The Long-run Relationship Between Pension Liabilities and Asset Prices: a Cointegration Approach

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Abstract

This paper explores the long-term relationship between salaries and asset prices and relates it to asset allocation decisions of defined benefit pension funds. In particular, it discusses whether a 100% bond allocation is indeed the minimum risk position for an ongoing scheme whose benefits accruing to active members rise in line with their wages. The main finding is that, while shorter-run correlation evidence is less consistent, there is indeed a long-run link between the evolution of salary-linked liabilities and a range of asset classes including equities. This long-run positive linkage is consistent with economic theory. The implications for effective practice in minimum risk asset allocation depend on portfolio rebalancing costs, how much short and medium-run asset allocations move around the long-term equilibrium as well as how significant active liabilities are as a proportion of total pension liabilities.

JEL Classification: G23, C22, J30

Keywords: pension funds, asset allocation, cointegration

1 Introduction

This paper explores the long-term relationship between salaries and asset prices and relates it to asset allocation decisions of defined benefit pension funds. In particular it discusses whether a 100% bond allocation is indeed the minimum risk position for an ongoing scheme, where benefits accruing to active members are rising in line with their wages.

This is a slightly different question from the issue of optimal asset allocation, or, in other words, from the discussion on whether pension funds should take any extra risk relative to their minimum risk position, usually referred to as the liability-matching portfolio. The latter debate, spurred by a prolonged bear equity market and by the Boots pension fund's decision in 2001 to move its entire portfolio into long-term bonds, implicitly assumes that the minimum risk position is a portfolio fully invested in bonds (ideally index-linked). Underlying this view is the idea is that a bond portfolio can be constructed to match projected liabilities by cash flow, term structure and duration. Ryan and Fabozzi (2002) in their discussion of corporate pension funds liabilities in the US use the Ryan Labs liability index as a liability benchmark, an index constructed using the term structure of zero-coupon Treasury securities, Treasury STRIPS.

One of the key issues in determining minimum risk portfolios for pension funds is to define the liability. If the liability is defined on an accrued benefit obligation (ABO) or accrued rights basis, this is somewhat different than if the liability is defined on a projected benefit obligation (PBO) or on the basis of accrued rights with projected salary increases (Bulow, 1985). The liability could also be defined on a prospective benefits basis and in other ways. Each definition of the liability conceivably can result in different minimum risk portfolios.

Under any definition of liabilities, it is important to recognise that there will be residual risk. To begin with, bonds carry inflation risk unless an index-linked term structure (which is not available in many countries) is used. Furthermore, there is reinvestment risk if bonds are not available of long enough duration. Moreover, liabilities also carry significant mortality risk.

The focus of this paper is on the relationship between salaries and asset returns, which is an issue pertinent if the focus is matching liabilities on an ongoing basis. In fact, where a scheme is still open to new entrants or accruals, the expectation is normally that it will continue to remain so and therefore one should try to minimise risks on the basis of the most likely scenario rather than a scenario with more limited liabilities.

While the implications for pension fund asset allocation of these results depend on the objectives and situation of the pension fund, the core results of this paper are about the relationship between salaries and asset returns. The key contribution of the paper is to challenge the assumption of no correlation between aggregate wages and asset markets in the long-run. If there is a long-run relationship between wages and assets other than bonds and companies have no control over the wage-setting process, a 100% bond portfolio would no longer be a perfect match for active members' liabilities. This leads one to argue that an asset liability modelling exercise should carefully consider, alongside the relationship between prices and asset markets (extensively studied since Fisher, 1930), a separate relationship between wages and asset prices. The Appendix explicitly relates long-run relationships to pension funds minimum risk portfolios and shows that the results apply not only to traditional final salary pension schemes but also to career average plans.

There are two main approaches that can be taken when evaluating the relationship between asset returns and wages: a micro approach using household data, typically applied to individual lifestyle asset allocation in a defined contribution setting, and a macro approach, which investigates the economic forces shaping the aggregate relationship between wages, the stock and the bond market. Although a significant divergence arises when it comes to specifying the nature of the relationship between asset prices and wages, both approaches generally find that a relationship indeed exists. Section 2 reviews the existing literature which finds that the hypothesis of zero correlation has very limited support in both economic theory and the empirical literature.

This paper takes a macroeconomic approach and focuses on the long-run. There are clear reasons for this. Short-run dynamics are complex and contemporaneous correlations may not reflect adjustment to long-run predictions, particularly in markets such as labour markets where adjustments take time. Economic theory, nevertheless, suggests that correlations should appear in the long-run. An extension of this approach, which lies beyond the scope of the present paper, would consider the dynamic relationship between wages and asset returns, explicitly taking into account deviations from the long-run equilibrium path and the implications such deviations have for shorter term asset allocation.

