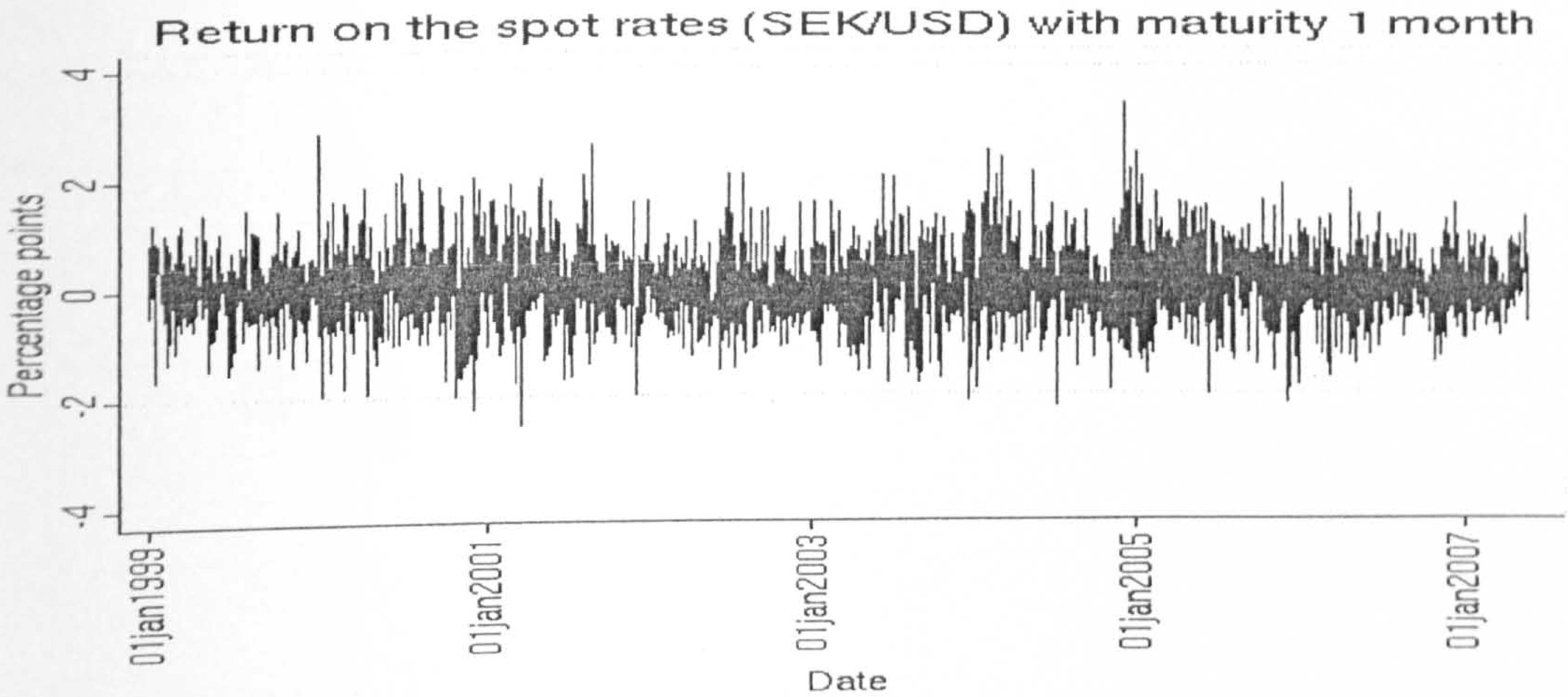


Figure 3.4: ACF of 1 month return on the spot rates filtered by MA(20) process

Figure 3.4A: AUD

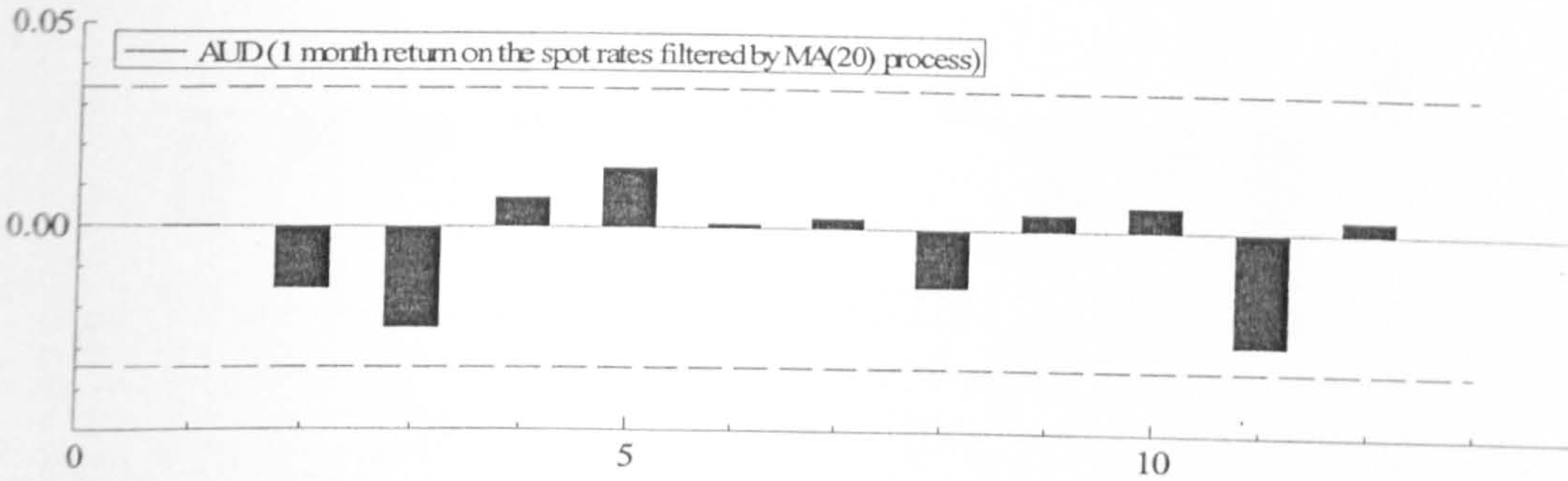


Figure 3.4B: CAD

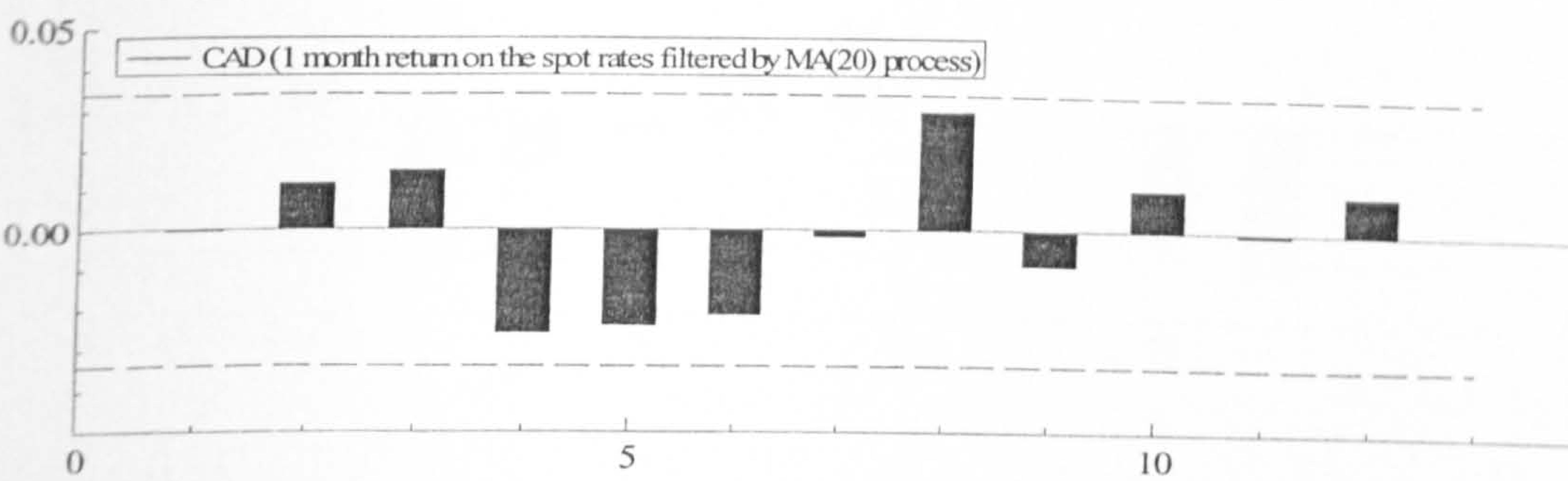


Figure 3.4C: CHF

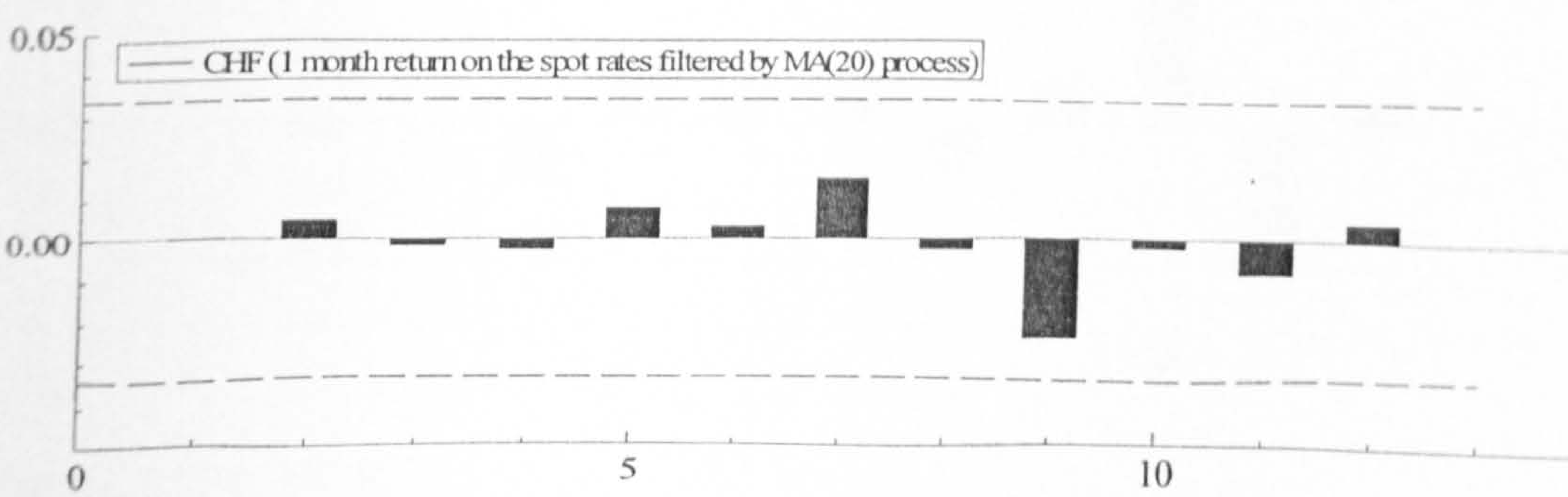


Figure 3.4D: DKK

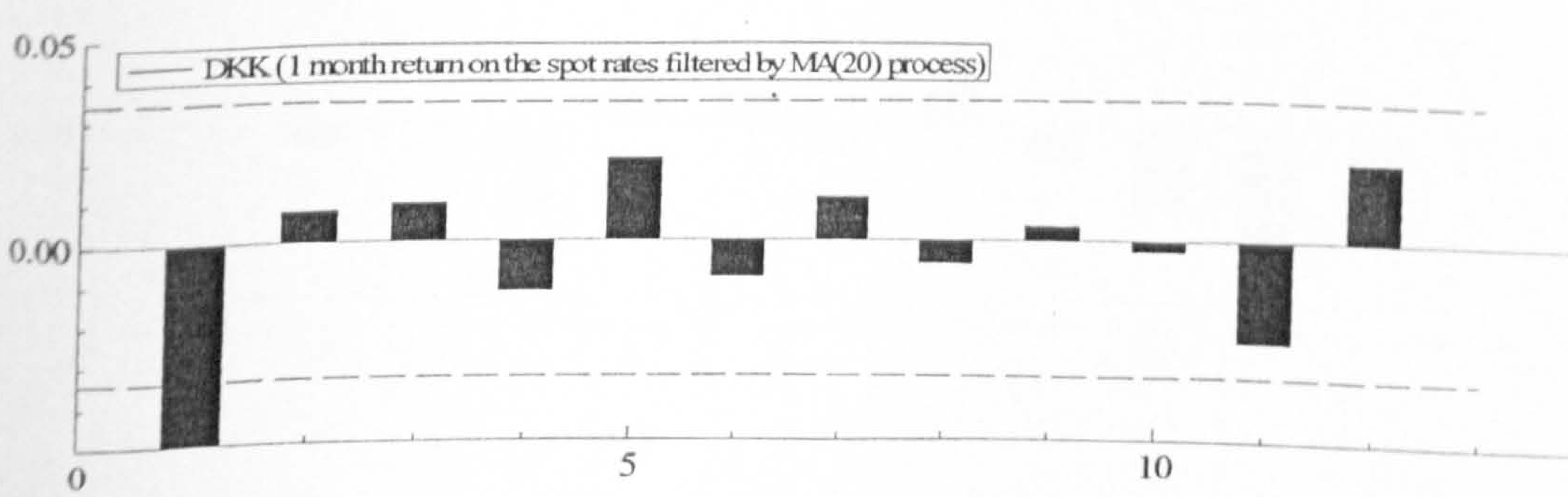


Figure 3.4E: EUR

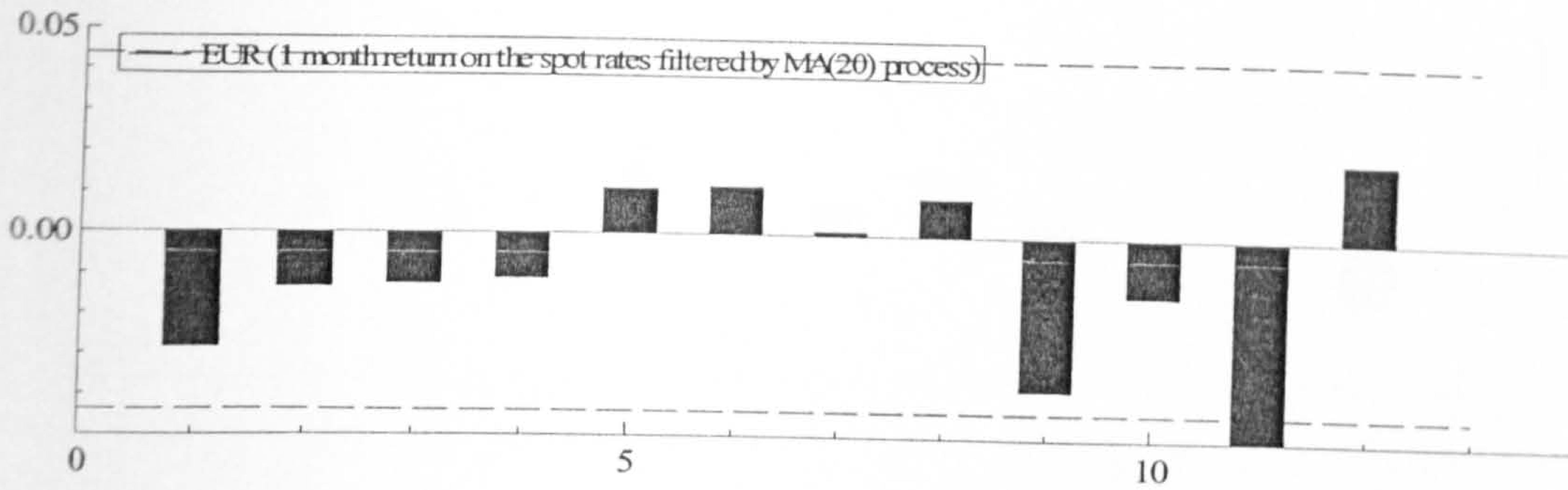


Figure 3.4F: GBP

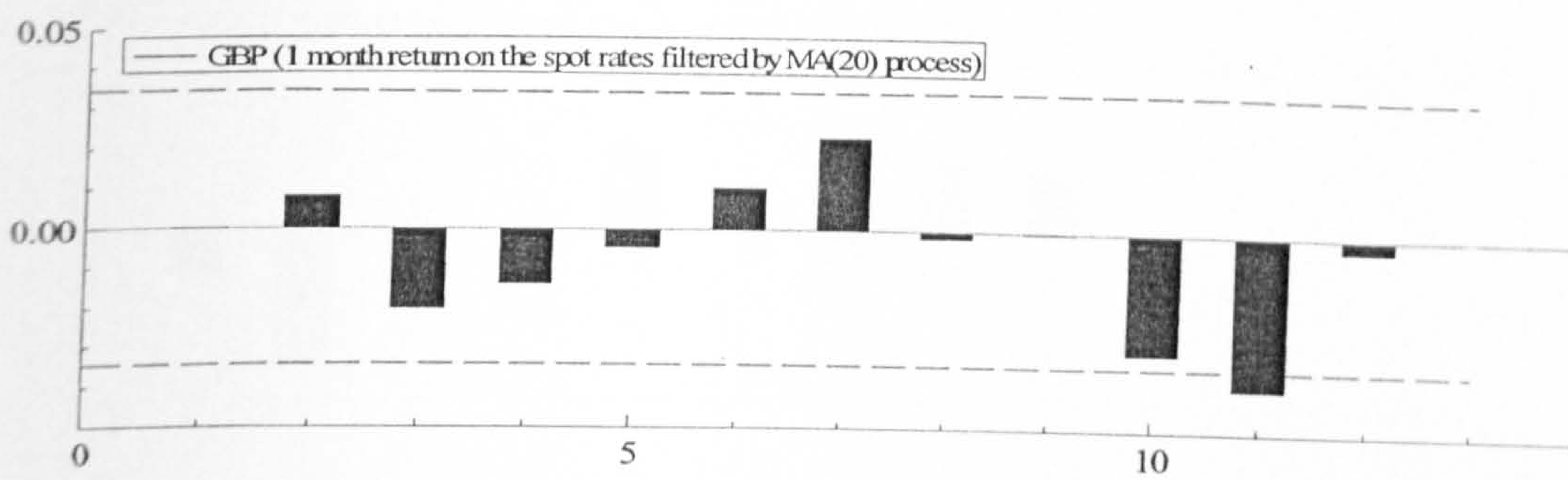


Figure 3.4G: JPY

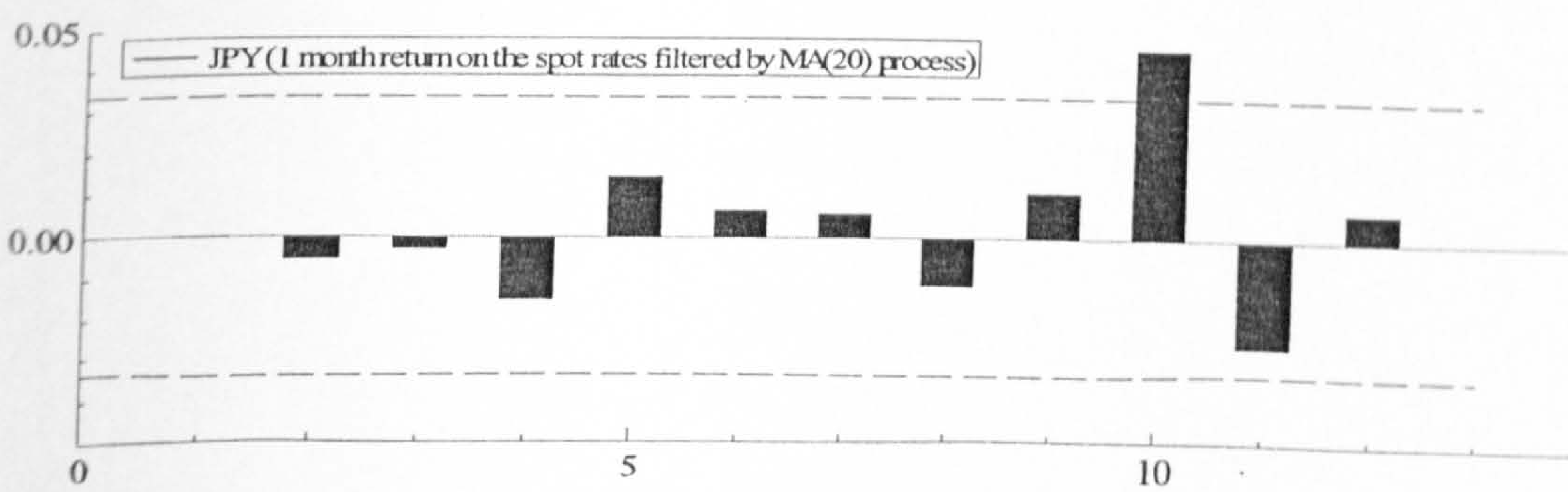


Figure 3.4H: NOK

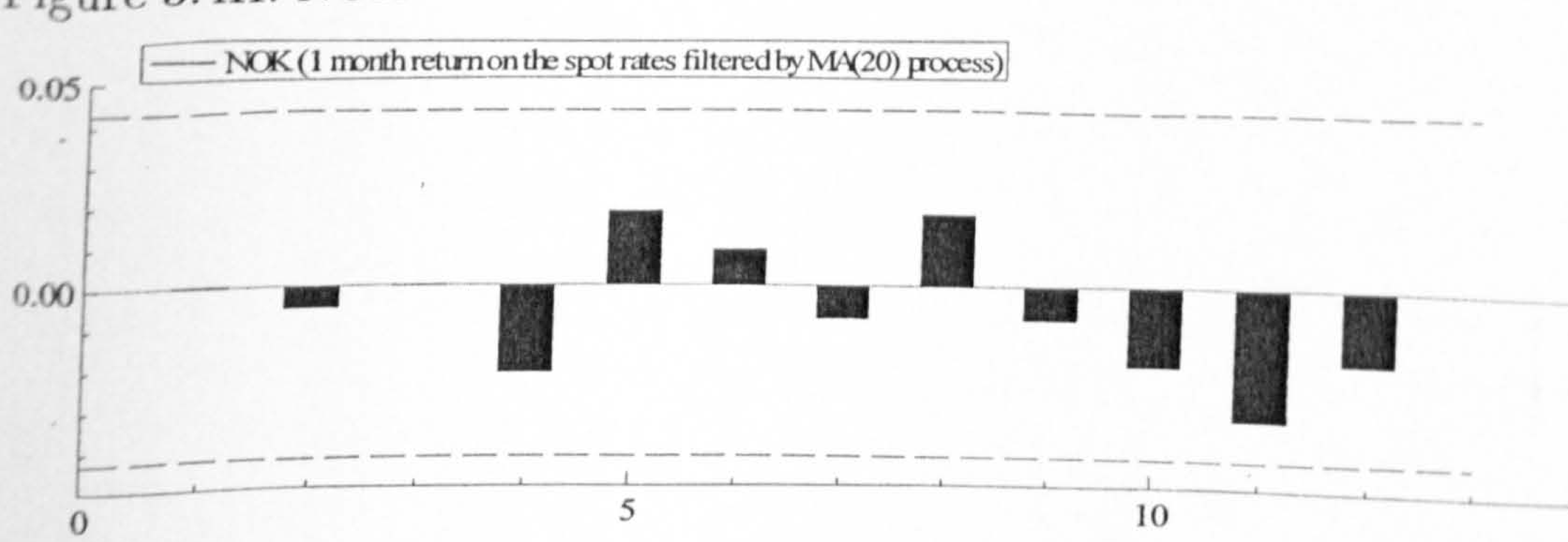


Figure 3.4I: NZD

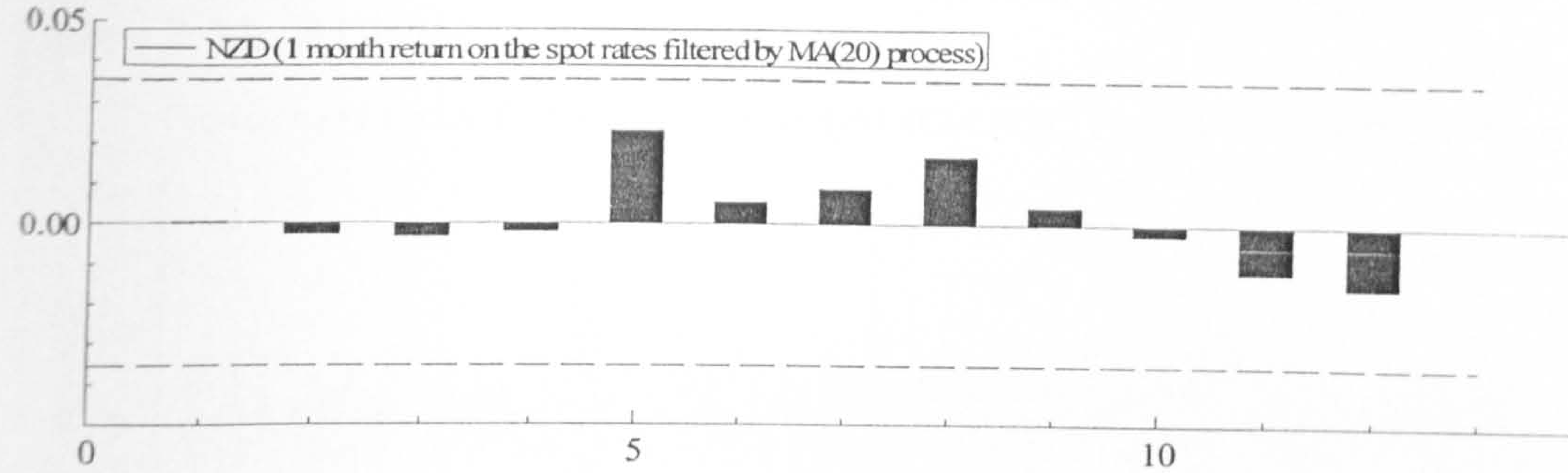


Figure 3.4J: SEK

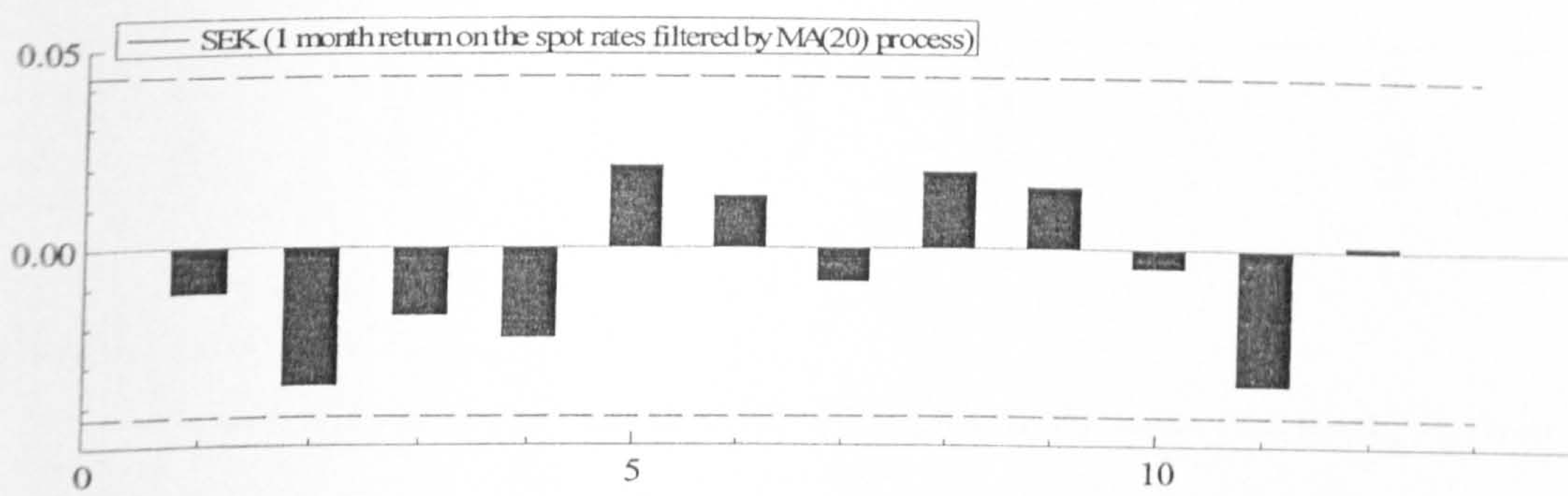


Figure 3.5: The forward rate forecast error filtered by MA(20) process.