With a shorter term focus it would also be possible to assess an actively managed strategy relative to an index-tracking strategy, while in the present paper the focus is on passive strategies. However the extent to which active strategies, exploiting shorter terms deviations from the equilibrium, are applicable in the context of pension funds can be disputed. The choice between passive and active strategies ultimately depends on the frequency of rebalancings. Here governance constraints can play an important role, as trustees might not have the time and expertise required to effectively supervise the implementation of active market-timing strategies, as highlighted by a recent survey performed by Watson Wyatt and Cranfield School of Management (Kakabadse and Robinson, 2001)

The empirical investigation in the paper is performed using macro data referring to the entire UK economy with the objective of shedding some light on the nature of the long-term relationship between the level of salary, the bond market, the stock market and other asset classes such as property. In order to relate economic theory and empirical results to the current debate on pension funds asset allocation, one should consider whether an aggregate salary index coupled with the use of salary scales is representative of the liability profile of pension funds.

A final salary or a career average pension fund may consider something in between the micro and macro approach as being most appropriate, because liabilities are affected by salary growth within the scheme members, which is obviously different from both individual wage profiles and the rise in aggregate earnings[1]. Even aggregate pension fund liabilities do not rise in line with aggregate earnings because aggregate earnings comprise companies and other entities that do not have a defined benefit pension scheme.

However, the bigger the population considered, the smaller will be the margin of error caused by using an aggregate earnings index. Moreover, aggregate earnings indices such as the Average Earnings Index calculated by the ONS in the UK are themselves based on a sample drawn out of all UK companies. There is a classical sampling theory issue here: to the extent that there is no systematic difference between individual wage profiles and demographic composition within the ONS sample and those at a specific firm or group of firms (i.e. one large company or all companies with a defined benefit scheme), an aggregate index is an acceptable proxy.

In particular, if the age composition of the scheme members deviates significantly from that of the general working population, there may be reasons to believe that wage growth within the scheme will follow a different path from aggregate earnings. If companies enter into long-term contracts with their employees and pay them less than their marginal product when they are young and more than their marginal product when they are old (Lazear, 1979), other things being equal, companies with a younger workforce will experience higher than average wage growth and companies with an older workforce will experience lower than average salary increases.

In addition to that, there might also be industry or regional patterns that will affect future earnings growth. While more disaggregated data would be useful in a pension liability context, the

[1] Differences are likely to be less severe if the scheme comprises members with heterogeneous income profiles and occupations as well as an age profile of the workforce similar to that of the aggregate index

constraint here is lack of data availability as well as comparability of historical data referring to a sub-group, as sub-groups have undergone significant transformation over time (mergers & acquisition, closure of DB funds etc.). An interesting line for future research would be to investigate in more detail whether there are systematic differences between salary increases within companies with defined benefit funds and the broader population, after correcting for other relevant factors.

Bulow (1985) discusses why the age distribution of earnings may be different in companies with final salary pension schemes. His argument is that because in defined benefit pension plans the employer gives higher pension benefits to older workers, this should be counterbalanced by lower salaries towards the end of the career than in firms without pension plans or with defined contribution arrangements. This would imply that companies with defined benefit schemes retain a margin of discretion over future wage increases and might be able to convince workers to accept lower than average salary increases to compensate for more generous pension benefits. However, whilst this is an important theoretical argument, there is little empirical evidence in support of this view at the present time.

From a company perspective, the conclusions of the paper on correlation between aggregate earnings and asset markets would be a key reference point, from which departures may be justified only to the extent that there are reasons to believe that salary increases within the firm will be systematically different from aggregate salaries (e.g. because the age structure of the workforce does not match the age structure of the salary index).

2 Review of the Literature

The traditional Markovitz (1959) portfolio choice framework hinges upon the two-fund separation theorem, which states that every investor holds the same risky assets portfolio, combined with the risk-free asset in different proportions, depending on individual risk aversion. Besides the problems arising when it comes to translating this result, based on a one-period optimisation, into long-term portfolio choice (see Campbell and Viceira, 2002), there is another simplifying assumption in this framework: the portfolio of financial assets is considered in isolation, without attempting to incorporate the labour income process. Bodie et al. (1992), Heaton and Lucas (1997), and subsequently Campbell et al. (1999), Heaton and Lucas (2000) and Davis and Willen (2000), have argued that individuals, in addition to their financial holdings, have an implicit holding in a non-tradable assets (human capital), whose dividends are typically paid every month in the form of salaries. In this context, portfolio choice can be significantly affected by the nature of the stochastic process for labour income.

Bodie et al. (1992) argue that labour supply flexibility, the risk profile of human capital returns and correlation between wages and risky asset returns influence individual investment strategy decisions. Davis and Willen (2000) show that the two-fund separation theorem breaks down if risky assets returns are correlated with the labour income process. They estimate occupational level income changes for 10 different occupations using US data from the Current Population Survey from 1967 to 1994. While they do not find significant correlation between aggregate earnings and a broad stock market index, they report significant evidence of co-movements between earnings in certain occupational categories and both the bond market and the Fama and French (1993) small cap stock portfolio (the SMB portfolio). Campbell et al. (2000) find that the correlation between occupational earnings changes and both long-term bonds and aggregate stock returns is higher for workers with higher education.

Using these micro approaches, the extent to which there is a pattern of correlation between asset returns and earnings innovations, or salary growth, is ultimately an empirical question. Macroeconomic approaches on the other hand look at the aggregate picture to discuss why there should be a correlation between the labour market and asset markets.

Textbook economics teaches that, given appropriate assumptions on the aggregate production function of an economy (i.e. constant returns to scale), the value of human capital should move in line with the value of the capital stock, unless factor income shares vary significantly over time. And, if asset prices reflect fundamentals and both future salaries and future dividends are discounted at similar rates, the outcome should be a fairly high positive correlation between salary growth and domestic stock returns (see for instance Baxter and Jermann, 1995).

However, complications arise when demand factors and savings-consumption decisions are incorporated into the model. Lucas (1978) assumed a long-run one-to-one relationship between dividends, consumption and output in the context of a pure exchange economy, where dividends are seen as a claim on the overall output of the economy. In a general equilibrium model of asset prices, productivity shocks are associated with the behaviour of asset prices, where assets are defined as "claims to all or part of the output" produced in a number of different productive units. In particular, Lucas (1978) showed that the relationship between asset prices and productivity shocks is complex even in a highly simplified economy with identical consumers and a single good produced. This is because a rise in productivity increases the demand for assets on the one hand, as individuals have more resources through higher wages, (the income effect), but on the other hand higher productivity carries new optimistic information about future dividends (the information effect), which may in principle have either a positive or a negative effect on asset prices[2]. The bottom line is that, even in the simplest economy one could think of, Lucas argues that "the relationship between real output and asset prices is far from simple and possible not even monotonic."

In a similar setting, Barsky (1989) attempted to shed more light on the relationship between the real economy and bond and stock prices. Barsky considers the effect on asset prices of (1) a fall in expected growth of productivity and (2) an increase in the dispersion (risk) of the future productivity growth probability distribution. The conclusions arising from the analysis are:

1. The real risk free interest rate falls when expected productivity growth decreases or its risk increases, for a given expected value. This is because individuals would save more in order to compensate for lower expected income or higher income risk in the future, raising prices and depressing yields on riskless securities (precautionary demand for saving). Barro and Sala-i-Martin (1990) reach a similar conclusion.
2. The effect on nominal bond returns is ambiguous because this depends on the stochastic process of money supply and money demand.
3. The effect on stock prices is also ambiguous because negative productivity shocks imply lower demand for stocks on the one hand, as future dividends are expected to be lower but, on the other hand, higher demand for stocks as a result of the precautionary saving motive, which tends to lower the required rate of return on stocks[3].

It appears therefore that there is indeed a link between the real economy and asset prices but it is a very difficult exercise to predict asset prices through productivity shocks. This might explain why most attempts to predict the behaviour of the stock market used valuation ratios such as the dividend yield or the price earnings ratio as predictors, rather than explicitly relating asset prices with trends in the real economy proxied by variables such as productivity or real salary growth.

Santos and Veronesi (2001) is one of the few empirical attempts to relate stock return predictability with the real economy. They estimate a model with a time-varying equity risk premium depending on the ratio between labour income and consumption. The intuition behind their approach is that the required rate of returns on stocks will be lower when labour income as a

[2] The positive effect arises because individuals would demand more assets in anticipation of higher dividends, while the negative effect has to do with the fact people would consume more and save less as they anticipate higher salaries in the future

[3] This is true only if productivity shocks do not increase the correlation between stock market returns and consumption because in this case the required rate of return would tend to increase to reflect the higher risk of investing in the stock market

percentage of total income is high because in this case the bulk of consumption is not related to stock market returns.

The macroeconomic empirical literature has thus far failed to reach a consensus on the nature of the relationship between asset prices and salaries. Fama and Schwert (1977a) found little evidence of correlation between income and stock returns with post WWII data because, while income provided a reasonable inflation hedge, in the same period stock returns and inflation had been negatively correlated. Subsequent studies, focusing on correlation between wages and domestic equities have come to different conclusions.

Bottazzi, Pesenti and van Wincoop (1996) even found a negative correlation pattern in most countries and argued that wages could move in the opposite direction to stock prices due to shocks affecting the bargaining power of workers or positive demand shocks asymmetrically affecting real wages and real profits, when wages are less flexible than prices. The problem with this approach is that, although it appears to represent an accurate description of certain observed patterns (i.e. the 1970s in Britain), a negative correlation becomes problematic if its long-run implications are considered. If stock prices grow in the long-term, the counterintuitive implication is that real wages should fall and this would be probably consistent only with a neo-Marxist interpretation of capitalist economies.