Figure 3.5A: AUD1M

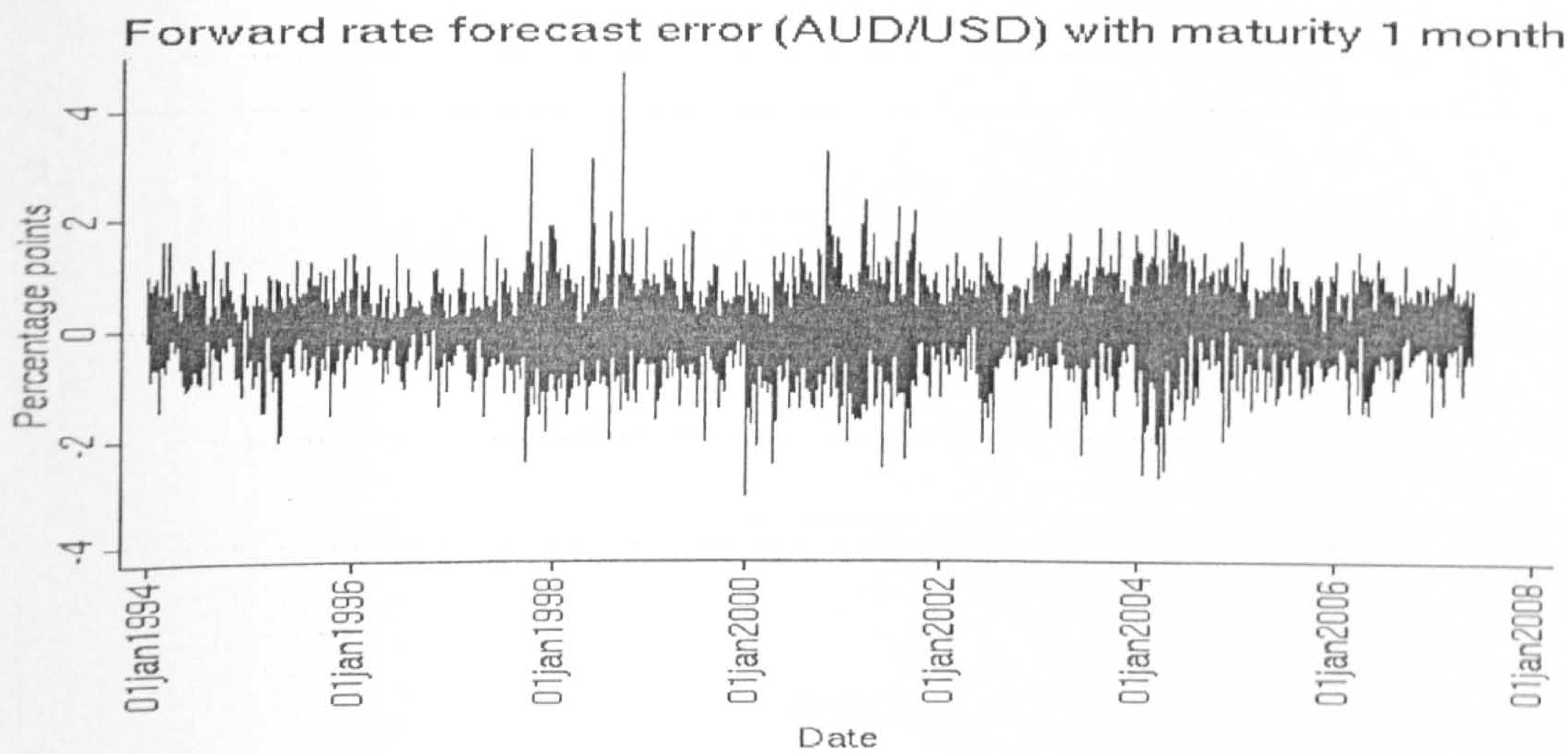


Figure 3.5B: CAD1M

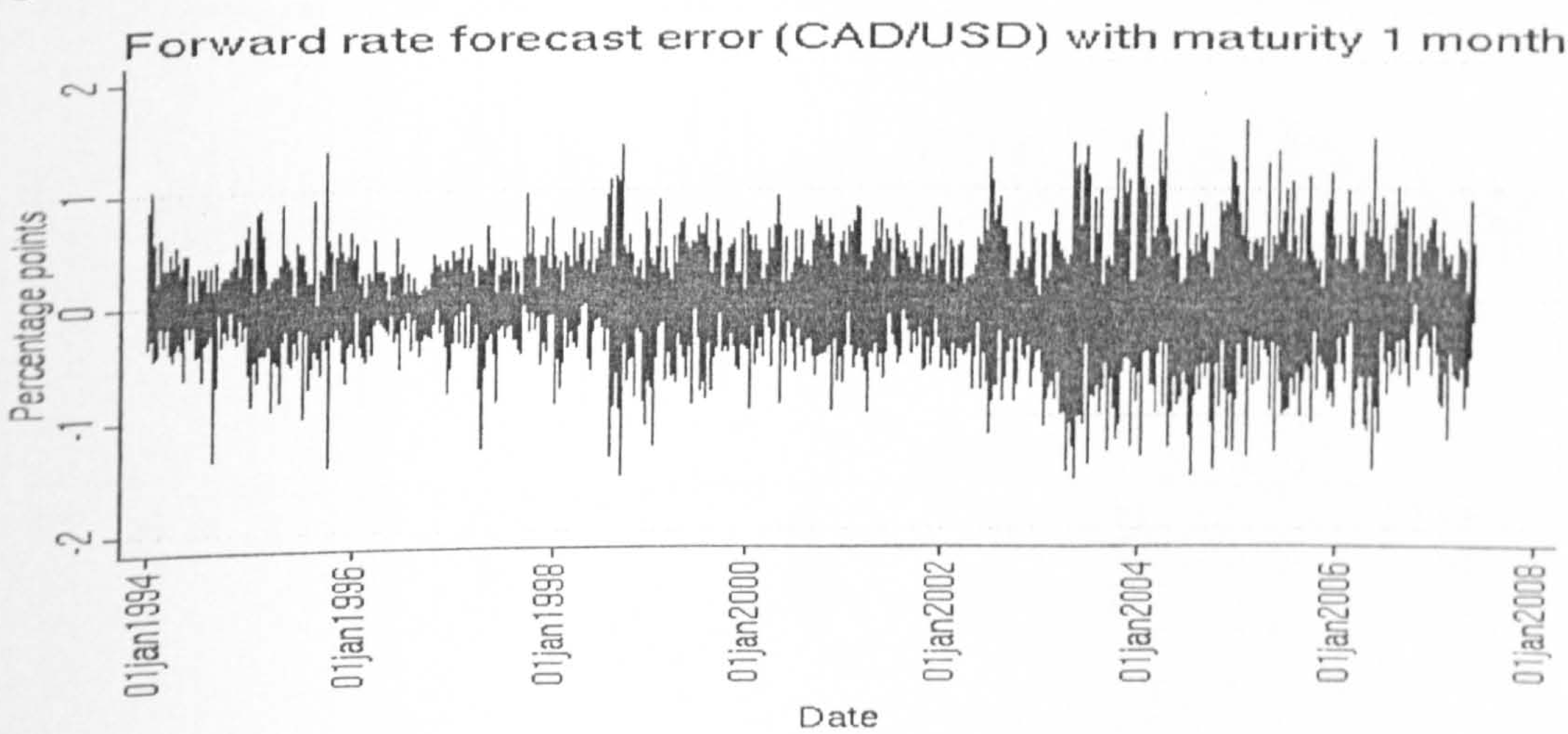


Figure 3.5C: CHF1M

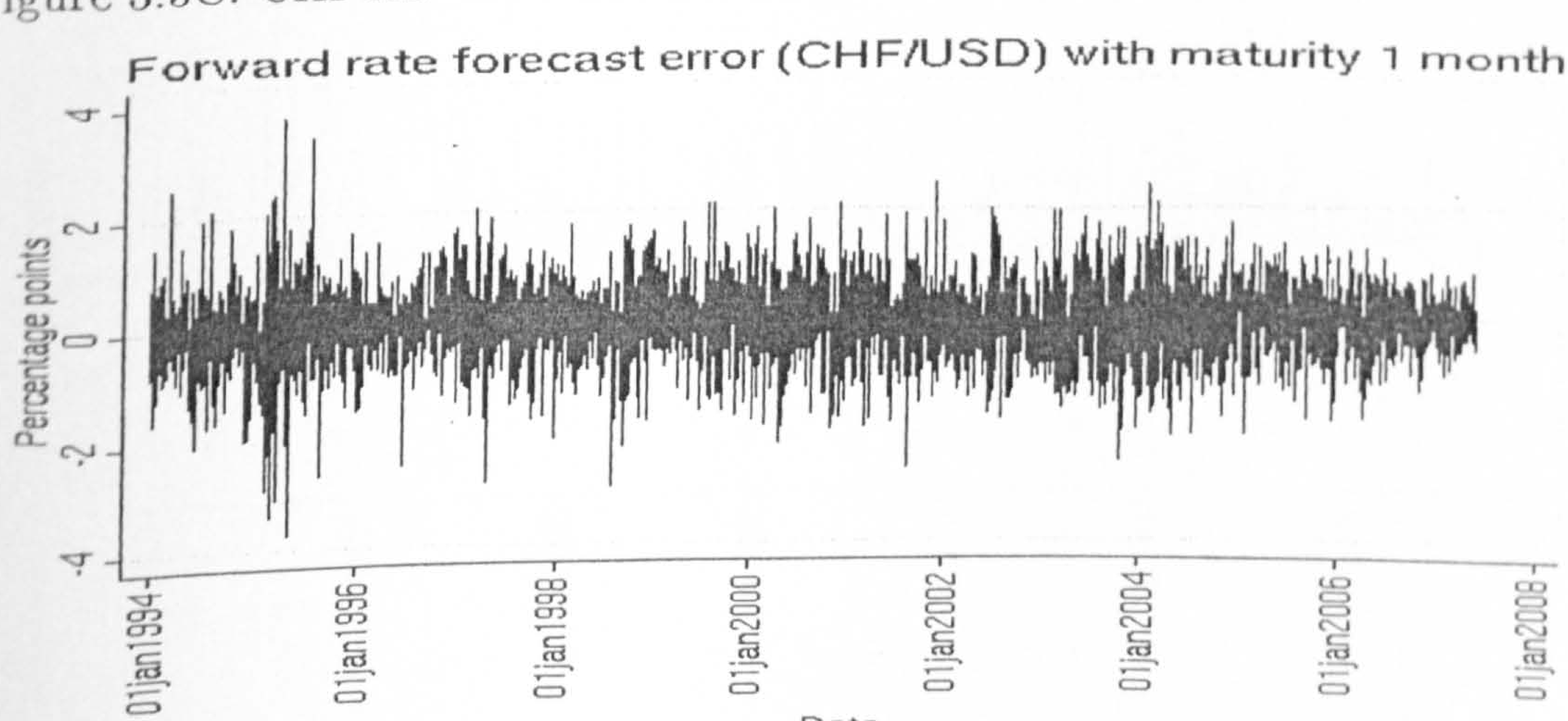


Figure 3.5D: DKK1M

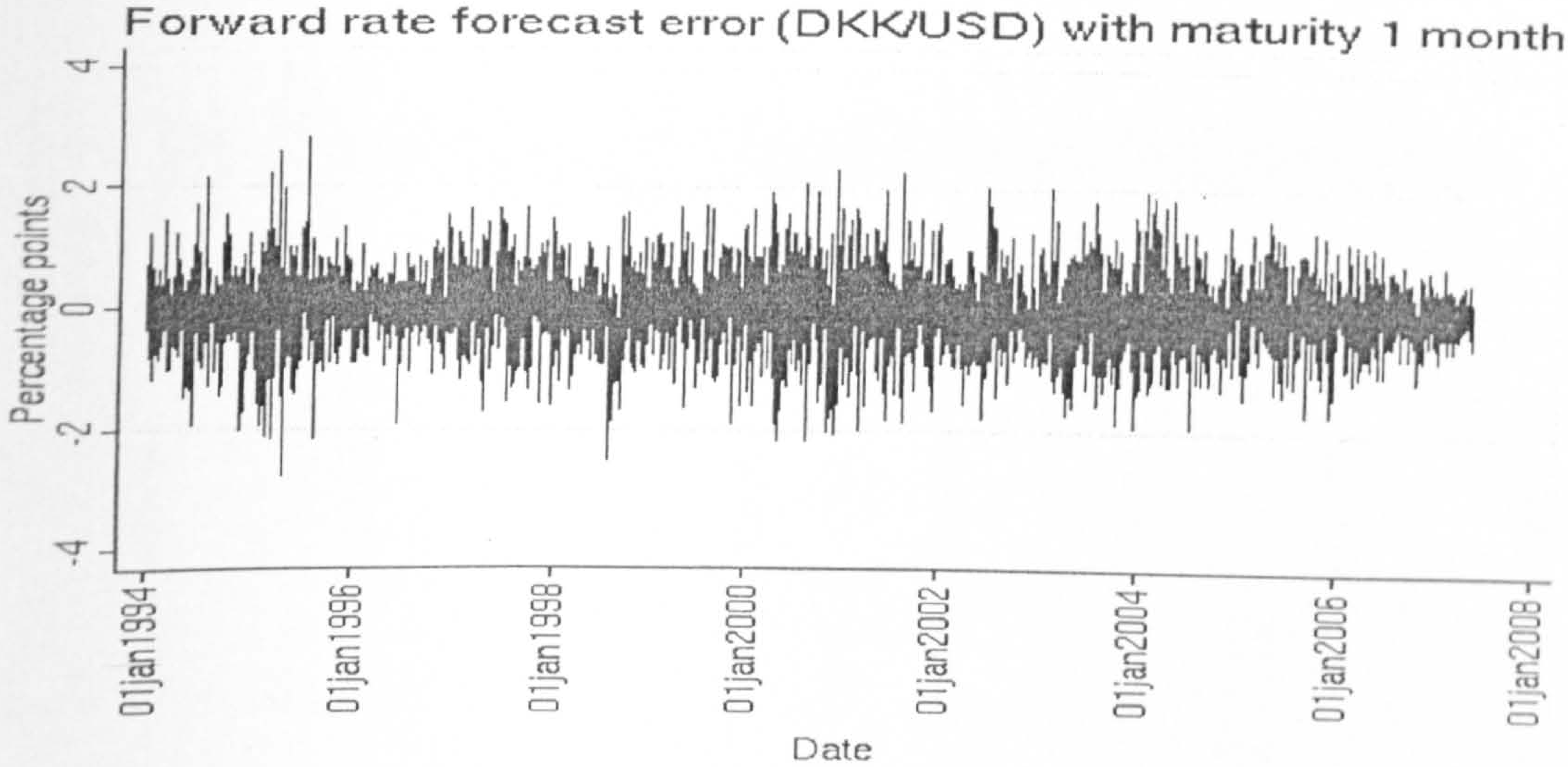


Figure 3.5E: EUR1M

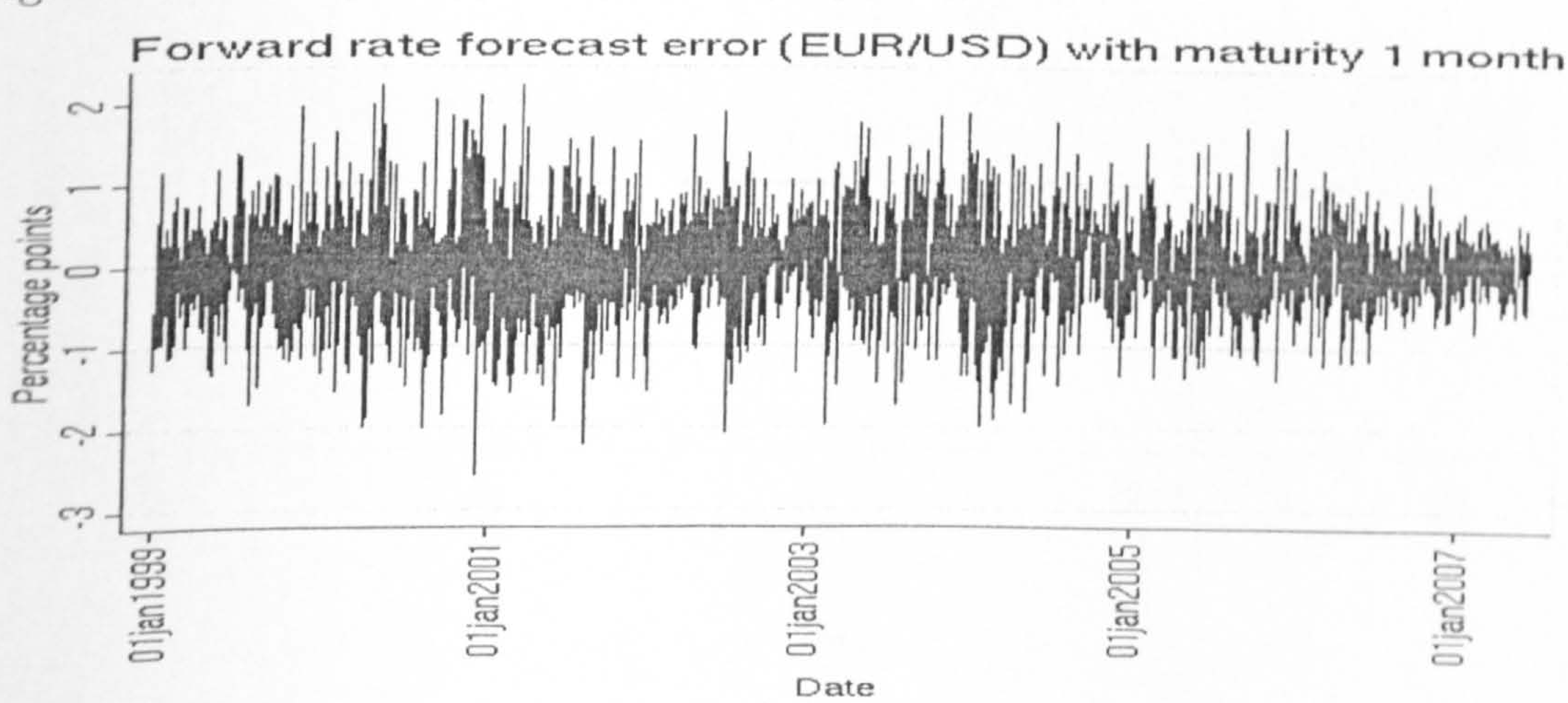


Figure 3.5F: GBP1M

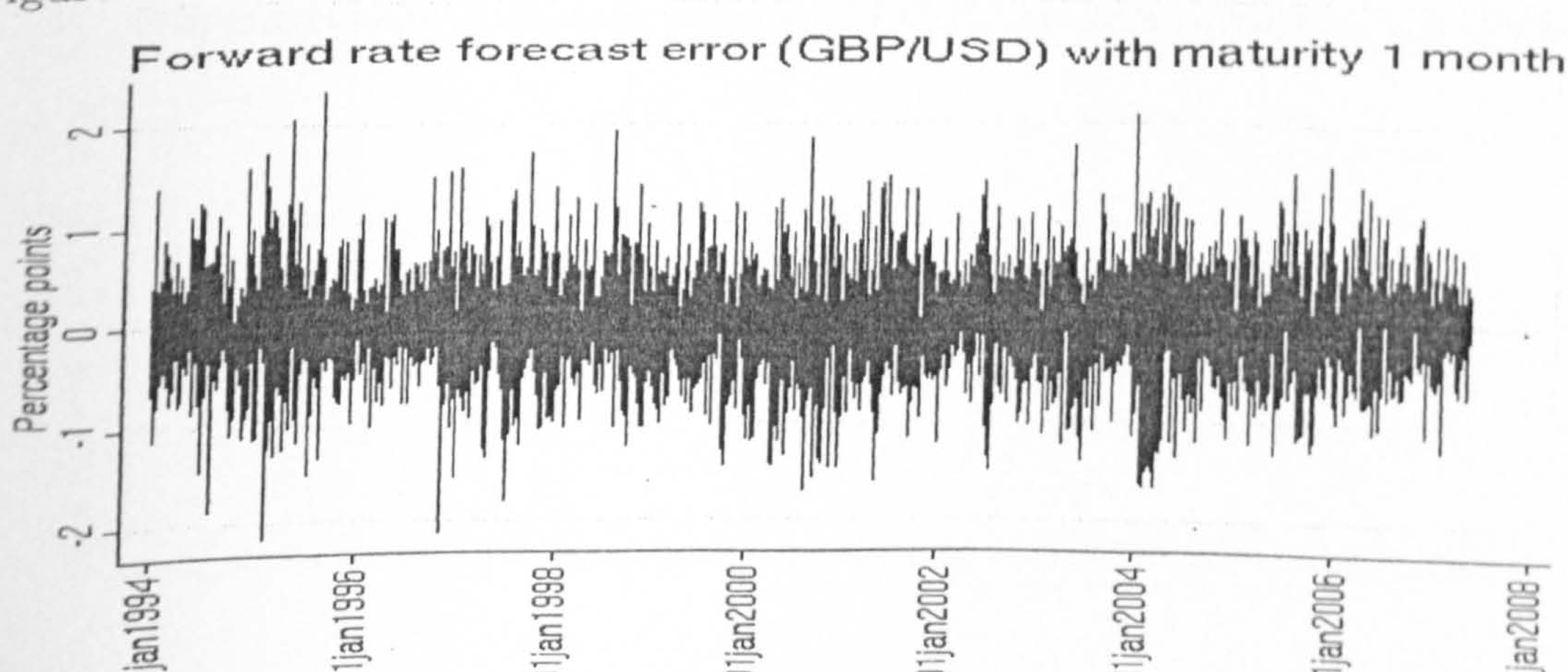


Figure 3.5G: JPY1M

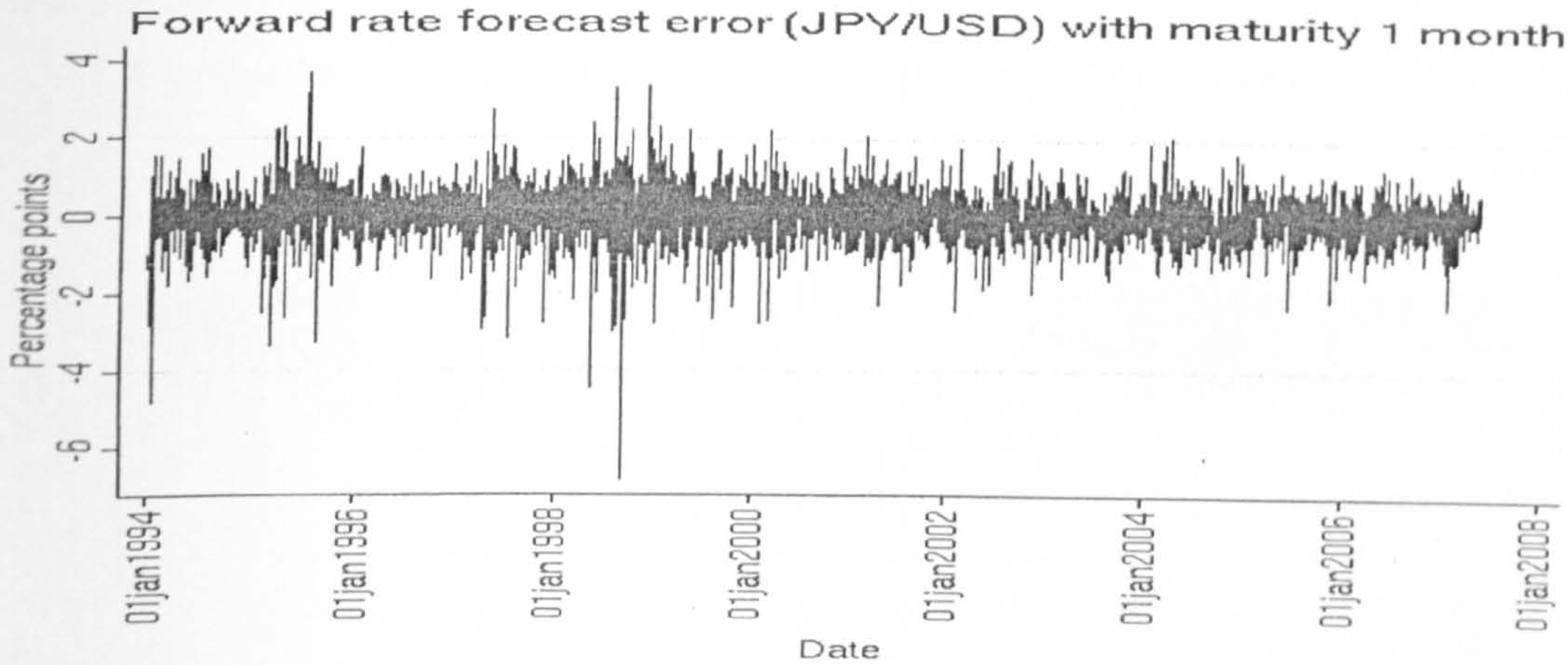


Figure 3.5H: NOK1M

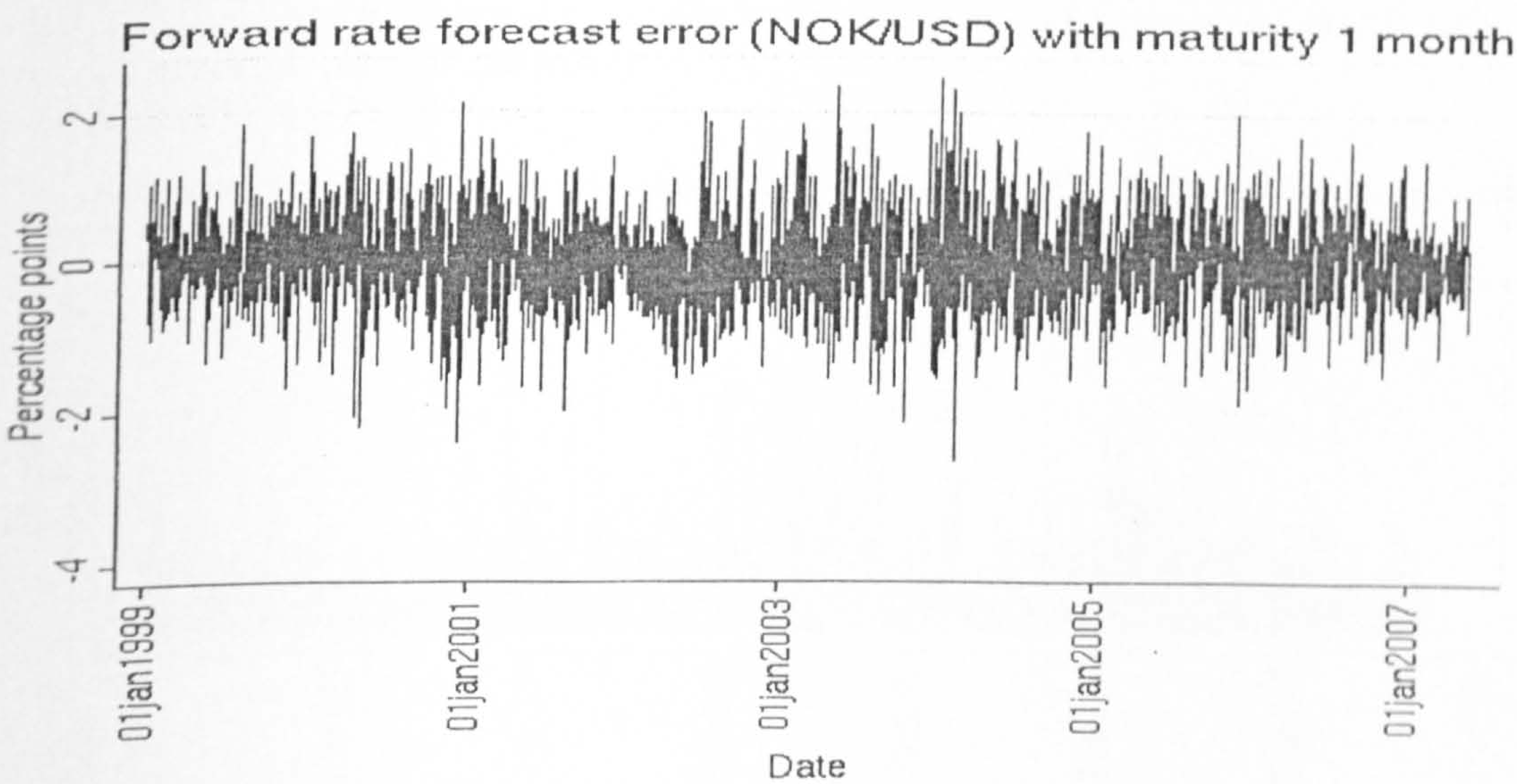


Figure 3.5I: NZD1M

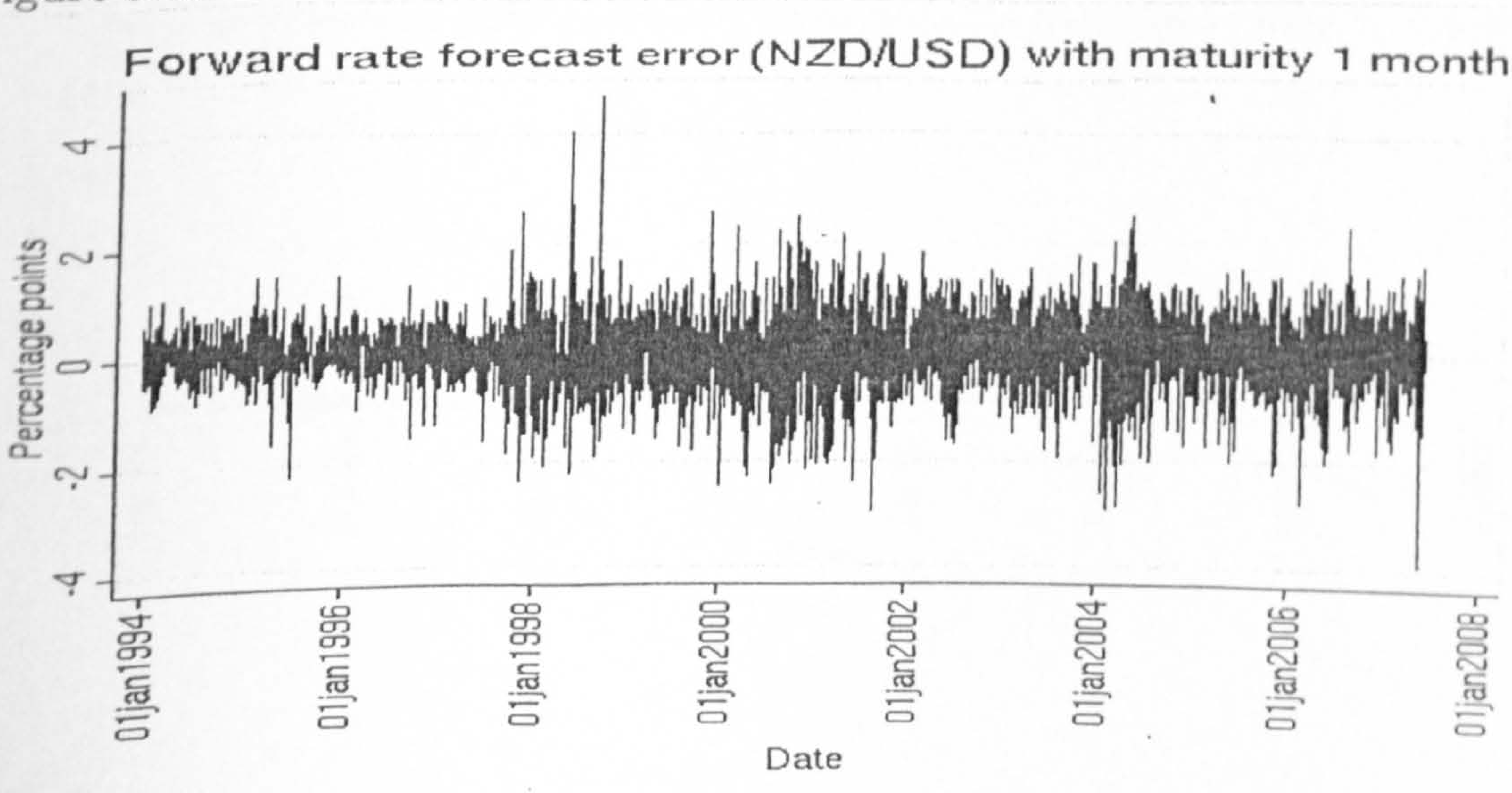
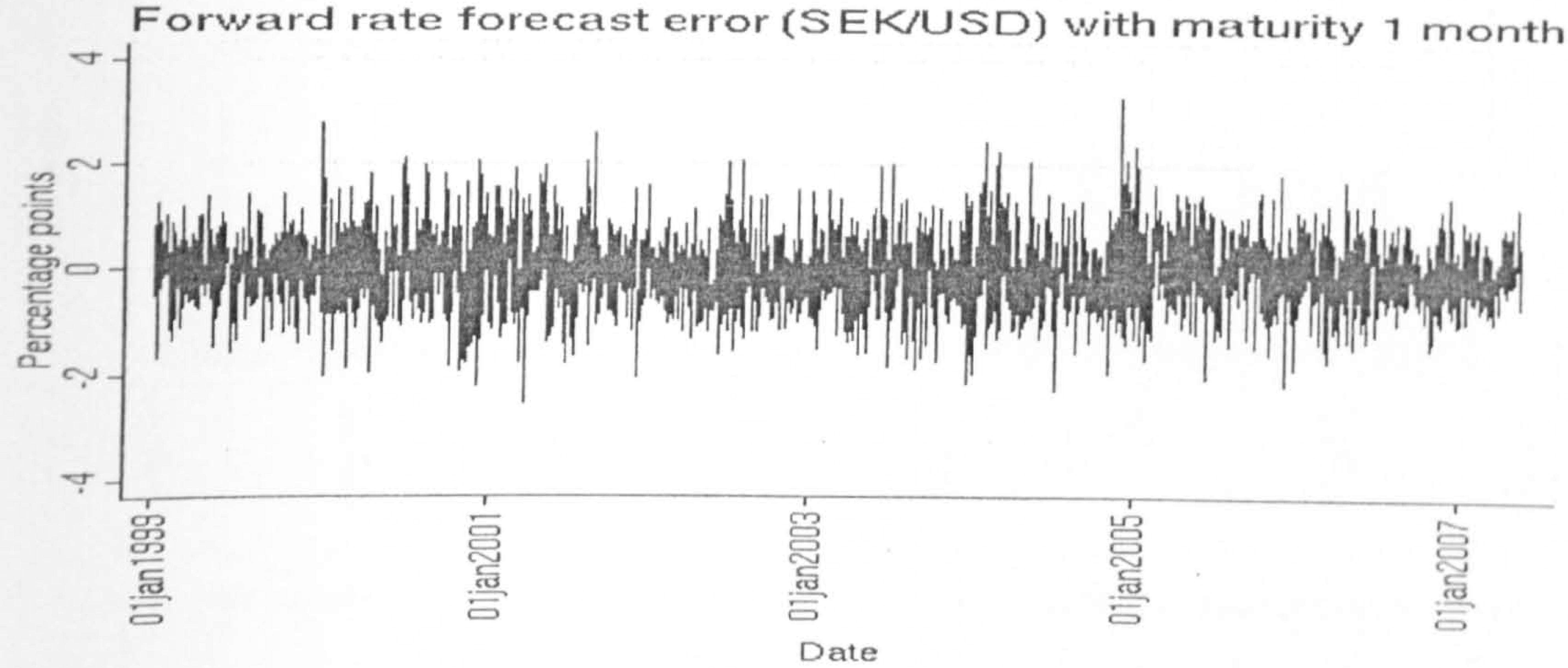
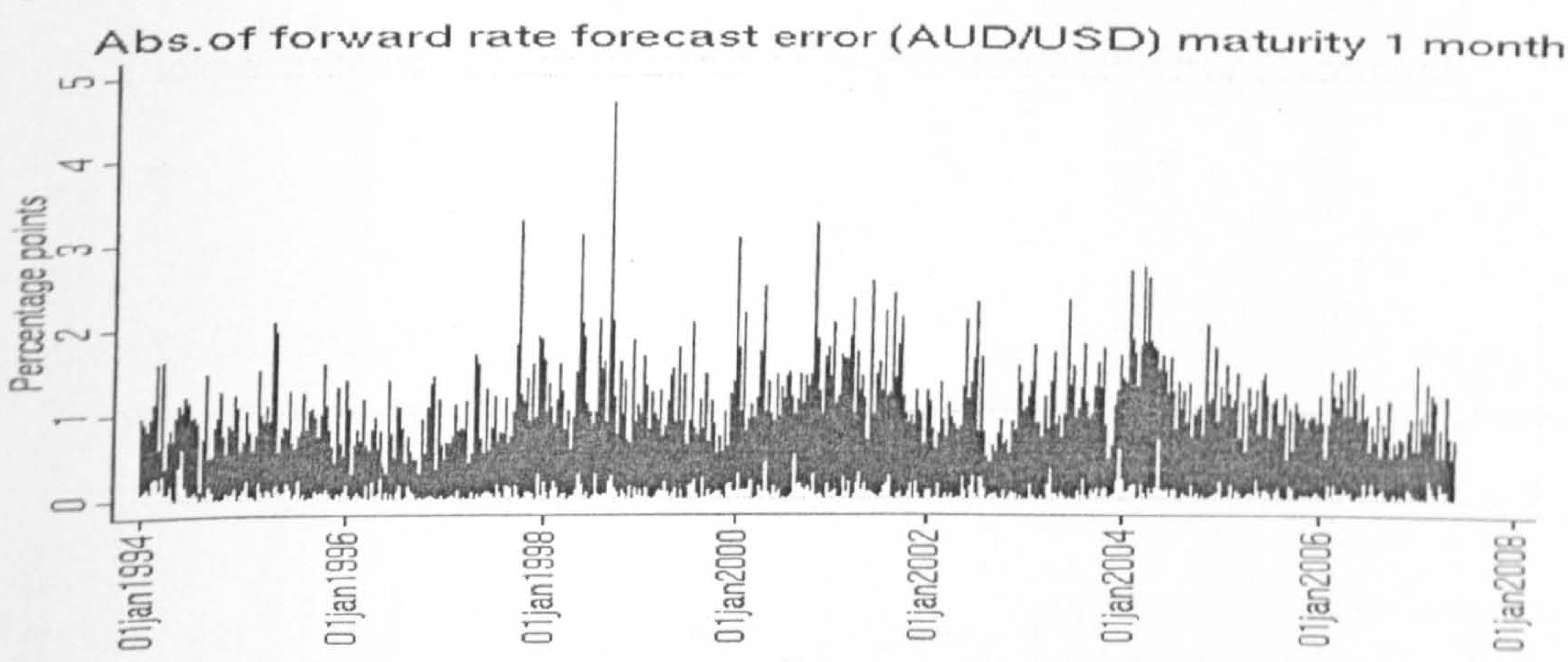


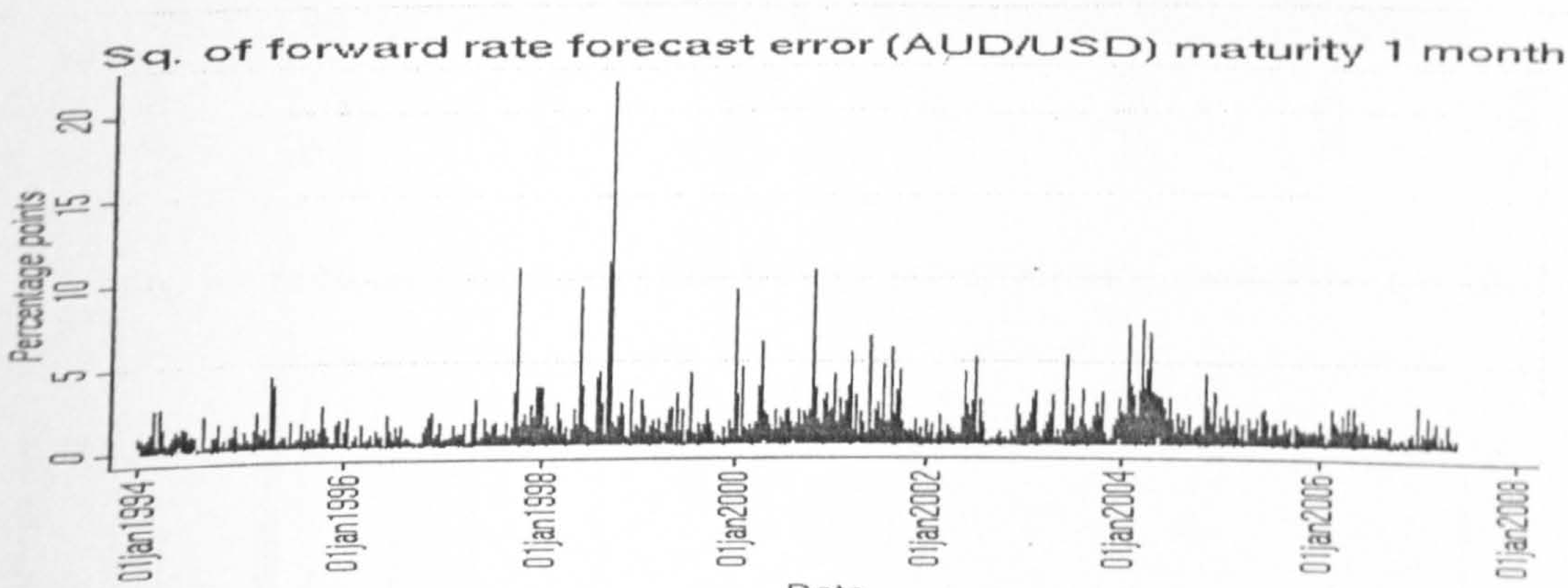
Figure 3.5J: SEK1M



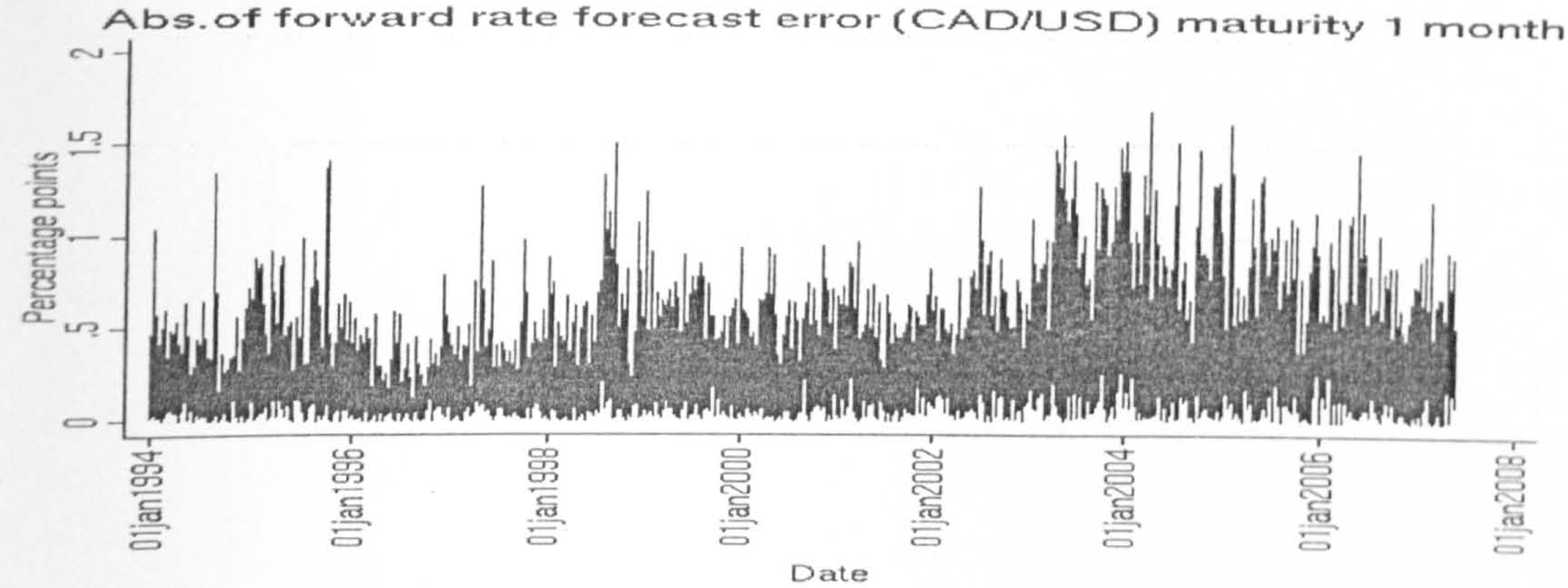
Figures 3.6: Plots of absolute and squared value of the filtered forward rate forecast error, 1 month maturity