It is probably more appropriate to consider these results as the product of relatively short-term dynamics, whereby demand factors (as in Barsky, 1989) play a key role or real wages and productivity are not aligned. Ball and Moffitt (2001), Blanchard and Katz (1997) and Stiglitz (1997) argue that real wages are tied closely to labour productivity in the long-run, although they can deviate substantially in the short-term because shifts in productivity may not be immediately reflected in shifts in wage aspirations. Conversely, in a Barsky (1989)-type of model, if a productivity shock causing present and future productivity to rise leads people to save less, depressing stock prices, this could eventually depress productivity, maintaining a long-term relationship between the equity market and productivity.

The implications of the discussion so far are the following: portfolio choices would be significantly affected if human capital returns are correlated with financial returns and economic theory would suggest a positive long-term relationship between stock returns and wage growth.

If this is indeed the case, a 100% investment in bonds may not be the minimum risk portfolio to hedge defined benefit liabilities. Sharpe (2002) shows that pension fund liabilities can be incorporated in an asset-only optimisation problem by taking into account the covariance between asset returns and liabilities. Within this framework, a positive allocation to stocks would probably have risk-reducing properties if there is a link between wage-related liabilities and equities.

There is an ongoing debate within the actuarial community on the correlation between equities and salaries and its implications for defined benefit plans asset allocation. Wilkie (1995) developed a stochastic model for actuarial use which does not include a relationship between equities and salaries in real terms, although he finds some evidence of correlation between dividends and wages. The paper also considers a vector autoregressive model of wages and prices and, although the estimated coefficient in the price-salary relationship is greater than 1, a one-to-one relationship model is recommended.

Exley, Mehta and Smith (1997) attempt to price a hypothetical National Average Earnings-linked bond and argue that wage growth has been historically better hedged by index-linked bond income as opposed to equity income (dividends). They also estimate a regression model to investigate the nature of the hedging portfolio for NAE growth if index-linked instruments are not available and find no substantial equity component. Smith (1998) derived a discrete representation of a continuous time stationary process driving salary increases and fitted an ARMA (1,1) model, finding a coefficient lower than one for prices and an insignificant relationship between salaries and dividends. This would prove the hedging portfolio for salaries is made up by a mixture of index-linked as well as conventional gilts, given that wage increases are muted at times of high inflation.

However, these contributions hardly provide a definite conclusion on the role of equities in minimum risk portfolios of defined benefit pension funds. Firstly, hedging portfolios have not been stable over time and may also depend on the mix of assets included in the analysis. Randall and Satchell (1999) show that the equity share rises to 25% if the period 1981-1996 is considered and the 1970s are excluded. It is worth noting that correlation between earnings and prices may also depend on institutional features of the labour market, such as union-sponsored pay deals or collective agreements with compulsory indexation, which are the norm in France and Germany (the other two countries considered by Exley, Mehta and Smith, 1997) and in the UK were more prevalent in the 1970s than today.

Secondly, hedging portfolios essentially rely on short-term correlation of returns. However, pension funds do not make payments indexed to the National Average Earnings Index and therefore do not need to match salary growth on a period by period basis. If one takes this into account the question becomes: what is the most appropriate mix of assets matching liabilities indexed to the long-term growth in the salary index? Clearly, the exact definition of long-term depends on the specific characteristics of the fund (maturity of active members' liabilities).

3 Data

Ideally, to be able to test the relationship between wages, productivity and the equity market, we should use indices referring to the same companies. In practice, this is not possible because aggregate earnings indices and the broadest stock market index do not include the same firms: some listed firms will not be accounted for in the wage index, if it is calculated from survey data, while the salary index will comprise several non-listed firms. This is an important methodological caveat and, when interpreting patterns in the data, it must not be forgotten that measurement errors can arise there. This is also why we focused on the broadest available indices, rather than considering sub-groups, in which the likelihood of measurement error is significantly higher.

Three UK data sets are considered in most of the analysis: a monthly and a quarterly one for the period 1963-2002[4], and a long-term annual data set (1850-2001). In the quarterly data sets two alternative salary indices are considered: the longest index based on survey data: the Average Earnings Index (AEI) with base 1995=100, supplied by the Office of National Statistics[5], and the total compensation of employees series out of UK national accounts (from now on referred to as wages). The latter enable to track the evolution over time of the share of total output accruing to workers. Figure 1 shows the two indices are broadly similar, although during the recession of the early 1990s the AEI index is smoother than the national accounts series. In the monthly data set only the Average Earnings Index (AEI) series is available.

In the annual data set, a wage index is derived using the historical index in Mitchell (1994), until 1963, consistently with other studies of long-run economic relationships in the UK (i.e. Chadha and Nolan, 2002), and the average earnings index thereafter. There are two series for real wages: real wages (1) created by deflating Mitchell (1994) money wages using the consumer price index in Global Financial Data[6], while real wages (2) uses the real wage index in Mitchell (1994) until 1913, and is equal to real wages (1) thereafter. The results are presented using real wages (2), which employs a consistent approach to track historical money wages and prices.

A historical stock market index developed by Global Financial Data tracks the performance of Bank of England stocks until 1917 and the broadest available panel of companies thereafter. For bonds a composite index was constructed using total return on consols before 1932 and 10-year

[4] The quarterly sample comprise data from the first quarter of 1963 to the third quarter of 2002 (1963:1 to 2002:3), while the monthly one is from January 1963 to October 2002 (1963:1 to 2002:10)

[5] Both the seasonally adjusted and the non-seasonally adjusted versions of the index are considered, but the results presented in the paper use the seasonally adjusted version

[6] This index uses a variety of sources including Mitchell (1998) for the 19th century and until 1915 and is equal to the retail price index calculated by the Central Statistical Office (Office of National Statistics after 1996) thereafter

government bonds thereafter, both supplied by Global Financial Data. This is because until the 1930s consols were the only segment of the UK bond market with substantial liquidity, as documented by Dimson, Marsh and Staunton (2002). In the quarterly and monthly data sets the equity index corresponds to the FTSE Actuaries General/All-Share index and the bond index is the 10-year gilt total return index. For index-linked bonds, a price index available from Global Financial Data and a total return index based on CAPS data were used since 1981, when the first index-linked bond was issued.

As a proxy for short-term investments we chose the bill total return index calculated by Global Financial Data, which uses the 3-month yield on commercial bills from 1800 to 1899 and the yield on treasury bills from 1900 onwards. As a proxy for international equity investments the paper uses the S&P 500 both converted in pound sterling with end of month $/Pound historical exchange rates from Global Financial Data.

Finally, a house price index, the Nationwide House Price index, which is available from the early 1950s is also considered in the quarterly and monthly data sets. Unlike the other indices, this is not a total return index but reflects only the capital appreciation component of property investments. A reliable historical total return index incorporating rental yields is however not available. In this context, house prices were chosen as a proxy for the behaviour of pension fund property portfolios, which are however largely made up of commercial property. However, indices of commercial property total return are not available on a consistent basis since the 1960s, as discussed by Cardinale (2003) in more detail.

Real salary, equity, bond, bill and house prices series are constructed using the Retail Price Index (RPI), calculated by the Central Statistical Office (Office of National Statistics after 1996), as a deflator.

In order to better understand the relationship between wages and the stock market the paper also considers measures of corporate profits, dividends and productivity, which are available on a consistent basis for the quarterly data set. There are two indices of corporate profits: earnings per share of listed companies, calculated as the ratio of FTSE All Share price index and its Price/Earnings ratio, both supplied by Global Financial Data, and gross profits of corporations from national accounts. The two series differ mainly because the second includes profits of non-listed companies and self-employed. Dividends are calculated as the product of the Dividend Yield series and the FTSE All Share price index, both supplied by Global Financial Data. For productivity the two most widely employed productivity measures calculated by the Central Statistical Office (Office of National Statistics after 1996) are employed: output per filled job and output per job.

Finally, the last section makes reference to a longer-term monthly sample (January 1920-March 2003), which uses a historical inflation and stock market total return indices from Global Financial Data, in conjunction with a historical wage index, calculated by Watson Wyatt combining historical aggregate wage indices from Central Statistical Office (interpolated between 1920 and 1934), and a longer term bond index, calculated by Watson Wyatt using consols until 1981 and over 15 years bonds thereafter (Global Financial Data uses 10-year bonds).

4 Time-series properties of wages and asset prices

This section focuses on basic time-series properties of wages and assets using the quarterly and the annual data sets described in the previous section. Figures 1-3 plot nominal and real series in the quarterly data set. Nominal wages, bonds and bills display almost uninterrupted increases, while UK and US equities as well as house prices appear to be characterised by breaks in their growth process. For UK and US equities this happened with the market downturn in the late 1990s, while for property the same can be said for the fall in house prices of the early 1990s. With variables in real terms the pattern is even less predictable. Real salaries, bonds and bills fell in the 1970s while real property display three clear 'boom and bust' cycles, consistently with the Grenadier (1995) hypothesis.

Figures 4-6 take a longer term view and plot nominal and real seal series from 1850 using the annual data set. We broke the sample in two for nominal variables (1850-1949 in Figure 4 and 1950-2001 in Figure 5) for convenience because wages and nominal assets have grown with unprecedented speed after WWII. Interestingly, no series displays a steady growth pattern if pre-WWII data are included in the sample. Equities displayed two 'boom and bust' cycles between the 1920s and 1930s, while salaries saw a spectacular increase in the 1910s, followed by a drop in the early 1920s, and nominal bonds suffered between 1900 and 1930. Interestingly, real bonds display a steady growth only after the 1970s, given that before they fell twice (mid 1920s and mid 1970s) and were stable in between. Real equities on the other hand display an unprecedented growth after the 1950s, in line with the pattern observed for the nominal series.