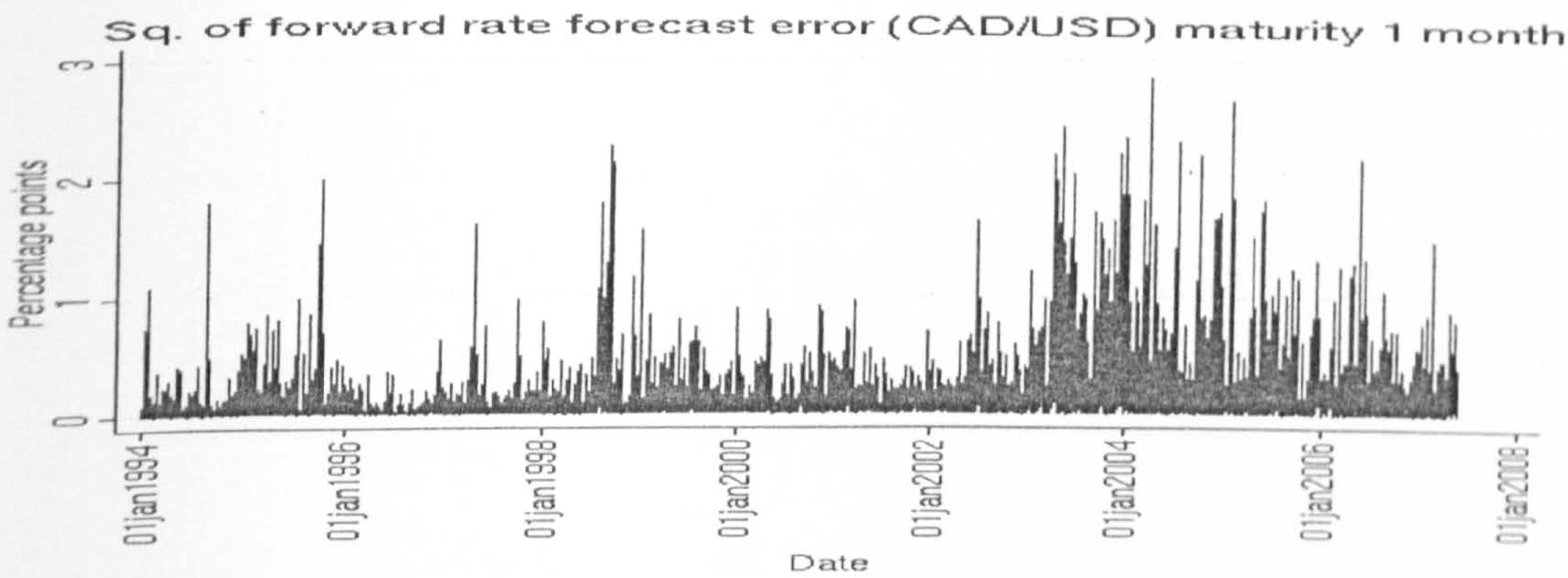
Figures 3.6A: AUD1M



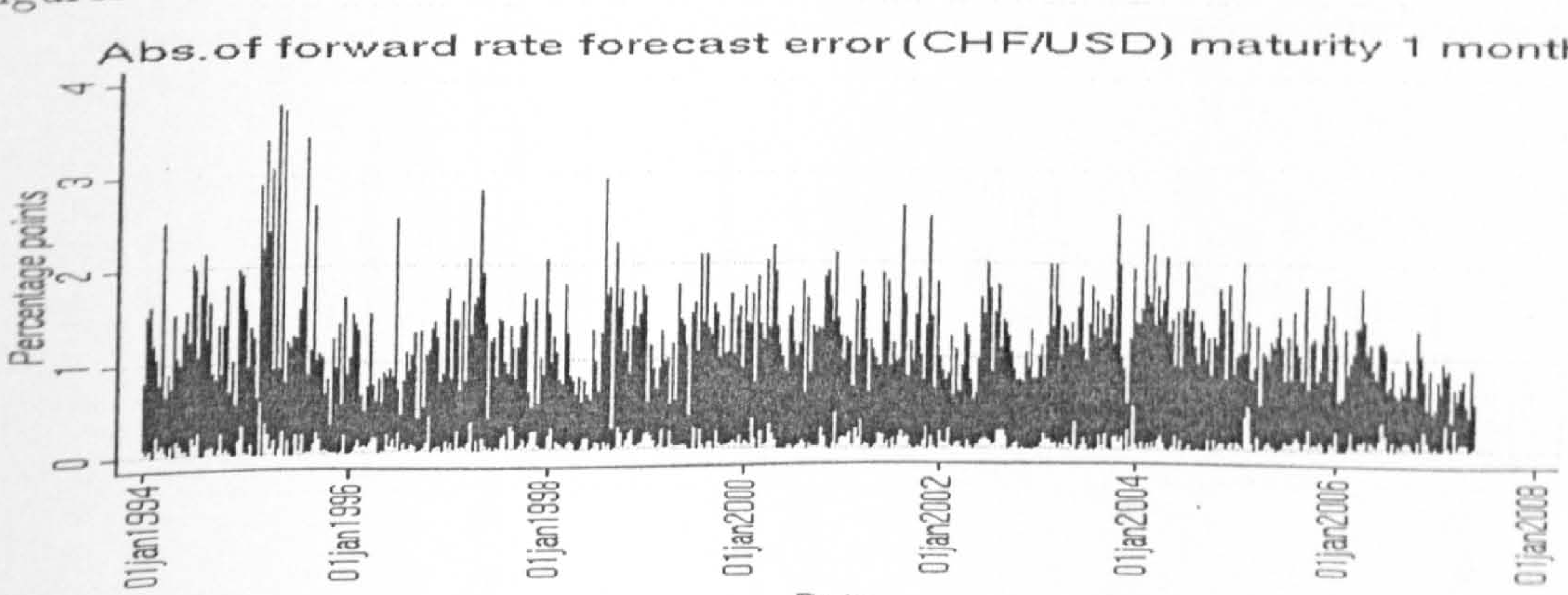


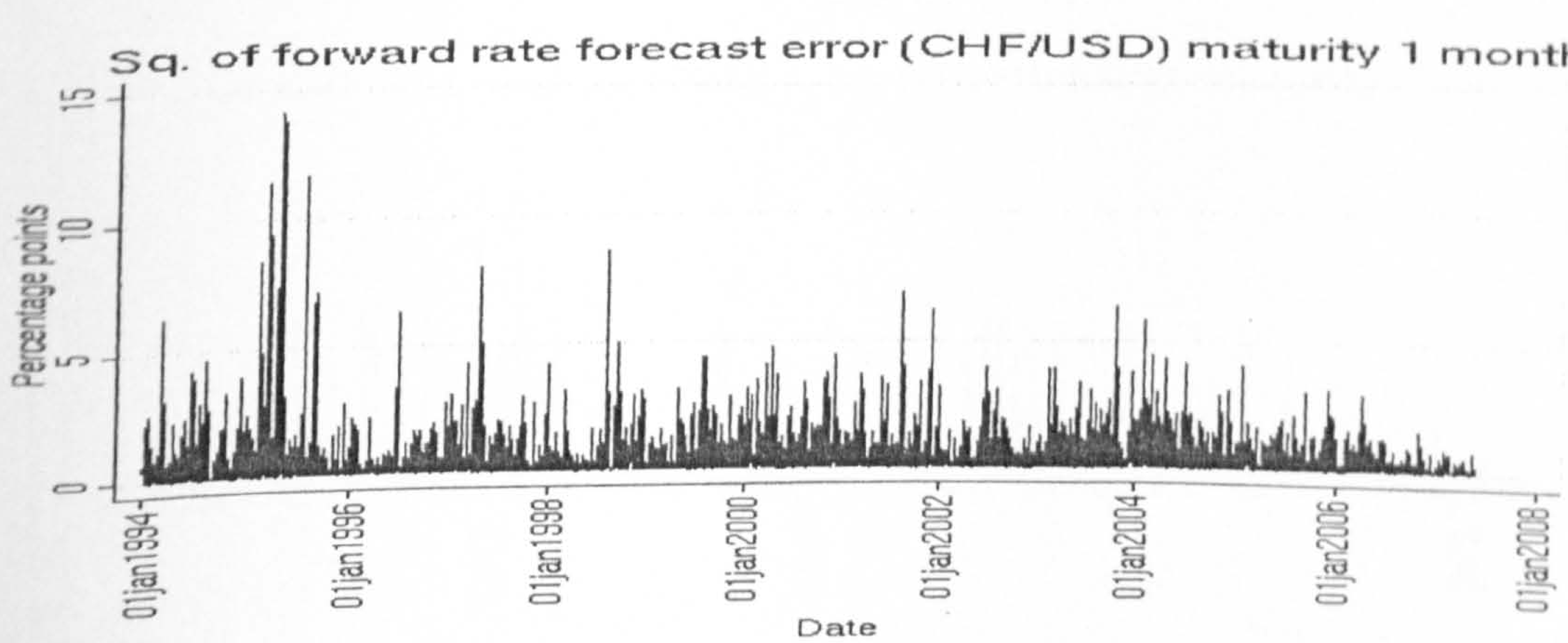
Figures 3.6B: CAD1M



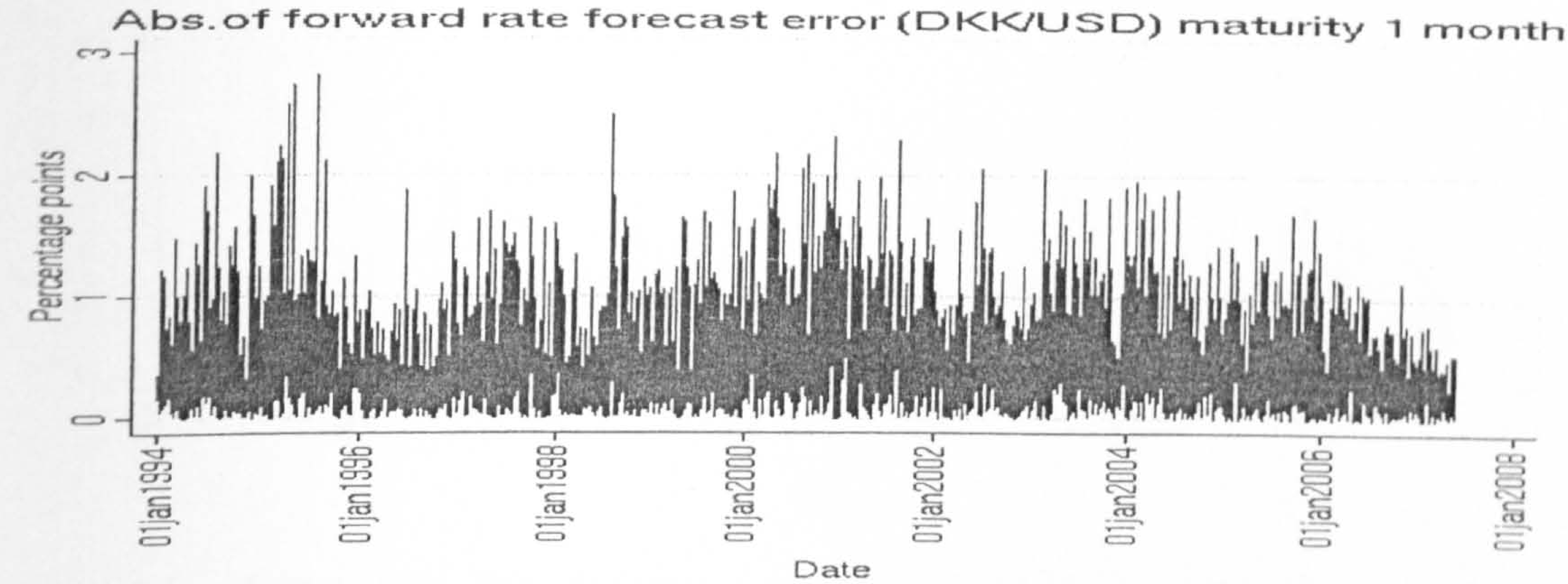


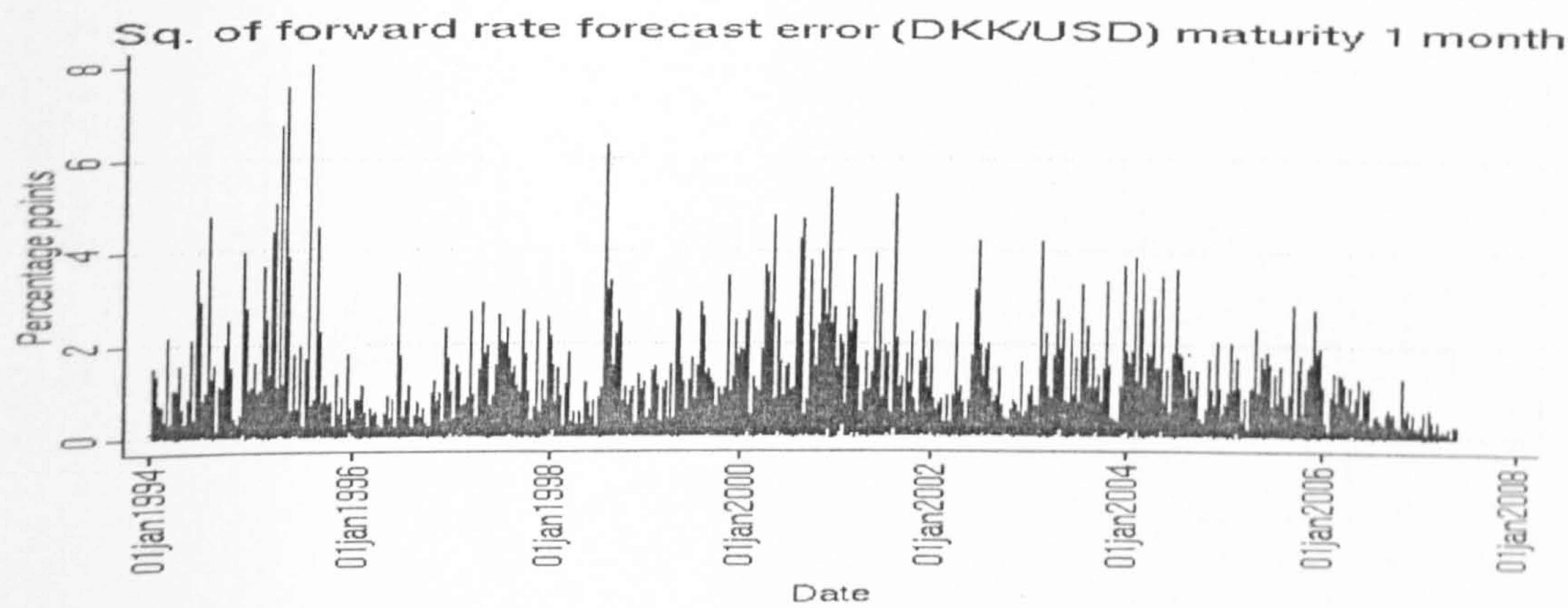
Figures 3.6C: CHF1M



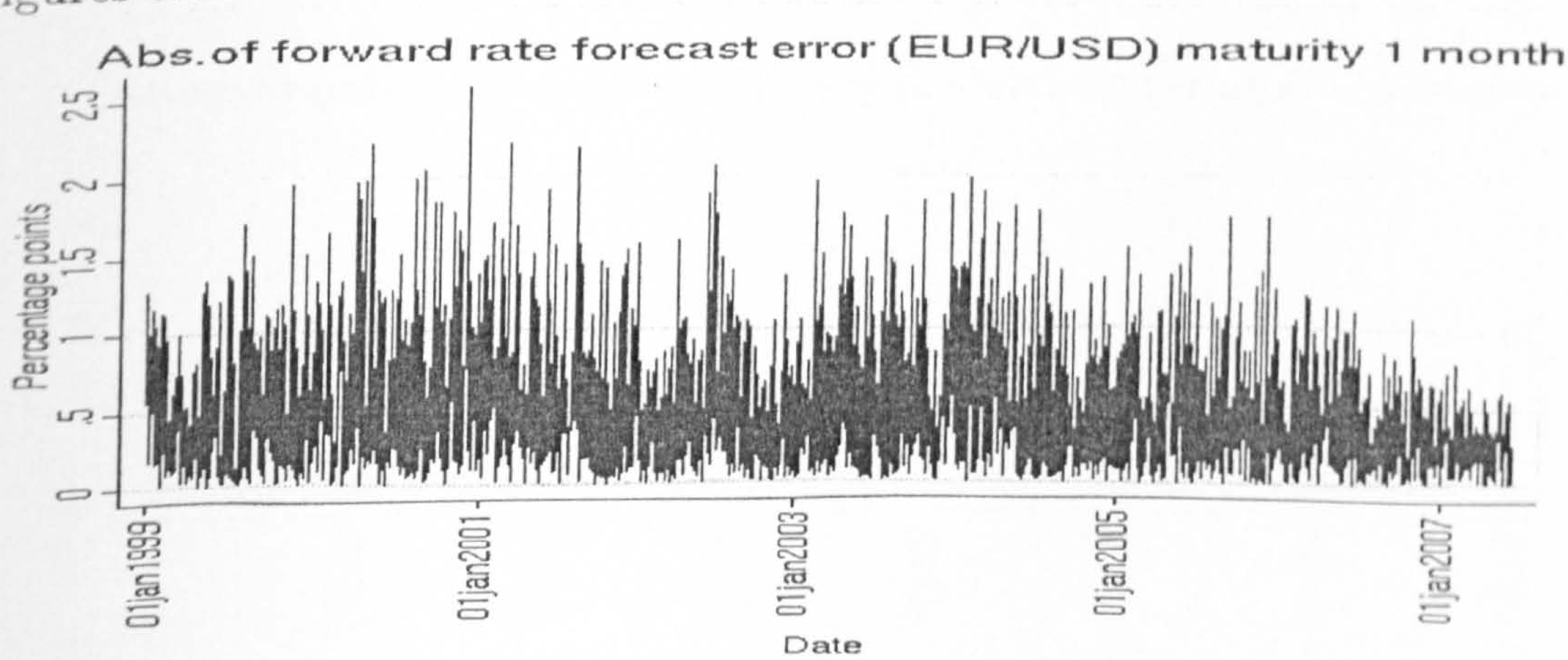


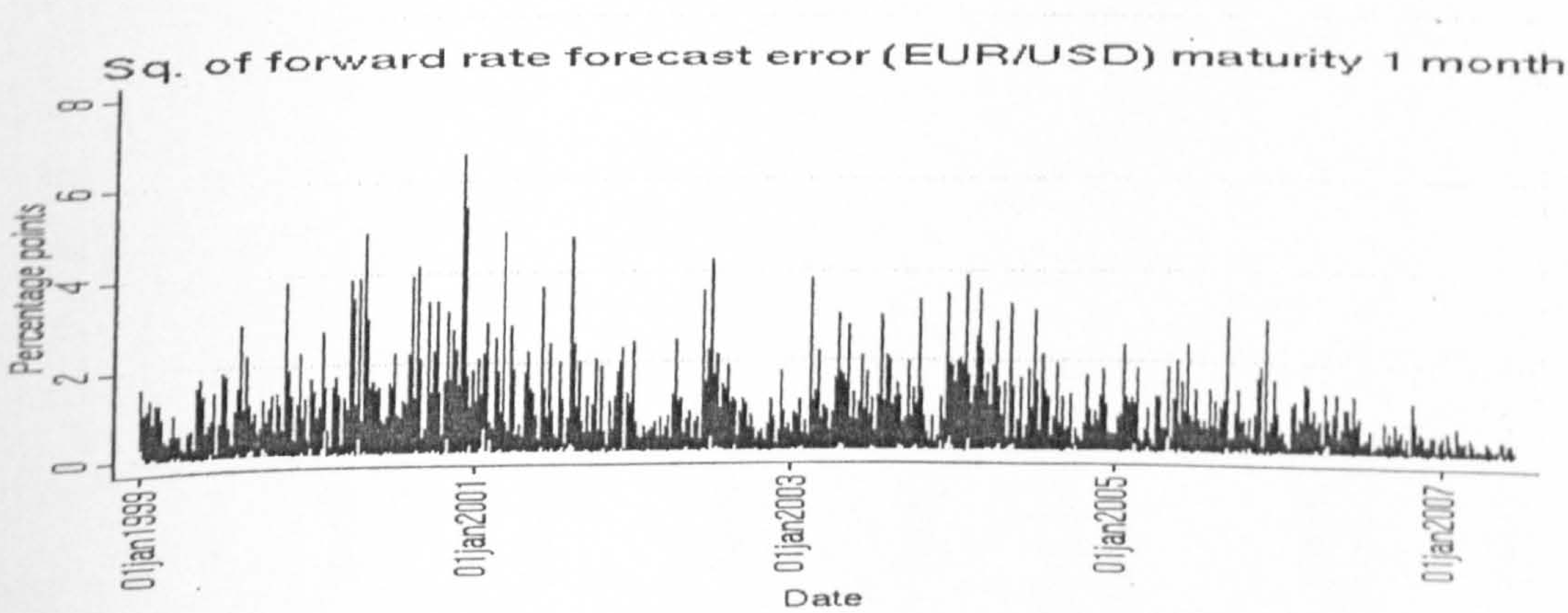
Figures 3.6D: DKK1M



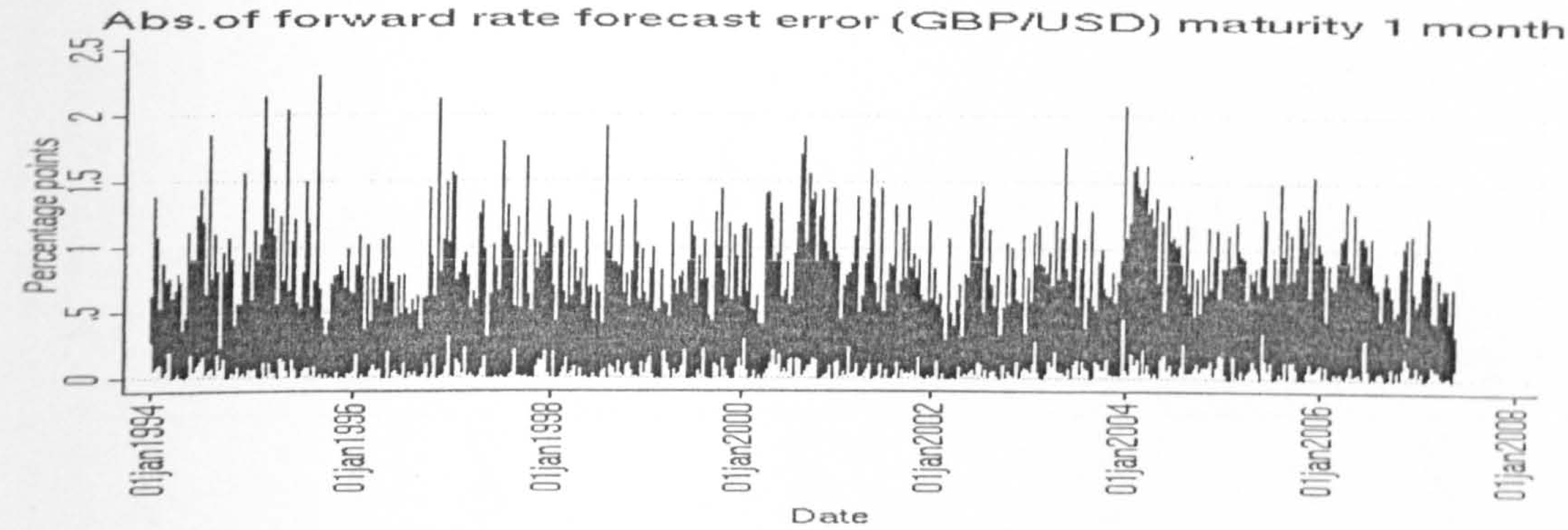


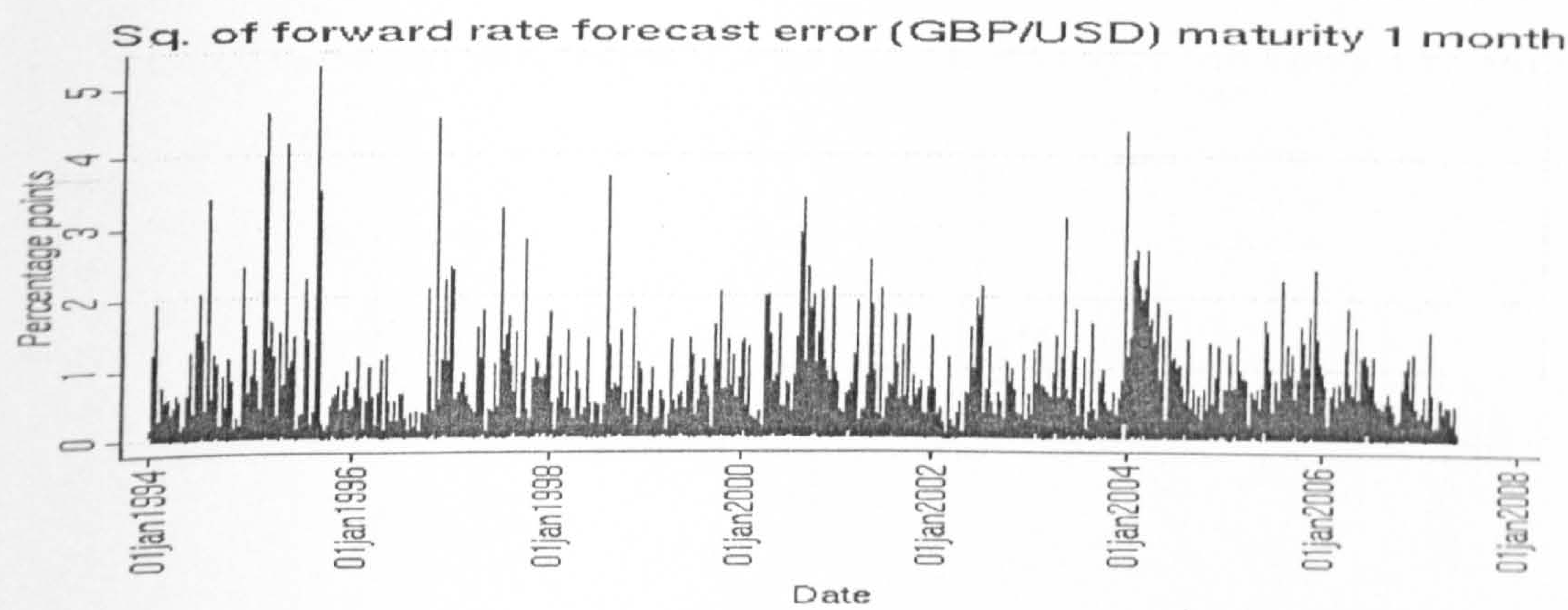
Figures 3.6E: EUR1M



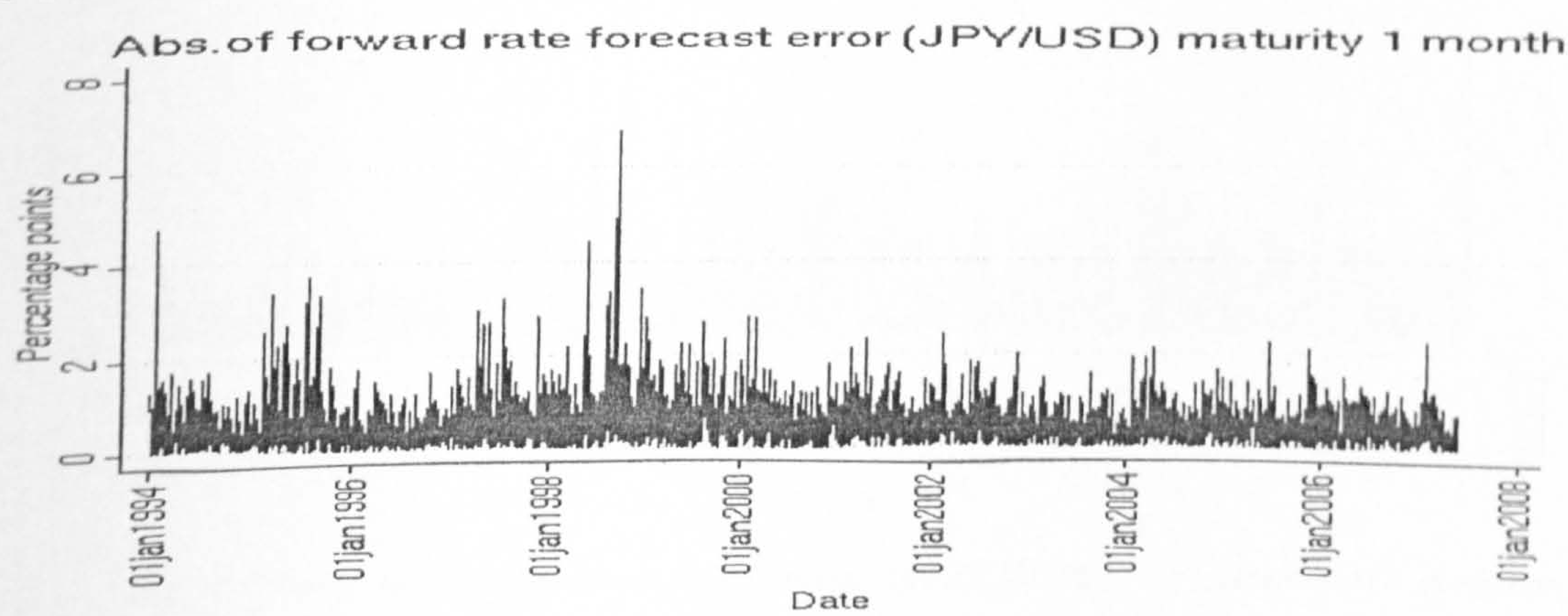


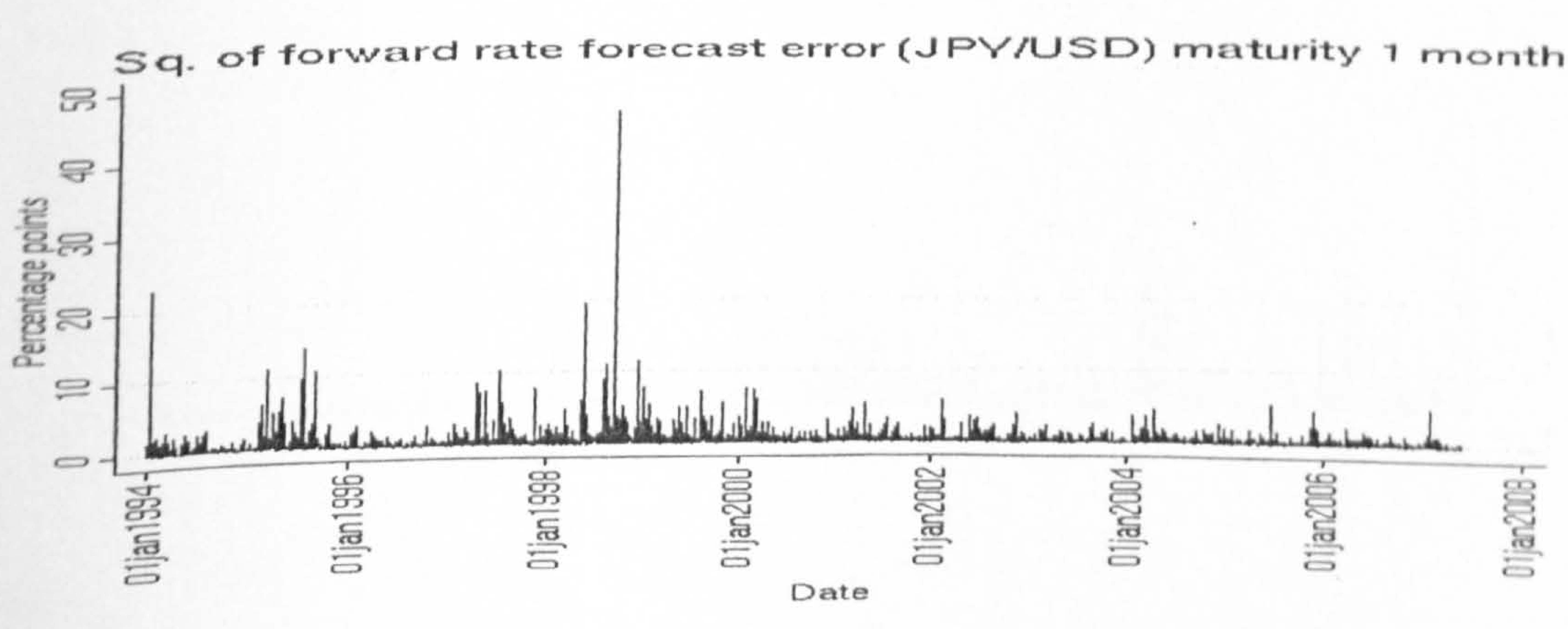
Figures 3.6F: GBP1M



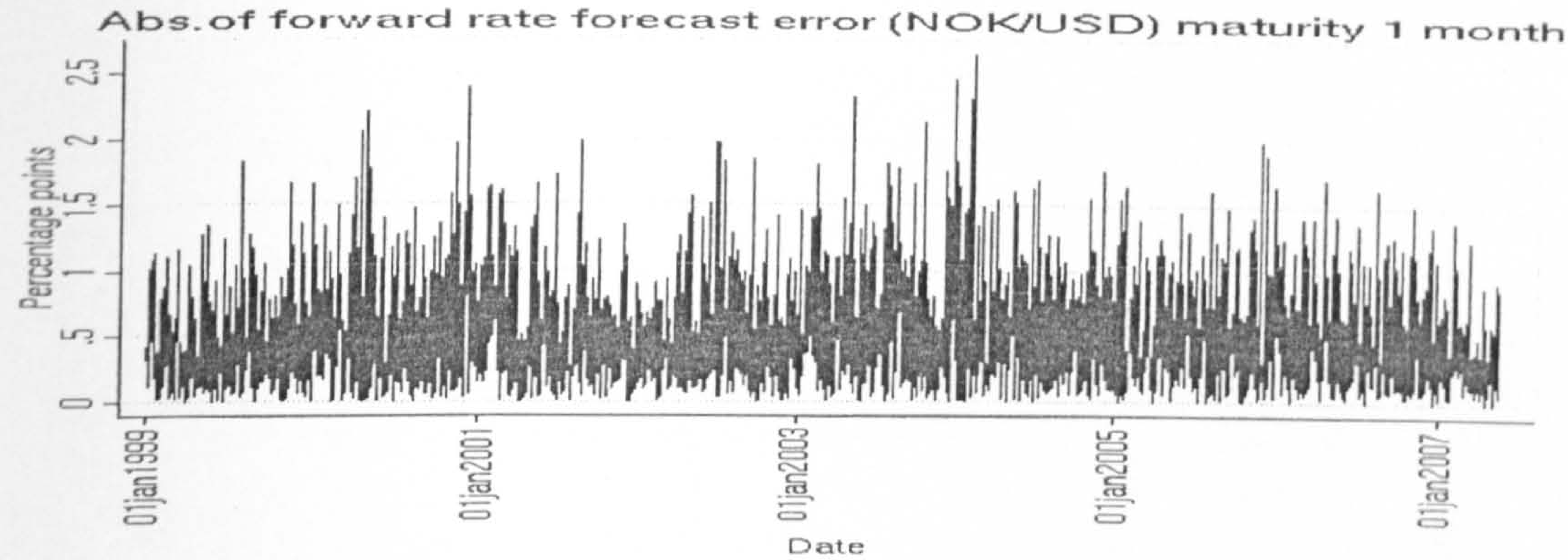


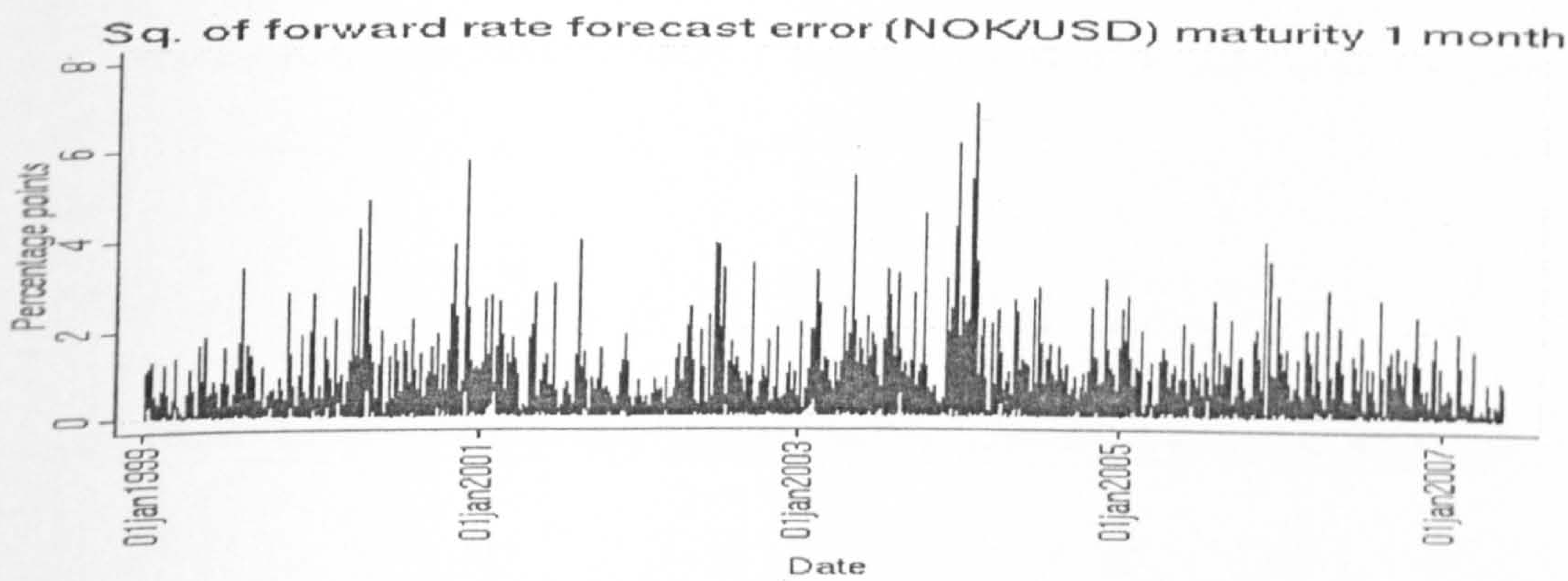
Figures 3.6G:
JPY1



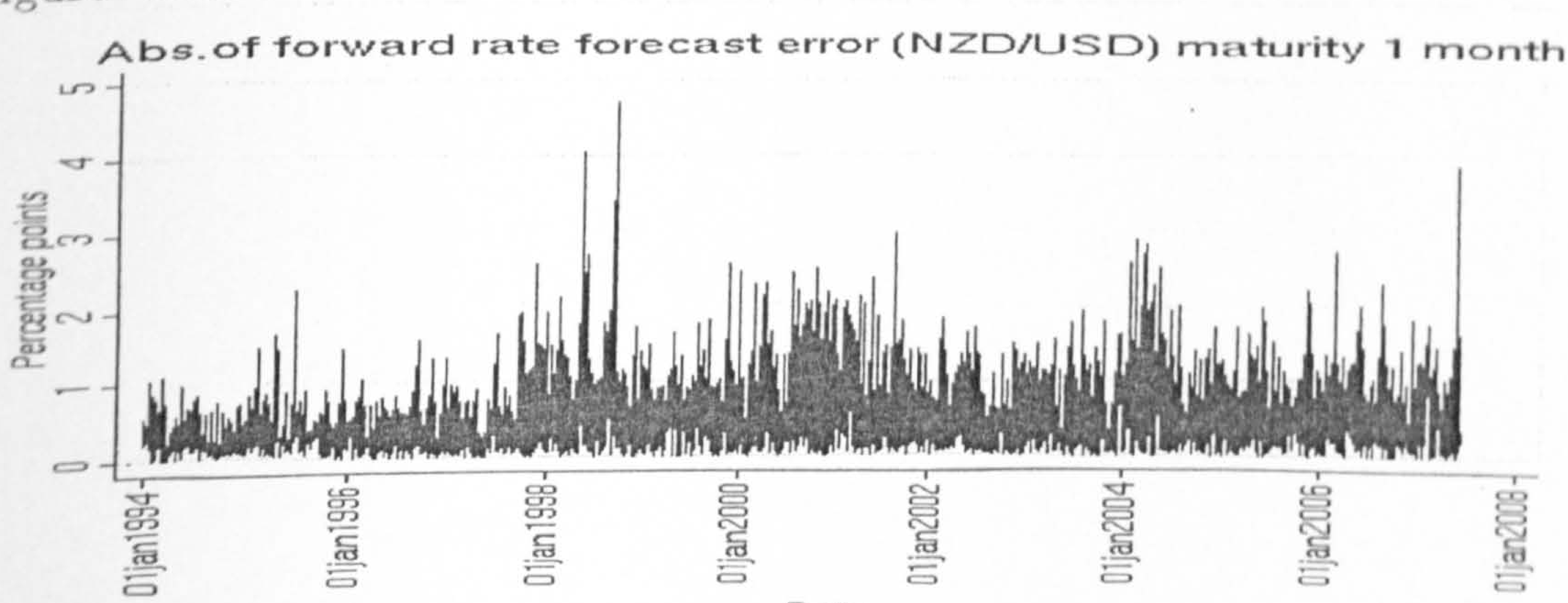


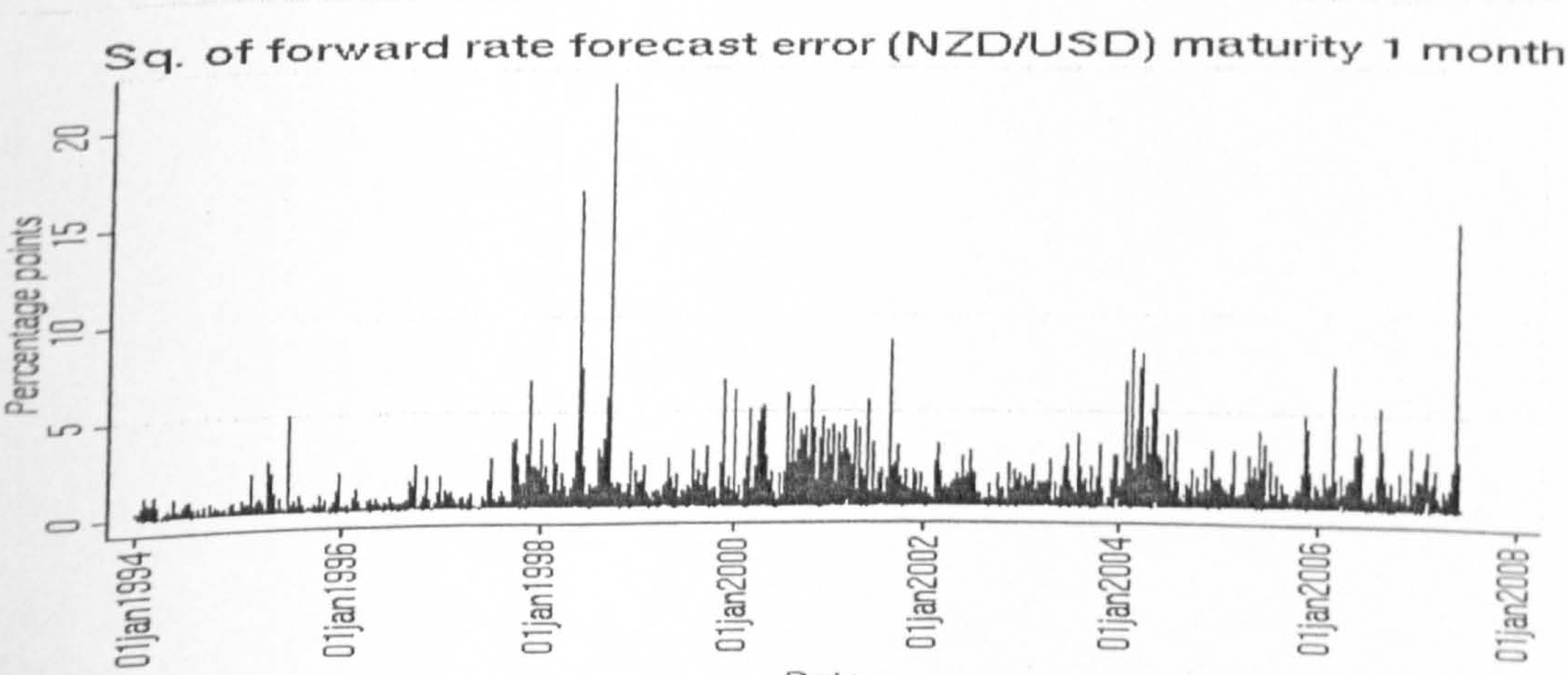
Figures 3.6H: NOK1M



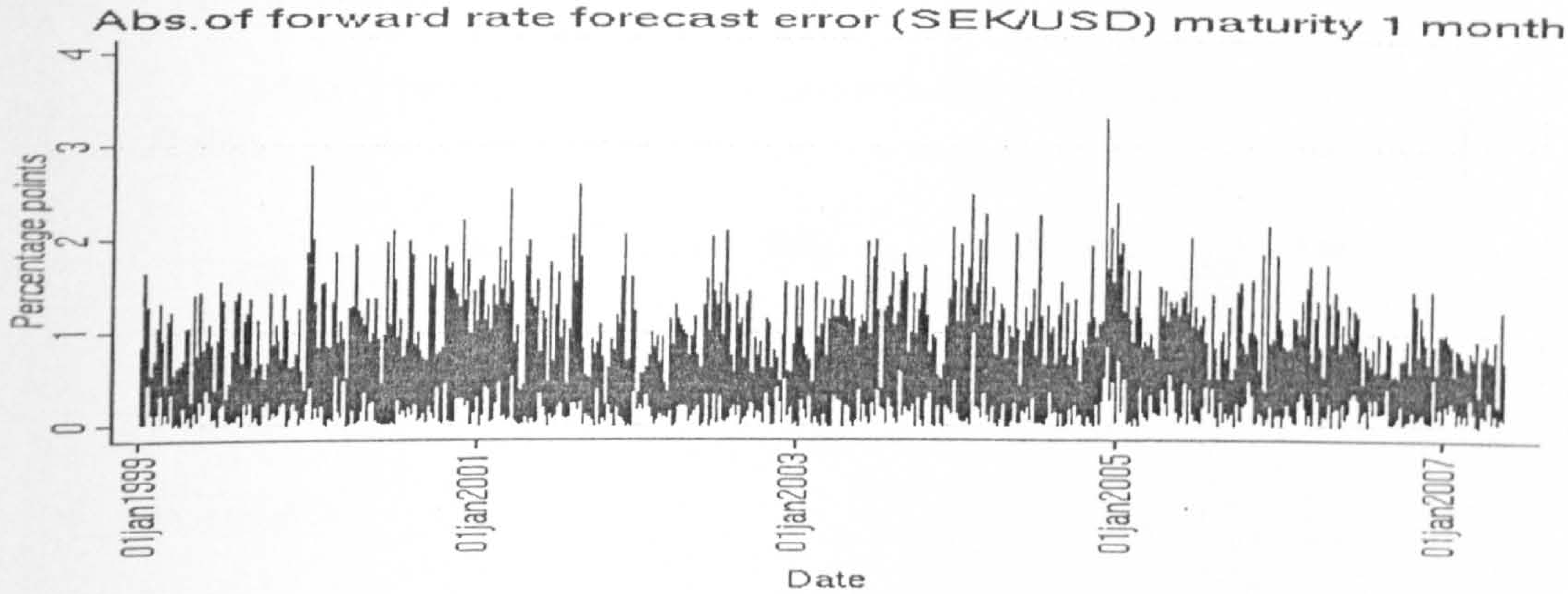


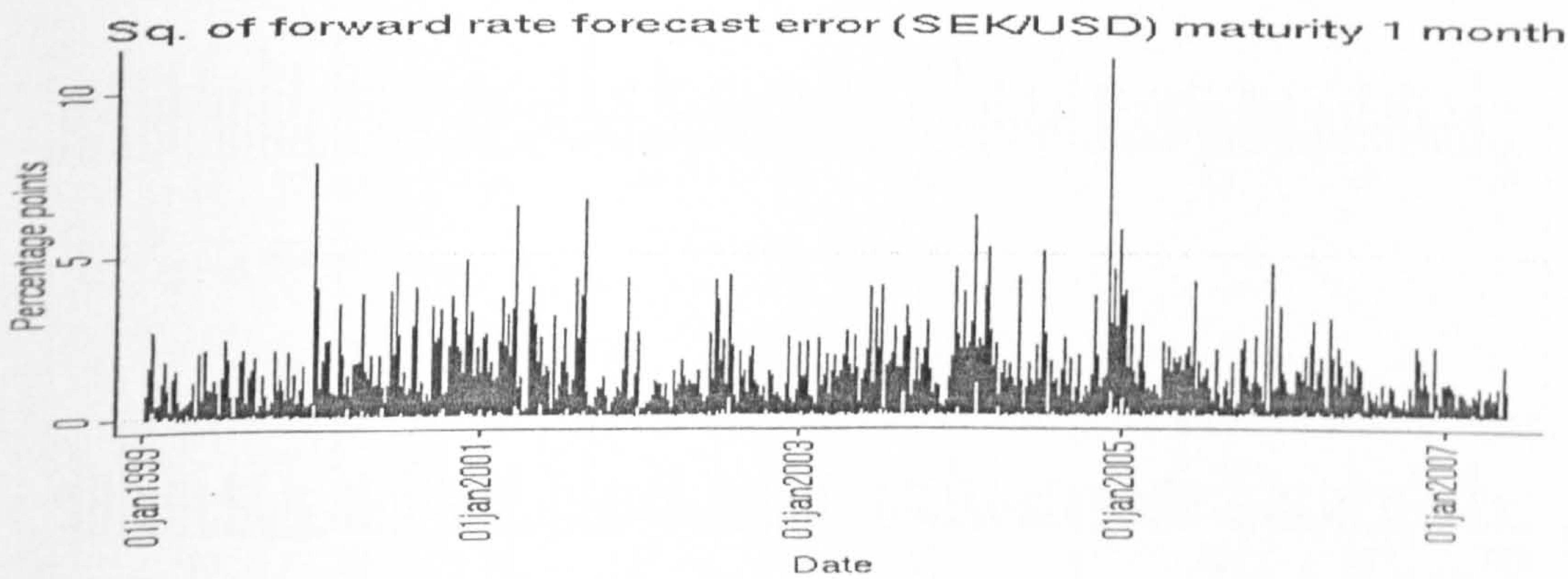
Figures 3.6I: NZD1M





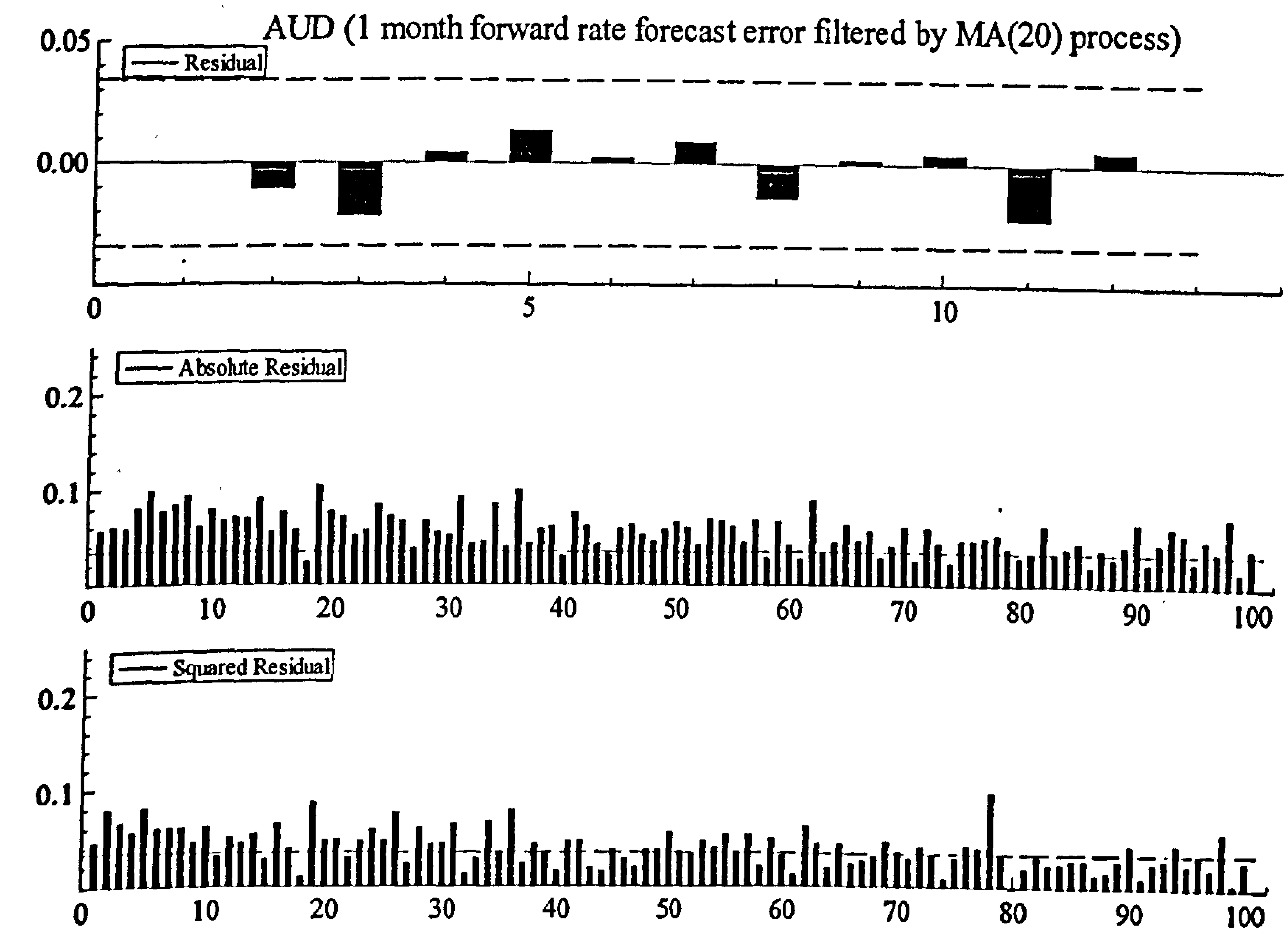
Figures 3.6J: SEK1M



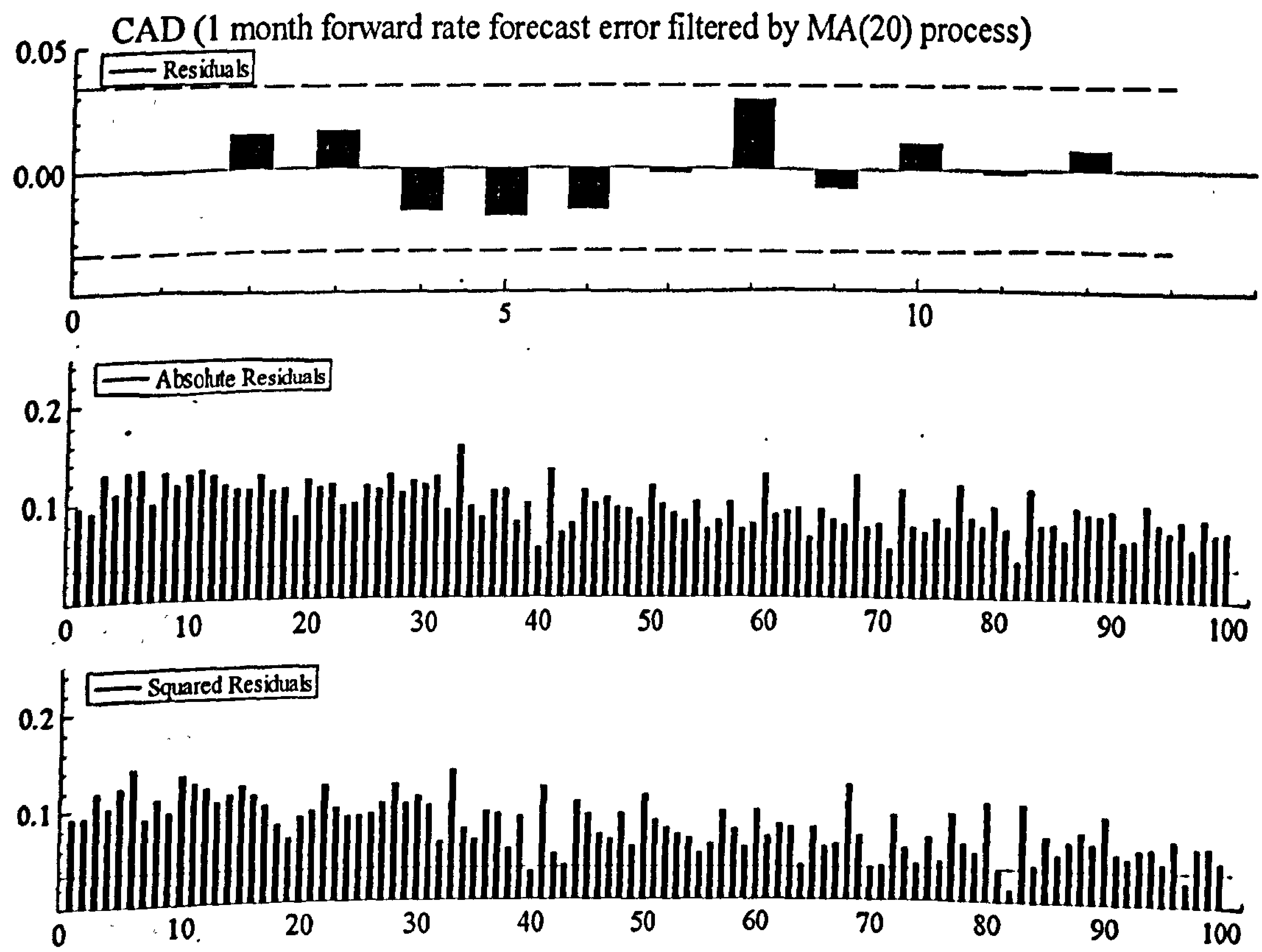


Figures 3.7: The autocorrelation plots of the forward rate forecast error series

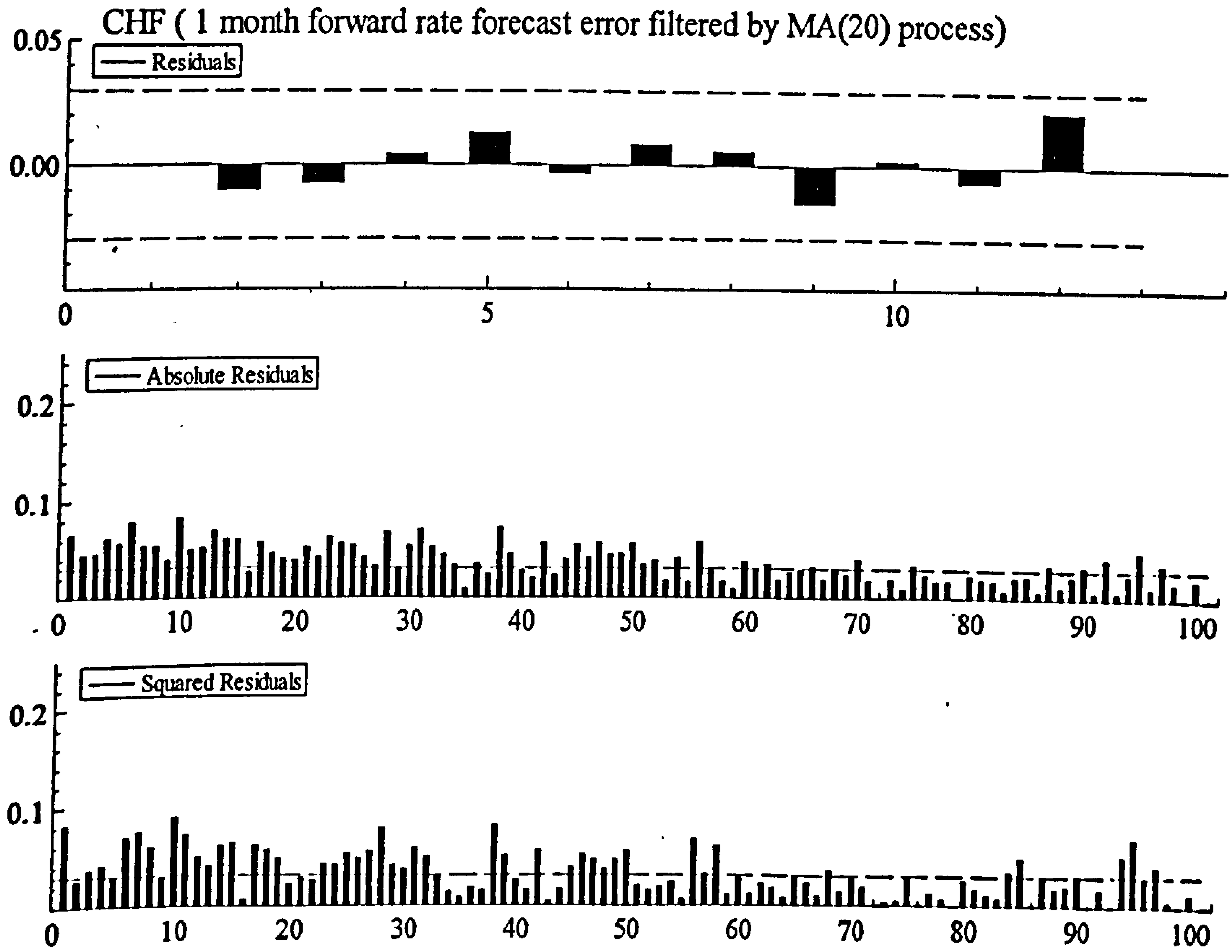
Figures 3.7A: AUD1M



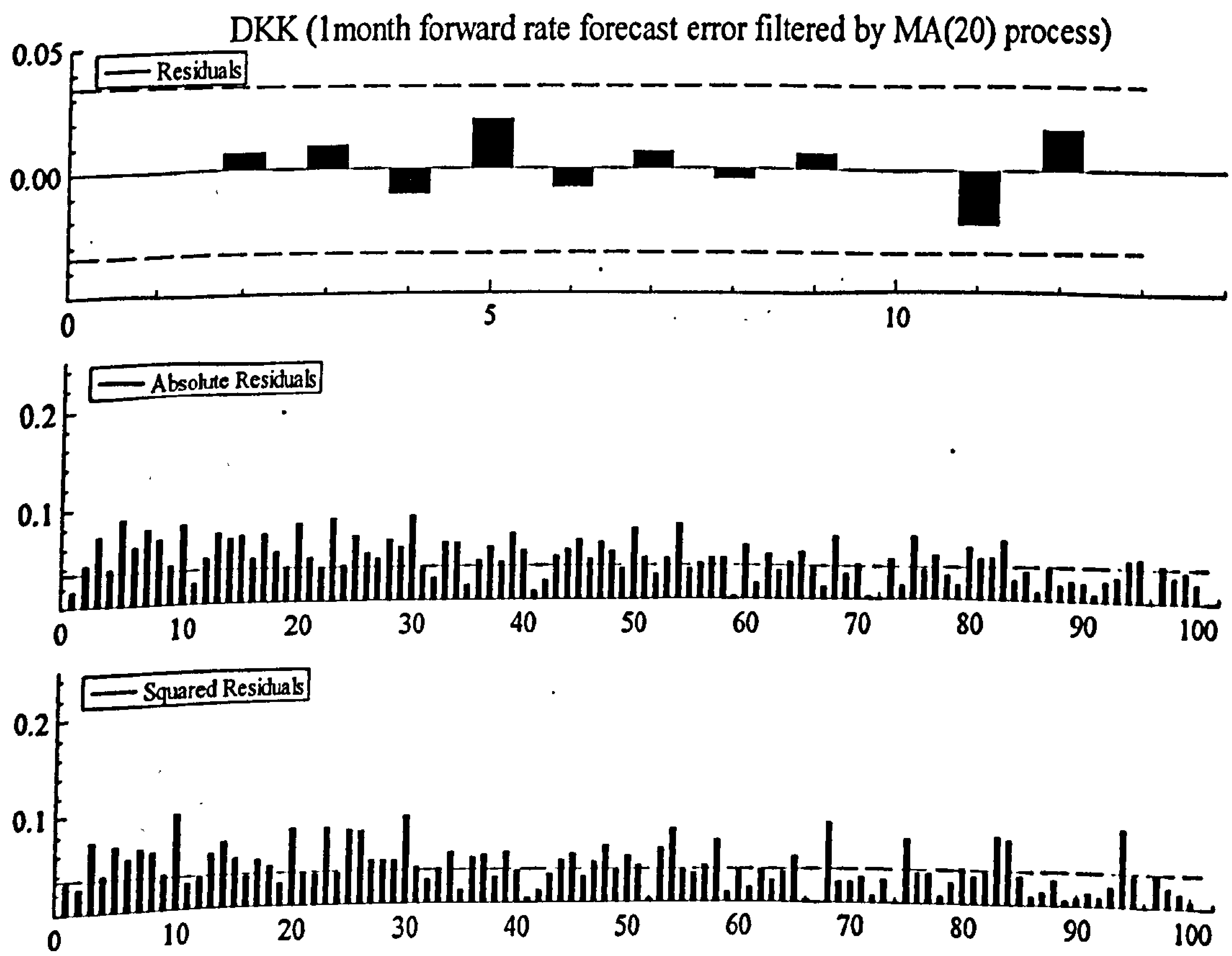
Figures 3.7B: CAD1M



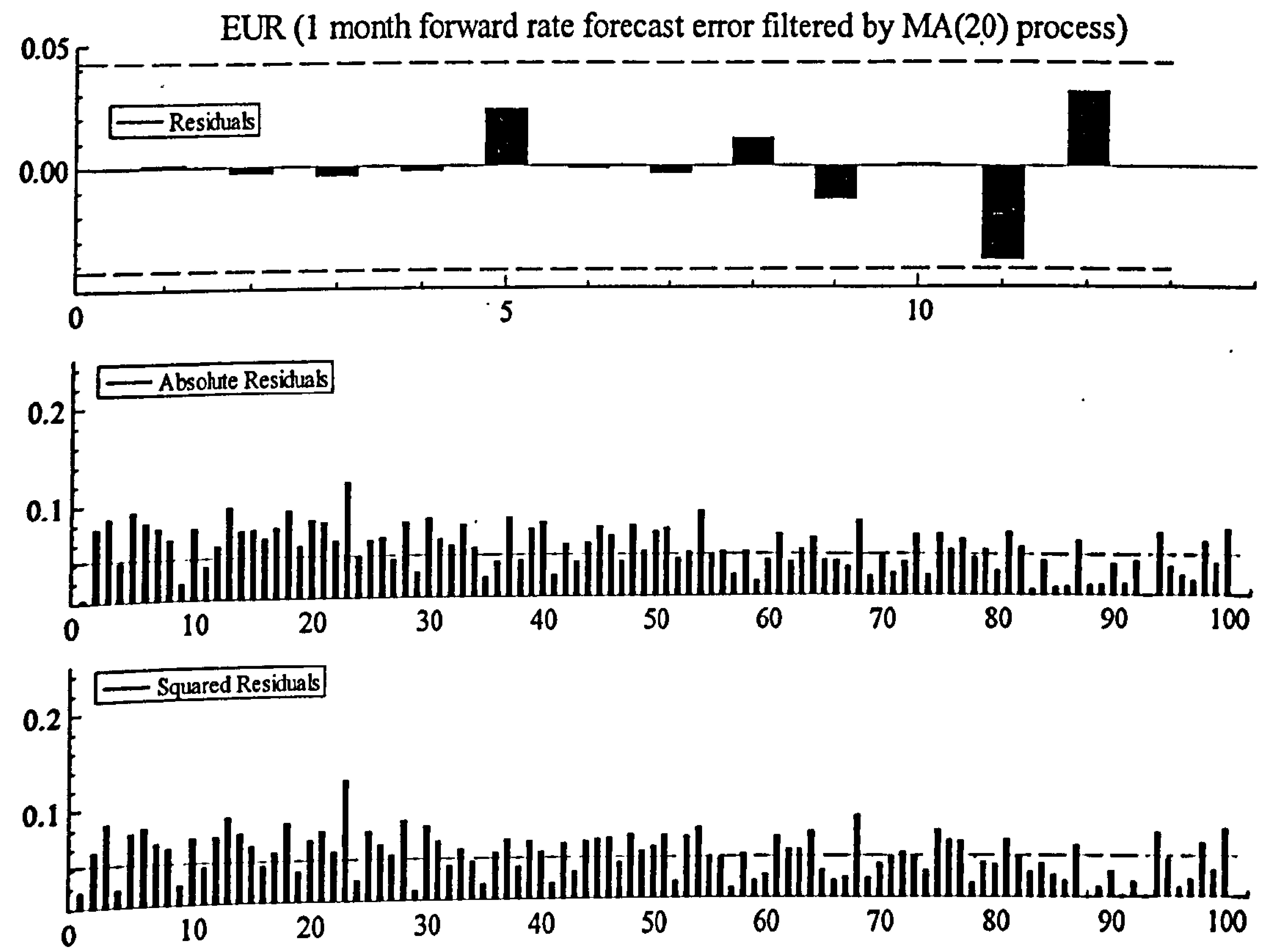
Figures 3.7C: CHF1M



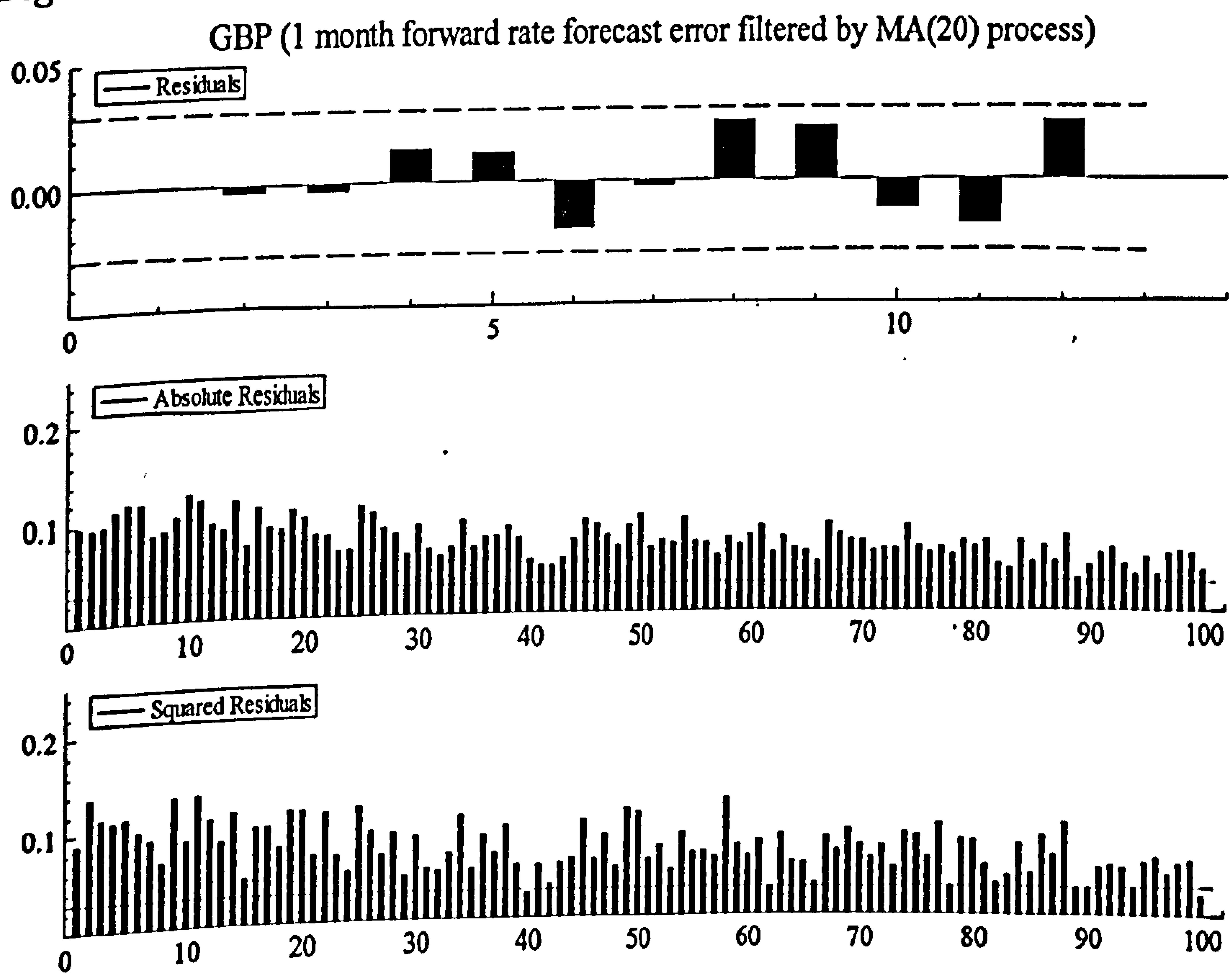
Figures 3.7D: DKK1M



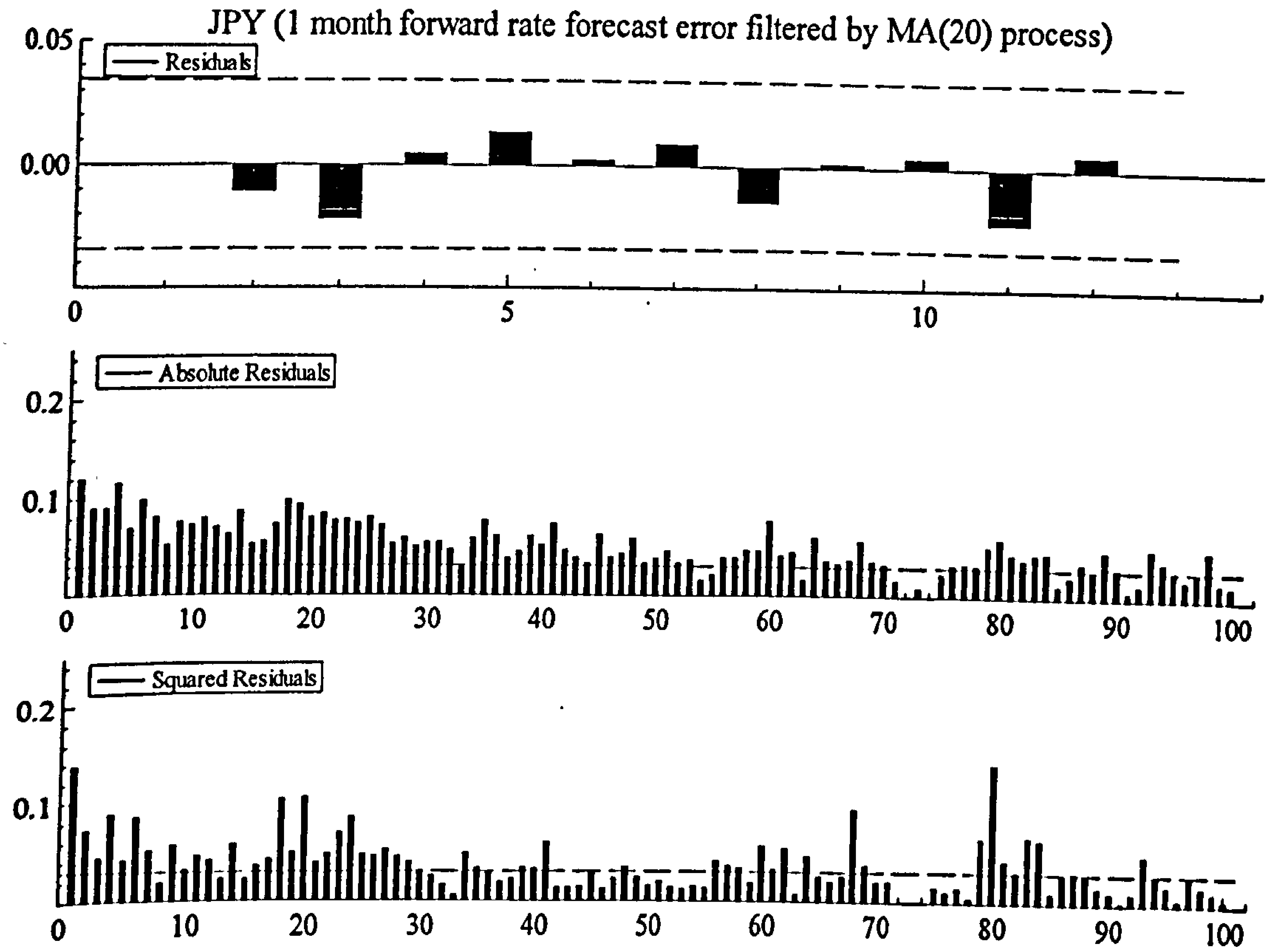
Figures 3.7E: EUR1M



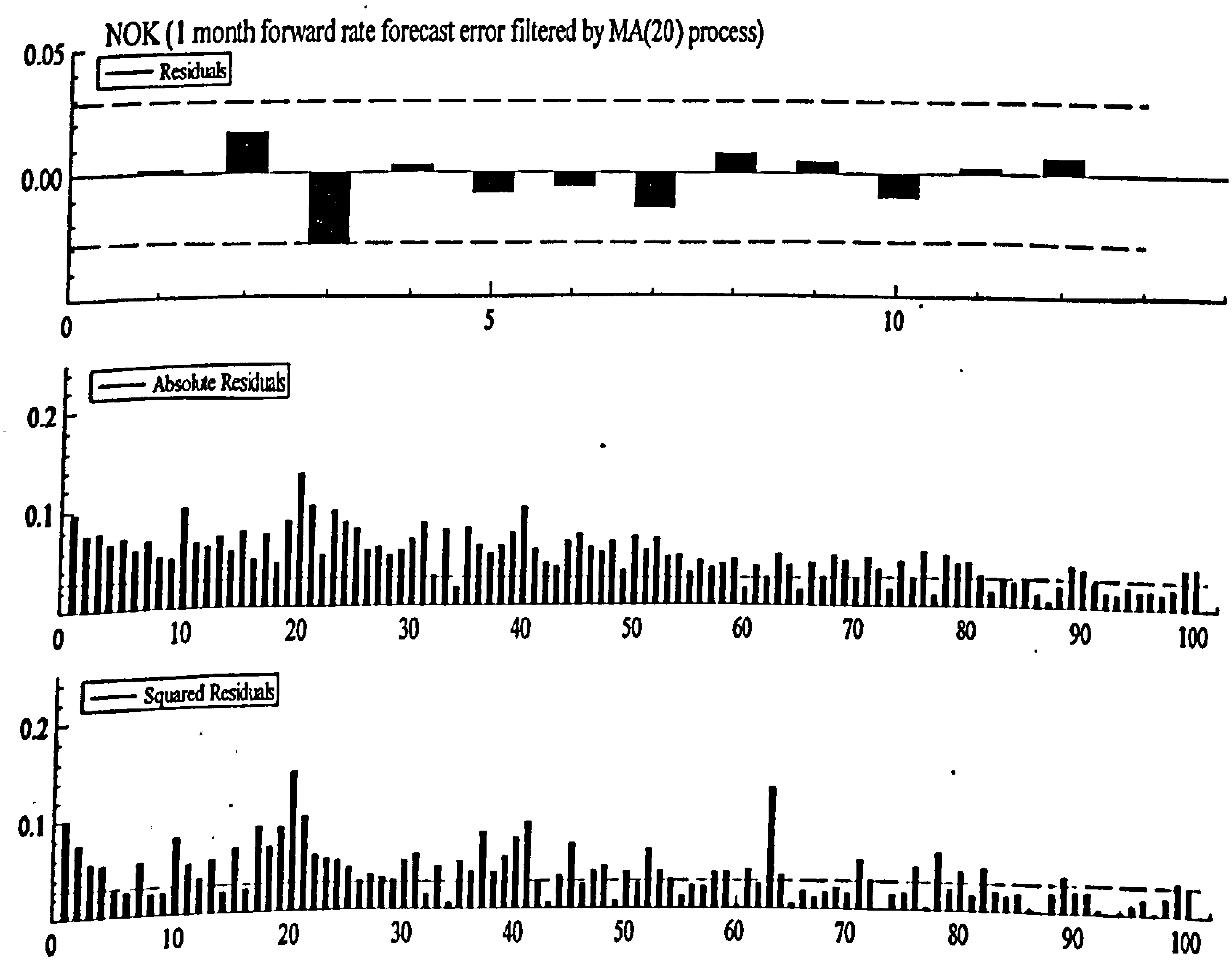
Figures 3.7F: GBP1M



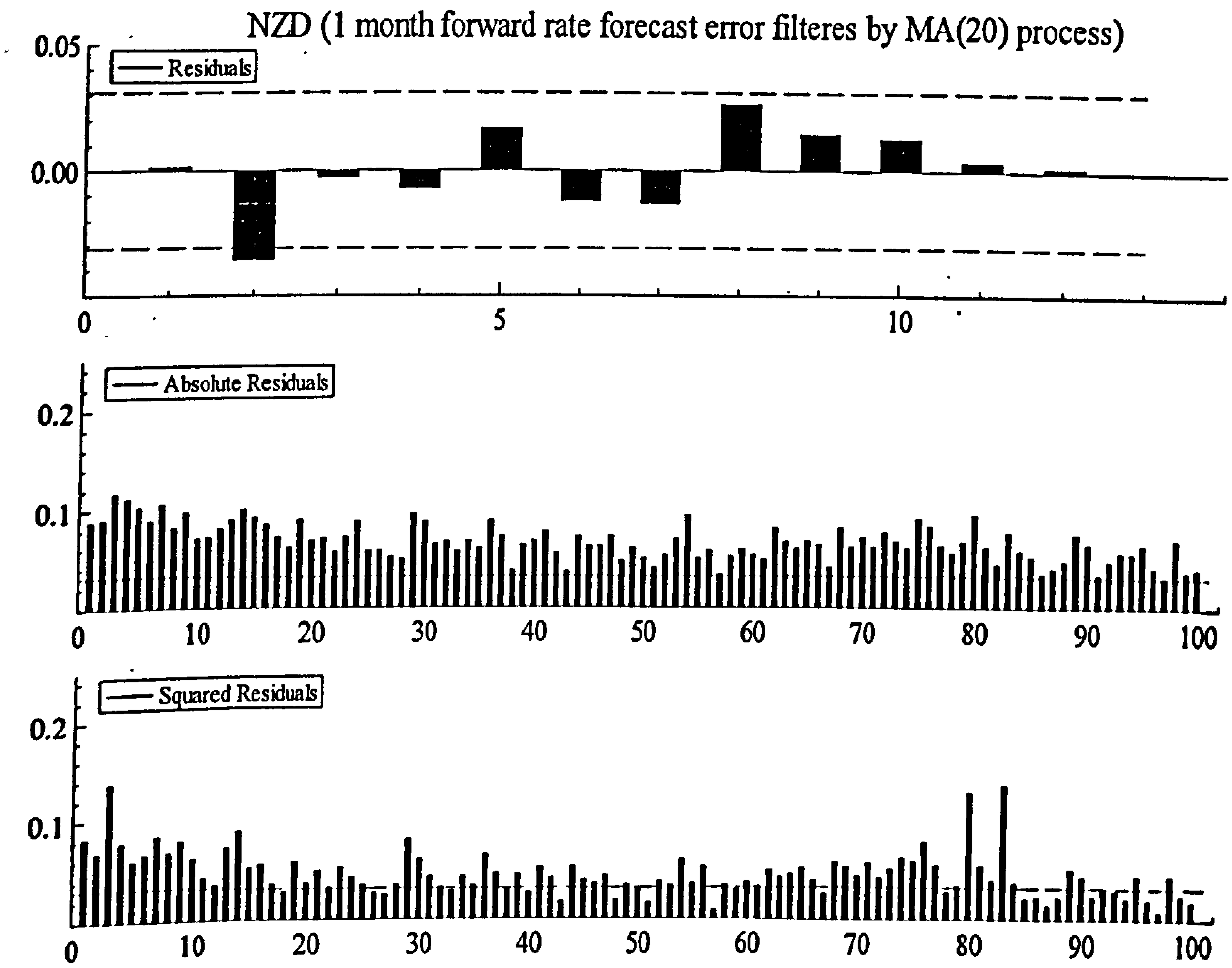
Figures 3.7G: JPY1M



Figures 3.7H: NOK1M



Figures 3.7I: NZD1M



Figures 3.7J: SEK1M

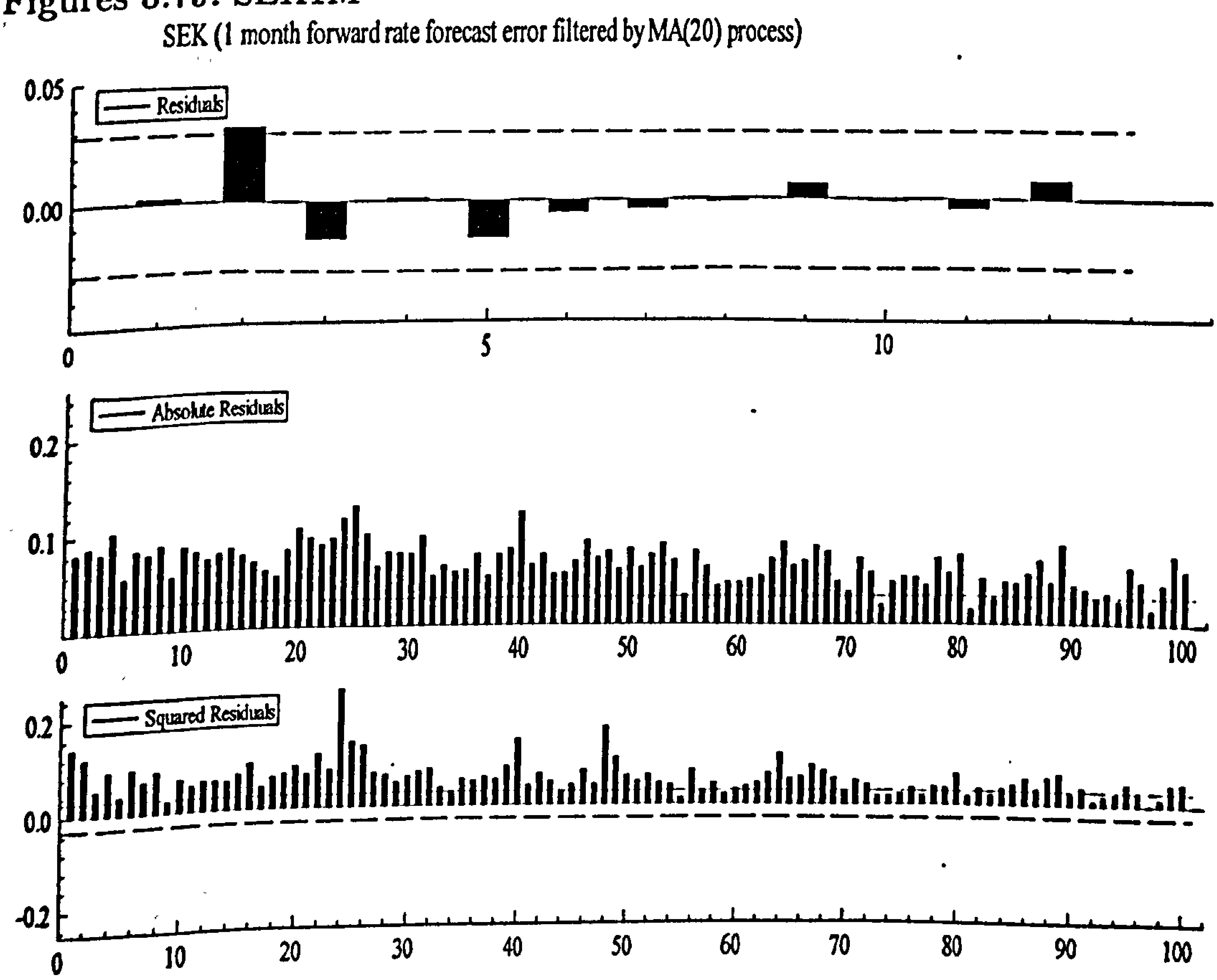


Figure 3.8: The rolling regression
Figure 3.8A: AUD1M

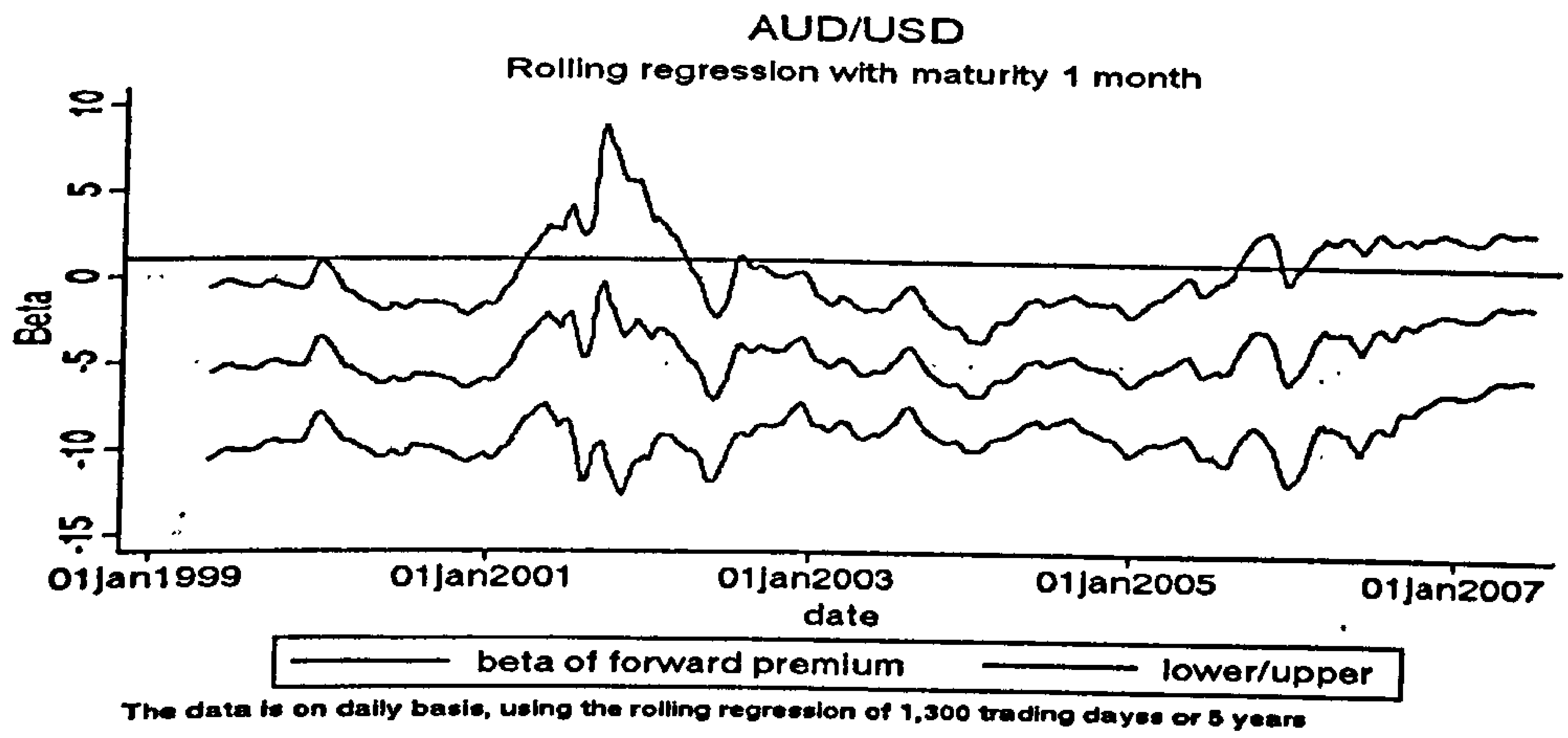


Figure 3.8B: CAD1M

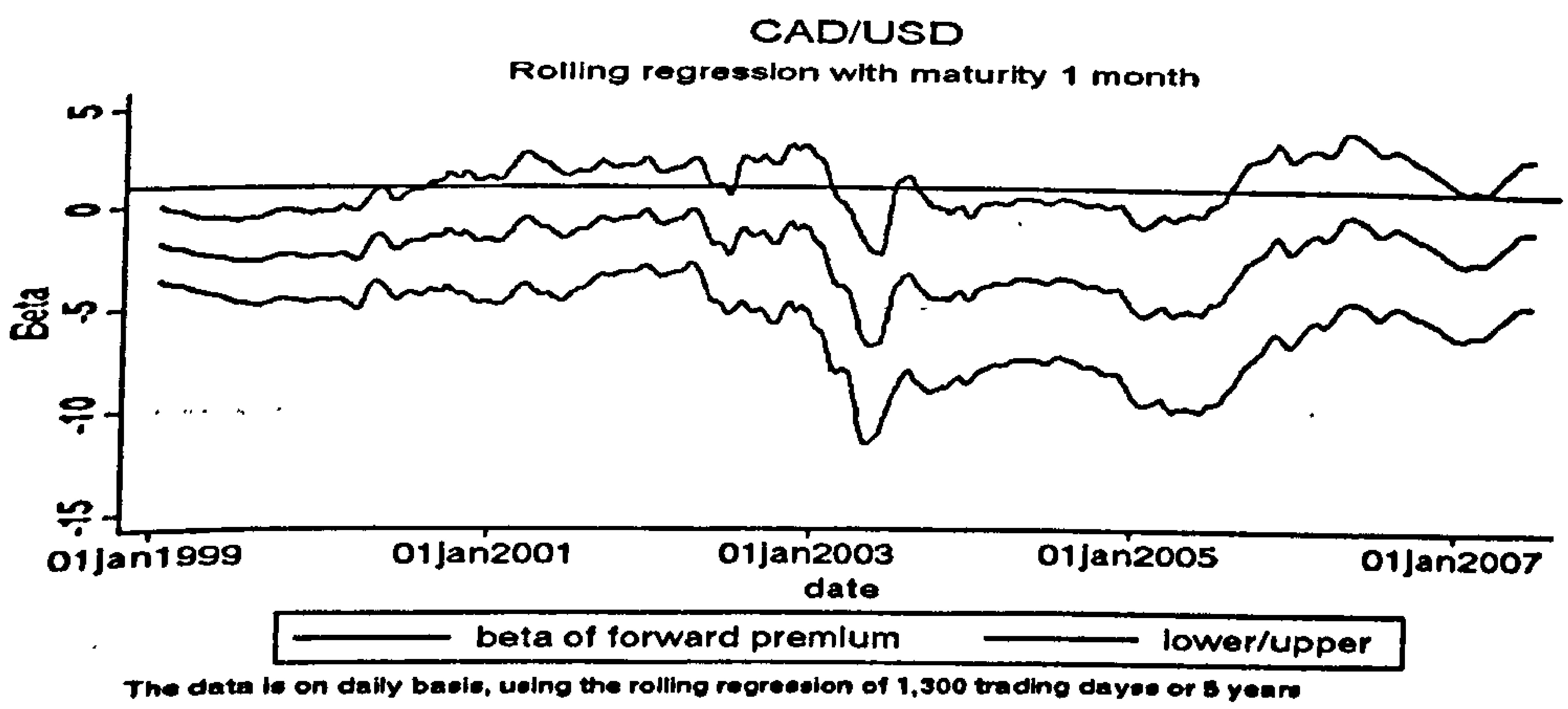


Figure 3.8C: CHF1M

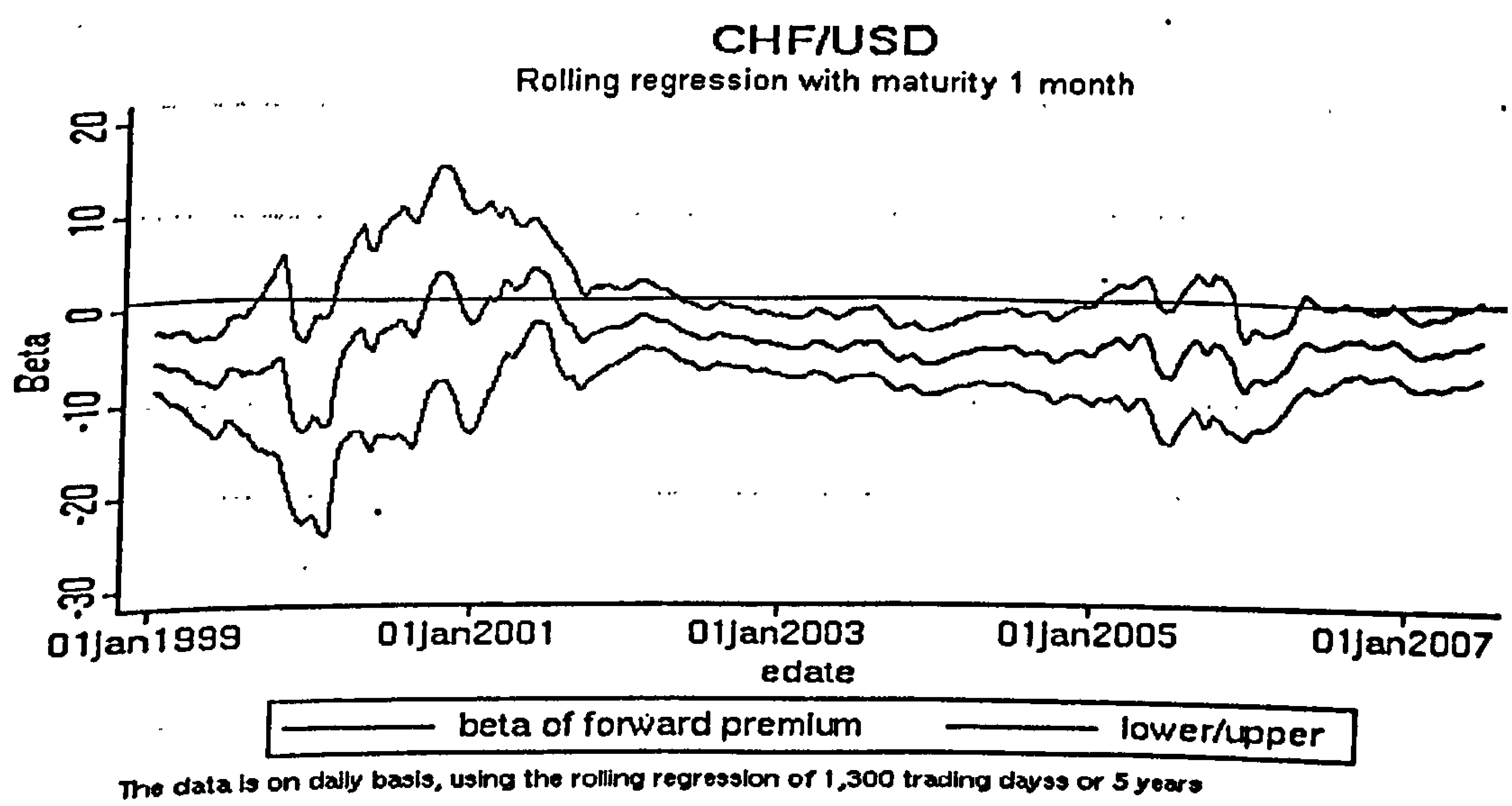


Figure 3.8D: DKK1M

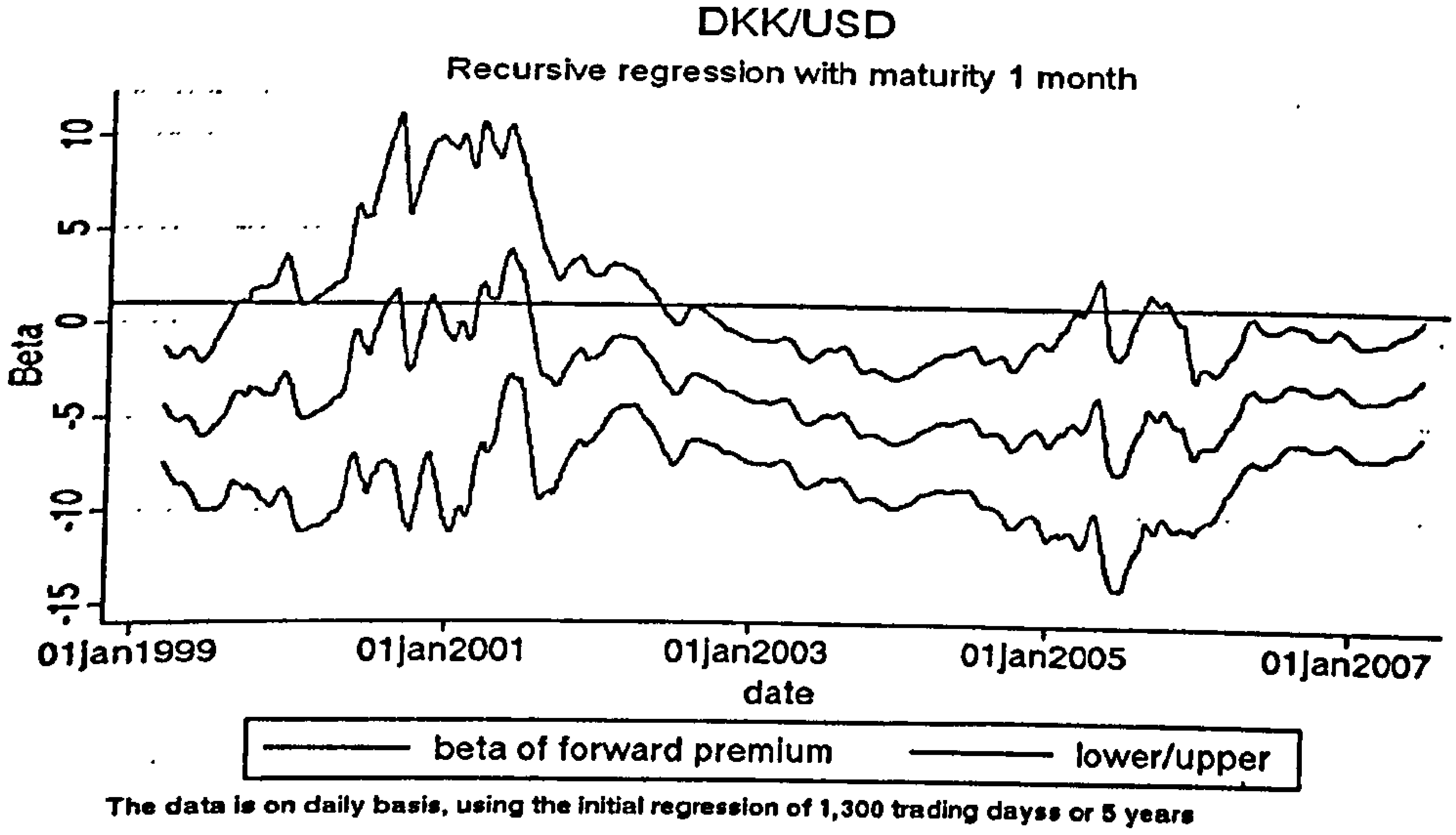


Figure 3.8E: EUR1M

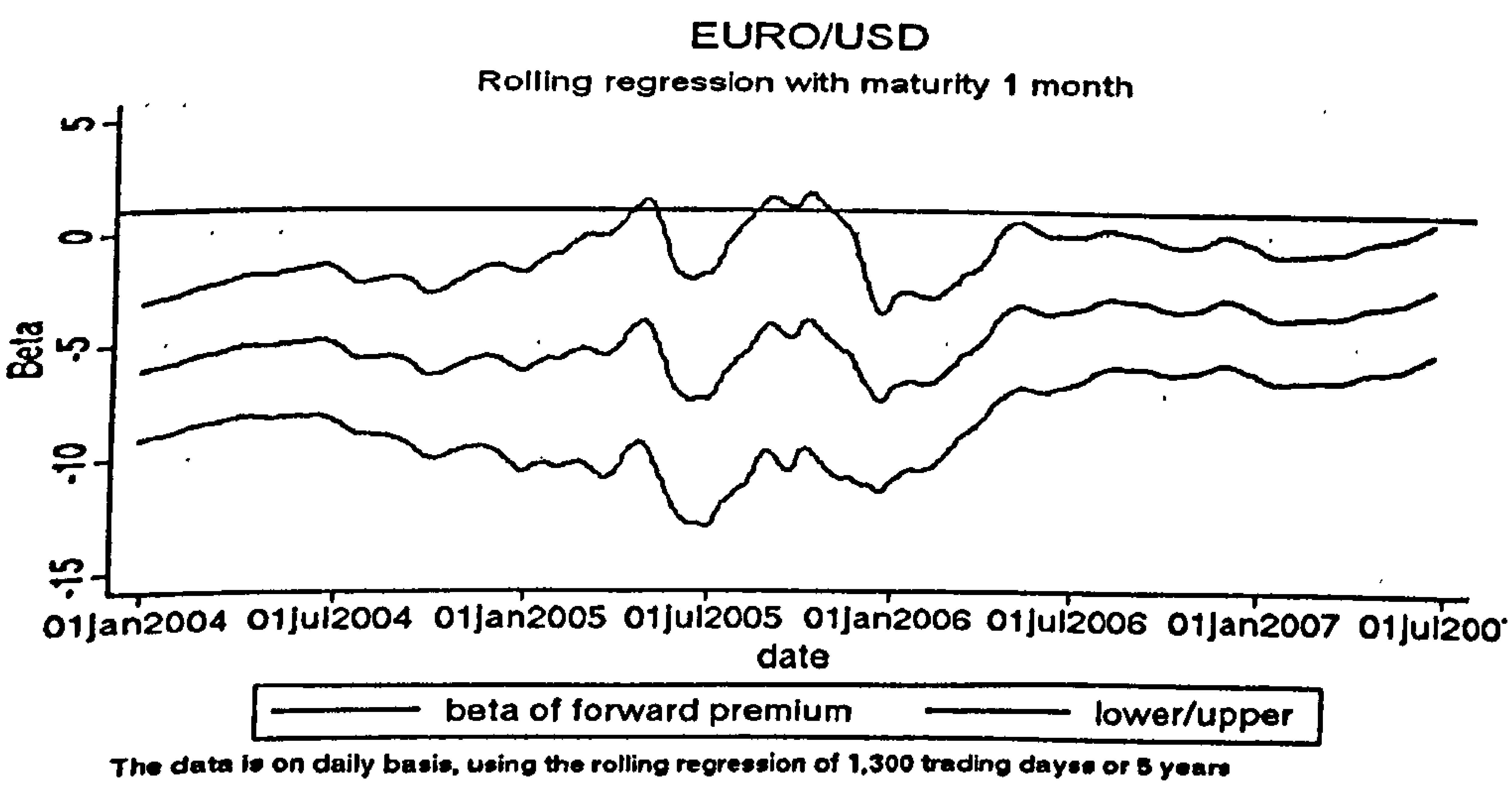


Figure 3.8F: GBP1M

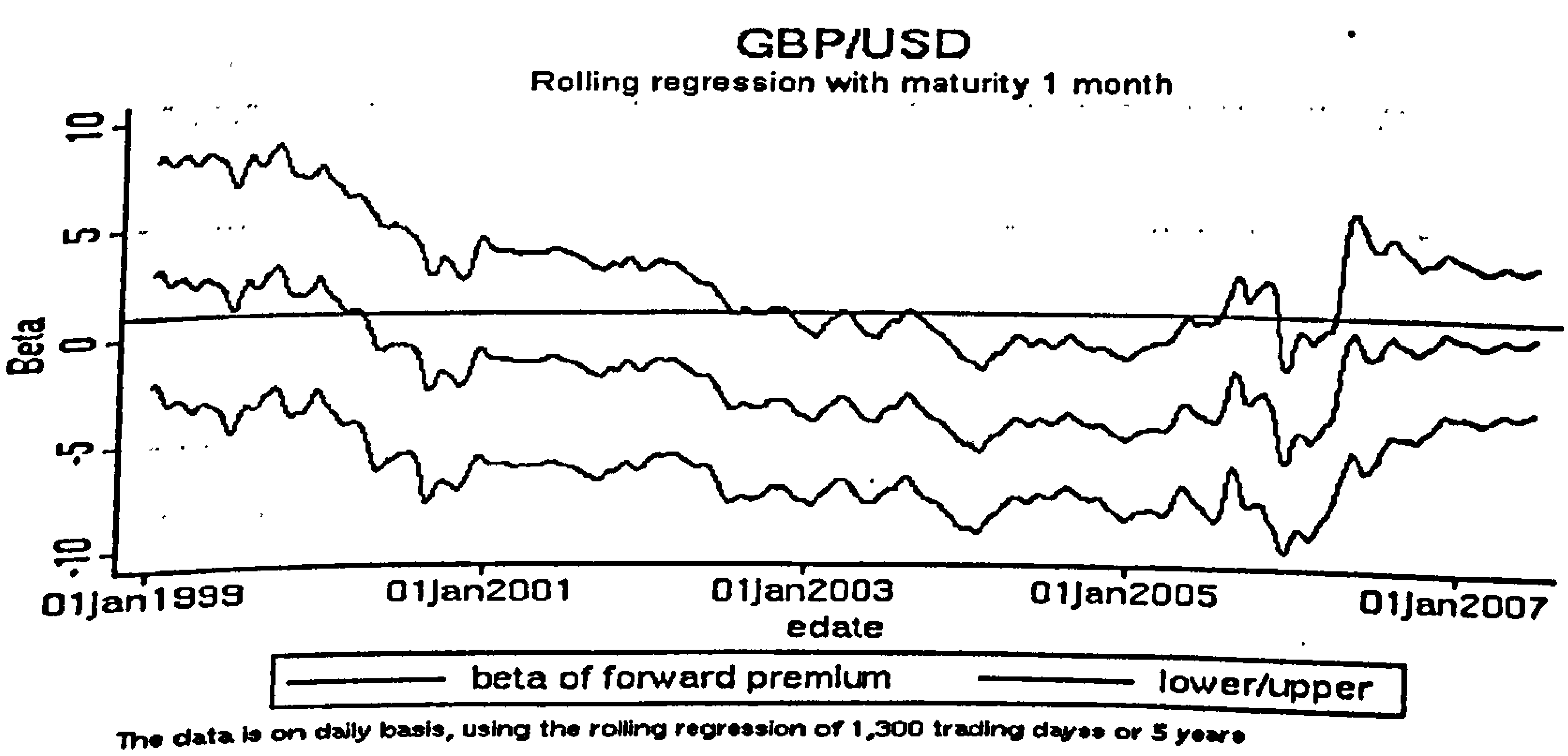


Figure 3.8G: JPY1M

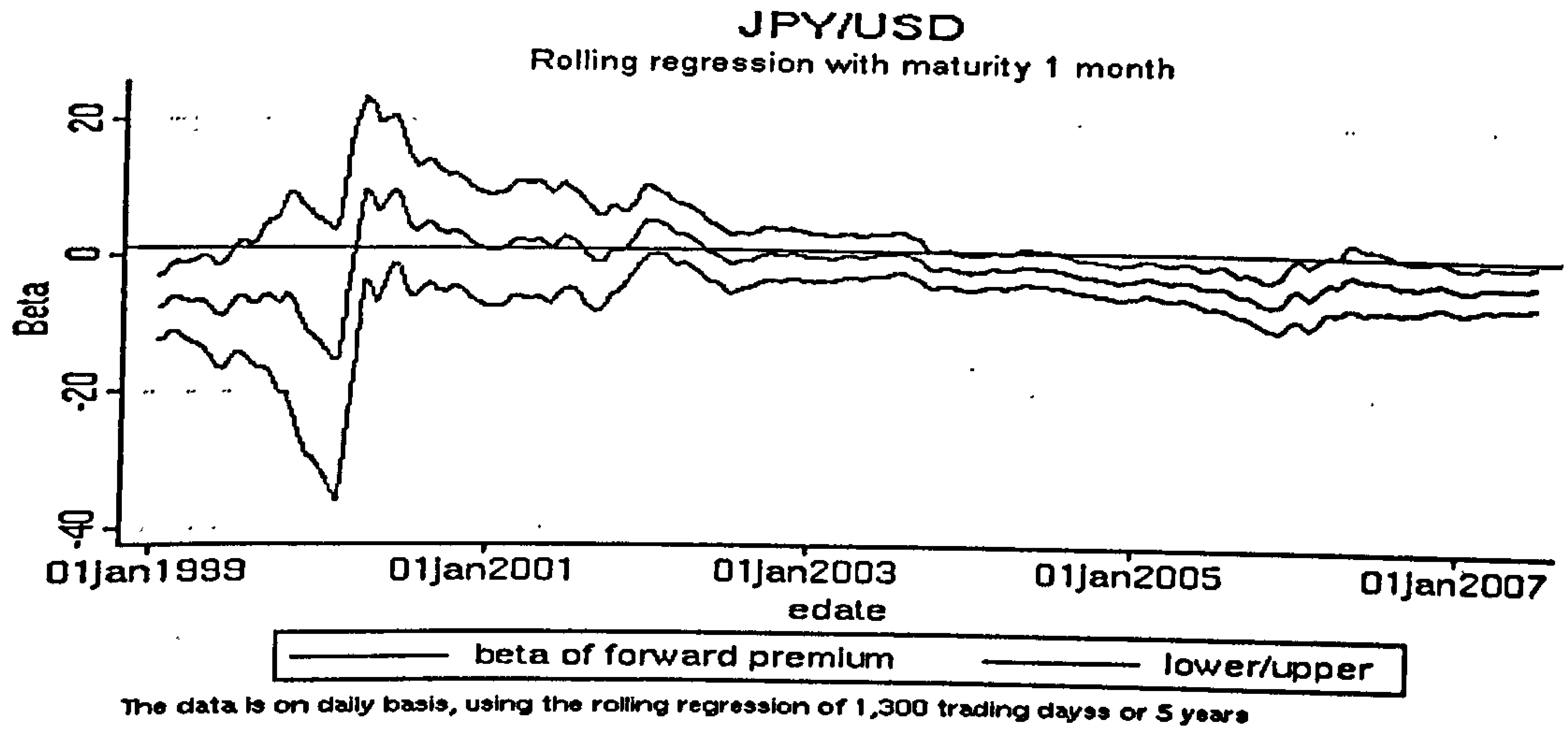


Figure 3.8H: NOK1M

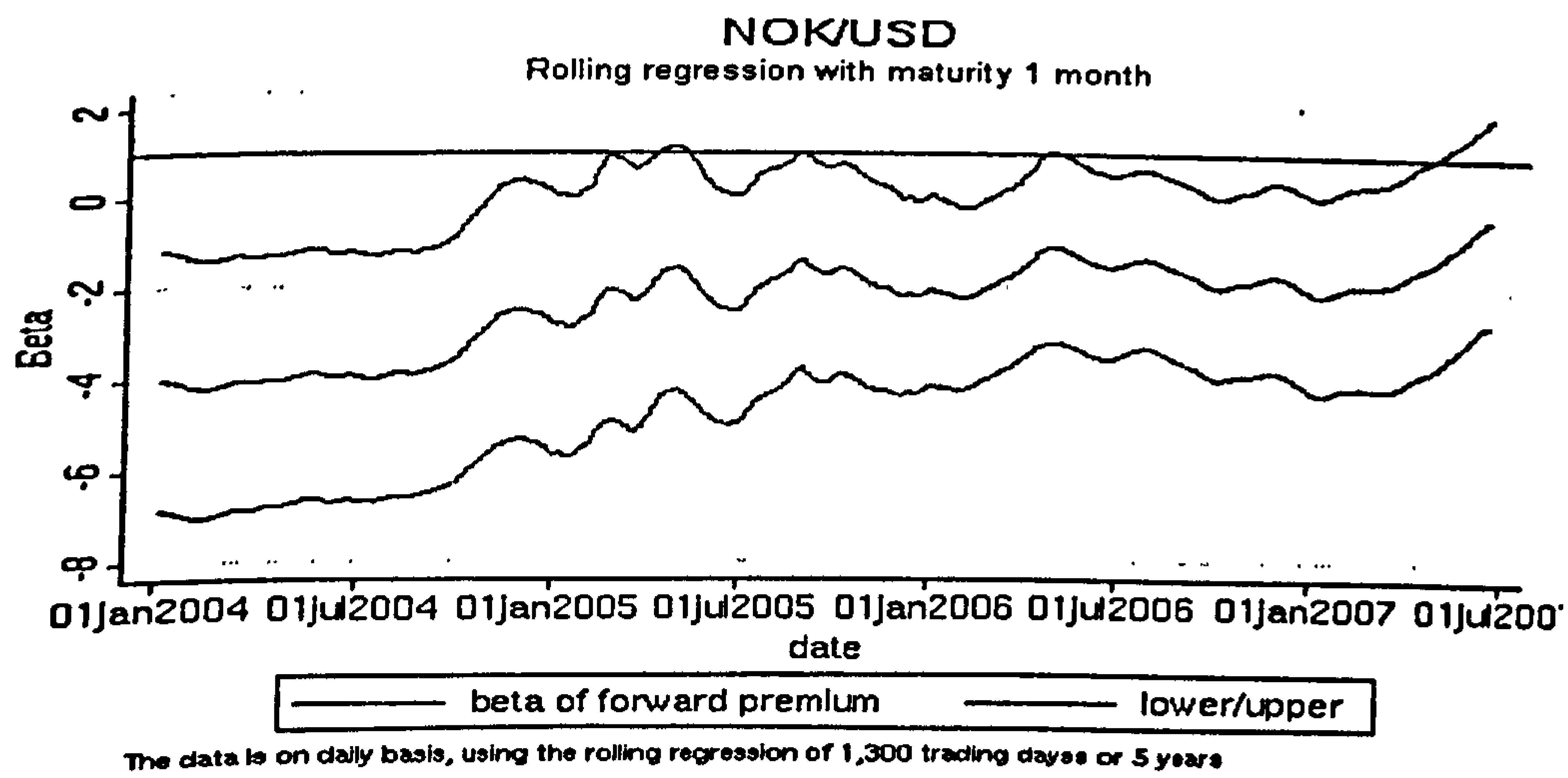


Figure 3.8I: NZD1M

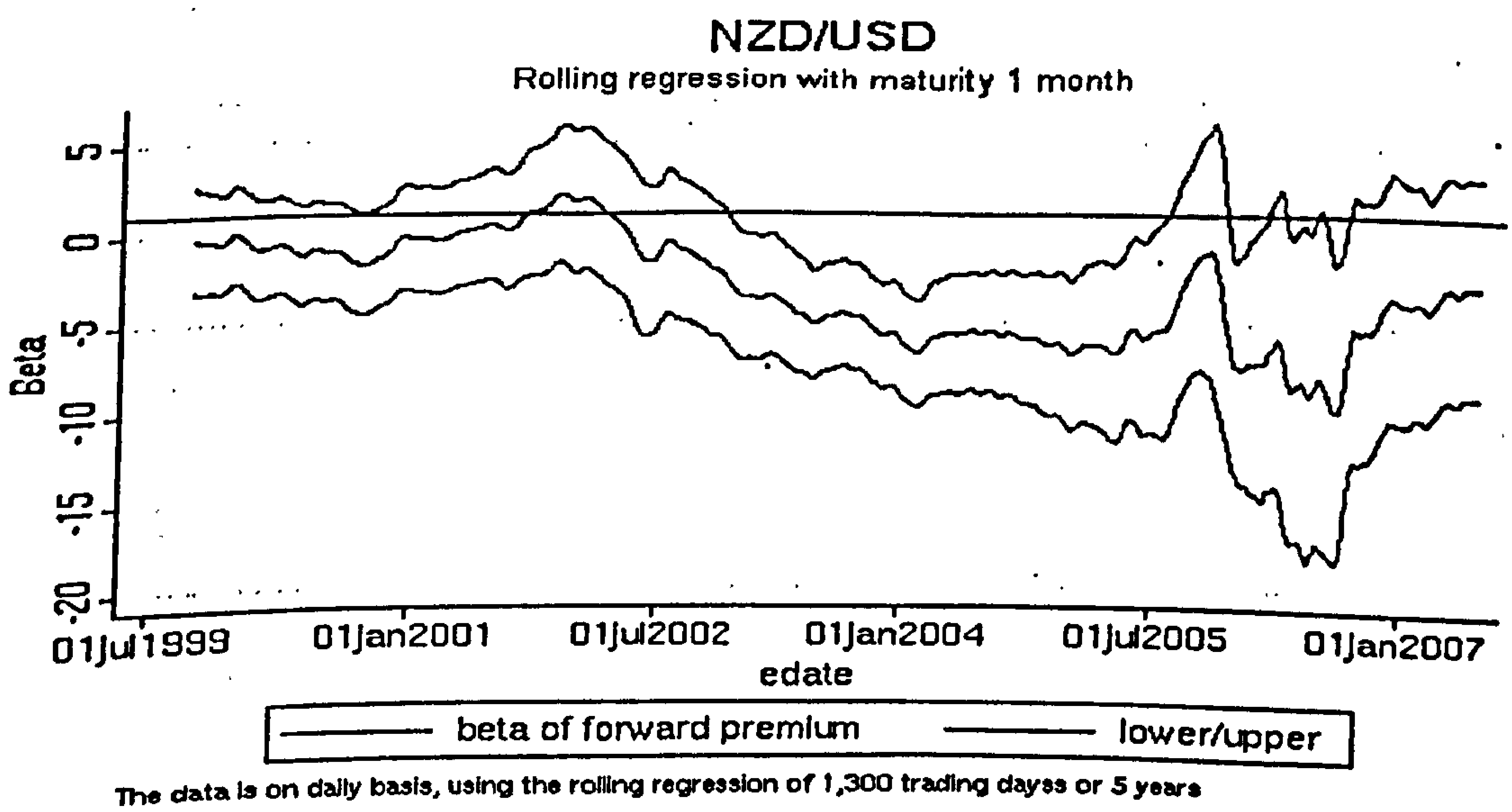


Figure 3.8.J: SEK1M

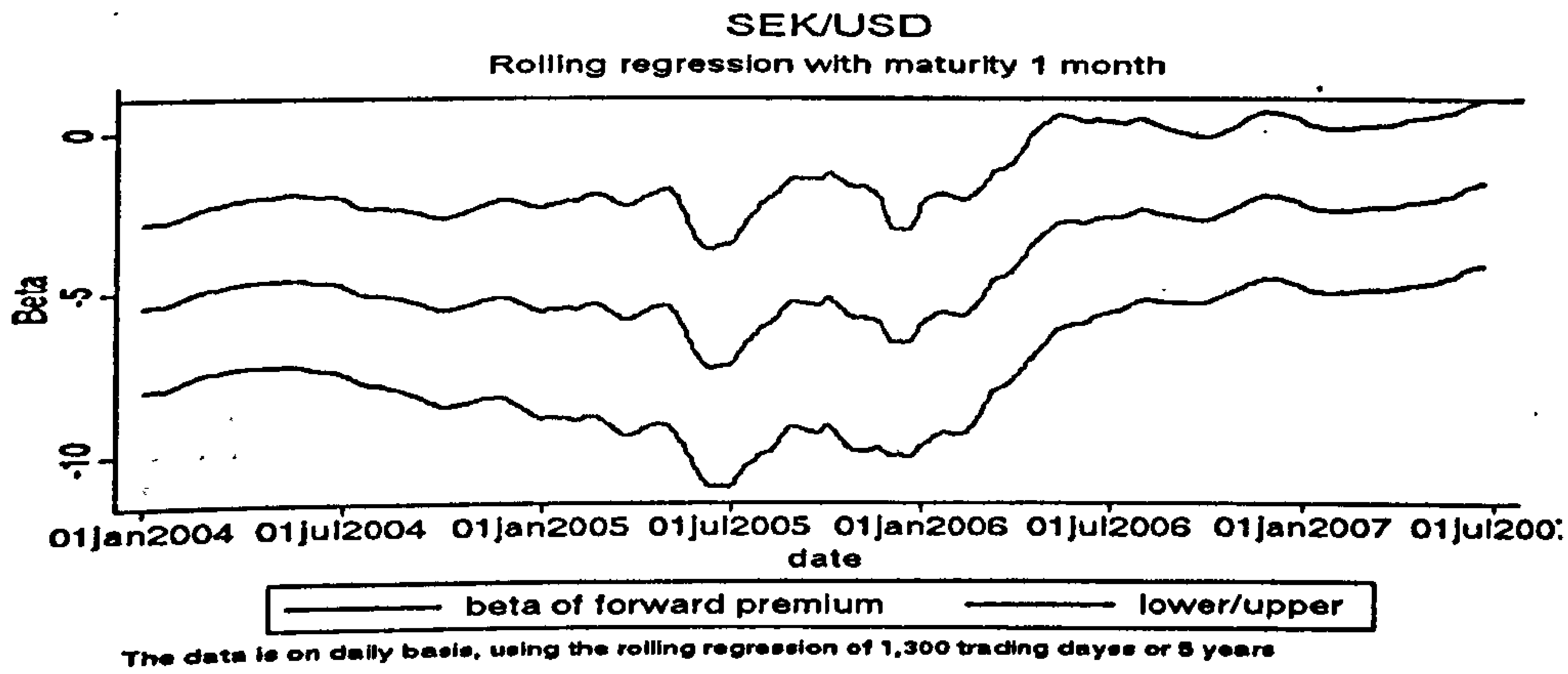


Figure 3.9: Forward premium series
Figure 3.9A: AUD1M

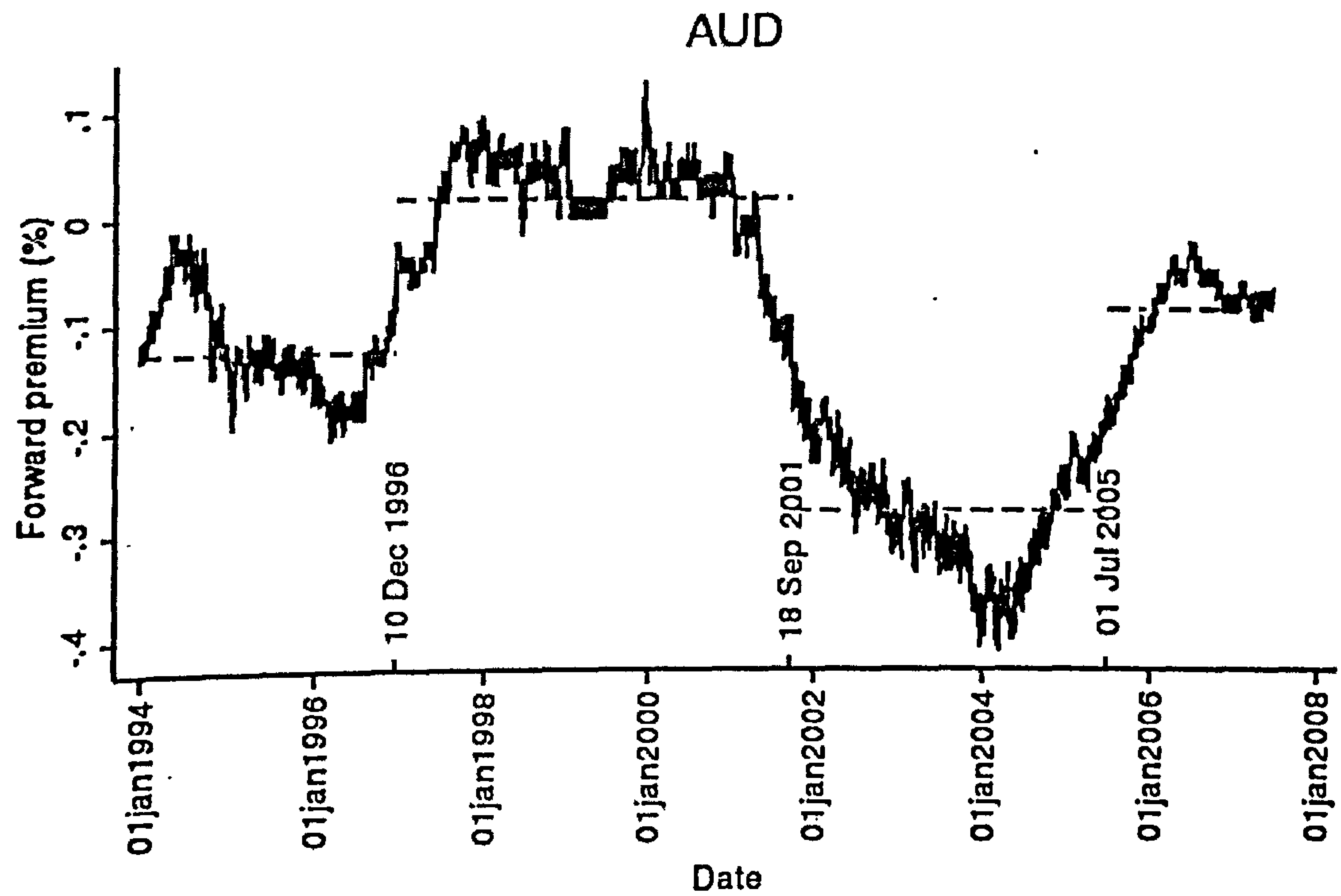


Figure 3.9B: CAD1M

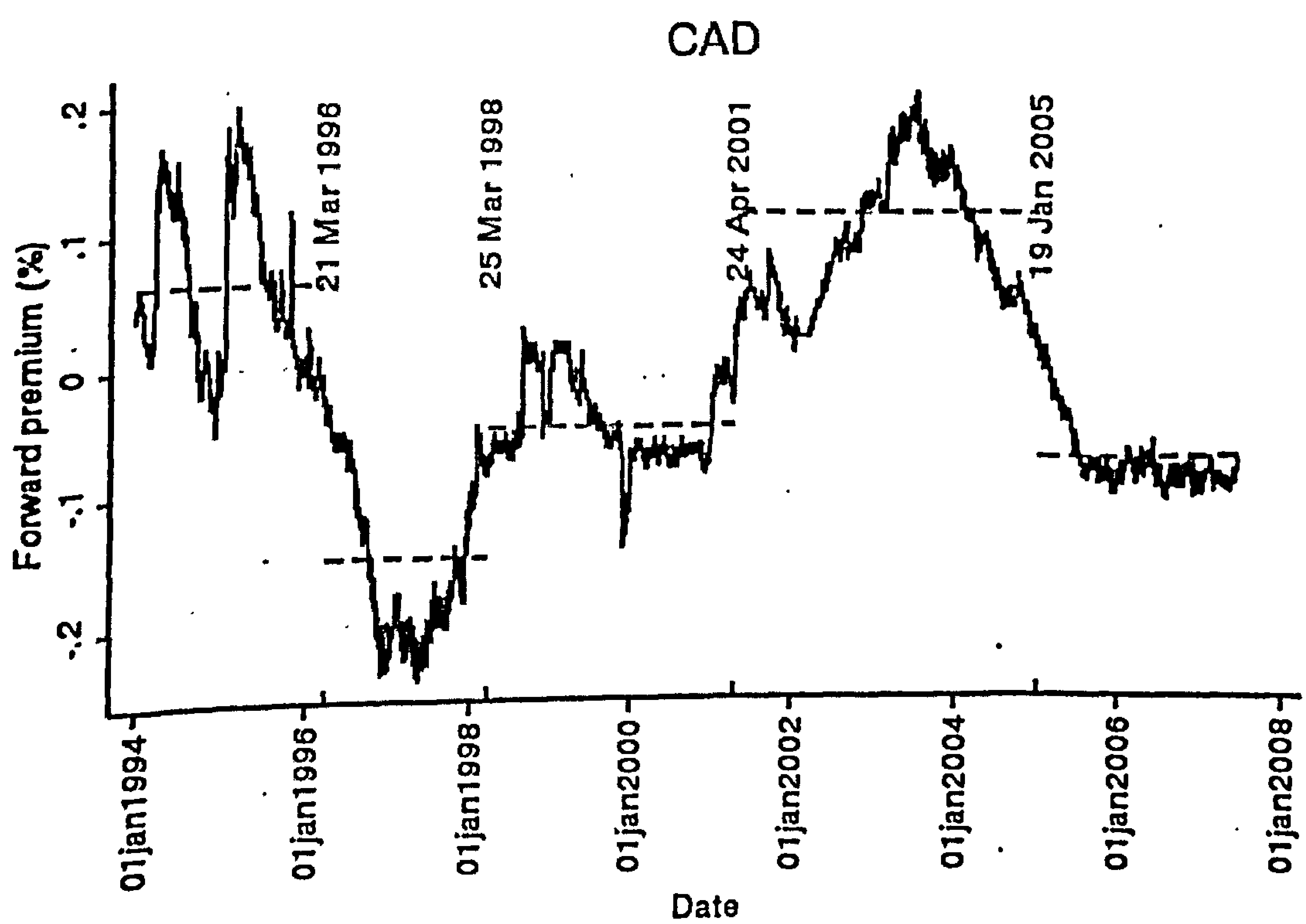


Figure 3.9C: CHF1M

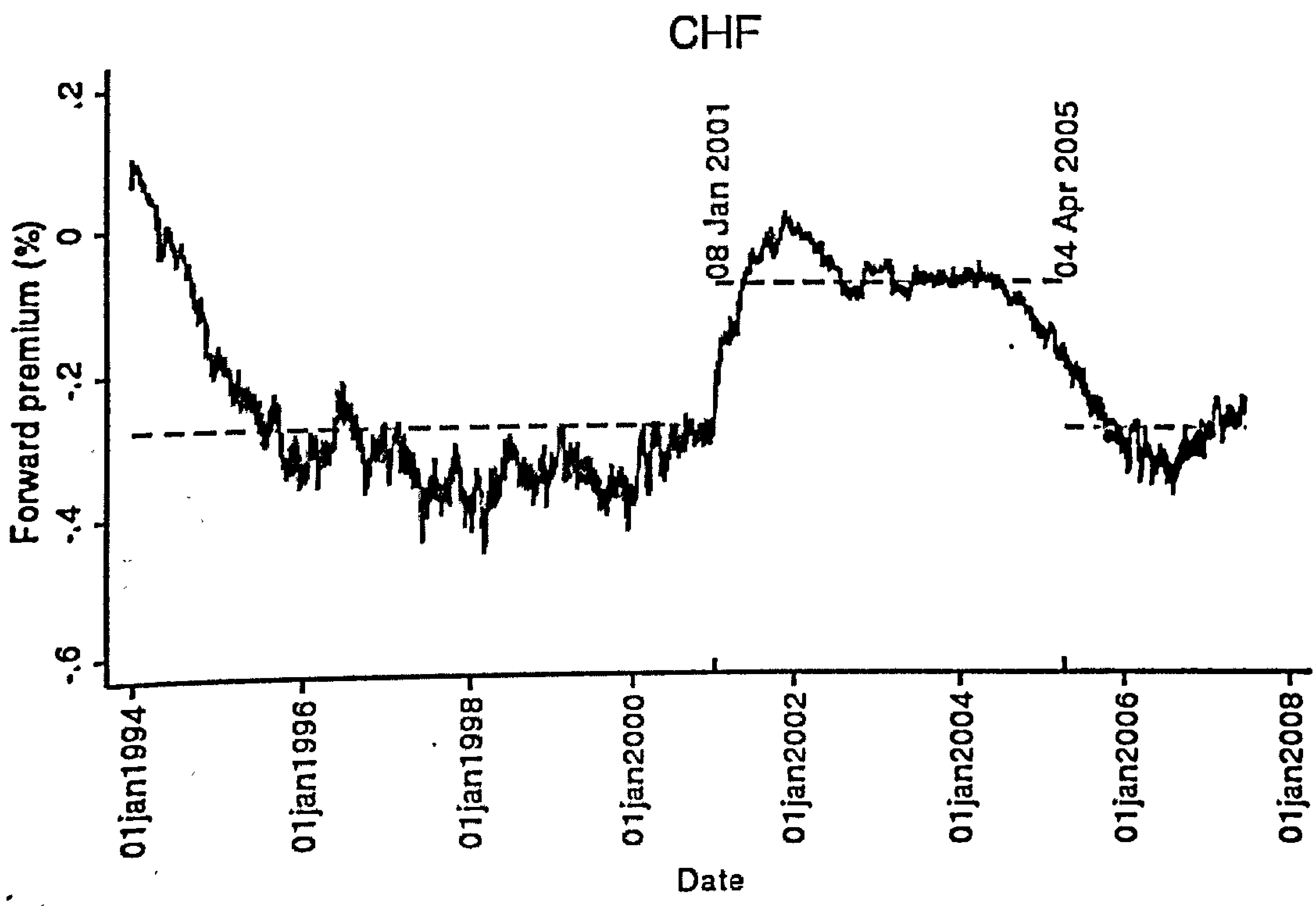


Figure 3.9D: DKK1M

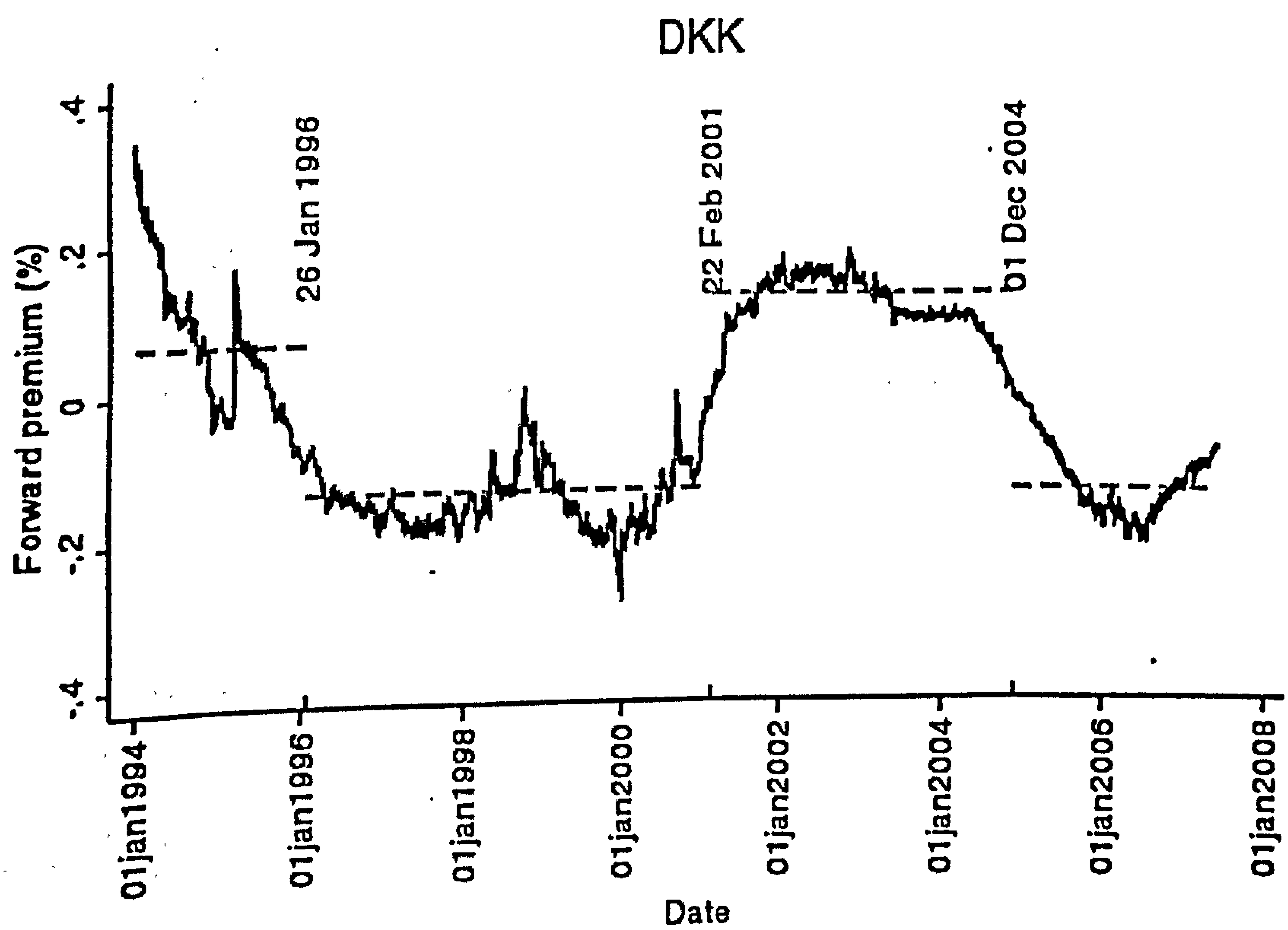


Figure 3.9E: EUR1M

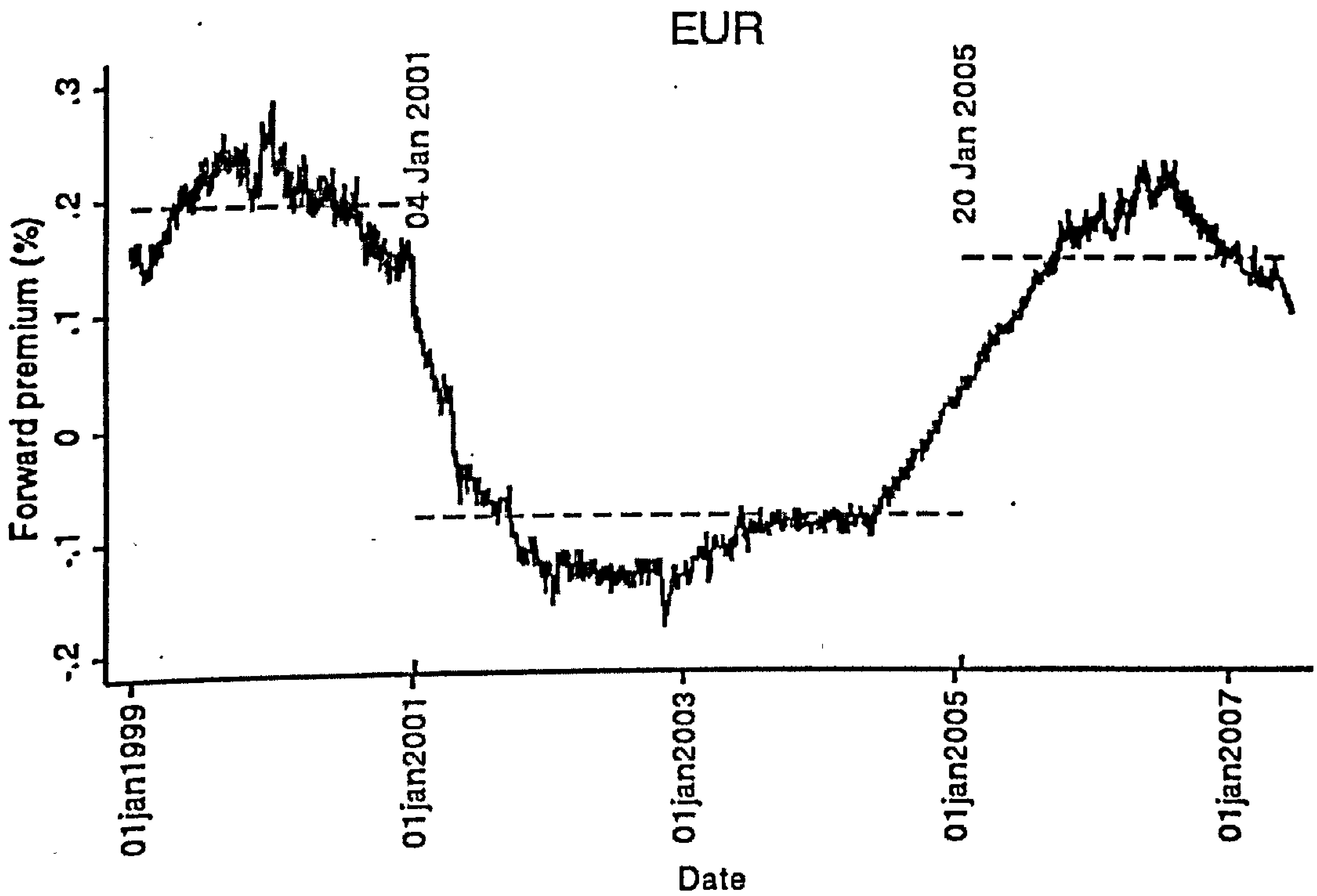


Figure 3.9F: GBP1M

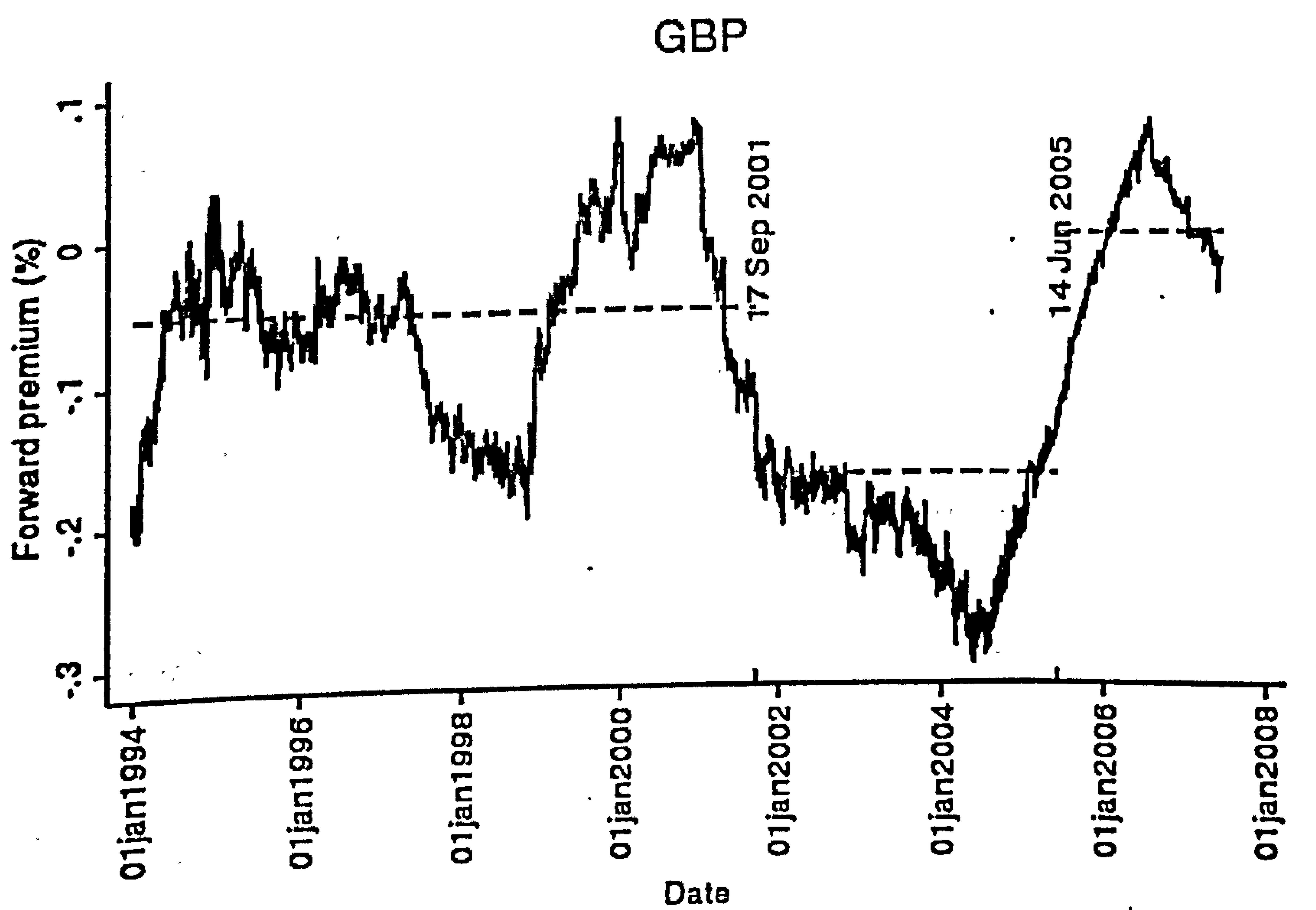


Figure 3.9G: JPY1M

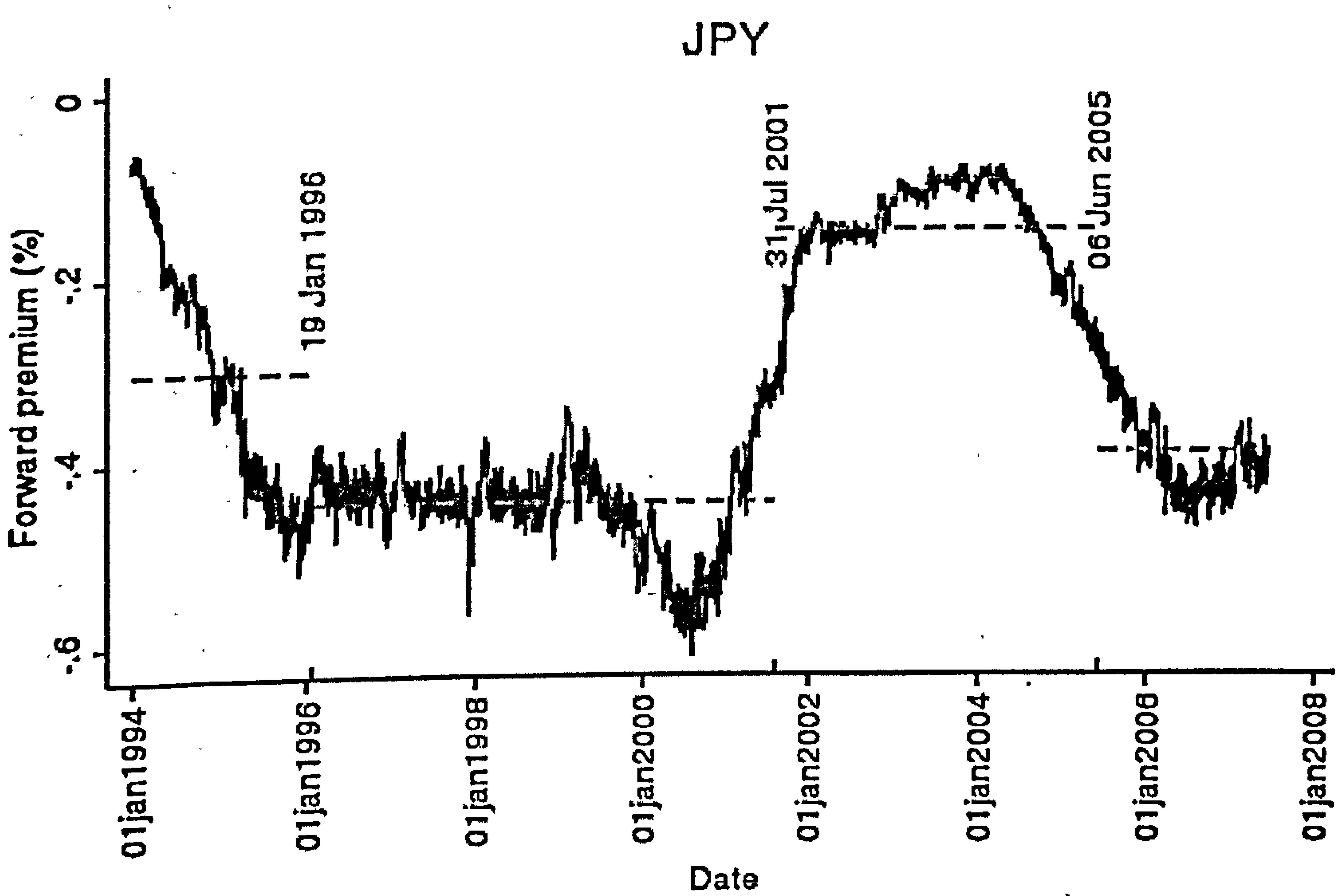


Figure 3.9H: NOK1M

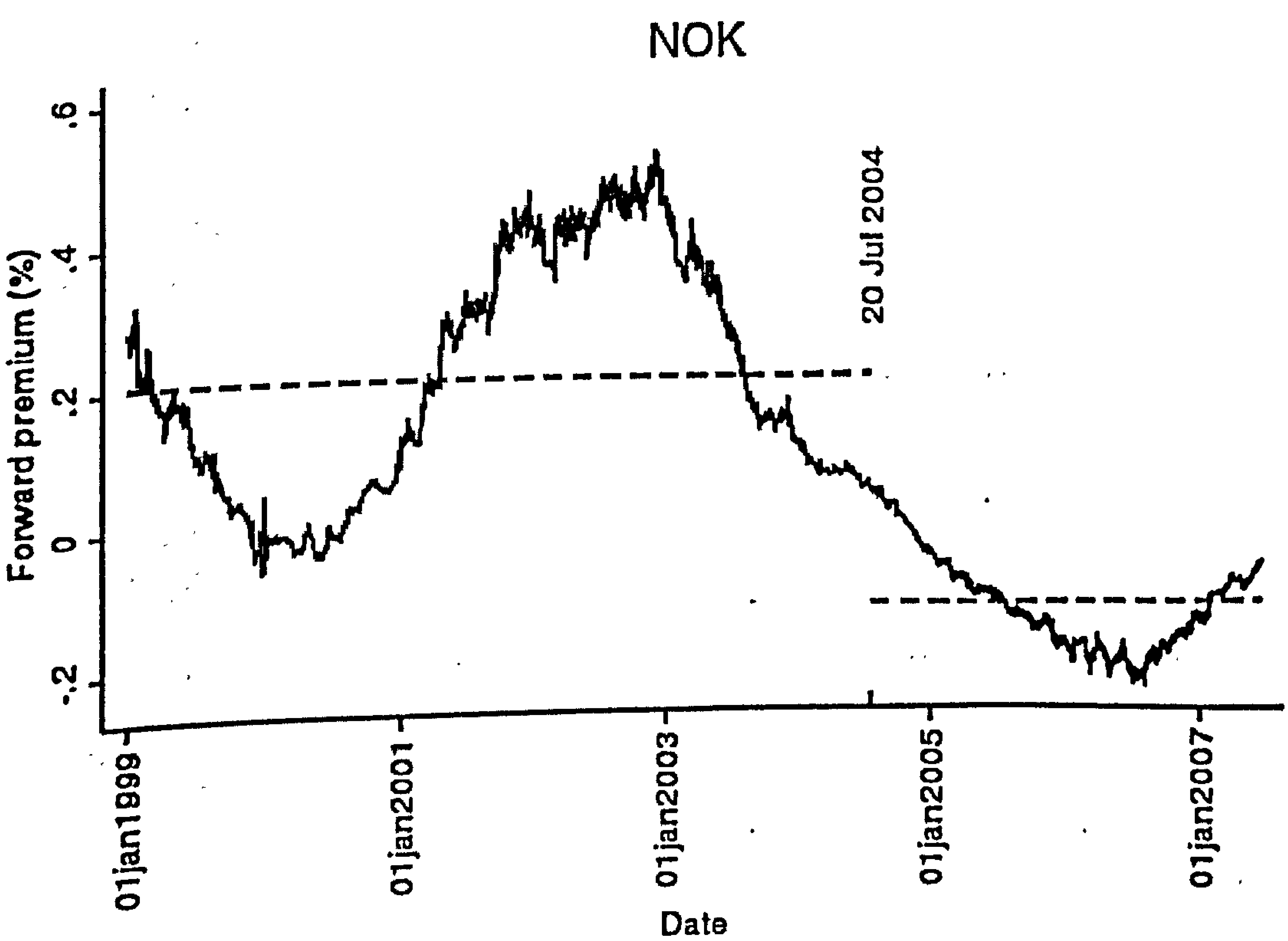


Figure 3.9I: NZD1M

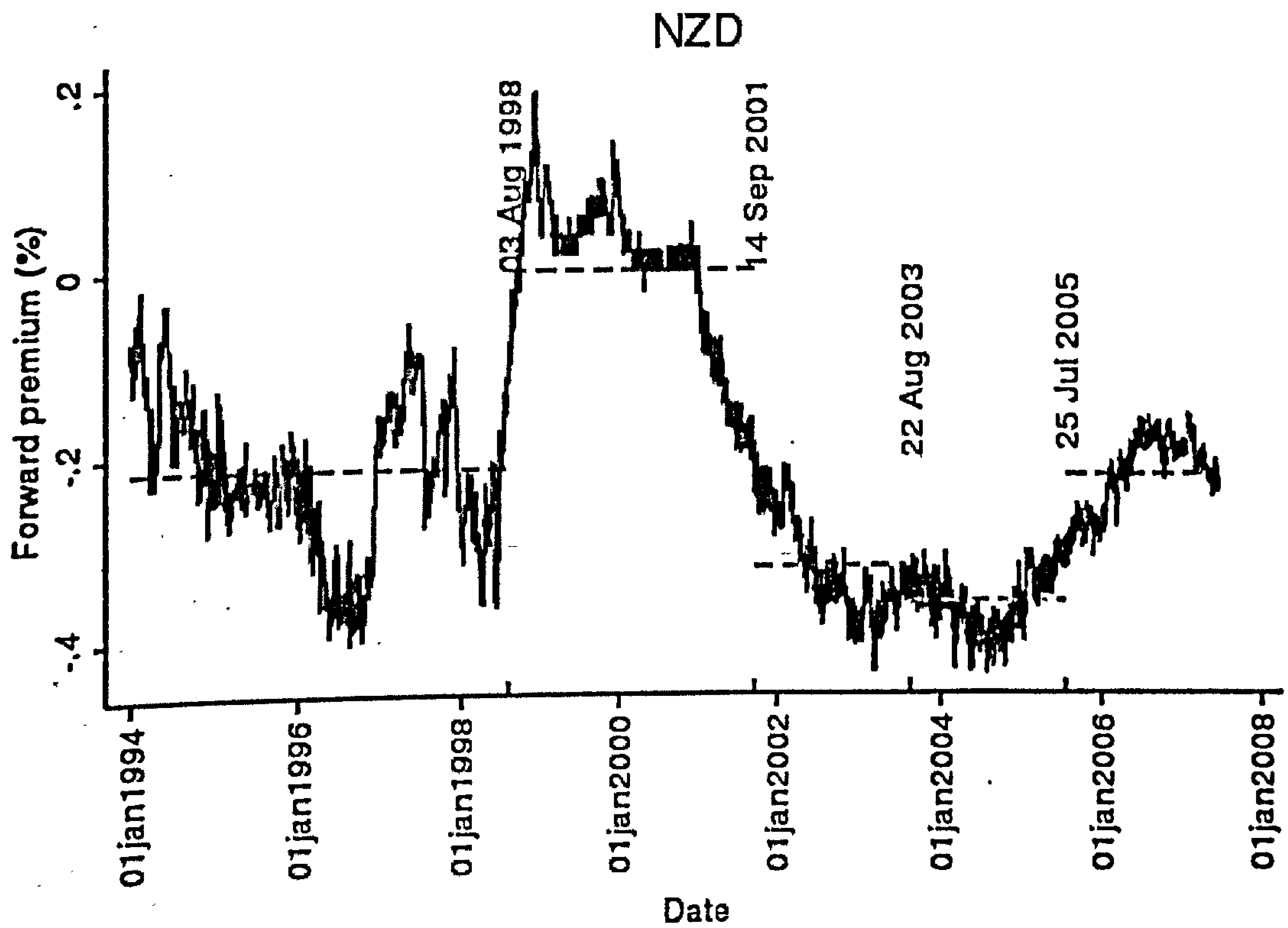
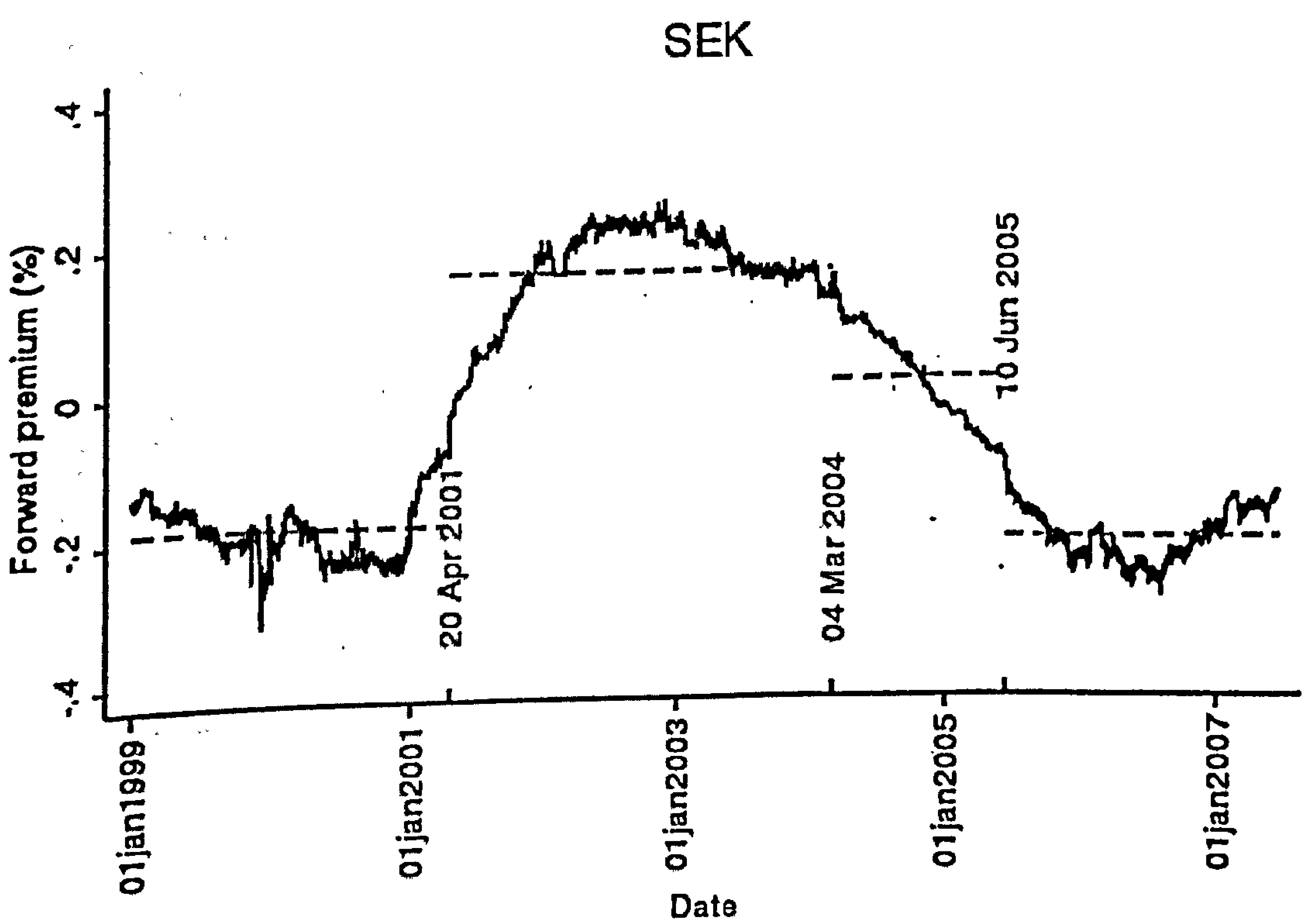


Figure 3.9J: SEK1M



Chapter 4
Funded and PAYG Pensions when Annuities are backed by Bonds[114]

4.1 Introduction

The ageing population in the UK and elsewhere has recently made pension finance an issue of great general interest. There are 2 main types of pensions, classified by the way they are funded. The first type is the fully funded scheme which is based on the savings of members. The contributions are invested in financial assets. The second type is the Pay As You Go (PAYG) scheme which is usually funded by the state. The state can either tax or issue debt to the current working generation to pay the pensions of the currently retired. According to a two period overlapping generations (OLG) model, Diamond (1965) suggests that a fully funded system provides better incentives to save and results in higher capital and is therefore superior. On the other hand, the state pension reduces individual's utility in the long run since it creates additional reductions in the productive capital stock arising from the substitution of government debt for physical capital in individual portfolios.

Traditionally, the principal function of the pension system is as a means for consumption smoothing (Samuelson, 1975 and Diamond, 1965). In this chapter, we argue that pensions take a role as a mechanism for both consumption smoothing *and* as a means of insurance. From the perspective of consumption smoothing, people maximise their well

114 This chapter is joint work with Edmund Cannon.

being over time, thus people will transfer consumption from their productive middle age to their retirement. From the perspective of insurance aspect, in a world of uncertainty, agents need to insure against the risk of outliving their pension savings and the risk of assets defaulting.

The contribution of this chapter is motivated by the observation that government debt is both a complement and a substitute to physical capital. We observe that ownership of physical capital is mediated largely through equity, which represent high risk high return financial assets, whereas government debt is mediated through bonds, which can insure against long term risk. So although government debt and physical capital compete for funds (suggesting they are substitutes), the financial assets which results are quite different (suggesting complementarity).

Empirically, the UK pension system can be approximately characterised by the two forms of assets being complements rather than substitutes. Agents choose to hold equity, and hence physical capital, during the first phase of their life, while accumulating their pension fund. During their retirement, while decumulating their pension, they hold an annuity, backed by government debt, since this is virtually risk free over the period of retirement. The UK pension system will be discussed in the next section (section 4.2) followed by review of the related academic literature (section 4.3).

We then turn to a small general equilibrium model of an economy where agents behave in this fashion and determine the optimal mixture of PAYG state pension and funded private pensions. Since funded pensions rely upon the supply of government debt to facilitate a functioning annuity market[115], the government influences both the state and private pensions, and since returns on government debt must be financed from taxation, taxes will also contribute to *both* sorts of pension. Section 4.4 explains the structure of the model and section 4.5 presents a simulation study. It generates 2 main analyses. First, it examines the optimal combinations of the state PAYG pension and the funded pension from the perspective of the steady state. Second, it studies the effects of two policy shocks which are a sudden reduction in national debt and the baby boom.

Lastly, we discuss two topical issues which face the UK at the moment: first, how to respond to the large bulge in the population as the baby-boom generation retires; second, the possible effects of cuts in government debt

4.2 The pension system in the UK

Before describing the UK pension system, we start with the observation that private physical capital is largely held through equity, while government debt is held in bonds. Equity and bonds are different sorts of financial instruments. We argue that a key reason for this difference is the time scales over which firms and nation states exist, although other factors such as incomplete information and contracts also play a part. In general, we observe that private firms are insufficiently long-lived to credibly provide long-dated debt and that cor-

115 In theory, pension liabilities are bond-like so pension funds could match their liabilities by switching out of shares and moving into bonds. Historically, British pension funds were in favour of equity for three main reasons. First, equities on average yield higher returns, firms that pay for pension schemes can pay less into the fund. The market will do part of the work for them. Second, it was assumed that over the long run, profits, dividends and share prices would rise at least in line with inflation, as would wages (and therefore pension payments). Third, equities may be volatile but they have unusually returned more than bonds over the long run. Pension funds could afford to be patient and collect this premium. However, in the late 1990s, many researchers argue that if pension liabilities are bond-like, then buying equities is a mismatch (Exley-Metha-Smith, 1997).

porate bonds rarely have terms to maturity much greater than five years, whereas the UK government provides bonds of 30 years maturity and in the past has sold debt with no date of redemption at all ("consols"), limited supplies of which still exist[117]. If it is true that government debt and private equity are not in direct competition, then we need to reconsider whether government debt will merely leads to crowding out.

The discussion is within the context of the United Kingdom. The reason for this is that the UK pension system has functioned successfully for a long time, and consist of both a minimal state pension system as well as a system of voluntary private funded pensions. As a result, it has a large private pension sector and the largest annuity market in the world. It is one of relatively few countries that can provide evidence for how a funded pension system operates in practice. It is also interesting to see how it will respond to the demographic challenges such as the retirement of the baby boom generation and the increased longevity.

There are three tiers of pension provision in operation in the UK. The first and second tiers are unfunded and compulsory while the third tier is provided by the private sector. The first two tiers represent the state pension, which is payable at 65 for men and 60 for women (UK pension commission, 2006).

The first tier is the Basic State Pension (BSP) which is "unfunded" and pays a flat-rate pension. This scheme has no underlying fund of assets and in principle provides a means of subsistence, arguably this is intended as a form of poverty alleviation. This scheme is a pay-as-you-go pension which represents intergenerational transfer between the working and the

117 The last issue was 3.5% war stock. Infinitely lived bonds such as this are a useful form of asset to back a pension, since the flow of payments is constant, whereas a conventional finite term bond has too large a payment at the longest maturity (when the probability of needing to make a pension payment is least).

retired population. The membership of this scheme is compulsory and the contributions are collected through the national insurance system.

The second tier is the Additional State Pension, otherwise known as the State second pension (S2P)[118]. This scheme is also unfunded, but pays a defined benefit pension which is related to average earnings over the employee's life. Membership is compulsory for all employees (but not the self-employed) unless the employee has contracted out into a private pension scheme in which case they are exempt.

The last tier is voluntary private pension provision, of which there are two types: occupational and personal pension schemes. Contributions into these schemes are to an extent subsidised by the government through tax-breaks. Occupational pension schemes are usually "funded" and require contributions throughout the employee's working life. These schemes are provided by employers and may pay on a "defined benefit" or a "defined contribution" basis. The former usually are defined in terms of some proportion of final year earnings, and are related to the number of years of employment (Tonks, 1999). Defined contribution (or money purchase) pensions are always funded and convert the value of the pension fund at retirement into an annuity. On the other hand, all personal pensions are defined contribution, and often lack any contribution by the employer. There are two distinct types, group schemes organised on a company basis, thus benefiting from low commissions and some pooling of annuity risk, and individual arrangements with life insurance companies.

118 The State Earning Related Pension Scheme (SERPS) was replaced in April 2002 with the 'State Second Pension' which is designed to give more to the lower paid and middle earners, carers and the long-term disabled with broken work records. Whereas with SERPS, the more you earn, the higher your pension, S2P operate a flat rate which means that high earners will be better off opting for private pension schemes.

The UK pension system is traditionally seen as offering a good example to other countries, having features such as low social security pension expenditures as well as a high coverage of well-financed voluntary private schemes. But recently it has been suggested that the UK pension system is less robust than previously believed. The state is currently planning to play a reduced role in pension provision because of demographic pressures. Policy has been based on the assumption that private provision will grow to offset this decline. Unfortunately, voluntary private pension provision has not grown as expected; instead, if anything, it has declined. Employers' willingness to voluntarily provide pensions is falling and initiatives to stimulate personal pension saving have not worked (UK pension commission, 2006).

In recent history, a pressing issue has been the under-funding of defined benefit occupational pension schemes. But the sector has also faced problems because of the bear market[119]; and there are also ongoing crises of mis-selling of personal pensions - for example, the failure of Equitable Life insurance company[120].

Annuity markets have attracted attention as part of a global debate on social security reform. There are proposals in many nations to replace or supplement defined benefit social security programs with defined contribution systems in which individuals would accumulate assets in individual accounts. In such systems, it is not clear how individuals would

119 The key current issue in occupational defined benefit funds is under-funding. At end-2002, estimates suggested there were pension fund deficits of 160-300 billion pound (CBI, 2003), relative to the accrued benefit obligation.
There are a number of factors underlying the deficits. Davis (2004) suggested that the most fundamental problem was the bear market which hit pension funds in the UK, given their large holdings in equities.

120 The near-collapse of Equitable Life in early 2000s, the world's oldest mutual insurer, has been one of the biggest financial scandals in recent times. It shocked not just the firm's policy holders, but the whole life insurance industry in the UK.

draw down their asset balances in retirement. Some proposals call for compulsory annuitisation at retirement, others would allow individuals to draw down their account balances in more flexible ways, either by choosing to purchase annuities from private insurance firms or by taking lump sum distributions. The relative attractiveness of these various options depends critically on whether reasonably priced individual annuities are actually available in the private annuity market. In case of the UK, the government requires that at least 75 percent of an individual private pension fund be used for pension purposes through the purchase of an annuity and these are purchased in the "compulsory" annuity market allowing pensioners largely to avoid selection effects (Cannon and Tonks, 2004; Finkelstein and Poterba, 2002, 2004).

Tonks (1999) pointed out that under defined contribution schemes, the pensioner bears the risk of fund underperformance, while the employer bears such risk under a defined benefit schemes. It is useful to note that in case the of defined benefit occupational pension schemes, the pension fund is overseen by trustees who are required to meet certain funding requirements to ensure that the pension fund will be able to meet its liabilities. Legislation to ensure that defined benefit pensions funds would be sufficient to meet liabilities was strengthened considerably after the Maxwell scandal (when pension funds were diverted to support Robert Maxwell's failing business empire) by the 1995 Pensions Act, which came into force in 1997, and required 100 percent funding of liabilities. This means that funds whose assets fell short of this requirement had to be made up within five years, with accelerated requirements if funding fell below 90 percent. In addition, pensions had to be indexed to prices up to a limit of 5 percent, insuring pensioners against low inflation (through not insuring them against high inflation).

These twin regulatory requirements of annuitisation for private pensions and minimum funding for occupational pensions has resulted in pensions providers having large near-certain liabilities which are most suitably backed by assets such as bonds, which are much less risky than equity. Since the UK government issues long-dated bonds whose coupons and principal are fully indexed to prices, the requirement to index pensions can easily be met while still using bonds[121].

For private pensions, it is clear that this requirement is only relevant in the "decumulation" phase, since workers could invest in any asset while accumulating their pension fund. Indeed a strategy of investing in equity during the accumulation phase, to obtain high rates of return, and in an annuity-backed bond during the decumulation phase could well be the utility maximising strategy: certainly this "life-cycle" type investment plan is recommended by many financial advisors in the UK[122]. For the defined-benefit occupational

121 UK inflation-indexed government bonds have been linked to an index of consumer prices (i.e. UK retail price index) since this is widely circulated and well understood and issued on a regular basis. Its first issue date was in 1981.