Assuming that the series are difference stationary (e.g., I(1)) [7], a reasonable assumption for many economic series, we can focus on the properties of the first difference series (the assumption will be formally verified in Section 7 with unit root tests). As the analysis is carried out in log

[7] Difference stationarity means that although the series in levels trend over time, the series obtained by taking the first difference of each element of the original series (if the original series is y_t the first difference series is obtaining by working out $y_t - y_{t-1}$ for each t) is stationary. Nelson and Plosser (1982) were the first to show that important macroeconomic variables are difference stationary rather than trend stationary

terms, first difference series will be equivalent to log returns (or log salary inflation) series. Log or continuosly compounded returns are widely used in finance because the assumption of normality is empirically more robust for log rather than simple returns. However log returns underestimate true underlying simple returns and the magnitude of underestimation is higher the higher is the variance in absolute terms (i.e. with annual returns). Fama (1976) is a key reference on this.

Basic sample statistics are presented in Table 4.1 (nominal returns) and 4.2 (real returns) for the quarterly sample and in Table 4.3 for the annual data set. As already observed before, stocks have grown at unprecedented speed after the 1950s[8]. In the 1963-2002 quarterly sample UK equity log returns have been on average 2.95% quarterly in nominal terms (11.8% annually) against an overall 7.2% average in the annual 1850-2001 sample. Similar results arise for bonds and wages, both in nominal and real terms, suggesting that the economy as a whole has experienced a faster rate of growth between the early 1960s and 2002.

Contemporaneous correlation matrices between nominal and real variables are presented in Table 4.4-4.6. Interestingly, in the 1963-2002 sample only bills and house price growth have been highly positively correlated with nominal wage growth, both using the AEI and the total compensation index. In real terms bonds were also positively correlated with wages, especially using the national accounts index. Equity returns on the other hand have been negatively correlated or uncorrelated with salary growth, both in real and nominal terms.

With the annual sample neither bonds nor stocks have been highly correlated with nominal wages while in real terms both real bonds and real stocks were highly correlated with real wages (1) (derived using the Global Financial Data price deflator). We also considered contemporaneous correlation in real terms using index-linked bonds (with monthly data), but the results are only referred to the sub-sample 1982-2002 and are not shown in the table. Index-linked bond returns are highly correlated with real bond returns using the deflated 10-year government bond series (around 50%) while correlation between index-linked returns and real salary growth is lower than with real bonds (less than 10%).

However, contemporaneous correlation is not an adequate yardstick when it comes to evaluate a long-run pattern of dependency because co-movements are not likely to be contemporaneous. We therefore considered the time-series properties of wages and asset returns, in particular autocorrelations and cross-correlations. Partial autocorrelation computed with quarterly data are presented in Table 4.7.

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[8] Cardinale (2002) shows this in detail by considering two centuries of stock and bond markets data

With the 1963-2002 quarterly data set nominal salary returns are highly positively correlated with returns in the previous quarter (around 60% correlation both with AEI and national accounts) and with returns with a lag of 4 or 5 quarters (over 20% partial autocorrelation). House price growth displays a similar pattern (62% positive correlation with previous quarter growth and partial autocorrelation of 21% with three quarter lagged growth). Bill returns are characterised by a very high positive correlation with returns in the previous quarter (94%) but they display signs of mean reversion at lag 2 (-24% partial autocorrelation).

Conversely, bond and equity quarterly returns are not significantly correlated with returns in the previous quarter and present a lower degree of correlation even at higher lags. Bond returns are only significantly negatively correlated with returns 8 and 9 quarters before, perhaps indicating a pattern of mean reversion with a length comparable to a business cycle. For equity returns there appears to be a positive dependency at lag 1 and a negative correlation at lag 4 but none of these autocorrelations is significant at conventional levels (5%) because they all lie within two standard error bounds.

In real terms, salary growth displays a high and significant negative correlation with two-quarters lagged growth (three-quarters with the national accounts index) and a positive correlation with four and eight-quarters lagged growth. With both indices autocorrelation at lag 1 is no longer significant.

Interestingly, real property and stocks returns are characterised by a very similar autocorrelation pattern to their corresponding nominal returns. Real bonds on the other hand display significant partial autocorrelation at lag 2 and 4 (positive) and 9 (negative). Finally real bills characterised by high positive dependency at lag 1, 2, 4 and 8.

With the annual data set (not shown), salary growth is significantly positively correlated with 1-year lagged previous year salary increases (58%), and with 5-year lagged salary growth (19%). Nominal stocks display positive dependency at lag 4 while signs of long-term mean reversion (negative partial autocorrelation) are visible only with a 15-year lag. Nominal bonds are characterised by positive and significant dependency at lag 1, 2, 5 and 8.

In real terms using both price deflators, wages display a pattern of long-run mean reversion at higher lags, similar to real stocks, while real bonds are only characterised by positive dependency at lag 1.

In summary, the analysis of autocorrelation functions suggests a similarity between the process driving nominal wages and house price growth in the quarterly data set. In the annual data set nominal bond returns appear to be characterised by a more similar process to nominal wages with respect to equities. In real terms however no asset return process closely resembled real salary growth in the quarterly 1963-2002 sample, although real bonds have been a better proxy than real stocks, while in the annual long-term data set real equities have been a better proxy than real bonds.