In order to construct precise protection against inflation, interest payments for a given period would need to be corrected for actual inflation over the same period. However, there is a possibility of lags between the movement in the price index and the adjustment to the bond cash flows that distort the inflation proofing properties of indexed bonds. Generally, the lags arise in 2 ways. First, inflation statistics can only be calculated and published with a delay. Secondly, in some markets, the size of the next coupon payment must be known before the start of the coupon period in order to calculate the accrued interest; this leads to a delay equal to the length of time between coupon payments. (see Choudhry 2001, page 98-99)

The point of view of this chapter is that ignoring a small measurement error, the index-linked securities are approximately index-linked.

122 One rationale for taking more risk during the accumulation phase is that one can offset poor investment performance if one still has labour income through increased saving: during retirement there no scope to hedge. Yaari (1965) shows that annuitising is optimal during retirement. Using the data set from Dimson, Marsh and Staunton (2002), the correlation between bond and equity returns in the UK has been 0.55 over the 20th century and 0.50 since 1950, so there is limited scope to use the two sorts of assets as hedges against each other. The correlation between the two asset returns is lower in most other countries in the data set, but still positive.

pension sector it is less clear whether pension funds hold government debt predominantly to back their liabilities for pensions currently being paid out or whether they also perceive them as being part of the assets to back liabilities of pensions still being accumulated. An extreme position is that occupational pension funds should hold bonds alone for both the accumulation and decumulation phases (Exley, Mehta and Smith, 1997) so that companies bear virtually no risk, through hardly any companies have acted on this advice in practice. Instead, occupational pensions have moved towards being defined contribution schemes, whereby the pension received by the employee is based upon an individual pension fund and where the rate-of-return risk is borne by the employee. Such occupational pensions then appear much closer to an individual private pension, with the only difference being that the employer is administering the fund.

Our brief characterisation of the UK pension system thus provides two reasons to believe that government debt and physical capital might be used in different phases of a pension plan: equity in the earlier phase, to benefit from the higher rates of return and bonds in the later phase to minimise risk.

The original contribution of this chapter is to ask what macroeconomic consequences would follow from such economic behaviour, where the decision rule is to invest in equity while accumulating the pension and in bonds while decumulating is taken as given. In particular we wish to know the optimal mixture of private and state pension systems, what effects would result from cuts in government debt and a baby-boom demographic shock.

4.3 Related Academic Literature:

The first and foremost research on pensions in the economics literature is by Samuelson (1958) and Aaron (1966), who model the difference between funded and unfunded schemes in an overlapping generations framework. The research showed that with a PAYG scheme, it is possible in principle for every generation to receive more in pensions than it paid in contributions, provided that the rate of growth of total real earnings exceeds the interest rate indefinitely[123]. This can happen when there is technological progress and/or steady population growth and capital accumulation. From this perspective, the introduction/expansion of unfunded public pension schemes in many industrial countries showed promising results in the post-war years.

However, this argument does not appear to be currently relevant. The old age dependency ratio in nearly all developed economies is substantially higher. In case the of the UK, life expectancy is increasing rapidly while low birth rates are predicted; this will produce a near doubling in the percentage of the population aged 65 years and over between now and 2050 (UK Pension Commission, 2005). This requires an adjustment to public policy and/or individual behaviour. The proposed options are either: 1) retirees become poorer relative to the rest of society; or 2) raising taxes/National Insurance contributions; or 3) Savings must rise;o r 4) average retirement ages must rise. This has generated a large literature on the reform of the pension systems (such as Feldstein, 1996; Feldstein and Samwick, 1998; Mitchell and Zeldes, 1996; Disney, 1996; Kotlikoff, 1996; Huang, Imrohoroglu and Sar-

123 In unfunded pension systems, the contributions of the working generation earn a return which is composed of the rates of growth of population, or what we called "a biological rate of interest" and of wages. For funded systems, the market rate of interest (the marginal productivity of capital) is relevant. It is feasible that the former is higher than the latter.

gent, 1997; Miles and Timmerman, 1999; Sinn, 1999; and Campbell and Feldstein, 2001). These works mainly concentrate on the appropriate proportion of the unfunded state pensions and the private pension system.

Among these, many researchers have expressed different opinions on the problem whether the pay-as-you-go state pension system should be replaced with a funded system. The first group suggested efficiency gains from a transition to a funded system (Diamond, 1965; Feldstein, 1977, 1995; Kotlikoff, Smetters, and Walliser, 1998; Feldstein and Samwickm 1998; Borsch-Supan, 1998; Homburg, 1990, 1997; and Miles, 1999). The second group argued that a Pareto improving transition to a funded system is not possible (Breyer, 1989, Fenge, 1995; Brunner, 1996; Sinn, 1997, 1998; and Geanakopolos, Mitchell, and Zeldes, 1998).

The first group argue that the pay-as-you-go schemes induce important labour market distortions due to the required income tax. Moreover, such schemes diminish the capital stock because they are a special form of government debt (Feldstein, 1977). On the other hand, funded pensions are dynamically more efficient, since they encourage saving, which in turn raises the capital-labour ratio and income per head. As a result, phasing out unfunded state pensions completely generates higher saving, a higher capital stock and lower real rates of return.

The second group argued that this comparison of rates of return does not imply that the abolition of the state pay-as-you-go pension system would lead to an intergenerational Pareto improvement since it is impossible to compensate the losers of the transition (namely, the first generation which does not receive the state pension) without making at least one of the later generations strictly worse off[124].

Among the work of Diamond, 1965; Feldstein, 1977; and others, the pay-as-you-go schemes diminish the capital stock because they are a special form of government debt. The logic of these models is that government debt will reduce welfare because it will divert savings away from productive capital, hence "crowding out" of private investment.

The next section describes a three-period Diamond's (1965) overlapping generations model.

4.4 The structure of the Model

This section outlines a three-period overlapping generations model and studies optimal pension provision. The model is a simplification of the complex economic phenomena in the real world by assuming a closed economy. All equilibria in this study are competitive. Additionally, the study deals with real economies, where there is no money involved and all exchange are barter exchange. Thus, the question of inflation and the interaction between real and nominal variable will not be addressed here. Individuals are assumed neither to leave bequests nor to receive inheritance. Finally, there is only one type of government debt in this model which is the long term government bond.

Our discussion begins by describing patterns and the key ingredients of the model. In the first period of life, the young work and save by purchasing equity to obtain high rate

124 In addition to this, there are important income redistribution consequences; as public pension systems often redistribute income from the rich to the poor (see Barr and Diamond, 2006).

of returns on investment in the next period. This is the asset accumulation phase. In order to prepare for the (final period) retirement asset decumulation, agents continue to work in middle age (the second period) and purchase government bonds at the end of their working life. In fact they would do this indirectly by joining a funded pension scheme. The elderly live off the proceeds of the bonds as well as from a state pension.

Agents contribute to both unfunded and funded systems. The first is through labour income tax. The latter is through contributing part of their earning to a private pension scheme when they are in their middle age. These private pension funds are then invested in an annuity market, which is assumed to be bond backed.

Investment income is capitalized each period to produce a fund at retirement. This simple three period life-cycle model allows government to choose optimal combinations of the funded (private) and unfunded (public) pension over agents' lifetime. The resource transfer occurs both via the market for capital assets (by funded pension) and via the tax system (through an unfunded pay-as-you-go system).

To characterise the model, we need to model the utility maximizing behaviour of agents and the productive sector of the economy.

4.4.1 Firms

Production

Consider a cohort of people who are young (age 1) at period t, middle-aged (age 2) in period $t+1$, and old (age 3) in period $t+2$. There is no mortality when young or middle-aged. N_t is the number of agents in each cohort, born at time t. At each period

$t \geqslant 1$, N_t, N_{t-1}, N_{t-2}individuals are alive, including N_t young born in t, N_{t-1} middle age born in $t-1$, and N_{t-2} old born in $t-2$.

We shall normalise the amount of labour provided by each young person to be unity (working full time). The simplest assumption to make would be that people are equally productive and have equal employment or participation rates when young and middle aged. However, we shall assume that the labour input of each middle-aged person is h times the labour input of the young. The parameter h can be interpreted as: first, how much more productive the middle-aged workers are compared to the young (productivity rising due to experience); and second, how much more labour (in terms of proportion of time) the middle-aged supply. Setting aside the issue of experience, high youth unemployment rates or long periods of university (such as is experienced in much of continental Europe) would suggest high values of h, while early retirement would suggest low values[125] of h.

The individual lifetime labor supply profile is $(1, h, 0)$. The young exogenously supply N_t units of labour, the middle age supply hN_{t-1} units. Hence at a given period t, the total unit of labour supplied is $N_t + hN_{t-1}$.

The production sector is characterized by a representative firm that uses capital, K_t and labor, L_t. The output is determined by a Cobb-Douglas aggregate production:

$$Y_t = QK_t^{\alpha} \left[A_t \left(N_t + hN_{t-1}\right)\right]^{1-\alpha}, \tag{4.49}$$

125 To be specific, $h < 1$ can be interpreted as part time work or early retirement, while $h > 1$ implies learning by doing (as workers are more productive when he gets older) or workers are unemployed when young.

De La Croix and Michel (2002), page 64, implies $h > 1$ as learning by doing alone. However, it is also plausible to interpret it as a case of youth unemployment.

where A_t is a "labor augmenting" or "Harrod neutral" technological progress [126] and Q is a constant. The technological progress (long run growth in output per worker) is assumed to be growing at a constant rate

$$A_t = gA_{t-1}, g > 1. \tag{4.50}$$

The population growth rate is

$$n_t = \frac{N_t}{N_{t-1}}, \tag{4.51}$$

where n_t varies from $n_t \in [0, +\infty]$.

The factor and output markets are assumed to be competitive. Factors are hired to the point where their marginal product equals factor payments and there is a one-period time-to-build capital.

Wages, Interest Rates and Pensions

Each young person saves a total amount S_{1t}, which is used to purchase productive capital, so

$$K_t = N_{t-1}S_{1t-1}. \tag{4.52}$$

In other words, productive capital K_t at time t is built from the savings of the last period's young generation, $N_{t-1}S_{1t-1}$. We assume a 100 percent depreciation rate.

The middle-aged continue to work and save in addition to the savings brought forward while young. The additional saving is through a private pension fund of S_{2t} per person. The private pension fund is assumed to be invested in government bonds.

126 This is identical to the labour augmenting technological progress in Solow growth model. The technological progress is exogenous and it occurs when "A_t" increases over time, for example, a unit of labour is more productive when the level of technology is higher.

Since we are not interested in the effects of continuous population growth, we can use lower case characters to denote variables scaled by technology alone[127] so that

$$y_t = Y_t/A_t, \quad w_t = W_t/A_t, \quad k_t = K_t/A_t, \quad s_{1t} = S_{1t}/A_t, \quad s_{2t} = S_{2t}/A_t, \qquad (4.53)$$

where W_t is the wage per unit of labour supplied. After rescaling, the production function in equation (4.49) can be rewritten as

$$y_t = QN_{t-1}\left(h+n_t\right)^{1-\alpha}\left(s_{1t-1}/g\right)^{\alpha}. \qquad (4.54)$$

Both capital and labour are paid their marginal products, so

$$\begin{aligned} R_{E,t} &= \alpha Q\left(\frac{K_t}{A_t\left(N_t+hN_{t-1}\right)}\right)^{\alpha-1}, \\ &= \alpha Q\left(\frac{s_{1t-1}/g_t}{h+n_t}\right)^{\alpha-1}, \qquad (4.55) \\ W_t &= (1-\alpha)\,Q\left(\frac{K_t}{A_t\left(N_t+hN_{t-1}\right)}\right)^{\alpha} A_t, \\ w_t &= (1-\alpha)\,Q\left(\frac{s_{1t-1}/g_t}{h+n_t}\right)^{\alpha}, \qquad (4.56) \end{aligned}$$

where R_{Et} is the rate of return on capital investment.

The old each receive a state pay-as-you-go pension, P_t, which the government sets as a proportion of wages, so that

$$P_t = p_t W_t. \qquad (4.57)$$

The government also issues bonds in each period t to be redeemed in period $t+1$ with a total redemption value $b_t W_t N_{t-2}$. These bonds are exclusively held by the elderly using their private savings and thus the private pension per person is

$$B_t = b_t W_t, \qquad (4.58)$$

127 Note that it is not the same as the common definition dividing by effective labour units.

where b_t can be interpreted as the government bonds (or funded pension) relative to the wage.

Government set the rate of return on bonds as follows

$$R_{B,t} = \frac{b_t W_t}{S_{2t-1}} = \frac{g_t b_t w_t}{s_{2t-1}} = (1-\alpha)\, g_t^{1-\alpha} b_t Q \left(\frac{s_{1t-1}}{h+n_t}\right)^{\alpha} s_{2t-1}^{-1}. \tag{4.59}$$

It is useful to continue the analysis by considering the government budget constraint.

4.4.2 The Government budget constraint

Government receipts are based on taxation labour income and sales of bonds. Outlays are the state pension and the value of the bonds issued the previous period which must be redeemed. Assuming that the government levies the same income tax rate, τ_t on workers at different ages at time t, the intertemporal government budget constraint can be written as follows;

$$N_{t-1} S_{2t} + (N_t + hN_{t-1})\tau_t W_t = (B_t + P_t)\, N_{t-2}, \tag{4.60}$$

where the government distributes a pay-as-you-go state pension of P_t per old individual. Since the government runs a balanced budget policy in every period, the income tax is set to compensate for the difference between government net worth $N_{t-1}S_{2t}$ and its liability $N_{t-2}\,(B_t + P_t)$.

Using (4.51) and (4.53), equation (4.60) can be rescaled as

$$s_{2t} + (n_t + h)\tau_t w_t = (b_t + p_t)\, w_t / n_{t-1}. \tag{4.61}$$

In most of our analysis, we assume that the government aims at a particular level of pension and adjusts the tax rate accordingly, so the equation that determines the tax rate is

$$\tau_t = \frac{1}{(h+n_t)}\left[\frac{(b_t+p_t)}{n_{t-1}} - \frac{s_{2t}}{w_t}\right]. \tag{4.62}$$

4.4.3 Households

An individual's preferences can be represented by a lifetime utility function U = U($C_{1,t}$, $C_{2,t+1}$, $C_{3,t+2}$), where $C_{1,t}$, $C_{2,t+1}$, and $C_{3,t+2}$ represent consumption of an individual in each period of life; young, middle age and old, respectively.

Assuming logarithmic utility function, the individual lifetime optimization problem is to maximize

$$\sum_{i=0}^{2} U(C_{i+1,t+i}) = \sum_{i=0}^{2} \beta^i \ln(C_{i+1,t+i}), \tag{4.63}$$

where β is a geometric discounting parameter and $0 < \beta < 1$. $\sum_{i=0}^{2} U(C_{i+1,t+i})$ expresses the lifetime utility of individual in generation t, as a function of consumption over the three periods of life. With a log-linear utility function, the implied indifference curves of individuals are convex. The utility function is twice continuously differentiable[128] on the set of strictly positive real numbers $\mathbb{R}_{++}$. It is also strictly increasing (implying no satiation) and concave (decreasing marginal utility).

The generation-specific budget constraints are written as follows

$$(1-\tau_t)W_t = C_{1,t} + E_t, \tag{4.64}$$

$$(1-\tau_{t+1})hW_{t+1} + R_{E,t+1}E_t = C_{2,t+1} + B_{t+1}, \tag{4.65}$$

128 See De La Croix and Michel (2002), page 5

$$R_{B,t+2}B_{t+1} + P_{t+2} = C_{3,t+2}, \tag{4.66}$$

where W_t is wage paid per unit of labour to the young and the middle aged at time t. Individuals receive total income W_t from inelastically supply 1 unit of labor when young, and will receive hW_{t+1}. when in middle age. In each period, the government sets a single income tax rate for all workers regardless of their age. At time t the individual is subjected to an income tax rate of τ_t and of τ_{t+1} at time $t+1$. The other parameters and the savings portfolio are described as follows.