The next step is to consider explicitly the correlation between asset returns and salary growth using both lags and leads. This is performed using cross correlograms and the results (using the average earnings index and the quarterly data set) are shown in Figures 7 and 8. Positive lags refers to the relationship between salary growth and past asset returns, while negative lags represent the relationship between asset returns and past salary growth.

In nominal terms the most striking patterns of dependence in the quarterly data set are those between salary growth and property returns and between salary growth and bill returns. With house prices, positive correlation between salary growth and up to 12-quarter lagged property returns is always higher than 20% (reaching 50% with a lag of 9 quarters) and is statistically significant. Conversely, correlation between property returns and past salary growth is not as strong. With bills, there is a stronger relationship between bill returns and past salary growth. This overall pattern might be interpreted as follows: housing booms lead to higher nominal wages and higher inflation, which put pressure on interest rates. Finally, with bond and equities there appears to be a positive, albeit weaker, relationship between bond and stock returns with past salary increases (especially at lag 2 or 3 and 10).

In real terms the pattern is different. There is evidence of negative correlations of bills, bonds and equity returns with past salary increases. This is evident at shorter lags for bills (2 and 3 quarters) and higher lags for stocks and bonds (7 quarters). This can be read as evidence of a negative effect of high real salary increases on subsequent real interest rates, possibly because of increasing demand for savings (income effect). There appears also to be further evidence of mean reversion for the real bond process, given that contemporaneous correlation with wages is positive and significant, while it is less straightforward to justify for equities given that correlation at shorter lags was not significant.

Real salary increases on the other hand appear to be positively correlated with past stock and bond returns 12 quarters before (16%), 2 quarters lagged bond returns (21%), 8 quarters lagged house price growth (17%) and real bills returns 4 and 8 quarters before (respectively 26% and 32%). This would suggest that healthy real returns in asset markets contribute to create the conditions for real salary growth, although this pattern of dependency does not follow symmetrical cycles in different asset markets.

With the annual long-term data set, cross correlogram analysis shows a stronger positive relationship between bond and equity returns with past nominal wages (with a peak corresponding to a 7-year lag, 37% correlation for bonds and 28% for equities). Nominal wages on the other hand display a moderate level of positive correlation with lagged bond returns (18% with a lag of 3 and 7 years) and no significant correlation with past stock returns.

In real terms there are little signs of a significant pattern of cross-correlation both between bonds and salaries and between equities and salaries. The only exceptions are negative correlation between real salary growth and 1-year lagged bonds and positive correlation between real bonds and 4-year lagged real salary growth.

In conclusion, cross-correlation analysis suggests a complex pattern of dependency and interactions between salary growth and asset markets and not a perfect relationship with any of the assets taken individually.

5 Long-term Economic Relationships: Cointegration

At the heart of economic theory is the notion of equilibrium and most theoretical models describe patterns of dependency or interaction between economic variables when markets are in equilibrium. Examples of equilibrium relationships are equations linking market prices (i.e. stock prices, bond prices, exchange rates) to "fundamentals", which are the factors believed to be relevant to determine prices.

Granger (1981) argues that many economic variables may drift apart in the short-run, but, if they continue to be too far apart in the long-run, economic forces (the market itself or in some cases even government intervention) will act to bring them together again. Examples of this could be interest rates on assets of different maturities, prices and wages or prices of a commodity in different parts of the same country.

However, assessing whether long-term relationships truly hold is essentially an empirical question. Markets may not reflect fundamentals because either they are irrational even in the long-term or because fundamentals are not correctly specified. Granger (1981) introduced cointegration as a way of statistically characterising equilibrium. Cointegration in itself does not imply equilibrium in any behavioural sense, it just describes the tendency of two or more economic variables to move towards a particular region of the possible outcome space. Furthermore, as shown by Campbell and Shiller (1988) deviations from long run relationships or equilibrium errors could causes changes in the variables of the model or could themselves result from agents' forecast of these changes (e.g. the long term interest may reflect agents' rational expectation on the future of the short interest rate).

The concept of cointegration is an extension of the theory of non-stationary time-series. The starting point is that most economic variables are characterised by the presence of a stochastic trend or, in other words, they exhibit systematic variations over time, which are hardly predictable (Maddala and In-Moo Kim, 1998). This leads to the famous problem of spurious regression first mentioned by Yule (1926), the fact standard regression analysis is not applicable to judge dependency between two non-stationary series. However, Engle and Granger (1987) showed that, if a linear combination of two or more non-stationary series (i.e. $y-\lambda x$) displays a mean-reverting behaviour, then there is a long-term equilibrium between the series as they share a common stochastic trend. It has also been showed that the cointegrating coefficient λ can be efficiently estimated using ordinary least squares[9] and that cointegration is consistent with both discrete time and continuous time models[10].