The individual young works and saves at time t by purchasing equity E_t, which gives him interest returns in the following period of $R_{E,t+1} = (1 + r_{Et+1})$ (this is presented in equation (4.64)). He continues to work in middle age and joins the defined contribution (or money purchased) scheme at time $t+1$. The scheme is always funded and converts the value of the pension fund at the retirement into an annuity which is bond-backed. Thus, the worker indirectly purchases government bonds of B_{t+1} and expect to receive returns on bonds in the following period of $R_{B,t+2} = (1 + r_{B,t+2})$. This is illustrated in equation (4.65). Generally, the rate of return on bonds, $r_{B,t+2}$ is lower than on equity, r_{Et+1} since government bonds are relatively safer bets. Equity and government debt are therefore not perfect substitutes in this framework.

Apart from receiving the fully funded pension benefit at retirement, the retiree also receives a pay-as-you-go state pension, P_{t+2} at time $t+2$. This is be illustrated in equation (4.66). Thus, it is useful to note here that labor income is not the sole determinant of life cycle saving and consumption in the three periods model. This is because individuals also receive capital income and the state pension for their retirement consumption.

A key aspect of the set-up is that there are inter-generational transfers between the working and the retired population since there is no underlying fund of assets for the flat rate pension P_{t+2}. Hence current workers partly pay the pensions of the retired through income tax.

Combining equations (4.64) to (4.66), the lifetime budget constraint can be rewritten as

$$(1-\tau_t)\,W_t+\frac{(1-\tau_{t+1})\,hW_{t+1}}{R_{E,t+1}}+\frac{P_{t+2}}{R_{E,t+1}R_{B,t+2}}\leqslant C_{1,t}+\frac{C_{2,t+1}}{R_{E,t+1}}+\frac{C_{3,t+2}}{R_{E,t+1}R_{B,t+2}}, \quad (4.67)$$

where the left hand side of equation (4.67) represents the present value of lifetime income of an individual born in period t. The right hand side of equation (4.67) is the present value of the lifetime consumption.

It is useful to separately consider the utility maximisation of the young and the middle aged. The results are as follows.

Utility maximization of the young

The optimisation problem of the young is

$$\max U = \sum_{i=0}^{2}\beta^i \ln\left(C_{i+1,t+i}\right) \quad (4.68)$$

$$s.t.$$

$$C_{3,t+2} \leqslant P_{t+2|t}+R_{B,t+2|t}\left\{\left(1-\tau_{t+1|t}\right)hW_{t+1|t}-C_{2,t+1}+R_{E,t+1|t}\left[\left(1-\tau_t\right)W_t-C_{1,t}\right]\right\},$$

where $P_{t+2|t}$ is the expectation at time t of pension at time $t+2$ and similar notation is used for the bond rates, rates of return on equity and the wage rates. This chapter models expectations of future variables with static expectations[129]: agents assume that trending variables will grow by their long-run trend and that non-trending variables will be constant so

$$\begin{aligned} P_{t+i|t} &= g^i P_t, \\ R_{B,t+i|t} &= R_{B,t}, \\ R_{E,t+i|t} &= R_{E,t}. \end{aligned} \tag{4.69}$$

An individual chooses an optimal lifetime consumption path, given his preferences and lifetime budget constraint. The optimal level of first period consumption is calculated by setting up Lagrangian, with the following result,

$$C_{1t} = \frac{(1-\tau_t) W_t + (1-\tau_{t+1|t}) hW_{t+1} (R_{Et+1|t})^{-1} + P_{t+2|t} (R_{E,t+1|t} R_{B,t+2|t})^{-1}}{[1+\beta+\beta^2]}. \tag{4.70}$$

From equation (4.70), consumption in the first period of life is a normal good since it rises with the present discounted value of labor income after tax in both periods of working life. There is also a positive relationship between the present value of the public pension and consumption in the first period. In contrast, an expected rise in returns on equity at time $t+1$ and on bond at time $t+2$ discourages consumption of the young at time t. This can be regarded as a substitution effect. When the future interest rates are expected to rise, the opportunity cost of current consumption is higher. As a result, people tend to save more and shift parts of their current consumption to the future.

129 The rationale for this assumption is that the time frame for each period of life is approximately 20-30 years. Agents can hardly predict the amount of public social security benefit and the returns on assets they will receive. Thus agent's consumptions and savings decision are based on current public information instead of the expected real value in the future.

The optimal consumption in the second and third periods of life are $C_{2,t+1} = \beta R_{E,t+1} C_{1,t}$ and $C_{3,t+2} = \beta^2 R_{E,t+1} R_{B,t+2} c_{1,t}$. Moreover, at a consumer optimum, the gross interest rates (on bonds and equity) are proportional to consumption growth which can be expressed as: $R_{Et+1} = \frac{1}{\beta}\left(\frac{C_{2t+1}}{C_{1t}}\right)$ and $R_{Bt+2} = \frac{1}{\beta}\left(\frac{C_{3t+2}}{C_{2t+1}}\right)$.

Utility maximization of the Middle-aged

The middle aged individual's optimization problem is to maximize the rest of his life-time utility subject to his budget constraint. The budget constraint of the middle aged becomes

$$C_{3,t+1} = P_{t+1} + R_{B,t+1}\left[(1-\tau_t)\, hW_t + R_{E,t}E_{t-1} - C_{2,t}\right]. \tag{4.71}$$

To simplify the calculations for the dynamics of saving and to give a better explanation for agent's behavior, it is useful to consider the optimization problem of an individual in his second period of life as follows:

$$\begin{aligned} \max U &= \sum_{i=0}^{1} \beta^i \ln\left(C_{i+2,t+i}\right) && (4.72) \\ & \quad s.t. \\ C_{3,t+1} &\leqslant P_{t+1|t} + R_{B,t+1|t}\left[(1-\tau_t)\, hW_t + R_{E,t}S_{1,t-1} - C_{2,t}\right], \end{aligned}$$

where E_{t-1} in equation (4.71) is replaced by $S_{1,t-1}$. The budget constraint of a middle aged individual is contingent on his current income, returns on saving when young (which determine the value of his private pension), and the public pay-as-you-go pension benefit. Again, with static expectations, the middle aged individuals makes his decision based on currently available information on public pension benefits and returns on bonds instead of the expected future value. The optimal level of second period consumption is derived to be

$$C_{2t} = (1+\beta)^{-1}\left[(1-\tau_t)W_t h + R_{E,t}E_{t-1} + (R_{B,t})^{-1}P_t\right]. \tag{4.73}$$

Agents' projection of future tax rates is determined by the government budget constraint. Since S_{2t} is the total pension fund of each middle-aged person, we have

$$S_{2t} = (1-\tau_t)hW_t + R_{E,t}S_{1,t-1} - C_{2,t}. \tag{4.74}$$

The next task is to derive the optimal level of savings for the young and the middle age generations. The results are presented in the following subsection.

4.4.4 Savings

Optimal level of saving for the middle age

To derive the optimal solution to equation (4.72), we assume that the demand for government bonds comes purely from the funded pension sector. Combining equations (4.53), (4.57), (4.69) and (4.74), the optimal saving function is

$$\begin{aligned} S_{2,t} &= \left(\frac{\beta}{1+\beta}\right)[(1-\tau_t)hW_t + R_{E,t}S_{1,t-1}] - \frac{gP_t}{(1+\beta)R_{B,t}}, \\ s_{2,t} &= \left(\frac{\beta}{1+\beta}\right)\left[(1-\tau_t)hw_t + \left(\frac{s_{1,t-1}}{g}\right)R_{E,t}\right] - \frac{gp_t w_t}{(1+\beta)R_{B,t}}. \end{aligned} \tag{4.75}$$

We then substitute for expressions for government budget constraint and factor market equilibrium such as wages (equation (4.56)), interest rates (equations (4.55) and (4.59)) and tax rates (equation (4.62)). The equation of motion for $s_{2,t}$ is therefore

$$s_{2,t} = L_{21}s_{1,t-1}^{\alpha} - L_{22}s_{2,t-1}, \tag{4.76}$$

where

$$L_{21} = \frac{\beta Q g^{-\alpha} (n_t + h)^{1-\alpha}}{[(1+\beta) n_t + h]} \left[(\alpha n_t + h) - \frac{(1-\alpha) h}{n_{t-1} (n_t + h)} (b_t + p_t) \right] \lessgtr 0, \tag{4.77}$$

and

$$L_{22} = \frac{(n_t + h)}{(1+\beta) n_t + h} \left(\frac{p_t}{b_t} \right) > 0. \tag{4.78}$$

Equation (4.76) shows that the savings of the previous middle aged generation, $s_{2,t-1}$ discourages saving of the current middle age generation, $s_{2,t}$. The intuition is as follows. If there is higher demand for government bonds from the middle age generation at time $t-1$, the resulting return on bonds maturing at time t will be lower. The new middle age generation at time t observes that the return on bonds is not as high as before. Thus, the opportunity cost of current consumption declines. If substitution effects outweigh income effects, then people will save less for retirement.

Optimal saving in the first period of life

To find an optimal solution for equation (4.68), we first substitute for optimal consumption (4.70) in the equation for budget constraint to get

$$s_{1,t} = \frac{(1-\tau_t) w_t}{(1+\beta+\beta^2)} \left[\beta + \beta^2 - \frac{gh}{R_{E,t}} \right] - \frac{g^2 p_t w_t}{(1+\beta+\beta^2) R_{E,t} R_{B,t}}. \tag{4.79}$$

The optimal level of saving for the young can be computed by substituting equation (4.76) into equation (4.79), which is

$$s_{1,t} = L_{11} s_{1,t-1}^{\alpha} - L_{12} s_{1,t-1} - L_{13} s_{2,t-1} - L_{14} s_{1,t-1}^{1-\alpha} s_{2,t-1}, \tag{4.80}$$

where

$$
\begin{aligned}
L_{11} &= \frac{\left(\beta+\beta^2\right) Q\left[(1-\alpha+\beta)\left(n_t+h\right)-\frac{(1-\alpha)(1+\beta)}{n_{t-1}}\left[b_t+p_t\right]\right]}{\left(1+\beta+\beta^2\right) g^\alpha\left(n_t+h\right)^\alpha\left[(1+\beta) n_t+h\right]} \lesseqgtr 0, \\
L_{12} &= \frac{h\left[(1-\alpha+\beta)\left(n_t+h\right)-(1-\alpha)(1+\beta)\left(n_{t-1}\right)^{-1}\left[b_t+p_t\right]\right]}{\alpha\left(1+\beta+\beta^2\right)\left(n_t+h\right)\left[(1+\beta) n_t+h\right]} \lesseqgtr 0, \\
L_{13} &= \frac{\left(\beta+\beta^2\right)}{\left(1+\beta+\beta^2\right)\left[(1+\beta) n_t+h\right]}\left(\frac{p_t}{b_t}\right)>0, \\
L_{14} &= \frac{(1+\beta) g^\alpha n_t}{\alpha\left(1+\beta+\beta^2\right) Q\left(n_t+h\right)^{1-\alpha}\left[(1+\beta) n_t+h\right]}\left(\frac{p_t}{b_t}\right)>0.
\end{aligned}
$$

Although we cannot show that all of the parameters L_{1i} are positive universally, they are positive for all plausible parameter values and would only be negative if $b+p$ was very high.

The savings of the previous period's middle aged generation not only discourages savings of the current middle age generation (as shown in equation (4.76)), but also the savings of the current young generation (as in equation (4.80)).

Equilibrium: Steady state values of savings:

To find steady state values of savings, we first perform the first order difference of $s_{1,t}$ and $s_{2,t}$ as follows;

$$
\begin{aligned}
\Delta s_{1,t} &= L_{11} s_{1,t-1}^\alpha-\left(1+L_{12}\right) s_{1,t-1}-L_{13} s_{2,t-1}-L_{14} s_{1,t-1}^{1-\alpha} s_{2,t-1}, & (4.81)\\
\Delta s_{2,t} &= L_{21} s_{1,t-1}^\alpha-\left(1+L_{22}\right) s_{2,t-1}. & (4.82)
\end{aligned}
$$

Let $\Delta s_{1t}=0$ and $\Delta s_{2t}=0$ to establish an equilibrium system. The first is called the constant first period of life saving, while the latter is called the constant second period of life saving.

Setting $\Delta s_{1t} = 0$, the resulting equilibrium saving is

$$s_2 = \frac{1}{L_{13}} \left[L_{11} s_1^{\alpha} - (1 + L_{12}) s_1 - L_{14} s_1^{1-\alpha} s_2 \right]. \tag{4.83}$$

Similarly, setting $\Delta s_{2t} = 0$ yields

$$s_1^* = \left[\frac{L_{11}(1 + L_{22}) - L_{13} L_{21}}{(1 + L_{12})(1 + L_{22}) + L_{14} L_{21}} \right]^{\frac{1}{1-\alpha}}, \tag{4.84}$$

$$s_2^* = \frac{L_{21}}{(1 + L_{22})} (s_1^*)^{\alpha}, \tag{4.85}$$

where s_1^* and s_2^* are steady state values of savings. There is a unique interior solution as long as the term in the square brackets is positive. Considering the definitions of the L_{ij} parameters, the steady state value of savings is determined by the independent variables (pension per GDP, p and bonds per GDP, b) and other parameters, namely, the geometric discounting factor in the utility function (β), the relative productivity of the middle aged (h), the relative proportions of capital and labor in the production function (α), the population growth rate (n), and total factor productivity (A).

Stability of equilibria

In a planar system, we can construct an autonomous first-order system of difference equations for (4.76) and (4.80) as follows;

$$s_{1,t} = f(s_{1,t-1}, s_{2,t-1}), \tag{4.86}$$

$$s_{2,t} = g(s_{1,t-1}, s_{2,t-1}), \tag{4.87}$$

where both f and g are assumed to be continuously differentiable, $f : \mathbb{R}^2 \to \mathbb{R}$ and $g : \mathbb{R}^2 \to \mathbb{R}$. The solutions to the system of equations in (4.86) and (4.87) (or in equations (4.80) and 4.76)) are steady states value of savings (s_1, s_2) in the dynamic system.

Since equations for savings in both periods of life are non-linear, we need to approximate them in a neighbourhood of the steady state with a linear system. Thus all the partial derivatives are evaluated at the steady state. The system of equations above transforms into the linear system as follows;

$$\begin{aligned} s_{1,t} &\approx \bar{s}_1 + \frac{\partial f\left(s_{1,t-1}, s_{2,t-1}\right)}{\partial s_{1,t-1}}\left(s_{1,t-1} - \bar{s}_1\right) + \frac{\partial f\left(s_{1,t-1}, s_{2,t-1}\right)}{\partial s_{2,t-1}}\left(s_{2,t-1} - \bar{s}_2\right), \\ s_{2,t} &\approx \bar{s}_2 + \frac{\partial g\left(s_{1t-1}, s_{2t-1}\right)}{\partial s_{1t-1}}\left(s_{1,t-1} - \bar{s}_1\right) + \frac{\partial g\left(s_{1,t-1}, s_{2,t-1}\right)}{\partial s_{2,t-1}}\left(s_{2,t-1} - \bar{s}_2\right). \end{aligned}$$

Applying Taylor's formula[130] to the system of equations above, we derive an approximation for a non-linear system near the steady state as

$$x_{t+1} = \bar{x} + Df\left(\bar{x}\right)\left(x - \bar{x}\right), \tag{4.88}$$

where $Df\left(\bar{x}\right)$ is an invertible Jacobian matrix[131], and $\bar{x}$ is a hyperbolic equilibrium of equations (4.86) and (4.87). Stability of the steady state (s_1, s_2) depends on the eigenvalues of the Jacobian matrix of partial derivatives, namely,

$$\begin{bmatrix} \frac{\partial f(s_{1,t-1}, s_{2,t-1})}{\partial s_{1,t-1}} & \frac{\partial f(s_{1,t-1}, s_{2,t-1})}{\partial s_{2,t-1}} \\ \frac{\partial g(s_{1,t-1}, s_{2,t-1})}{\partial s_{1,t-1}} & \frac{\partial g(s_{1,t-1}, s_{2,t-1})}{\partial s_{2,t-1}} \end{bmatrix},$$

which is simply the coefficient matrix of equation for Taylor approximation (4.88). From equations (4.81) and (4.82), this matrix can be rewritten as

$$\begin{bmatrix} \alpha L_{11} s_1^{\alpha-1} - L_{12} - (1-\alpha) L_{14} s_1^{-\alpha} s_2 & -L_{13} - L_{14} s_1^{1-\alpha} \\ \alpha L_{21} s_1^{\alpha-1} & -L_{22} \end{bmatrix}. \tag{4.89}$$

130 From Azariadis (1993, page59), Taylor's formula is

$$f(x) = f\left(\bar{x}\right) + Df\left(\bar{x}\right)\left(x - \bar{x}\right) + O\left(\left\|x - \bar{x}\right\|\right)$$

where the reminder term $O(\cdot)$ is small in a well defined sense. We thus assume that this term is close to zero.

131 From Azariadis (1993, p59), if the Jacobian matrix $Df\left(\bar{x}\right)$ is invertible, there is a neighbourhood of $\bar{x}$ in which a non linear system of equations (4.86) and (4.87) is topologically equivalent to the linear system (equation (4.88)).

The matrix in equation (4.89) can be rearranged by substituting for the steady state values of savings in both periods of life (equation(4.84) and (4.85)) and yields

$$\begin{bmatrix} \begin{matrix}[\alpha L_{11}\left(\frac{L_{11}(1+L_{22})-L_{13}L_{21}}{(1+L_{12})(1+L_{22})+L_{14}L_{21}}\right)^{-1} \\ -L_{12}-(1-\alpha)\frac{L_{14}L_{21}}{(1+L_{22})}]\end{matrix} & [-L_{13}-L_{14}\left(\frac{L_{11}(1+L_{22})-L_{13}L_{21}}{(1+L_{12})(1+L_{22})+L_{14}L_{21}}\right)] \\ [\alpha L_{21}\left(\frac{L_{11}(1+L_{22})-L_{13}L_{21}}{(1+L_{12})(1+L_{22})+L_{14}L_{21}}\right)^{-1}] & [-L_{22}] \end{bmatrix}.$$

The stability of equilibria can be determined by examining the eigenvalues of the Jacobian. The equilibrium will be stable (i.e. a sink) if both the eigenvalues are less than unity. Since the unstable equilibrium is viewed as uninteresting, the analysis below will be confined to the cases where this condition is met.

The Jacobian matrix cannot be solved analytically; instead we will present numerical solutions to the model in the later sections. It is useful to first discuss the dynamics of savings and the comparative static analysis. Then, we will address the questions that we posted earlier, namely which sort of pension allocation is optimal and the effect of shocks to the economy, in the following sections.

The dynamics of savings

To construct a phase diagram for the non-linear system in equations (4.86) and (4.87) (or equivalently equations (4.80) and (4.76)), it is useful to discuss the sign of $\Delta s_{1,t}$ and $\Delta s_{2,t}$ implied by equations (4.81) and (4.82). The results are

$$\Delta s_{1,t} > 0 \Longleftrightarrow \frac{L_{11}s_{1,t-1}^{\alpha}-(1+L_{12})\,s_{1,t-1}}{L_{13}+L_{14}s_{1,t-1}^{1-\alpha}} > s_{2,t-1}, \tag{4.90}$$

$$\Delta s_{2,t} > 0 \Longleftrightarrow \frac{L_{21}}{(1+L_{22})}s_{1,t-1}^{\alpha} > s_{2,t-1}. \tag{4.91}$$

The next task is to discuss the effect of pension and government bonds on savings by comparative static analysis.

4.4.5 Comparative static analysis:

This section presents a comparative static analysis for the unique interior solution of equation (4.84)). To clarify the analysis, we concentrate on two measures, namely $(p+b)$, which is a measure of total pension received by the elderly relative to the wage, and p/b as a measure of the composition of the public/private pensions.

First, we study the impact of an increase in the private-public pension ratio on the equilibrium capital stock (i.e. the saving of the young). The results show that an equilibrium capital stock positively relates to the ratio of private to public pension.

$$\frac{ds_1}{d(b/p)} > 0. \tag{4.92}$$

The pay-as-you-go state pension (p) negatively relates to capital accumulation. The role of the PAYG pension is to redistribute across generations. It is contractionary in nature since the state taxes the working population to pay the pensions of the retired generation. The income tax generally creates disincentive effects on labour supply. If the substitution effects outweighed the income effect, the working generation will work less. Aggregate output and savings will therefore decline.

The fully funded pension (private pension, b) works in the opposite way. The fully funded scheme is based on savings and is a method of accumulating financial assets, which are exchanged for goods at some later date. As a result the equilibrium capital stock rises with private pensions.

The relationship between s_1 and the total pension $(b+p)$ received by the elderly relative to the wage shows an ambiguous sign $\left(\frac{ds_1}{d(b+p)} \gtrless 0\right)$ in our study.

We further analyse the impact of the private to public pension ratio on output,

$$\frac{dy}{d(b/p)} = \alpha g^{-\alpha} Q \left(\frac{n+h}{s_1}\right)^{1-\alpha} N \frac{ds_1}{d(b/p)}. \tag{4.93}$$

We find that $\frac{dy}{d(b/p)} > 0$ since $\frac{ds_1}{d(b/p)} > 0$ from equation (4.92). Thus, a higher level of output can be achieved by increasing the private to public pension ratio. This argument is an objection to the simple Diamond (1965) model, which argues that government borrowing crowds out investment and lowers output in the long run. In our three-period model, the funded pension (which is bond-backed) generates a positive impact on national savings and income. However, we should note that this conclusion obtains under an assumption of no severe shrink in population growth. If there is a persisting slump in the fertility rate (or population growth trend) ($n < 0$), the ratio of workers to pensioners will fall even though the life expectancy is unchanged. In that scenario, if the fertility rate falls until $|n| > h$, there is a possibility that private pensions will crowd out saving and reduces output.

The next question is the impact of alternative pension schemes on welfare. Equations (4.92) and (4.93) show that the private pension positively affects capital accumulation and output while the PAYG pension has a contractionary impact. However, it is not sufficient to automatically infer parallel responses of welfare since the latter cannot be gauged by output alone. There are some counterbalancing factors that need to be netted out. In order to assess this fully, we first posit steady state welfare as

$$\tilde{u} = \ln c_1 + \ln c_2 + \ln c_3, \tag{4.94}$$

where $c_i \equiv C_i/A$ is consumption adjusted for technical progress. The impact of the pension ratio on consumer welfare can be investigated using the total derivative,

$$\frac{d\tilde{u}}{d(b/p)} = \tilde{u}_Y \frac{dc_1}{d(b/p)} + \tilde{u}_M \frac{dc_2}{d(b/p)} + \tilde{u}_O \frac{dc_3}{d(b/p)}, \tag{4.95}$$

where $\tilde{u}_Y = \frac{\partial \tilde{u}}{\partial c_1}, \tilde{u}_M = \frac{\partial \tilde{u}}{\partial c_2}$,and $\tilde{u}_O = \frac{\partial \tilde{u}}{\partial c_3}$. The first order conditions for each period consumption implies that $\tilde{u}_Y = \left(\frac{c_2}{\beta c_1}\right)$ and $\tilde{u}_M = \left(\frac{c_3}{\beta^2 c_1}\right) \tilde{u}_O$, thus

$$\frac{d\tilde{u}}{d(b/p)} = \tilde{u}_Y \left[\frac{dc_1}{d(b/p)} + \left(\frac{\beta^2 c_1}{c_2}\right) \frac{d(R_E c_1)}{d(b/p)} + \left(\frac{\beta^4 c_1}{c_3}\right) \frac{d(R_B R_E c_1)}{d(b/p)}\right]. \tag{4.96}$$

From the first-order conditions of the consumer's optimization problem, $c_1 = \frac{c_2}{\beta R_E} = \frac{c_3}{\beta^2 R_B R_E}$ and u_Y can be substituted by $\frac{1}{c_1}$. The impact of the pension ratio on welfare thus depends on three factors; namely, the impact of pension on returns on both types of assets, and on first period consumption. The relationship can be written as follows,

$$\frac{d\tilde{u}}{d(b/p)} = \theta_1 \frac{dc_1}{d(b/p)} + \theta_2 \frac{dR_E}{d(b/p)} + \theta_3 \frac{dR_B}{d(b/p)}, \tag{4.97}$$

where

$$\begin{aligned}
\theta_1 &= \frac{1}{c_1} + \frac{\beta^2 R_E}{c_2} + \frac{\beta^4 R_B R_E}{c_3} > 0, \\
\theta_2 &= \left(\frac{1}{c_2} + \frac{\beta^2 R_B}{c_3}\right) \beta^2 c_1 > 0, \\
\theta_3 &= \frac{\beta^4}{c_3} R_E c_1 > 0.
\end{aligned}$$

Equation (4.97) implies that the welfare effects of the pension ratio (b/p) depend on the relationship between the pension ratio and the consumption of the young, and the returns on capital and on bonds.

We can explore this further by examining equations for the steady state of savings in both periods of life (equations (4.84) and (4.85)). The results are

$$\begin{aligned}
\frac{dR_E}{d(b/p)} &= -\alpha(1-\alpha)Q\left[g(n_t+h)\right]^{1-\alpha} s_1^{\alpha-2}\frac{ds_1}{d(b/p)} < 0, && (4.98)\\
\frac{dR_B}{d(b/p)} &= -\alpha^2 Q(n_t+h)^{-\alpha}(s_1/g)^{\alpha-1} s_2^{-1} b\frac{ds_1}{d(b/p)} < 0, && (4.99)\\
\frac{dc_1}{d(b/p)} &= \left\{\frac{\alpha(1-\alpha)Q}{[g(n_t+h)]^{\alpha}}\left[1-\frac{b+p}{n_{t-1}(n_t+h)}\right]s_1^{\alpha-1}\right\}A\frac{ds_1}{d(b/p)}\\
&\quad+\left[\left(\frac{\alpha}{1-\alpha}\right)\left(\frac{1}{n_t+h}\right)s_1^{-1}s_2-1\right]A\frac{ds_1}{d(b/p)}\\
&\gtrless 0. && (4.100)
\end{aligned}$$

An explanation for equation (4.98) is straight forward. Referring back to equation (4.92), an increase in the private to public pension ratio raises the capital stock, which in turn will pressure down its return.

Equation (4.99) shows that when the debt stock increases, its returns decline. Lastly, the relationship between consumption of the young and the pension ratio exhibits an ambiguous sign. Using equation (4.98) - (4.100) in equation ((4.97)) then

$$\frac{d\tilde{u}}{d(b/p)} = (\Omega+\Pi)\frac{ds_1}{d(b/p)}, \tag{4.101}$$

where $\Omega = \frac{\alpha(1-\alpha)Q}{[g(h+n_t)]^{\alpha}}s_1^{\alpha-1}\left[\theta_1 A\left[1-\frac{b+p}{n_{t-1}(n_t+h)}\right]-g\left[\left(\frac{\alpha}{1-\alpha}\right)\theta_2 b s_2^{-1}+(n_t+h)s_1^{-1}\theta_3\right]\right]$ and $\Pi = \theta_1 A\left[\left(\frac{\alpha}{1-\alpha}\right)\left(\frac{1}{n_t+h}\right)s_1^{-1}s_2-1\right]$. However, $(\Omega+\Pi)$ is ambiguous signed, hence it is not possible to infer a clear welfare effect of the private-to-public pension ratio. However, the size of pensions relative to wages $(b+p)$ and savings (s_1, s_2) play a major role in the welfare analysis. It is useful to illustrate the solutions numerically.

The next section presents the results of numerical simulaitons of the model. We address the question we posted earlier, namely which sort of pension is optimal and the effects of shocks to the economy.

4.5 Simulation Results

This section asks two sets of questions. First, we ask what is the optimal combination of the alternative pension schemes. Second, we study the effects of two different shocks to the economy and welfare: first, a reduction in the size of the national debt and second, the effect of a baby boom.

In the simulations, we assume that each time period lasts 20 years, so that the young are aged between 21 and 40, the middle-aged between 41 and 60 and the elderly 61-80. These figures crudely correspond to labour market and mortality conditions in the UK. The discount factor in the utility function (β) is set equal to 0.6676 implying a discount rate of 2 percent per annum. On the production side, the parameter for middle aged labor productivity (h) is set alternatively at 0.8, 1, and 1.2. Exogenous technical growth is set alternatively at 0 and 2.5 percent per annum (implying that g is set alternatively at 1 and 1.638). Labour augmenting technological progress (A) equals 6, the capital share in the production function (α) is 0.3333 and the arbitrary constant Q is set equal to 3.

4.5.1 Optimal Pensions in the Steady State

The objective of this section is to describe the optimal combination of private (b) and public pensions (p) in the steady state by allowing variation in the population (n_t) and technological (g) growth rates, and the parameter of the middle age productivity (h). We first consider the optimal pensions in the case where the population growth rate (n_t) is assumed to be in the range of -0.5 to 0.5 percent annually. We then consider the optimal pensions in the case of UK where the population grew by an annual rate of 0.3 percent.

The first part of this subsection describes the optimal combination of private (b) and public pensions (p) in the steady state where the population growth rate (n_t) is assumed to be in the range of -0.5 to 0.5 percent annually[132]. This corresponds to values for n in the range of $n = (0.90461, 1.10489)$. The solutions to the model are presented in table 4.1.

	Optimal combinations of (p^*, b^*)					
	$h = 0.8$		$h = 1$		$h = 1.2$	
n	$g = 0\%$	$g = 2.5\%$	$g = 0\%$	$g = 2.5\%$	$g = 0\%$	$g = 2.5\%$
-0.5%	0, 0.500	0, 0.495	0, 0.545	0, 0.545	0, 0.595	0, 0.590
0%	0, 0.580	0, 0.580	0, 0.635	0, 0.630	0, 0.685	0, 0.685
0.5%	0, 0.680	0, 0.675	0, 0.735	0, 0.735	0, 0.795	0, 0.790

Table 4.1. Optimality of different pension schemes

From table 4.1, p^* is the optimal public PAYG pension (expressed as percentage of the average wage), b^* is the optimal private pension (again, expressed as a percentage of wages), g is technological growth per annum, h is the labour productivity of the middle aged, and n is population growth rate per annum.

The simulation results in table 4.1 suggest that the optimal solution to the model is always to have no state PAYG pension at all[133]. The results are robust across different values

132 A scenario of negative population growth is plausible. From the projection of population growth during 2000-2030 by Population Reference Bureau, population growth rates are negative in many European countries, including Russia (-0.6%), Estonia (-0.5%), Hungary (-0.4%), and Ukraine (-0.4%). For developed countries in Europe and North America, as well as Japan, Australia, and New Zealand, populations are growing by less than 1 percent annually.

133 Linear interpolation is adequate to calculate the optimal values of p and b for intermediate values of the

of technological growth (g), population growth (n) and middle-aged productivity (h). The results imply that the representative agent provides for his pension through voluntary savings to achieve his optimal time path of consumption and through an annuity to secure his retirement consumption with no need for government involvement. The simulation results are consistent with previous arguments that the transition to a funded system will generate efficiency gains, such as provided by Breyer and Straub, 1993; Brunner, 1996; Feldstein, 1995; Fenge and Schwager, 1995, Kotlikoff, Smetters, and Walliser, 1998; Feldstein and Samwick, 1998; Borsch-Supan, 1998; and Homburg, 1990, 1997.

It is useful to further discuss the simulation results. The intuition for the optimality of the zero unfunded state pensions is as follows.

We focus the explanation of the effects of social security on capital accumulation and welfare. Any social security program that affects the path of income received by individuals is likely to have an effect on savings and thus on capital accumulation. In a fully funded system, the rate of return on the social security contribution is the interest rates, while the rate of return on the contribution for the case of pay as you go system equals to population and economic growths.

The fully funded social security has no effect on total savings and capital accumulation. Usually, an increase in social security savings (funded pension) is exactly offset by a decrease in private saving. In contrast, the social security in the pay-as-you-go system is a pure transfer scheme, which does not save at all, the only source of capital for the econ-

chosen parameters. In all cases, we retain the result that the optimal pensions policy is to have an entirely funded system.

omy in case of unfunded system is the private saving. Thus, social security contribution decreases private savings.

In the comparative static analysis part (section 4.4), we found that the pay as you go pension negatively relates to the capital accumulation. The PAYG is contractionary in nature, since the state tax the working population to pay the pensions of the retired generation. The fully funded pension appears to work in opposite way. The fully funded schemes base on savings and it is a method of accumulating financial assets, which are exchanged for goods at some later date. As a result, an equilibrium capital stock rises with private pension. The model also shows that the funded pension generates positive impacts on national savings and income. With higher savings and income, the consumption is increasing so as the welfare.

Lastly, in our model, pensions are allowed through the utility maximisation (through the optimal saving decision). We do not impose fixed pensions; instead representative agents are allowed to select the optimal proportions of alternative pension schemes. Thus, agents find that the pay-as-you-go pension is bad and choose not to hold it.

There is no central government intervention in our model. Thus the redistributional purpose is omitted in this study. However, it is useful to discuss that in the real world, it may be potentially dangerous to totally phase out the public PAYG pension due to the existence of imperfect information, missing markets, risk and uncertainty. Moreover, public policy generally has additional objectives to improving consumption smoothing and insurance, namely, poverty relief and redistribution (Barr and Diamond, 2006). On the poverty relief grounds, the state pension can target the poor who may be unable to save enough. On income redistribution grounds, the state pension can serve a role for both intra-generational and the inter-generational transfer. The latter is through subsidising the consumption smoothing of low earners.

The simulation results in table 4.1 also suggest that the optimal level of private pension is positively related to the population growth rate and middle age productivity. From the table, greater population growth rates and middle age productivity leads to higher optimal level of funded pension, while the change in technological progress has a very little impact on optimality.

The second part of this subsection restricts the demographic growth rate to that found in the United Kingdom. The Office for National Statistics, reports that the UK still has a growing population[134]. Between mid-1991 and mid-2003 the population grew by an annual rate of 0.3 percent (corresponding to $n = 1.0617$ in our model). Thus, we apply this population growth figure to the simulation study. Exogenous technological growth is again assumed to be 2.5 percent per annum ($g = 1.638$), while the middle age productivity (h) again varies from $(0.8, 1, 1.2)$. Using this parameterisation we investigate the shape of social-welfare functions in figures 4.1-4.3 and the steady state values of other variables described in the model.

Figures 4.1-4.3 illustrate the social welfare contours. Figure 4.1 illustrates the social welfare contour for $h = 0.8$ perhaps suggestive of a scenario of part time work in middle age or early retirement. Figure 4.2 corresponds to the case of $h = 1$, or constant labor

134 The UK population grew by 232,100 people in the year to mid-2003, and the growth was 0.4 per cent in each of the years since mid-2001. The UK population increased by 6.5 per cent in the last thirty years or so, from 55.9 million in mid-1971. Growth has been slightly faster in more recent years.

productivity. Lastly, the case of $h = 1.2$ in figure 4.3 perhaps captures learning by doing or youth unemployment. These 3 cases show similar results: welfare increases as one moves up to the north western direction. The optimal solution is to have private pensions equal to 63.5%, 69% and 74.5% of the wage for $h = 0.8, 1$ and 1.2, respectively and no state pension at all. The resulting economy can be expressed in the table 4.2 below, where rates of return are expressed at annual rates:

Variables	$h = 0.8$	$h = 1$	$h = 1.2$
(p^*, b^*)	$(0, 0.635)$	$(0, 0.690)$	$(0, 0.745)$
Debt/GDP $\left(\frac{bw}{y}\right)$	24.1431%	23.6894%	23.315%
Utility	0.0328	0.1371	0.2050
R_E (% annually)	9.0196%	9.6610%	10.2394%
R_B (% annually)	2.7085%	2.2795%	1.9333%
Tax rate	3.1850%	3.1219%	3.0699%
Wage (W)	6.0070	5.7246	5.4312
Output (Y)	15.9660	16.6739	17.3540
s_1	0.3947	0.3666	0.3435
s_2	0.6166	0.6870	0.7532
c_1	0.6948	0.7096	0.7039
c_2	1.5916	1.6883	1.7788
c_3	0.9344	0.9574	0.9804

Table 4.2. Simulation results for the UK economy

The results in table 4.2 show that as before it is optimal to have no state pension. The table also shows that the optimal level of debt is non-negligible in all cases. Increasing the productivity of the middle age raises social welfare and aggregate output as would be expected.

Comparing the first two columns, the labour supply of the middle aged determines aggregate output and capital accumulation; if the middle aged work full time until retirement ($h = 1$), it will generate a larger economy than in the case of early retirement ($h = 0.8$).

Moreover, it also leads to a slightly better government fiscal stance, holding government expenditure constant. The productivity of the middle age (h) appears on the government revenue side of the government budget constraints equation (in equation (4.60)) because it contributes to government income tax revenues. Empirically, the fiscal burden increases with early retirement. This supports the current argument to increase the retirement age in order to relieve the fiscal deficit. From the pensions white paper 2006, the state pension age is going to rise to 66 years old in 2024, to 67 in 2034 and 68 in 2044. Lastly, the scenario that the worker develop learning by doing ($h = 1.2$) leads to the best outcome to the economy and welfare.

Middle age productivity also positively affects aggregate consumption and social welfare. As the workers become more productive, not only can they set aside more resources for their old age consumption and they also enjoy higher present consumption.

From the simulation results, the lifetime consumption pattern is such that consumption during the middle age of life is highest, followed by retirement consumption and consumption when young. This corresponds to the rate of return arguments and the life cycle theory of consumption (Modigliani, 1986): agents' consumption decisions do not depend solely on current income, but also on expected future income and financial wealth. From the simulation results, the rate of return on equity is high, so young individuals consume less and accumulate more capital during the asset accumulation phase, Consumption is highest in the second period of life as individuals enjoy high returns on assets carried forward from the first period of life. Middle-aged individuals choose to invest in safe assets, which give lower returns than equity, for their retirement consumption. Old age consumption is thus relatively lower than in the middle age, even though he receives the public pay-as-you-go pension too.

4.5.2 The effects of shocks

This section analyses the impact of two types of shock, namely: a reduction in the supply of bonds and the effects of the baby boom in 1941-1961. Throughout this section, middle aged productivity is assumed to be the same as that for the young ($h = 1$).

A reduction in National debt

A reduction in national debt corresponds to a reduction in the supply of bonds, in our model the parameter b. In our simulation, we consider a scenario where the public pension relative to wage is nonzero and is constant across periods with the value of $p = 0.3$. The private pension is initially set at $b = 0.4$ and falls to $b = 0.3$ at the beginning of year 2001. Exogenous technical growth is again assumed to be 2.5 percent per annum, while the population is held constant. The effects of this policy change are presented in figures 4.4-4.6. The results are described as follows.

Figure 4.4 presents the rate of return on equity following a fall in the national debt in 2001. The results show that there is almost no effect on the rate of return on equity; the maximum change in the return is only from 9.08 percent in 2021-2040 to 8.83 percent in 2041-2060.

The path for the rate of return on bonds following a fall in the national debt in 2001 is presented in figure 4.5. The results show that such a shock generates an excess demand for bonds, which in turn significantly depresses the return on bonds within the period that the shock occurred and the following periods. The return on bonds hits its minimum at the rate of -0.36 percent annually in 2001-2020, though there is a small rise in the return on bonds in the following periods. The steady state value for the return on bonds before the shock was initially 1.09 percent annually. The new steady state value becomes 0.06 percent per annum, thus the return on bonds becomes much lower in the long run after a sudden fall in the national debt. The tax rate is, initially at 9 percent, falls immediately to 7 percent and 5 percent in the next period and reaches its stable value at 6 percent.

The effect of a fall in national debt on utility is even more striking. Figure 4.6 illustrates the welfare effects of the policy change following the transition. The utility of each generation can be derived from lifetime consumption, discounted by factor β (at the rate of 2 percent annually) over each period. It reflects welfare of each cohort, for example, the utility figure during 1981-2000 implies the utility level of a generation born in 1981, which is -0.540. The steady state values of our social welfare metric are -0.453 and -0.510 respectively, so this implies a small worsening in the long run.

The first group that are affected by a reduction in national debt are the retired in 2001-2020 (i.e. the generation that entered economic life in 1961-1980). According to figure 4.5, the return on bonds hits its low in 2001-2020, when this generation receive returns on their pension fund. However, they are unable to change their saving decisions and are faced with a much lower return on their pension fund. Hence their retirement consumption drops dramatically.

As the supply for bonds declines (holding other thing constant), the bond price will rise with corresponding reduced returns. The middle aged in 2001-2020 (i.e. the generation that entered economic life in 1981-2000) suffer most from the policy shock since this generation began to secure their old age consumption by investing in an annuity market, which is assumed to be invested in government bonds. This generation faces higher bond prices, and the return on bonds also hits its low this period. Additionally, this generation also get much lower returns on their saving when they get old, which in turn makes their old age consumption in the following period decline to even further than that of the retired in 2001-2020.

Apart from people who are alive in 2001, the next 3 generations also suffer considerable declines in utility and the steady state is not reached until new generations enter economic life in 2081-2100. The new steady state value for social welfare become -0.510, this means people are worse off in the long run after the reduction in national debt.

We conclude that when agents are relying upon government debt to finance their pension, a sudden fall in national debt does not only affect current generation, but also hurts future generations.

The effects of a baby boom

This section investigates the impact of a demographic shock on welfare. The population size is assumed to be constant until a demographic shock occurring in the period 1941-1960. We consider the effects of the generation entering economic life in 1941-1960 being 10 percent larger than the generations before and after. To isolate the demographic shock from all other effects, we assume a constant state pensions and bonds (proportional to wages) at the level of $p = 0.3$ and $b = 0.4$.

The effects are shown in Figures 4.7-4.9: The steady state values of our social welfare metric are -0.460 and -0.453 respectively, so the demographic shock implies a small improvement in the long run. From an initial level of 9 percent tax rates fall to 7 percent in 1941-1960 but rise to 13 percent in 1981, and thereafter remain between 8 and 9 percent.

The baby boom affects the rates of return on both assets (figures 4.7 and 4.8), although the effects are not as remarkable as on the social welfare (figure 4.9), but nonetheless are significant when compounded over long time periods. The group that initially suffers is the baby-boom generation itself. The generation immediately preceding (which retires in 1961-80) benefits from additional labour in the baby-boom generation (which increases the return to private capital and boosts this generation's funded pension scheme), whereas the generation immediately following (which enters economic life in 1961-80) receives the highest benefit from higher wages due to the large quantity of capital saved by the baby-boom generation.

However, although the baby-boom generation has relatively low utility, the two generations, which suffer the most, are the second and third generations later. The second generation (born in 1981-2000) suffer the most with the utility of -0.497. This is because the baby boomers retire in this period. The dependency ratio reaches its peak. This second generation pays the highest tax rate of 13 percent. This is partly to finance the pay as you go pension of the baby boom generation and also to finance the government deficit as the baby boomers start to redeem their return from holding bonds. This also results in lower saving and capital accumulation, which not only adversely affects the second generation directly but also the third generation (born in 2001-2020) who have to work with a lower capital stock and thus have lower labour income. From figure 4.7, the rate of return on equity reaches its peak in 2001-2020. This is because the second generation (born in 1981-2000) reduces savings and this results in a slump in an equity market. With an unusually low demand for this type of asset, asset prices fall considerably, which results in a high rate of return on equity the following period.

4.6 Conclusion

It is an open question whether pension funds should be held in the form of public debt or private equity. Indeed recent crises have led to strong arguments solely in favour of the former: so does this chapter.

It has been suggested that an optimal 'lifestyle investment plan' is to invest in equity during the asset accumulation phase (to obtain a high return) and bonds in the decumulation phase (to reduce risk). In this chapter, we analyze the effects of that savings behaviour on the macro economy. To do this we apply a three-period Diamond's (1965) overlapping generations model to study the optimal pension provision, allowing for demographic change.

In this chapter, we have taken the Diamond (1965) OLG model and extended it to model more explicitly the difference between the accumulation and decumulation phases of private pensions. We analyze a similar question to that of Diamond, namely the effect of government debt, but change the logic considerably by noting that government debt has different characteristics from private assets (i.e. equity) and that the two may be less close substitutes than is usually assumed. Our analysis of UK institutions suggests that this is a reasonable characterization of the pensions system.

Rather than crowding out private investment, government bonds provide an important part of the funded system. From this it follows almost automatically that it cannot be optimal to have no public debt. However, it is less clear what consequences this will have for the provision of pensions through a PAYG system: perhaps it could now be optimal to have positive state pensions.

We conclude that it is still socially optimal to rely entirely upon a funded pension. This arises because the taxation needed to fund a PAYG pension might change labour market behavior since labour is supplied inelastically; furthermore the optimality of government debt does not arise due to government using debt to provide public goods, since there are none in this model. Therefore the result here is a separate argument for preferring funded to PAYG pensions.

Thus our main conclusion can be quite simply stated that the usual preference for a funded to a PAYG pension holds even when the former is sustained through government debt.

Lastly, this chapter studies the effects of two alternative shocks to the economy and welfare: first a reduction in the size of the national debt and second, the effect of a baby boom. The results are as follows. When agents are relying upon government debt to finance their pensions, a sudden fall in national debt does not only affect the current generation, but also hurts future generations as well. The current middle age and retired are directly affected by lower returns on savings (in the bonds market) which in turn leads to lower levels of consumption and utility. Apart from people who are alive during the shock, the next three generations also suffer considerable declines in utility. After all, the new steady state value for social welfare declines too, hence people are worse off in the long run after the reduction in national debt. In contrast, a baby boom generates a small improvement to welfare in the long run. The generations immediately preceding and following the baby boomers obtain benefits from higher return to factors of production. On the other hand, the baby boom generation itself has relatively low utility. However, the two generations which suffer most are the second and third generations later. The second generation suffers the most since they need to pay the highest tax rates to support the retired baby boomers which results in lower saving and capital accumulation. This also affects the third generation since they have to work with a reduced capital stock and thus a lower labour income.

Figure 4.1: Social welfare contour map for early retiring case. (h=0.8)

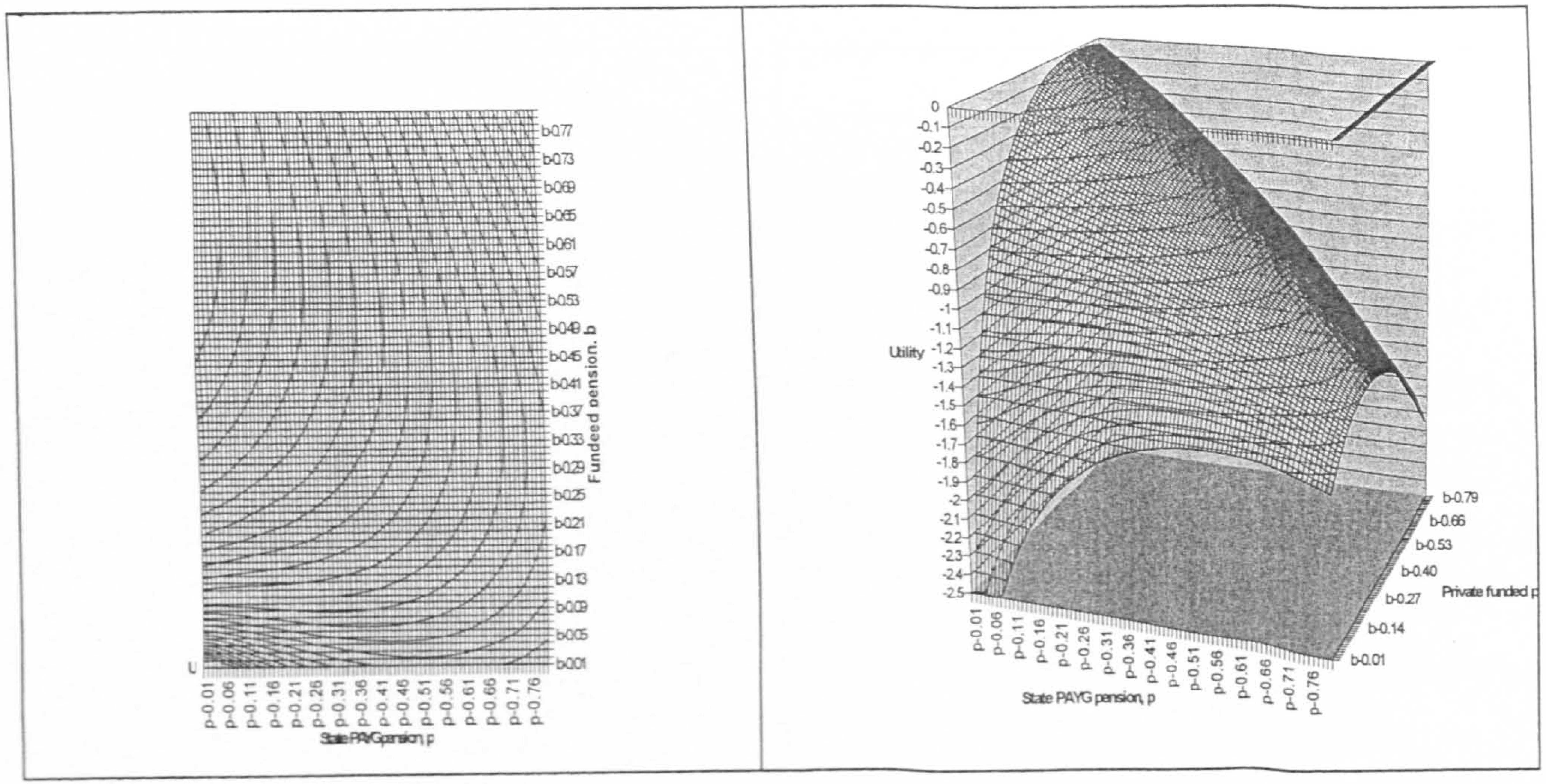


Figure 4.2: Social welfare contour map for working full time until the retirement age. (h=1)

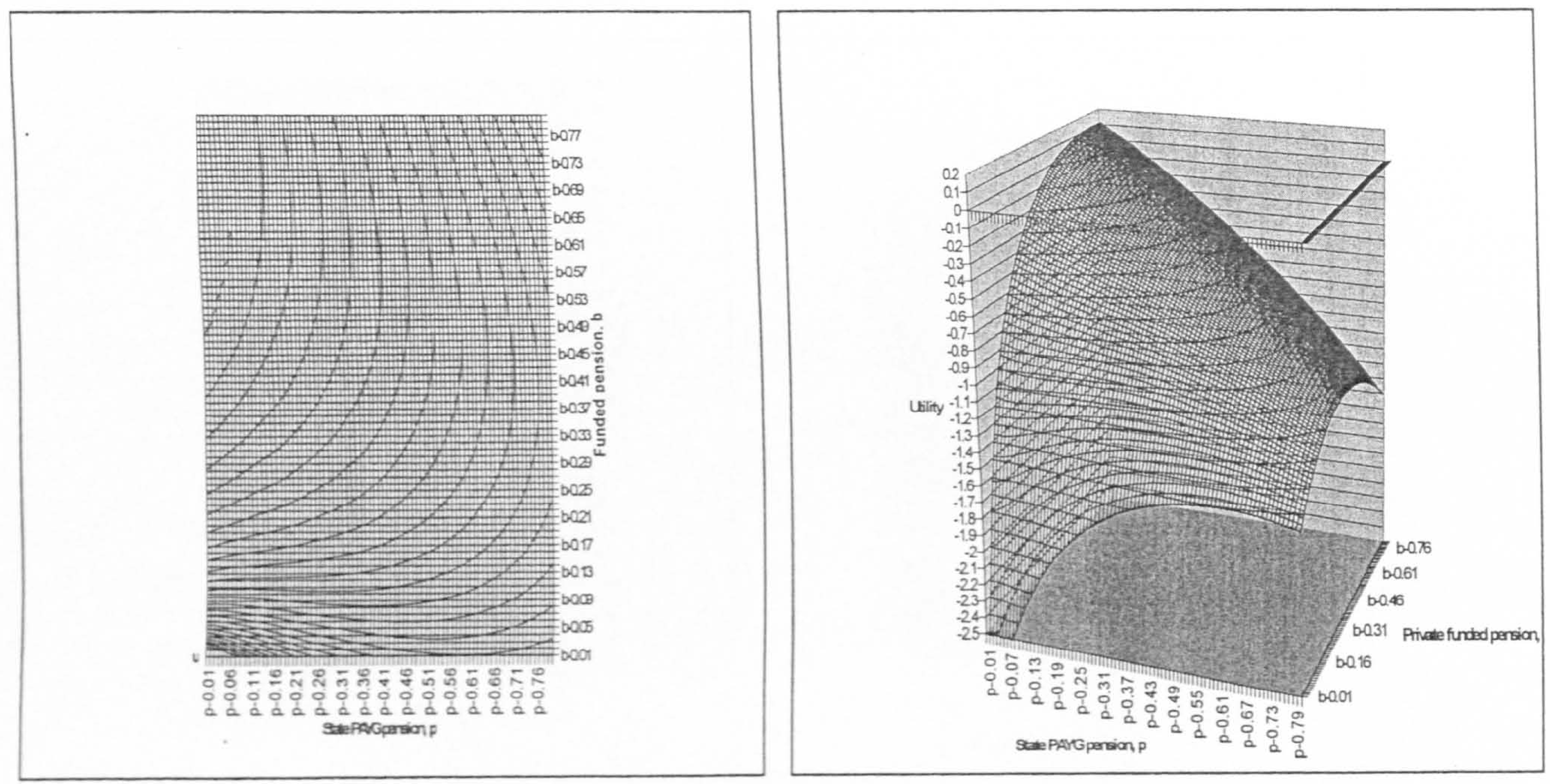


Figure 4.3: Social welfare contour map for the case of learning by doing. (h=1.2)

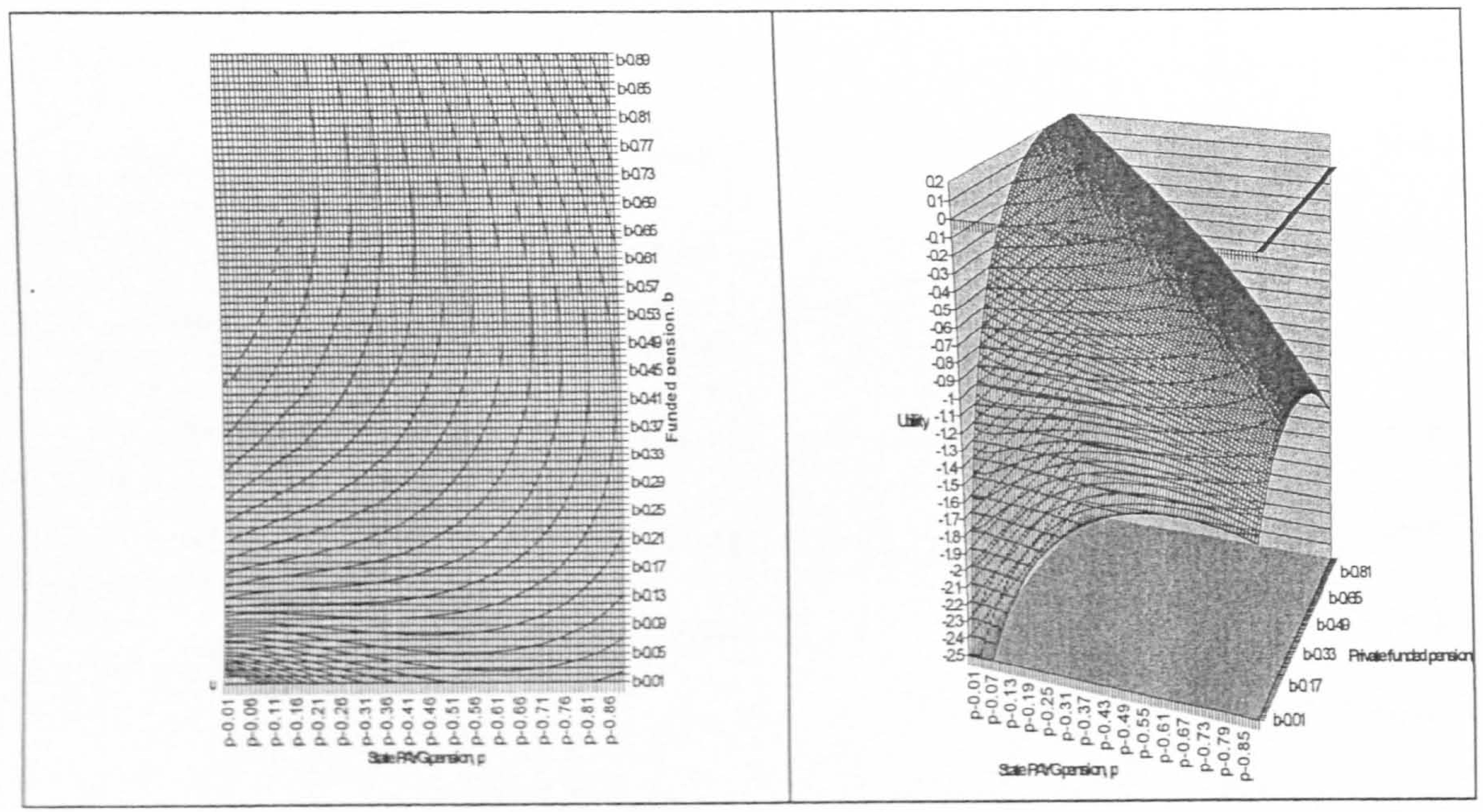


Figure 4.4: Annual Rates of Return on Equity with a fall in National Debt

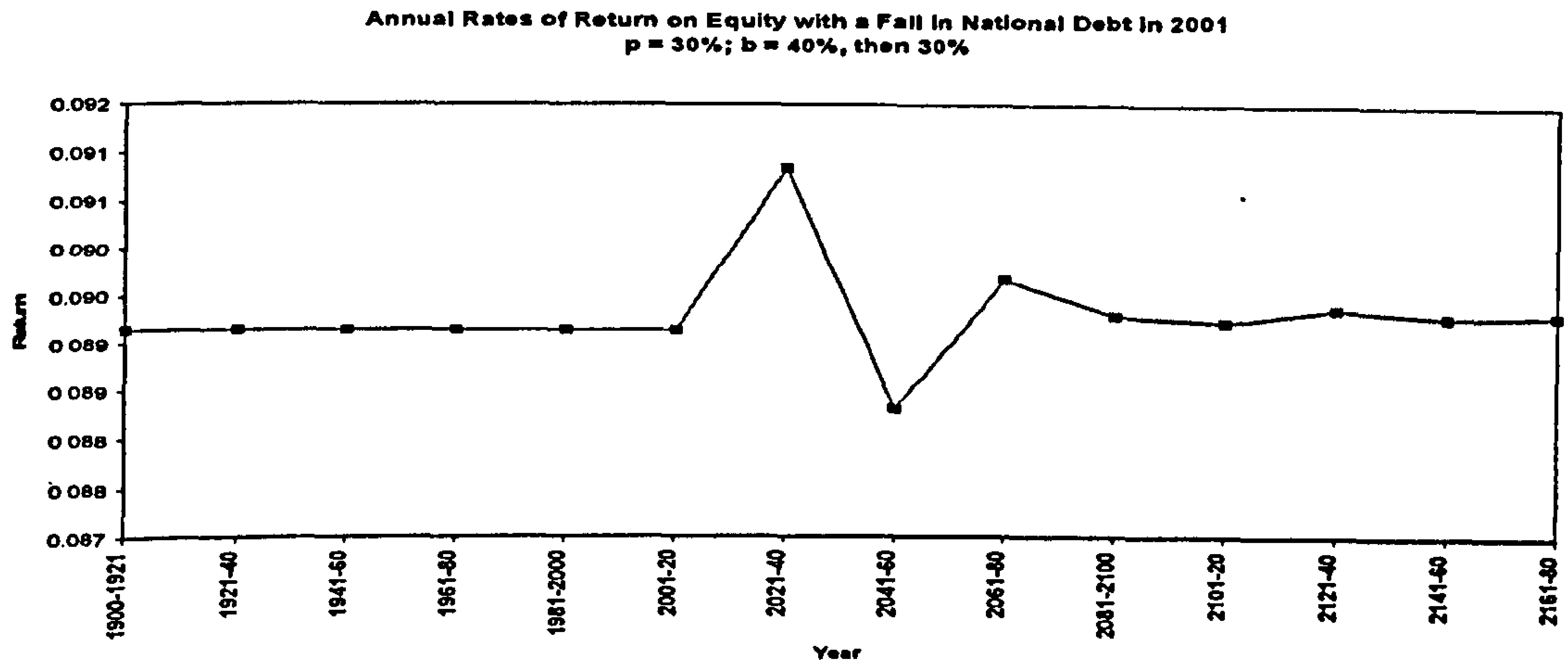


Figure 4.5: Annual Rates of Return on Bonds with a fall in National Debt

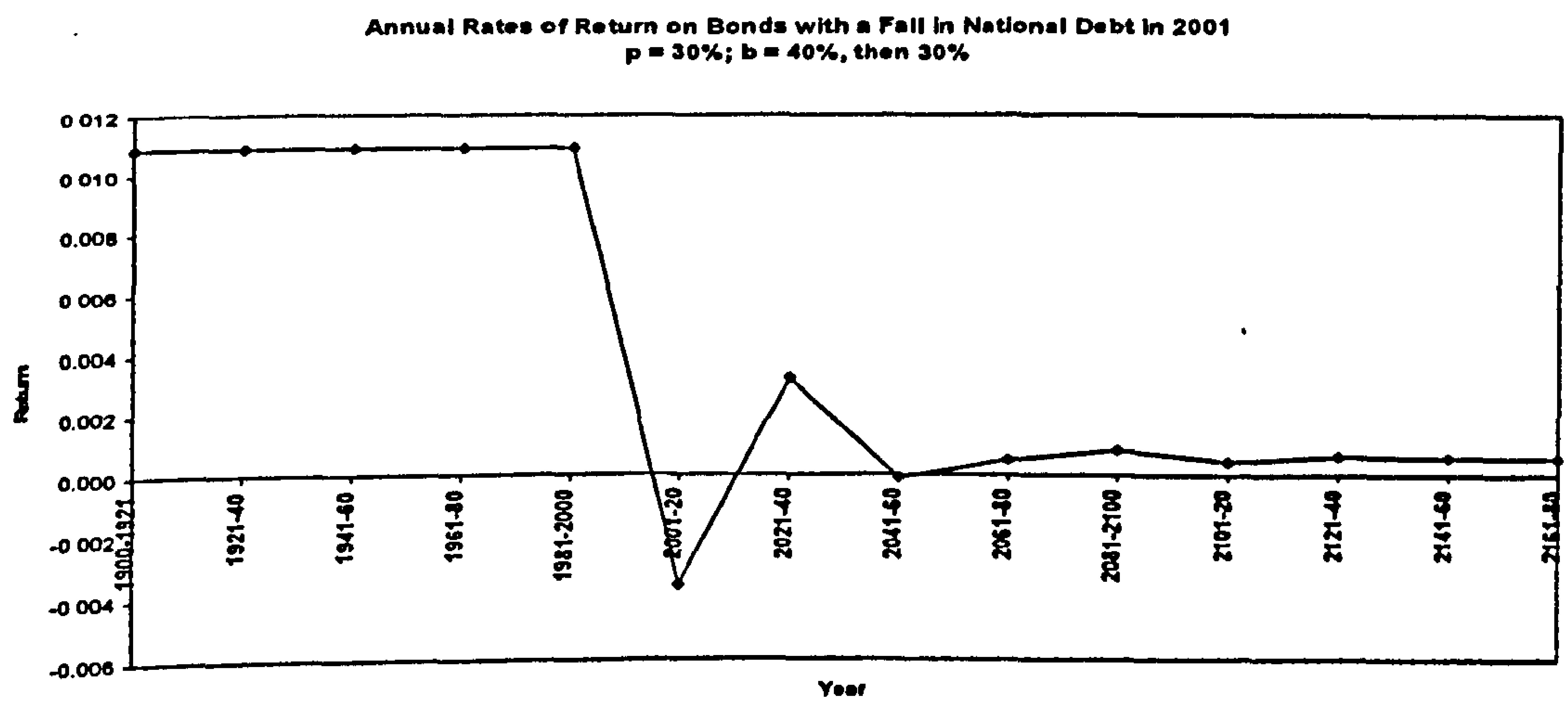


Figure 4.6: Utility with a fall in National Debt

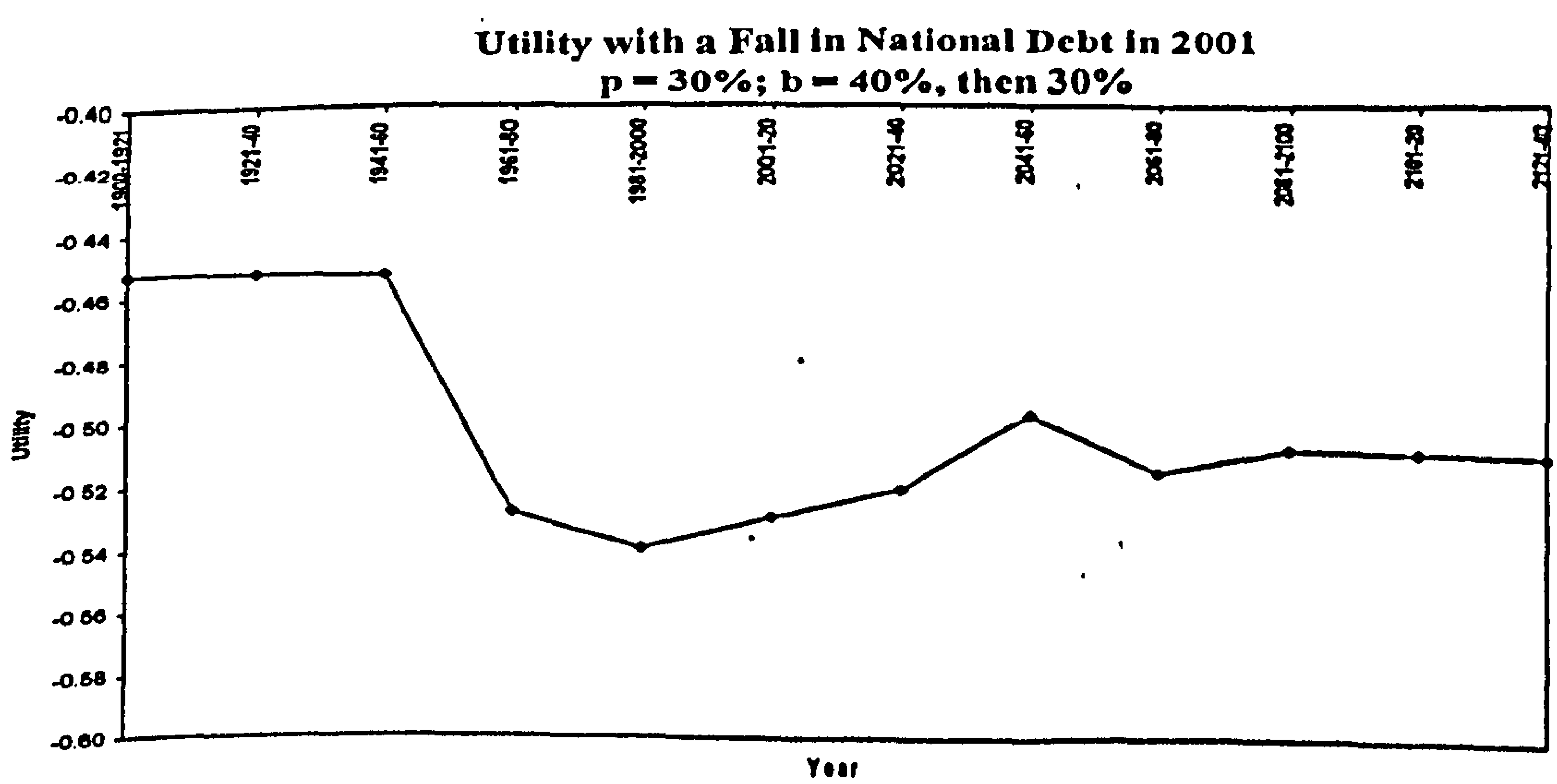


Figure 4.7: Annual Rates of Return on Equity with a Baby-Boom

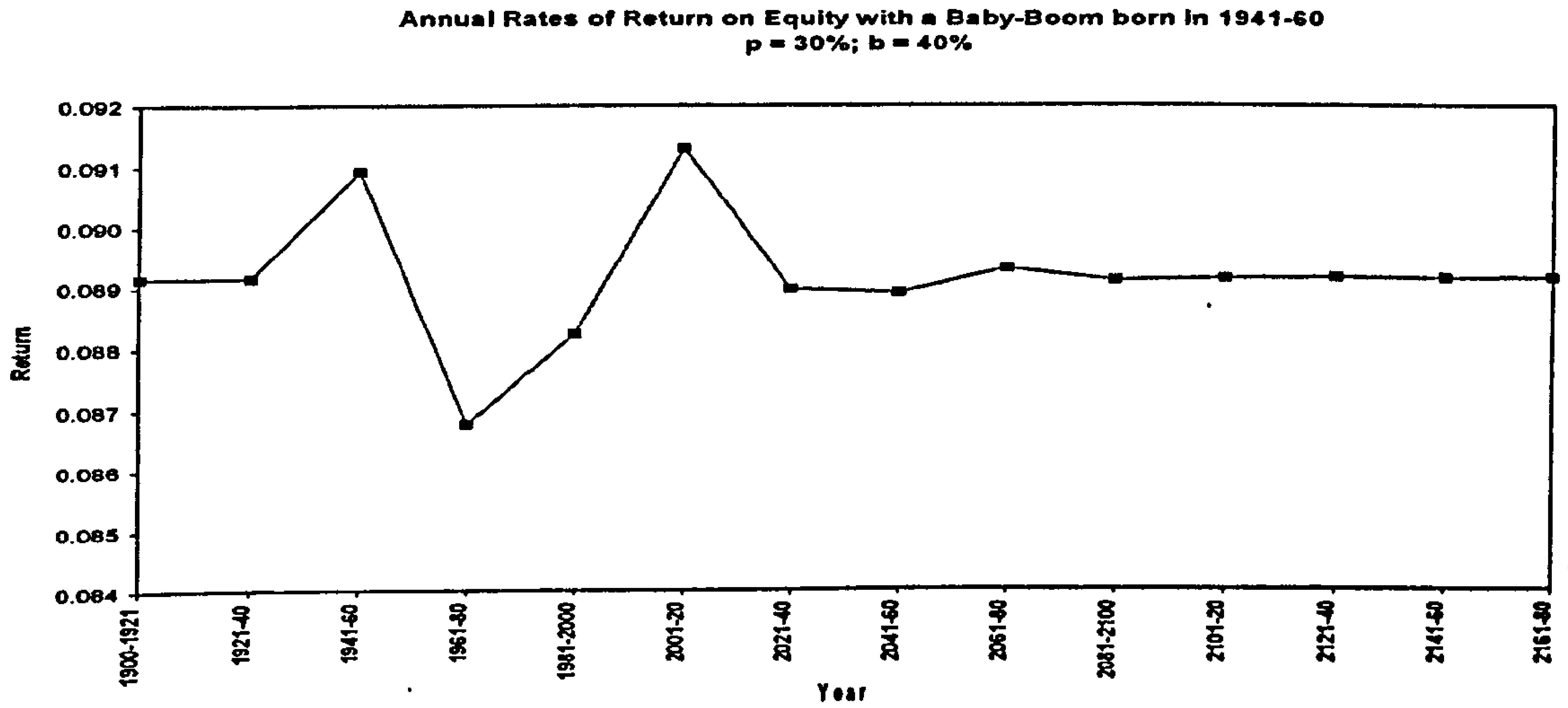


Figure 4.8: Annual Rates of Return on Bonds with a Baby-Boom

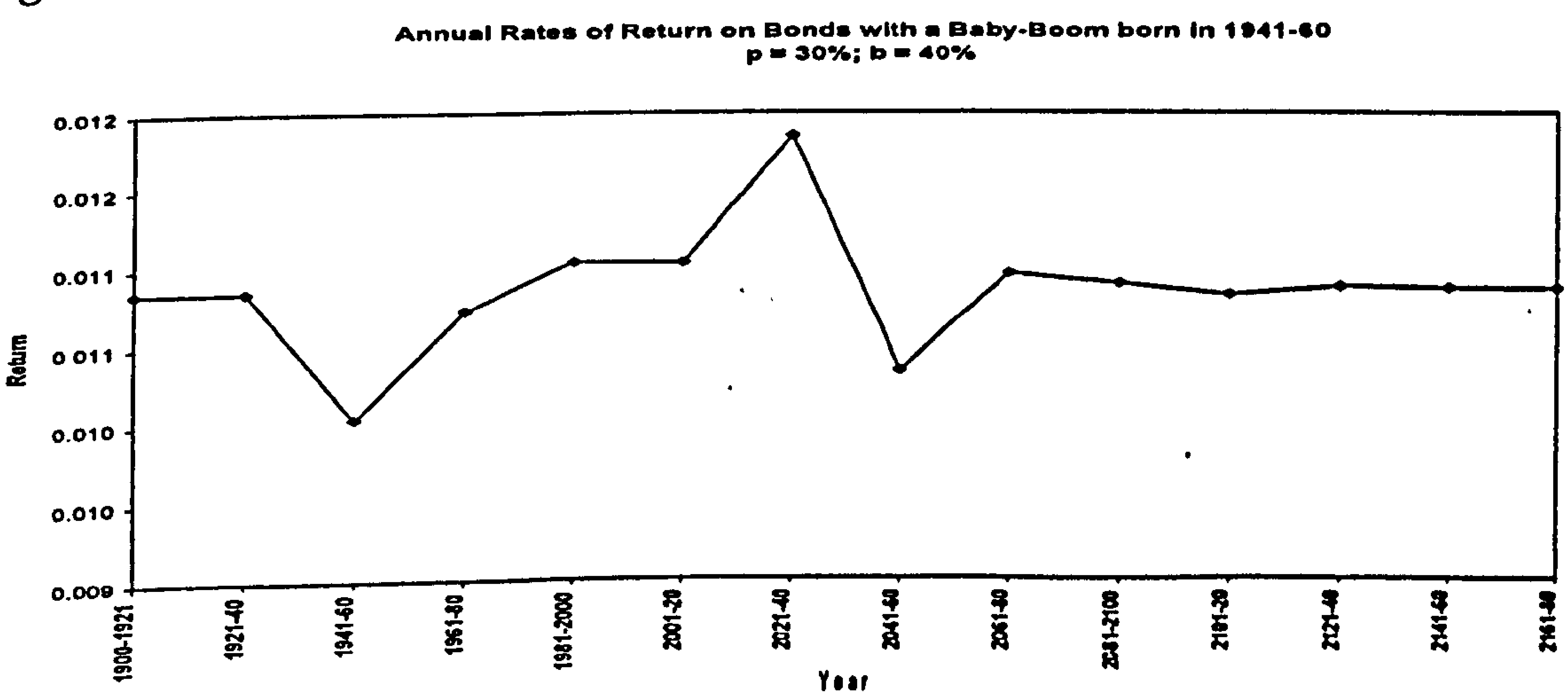
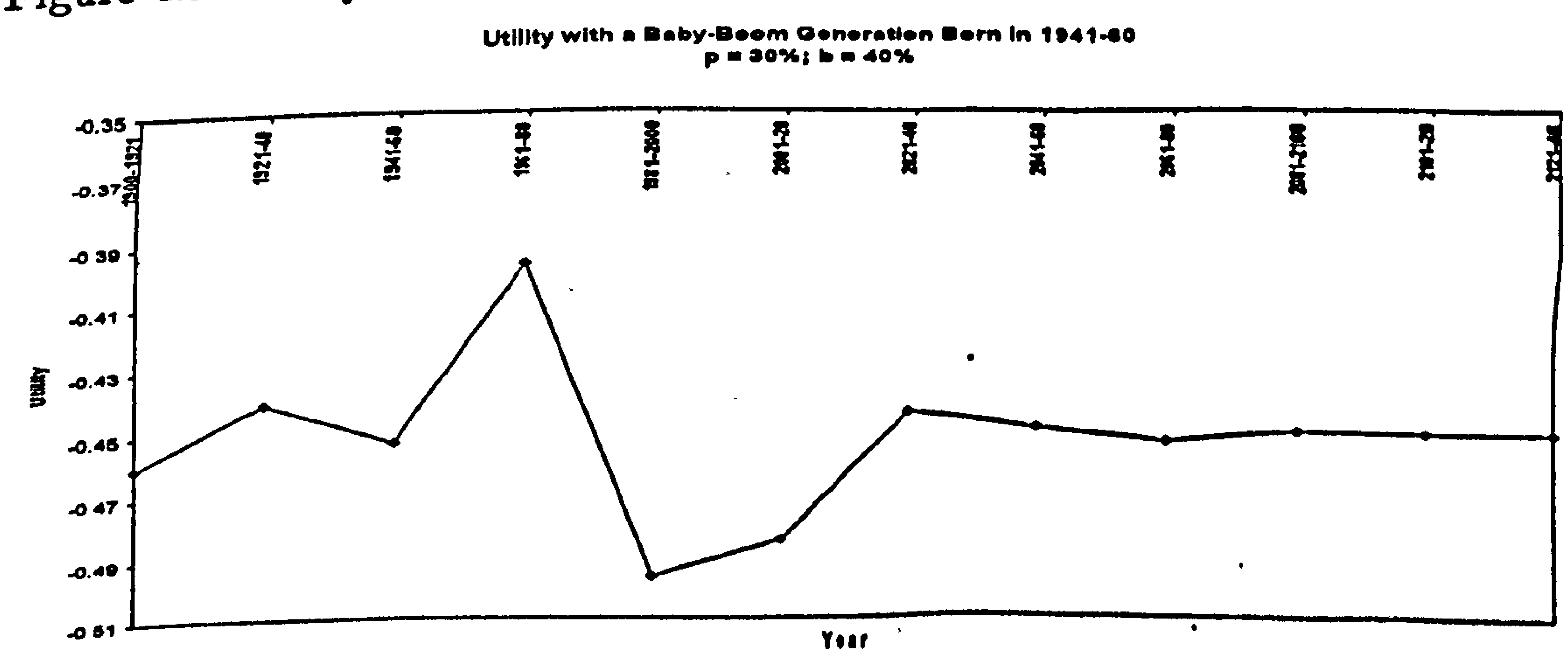


Figure 4.9: Utility with a Baby-Boom



Chapter 5
Conclusion

The first main chapter of the thesis is called "Determinants of the time varying risk premia". This chapter generates monthly risk premia data using zero coupon government treasury bills for 43 countries over the period of 1994-2006. The measure of risk premia is based on the ARCH-in-Mean (ARCH-M) model introduced by Engle, Lilien and Robins (1987). We show that the risk premia are time varying and also vary considerably between countries. This study also examines the macroeconomic and political determinants of the risk premia by using cross section regressions and dynamic panel regression analysis.

The cross section regression shows that on average through 1994-2006, the risk premia for holding government assets required by risk averse investors is positively influenced by the level of inflation and the budget deficit as a percentage of GDP and is negatively determined by the country's economic growth. Additionally, lower income countries are estimated to have risk premia about 19 percent more than in the high income countries outside the Eurozone, holding other variables constant. In the high income countries outside the Eurozone, the risk premia on holding government assets is predicted to be 10 percent more than those in Eurozone.

Using panel regression analysis, we found that economic growth and the volatility of the real effective exchange rate are the main determinants of risk premia in the full sample regression. Risk averse investors require lower risk premia for holding government assets in countries with good economic performance e.g. high economic growth and a stable external price competitive position e.g. low volatility of real effective exchange rate. The institutional variables and the government fiscal conditions have limited power in explaining the risk premia in this study.

Further analysis splits the sample by income group, namely, high income and low income groups. The results for the first group show that the main determinant of the risk premia is the real effective exchange rate; devaluation of the currency leads to better price competitiveness which in turn reduces the country risk premia. The opposite relationship is found in the regression of lower income countries. The possible explanation is that in financial vulnerable countries, weaker local currency can exacerbate the external debt service difficulties which result in economic contraction. This in turn raises the country risk premia. However, the impact of the level of real effective exchange rate is less strong in the low income group. For lower income countries, the volatility of the real effective exchange rate which reflects uncertainty in the exchange rate market plays an important role in determining the risk premia and there is a positive relationship between these two.

The next chapter, entitled "Fractional integration and the forward premium puzzle", corroborates earlier findings that a principal reason for rejection of the forward rate unbiasedness hypothesis is differences in persistence between the forward premium series and the spot rate return and the existence of the time varying exchange rate risk premium. First, the results show that the forward premium is best characterized by a (nonstationary) fractionally integrated process while the spot return series are found to be stationary, following an I(0) process in all sample currencies. This implies imbalance in the traditional regression of the return on the spot rate on the forward premium. Additionally, correcting for multiple structural breaks in the forward premium data (using Bai and Perron, 1998,2003a) shows that it reduces the persistence of the forward premium series, but the process is still found to be fractionally integrated.

Second, the research associates the exchange rate risk premia to the forward rate unbiasedness hypothesis. The issues of persistency in the foreign exchange data are also carefully dealt with. This chapter argued that the risk premia data should be proxied by the long memory in the conditional variance of the forward rate forecast error series. The exchange rate risk premium in this study is generated from the Fractionally integrated GARCH (FI-GARCH) model of the forward rate forecast error series. We argue that this model fits the data better than the multivariate GARCH model as used by Baillie and Bollerslev (1990) and the ARCH-in-mean model as used by Domowitz and Hakkio (1985)). We find strong evidence of statistically significant long memory parameters in both the conditional mean and conditional variance of the forward rate forecast error series. The exchange rate risk premium is found to be nontrivial.

The last main chapter of this thesis is entitled "Funded and PAYG Pensions when Annuities are backed by Bonds". In this chapter, we have taken the Diamond (1965) Overlapping Generation Model (OLG) model and extended it to model more explicitly the difference between the accumulation and decumulation phases of private pensions. We analyze a similar question to that of Diamond, namely the effect of government debt, but change the logic considerably by noting that government debt has different characteristics from private debt (i.e. equity) and that the two may be less close substitutes than is usually assumed.

Our analysis of the UK institutions suggests that this is a reasonable characterization of the pensions system.

Rather than crowding out private investment, government bonds provide an important part of the funded system. From this it follows almost automatically that it cannot be optimal to have no public debt. However, it is less clear what consequences this will have for the provision of pensions through a PAYG system: perhaps it could now be optimal to have positive state pensions. We conclude that it is still socially optimal to rely entirely upon a funded pension.

Lastly, this chapter studies the effects of two alternative shocks to the economy and welfare: first a reduction in the size of the national debt and second, the effect of a baby boom. The results show that when agents are relying upon government debt to finance their pensions, a sudden fall in national debt does not only affect the current generation, but also hurts future generations as well and people are worse off in the long run. The baby-boom also affects different generations and it generates a small improvement to the welfare in the long run.

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Appendix A
Appendix to Thesis

A.1 Regression results from LSDVC for whole sample;

A.1.1 Coefficients of the volatility of real effective exchange rate, $VREER$

Model 1 results: "as $\ln(VREER)$ increase by 1 unit, the $\ln(1+\frac{RP}{10})$ increase by 0.01 unit"

Antilog: as $VREER$ increase by 2.7183 unit, the RP increase by 0.1005 unit

thus as $VREER$ increase by 1 unit, the RP increase by $\frac{0.1005}{2.7183}$ =0.037 unit

(If $\ln(1+\frac{a}{10})=0.01$, then $a=0.1005$)

A.1.2 Coefficients of economic growth, $GGDP$

Model 1 results: "as $GGDP$ increase by 1 percent quarterly, the $\ln(1+\frac{RP}{10})$ decline by 0.461 unit"

Antilog: as $GGDP$ increase by 1 percent quarterly, the RP increase by 5.8566 unit

(If $\ln(1+\frac{a}{10})=0.461$, then $a=5.8566$)

Note: Annual growth rate is $GGDP_t = \log GDP_t - \log GDP_{t-4}$, but GDP growth percent quarterly is approximately $(\log GDP_t - \log GDP_{t-4month})/4$.

By approximation, $GGDP$ increase by 1 percent yearly, the RP increase by 1.4642 unit

Model 2 results: "as $GGDP$ increase by 1 percent quarterly, the $\ln(1+\frac{RP}{10})$ decline by 0.481 unit"

Antilog: as $GGDP$ increase by 1 percent quarterly, the RP increase by 6.1769 unit

(If $\ln(1+\frac{a}{10}) = 0.481$, Solution is: 6.1769)

Note: By approximation, $GGDP$ increase by 1 percent yearly, the RP increase by 1.5442 unit

A.2 Regression results from LSDVC for rich country;

A.2.1 Coefficients of real effective exchange rate, $REER$

Model 1 results: "as $\ln(REER)$ increase by 1 unit, the $\ln(1+\frac{RP}{10})$ increase by 0.059 unit"

Antilog: as $REER$ increase by 2.7183 unit, the RP increase by 0.60775 unit

thus as $REER$ increase by 1 unit, the RP increase by $\frac{0.60775}{2.7183} = 0.22358$ unit

Model 2 results: "as $\ln(REER)$ increase by 1 unit, the $\ln(1+\frac{RP}{10})$ increase by 0.058 unit"

Antilog: as $REER$ increase by 2.7183 unit, the RP increase by 0.59715 unit

thus as $REER$ increase by 1 unit, the RP increase by $\frac{0.59715}{2.7183} = 0.21968$ unit

A.2.2 Coefficients of the volatility of real effective exchange rate, $VREER$

Models 1 and 2 results: "as $\ln(VREER)$ increase by 1 unit, the $\ln(1+\frac{RP}{10})$ increase by 0.005 unit"

Antilog: as $VREER$ increase by 2.7183 unit, the RP increase by 0.050125 unit

Thus as $VREER$ increase by 1 unit, the RP increase by 0.01844 unit

A.2.3 Coefficients of economic growth, $GGDP$

Model 1 results: "as $GGDP$ increase by 1 percent quarterly, the $\ln(1+\frac{RP}{10})$ increase by 0.337 unit"

Antilog: as $GGDP$ increase by 1 percent quarterly, the RP increase by 4.0074 unit

By approximation: as $GGDP$ increase by 1 percent yearly, the RP increase by 1.0019 unit

Model 2 results: "as $GGDP$ increase by 1 percent quarterly, the $\ln(1+\frac{RP}{10})$ increase by 0.319 unit"

Antilog: as $GGDP$ increase by 1 percent quarterly, the RP increase by 3.7575 unit

By approximation: as $GGDP$ increase by 1 percent yearly, the RP increase by 0.93938 unit

A.2.4 Coefficients of Inflation, $INFL$

Model 1 results: "as $\ln(1+\frac{INFL}{10})$ increase by 1 unit, the $\ln(1+\frac{RP}{10})$ increase by 0.029

unit"

Antilog: as $INFL$ increase by 17.183 unit, the RP increase by 0.29425 unit

Thus, as $INFL$ increase by 1 unit, the RP increase by 0.017124 unit

Model 2 results: "as $\ln(1+\frac{INFL}{10})$ increase by 1 unit, the $\ln(1+\frac{RP}{10})$ increase by 0.027 unit"

Antilog: as $INFL$ increase by 17.183 unit, the RP increase by 0.27368 unit

Thus, as $INFL$ increase by 1 unit, the RP increase by 0.015927 unit

A.2.5 Coefficients of Government Budget Deficit (%GDP), $DEFGDP$

Models 1 and 2 results: "as $\ln(1+\frac{DEF}{30})$ increase by 1 unit, the $\ln(1+\frac{RP}{10})$ increase by 0.021 unit"

Antilog: as DEF increase by 51.548 unit, the RP increase by 0.212 22 unit

Thus, as DEF increase by 1 unit, the RP increase by 0.004117 unit