Cointegration between two variables implies that, if the system is to return to its long-run equilibrium, at least one of the two responds to the magnitude of the disequilibrium. For instance, if we believe wages and the stock market are cointegrated, then, when the positive gap between the two is large relative to the long-run relationship, at least one of the following must be true: 1) wages decrease and/or stock prices increase, 2) wages decrease more than stock prices, 3) stock prices increase more than wages.

This intuition can be formalised with a full error correction model of the form:

$$\Delta y_t = \alpha_1 + \alpha_y (y_t - \lambda x_t) + \sum_{i=1}^{L} \alpha_{11}(i) \Delta y_{t-i} + \sum_{i=1}^{L} \alpha_{12}(i) \Delta x_{t-i} + \varepsilon_{yt}$$

where α_y is the adjustment coefficient capturing the speed at which variable y converge to its long-term equilibrium position. Bearing in mind the example above, it is clear that not all adjustments coefficients need to be significantly different from zero, that is not all the variables in the system necessarily respond to deviations from the equilibrium (if they do not they are said to be block exogeneous). An important corollary to the error correction representation of two cointegrated

[9] Stock (1987) showed that not only least squares is consistent for the true cointegrating coefficient but also that it converges to its true value faster than a coefficients estimated with stationary variables because of the infinite variance of all other linear combinations

[10] Phillips (1988) showed that the long-run parameter of a continuous time model can be estimated from discrete data by formulating and estimating the corresponding discrete time error correction model

series is that Granger causality (Granger, 1969) must run at least in one direction. This means that past values of one variable must help forecast the other[11].

Since the 1980s when the concept of cointegration was formalised, a significant empirical literature has emerged applying the theory to a vast number of economic relationships. Some significant papers were Friedman and Kuttner (1992), who looked at the long-term relationship between monetary policy variables (money, income, prices and interest rates), Kremers (1989) who argued that there is a long-term equilibrium between government debt and GNP, Johansen and Juselius (1990), who introduced a new procedure to test for cointegration and applied it to a demand for money equation with Danish data and Diebold, Gardeazabal and Yilmaz (1994), who used cointegration to study exchange rate dynamics.

There has been recently a growing interest in cointegration models to study the behaviour of financial markets. An early reference in this area is Campbell and Shiller (1986), who estimated a long-term relationship between long-term and short-term interest rates as well as between stock prices and dividends. Tokat, Rachev and Schwatz (2003) estimate a long-run cointegrating relationship between the S&P 500 price index, inflation, the dividend yield (under the assumption that it is non stationary), Treasury bill and bond rates. Bessler and Young (2003) extended Kasa (1992) work using cointegration and error correction models to estimate dynamic relationships between nine major stock markets. Finally, Cassola and Morana (2002) investigate, among others, the relationship between stock market and economic growth in the Euro area.

In the actuarial literature Sherris, Tedesco and Zenwirth (1999) worked on cointegration with Australian data, exploring whether there was evidence of a long-run equilibrium relationship between short and long-term interest rates, dividends and consumer prices as well as the stock market and consumer prices. Kuo, Tsai and Chen (2003) applied cointegration to the study of insurance policy lapse ratios to capture the dynamics between lapse rates, interest rates and unemployment in the US.

Little work has been carried out so far on the implications of cointegration between asset prices for asset allocations of long-term investors. One example here is André Lucas (1997) which illustrates implications of cointegration for strategic and tactical asset allocation using the example of foreign exchange management, relating strategic asset allocation to the long-run equilibrium relationship and tactical asset allocation to the error correction model.

[11] See Granger (1991) for a proof

5.1 *Key properties of the series: trends and unit roots*

Cointegration theory is built on the premise that each individual series is non-stationary and cointegration tests (Section 6 will expand on this) provide different results depending on assumptions about the data generating process of the individual series (i.e. presence of an intercept and/or of a linear deterministic time trend). This section will focus on the key properties of salary and asset return series. All series considered are in log terms.

In order to test the hypothesis of non-stationarity we apply to all series of interest the two most widely used unit root tests: Augmented Dickey Fuller (ADF) and Phillips-Perron[12]. Both tests require the estimation of parameters in the regression below (1) using an appropriate number of lags to eliminate residual autocorrelation and including both intercept and trend.

$$\Delta x_t = \alpha_0 + \alpha_1 t + \gamma x_{t-1} + \sum_{i=2}^{p} \beta_i \Delta x_{t-1+i} + \varepsilon_t \quad (1)$$

where Δx_t is the series constructed taking first differences of the variable of interest (x_t-x_{t-1}), t is a time trend and p is the number of lags chosen (in general we chose four lags with quarterly and one lag with annual data as no significant residual autocorrelation appeared).

Using the ADF or Phillips-Perron regression, testing the hypothesis of non-stationarity is equivalent to testing the hypothesis γ=0 using ad-hoc critical values[13]. At the same time, the ADF and Phillips-Perron framework enable to test whether, under the hypothesis of non-stationarity, the process is likely to contain an intercept and a linear deterministic trend component (this involve testing the hypotheses a_0=a_1=0 and a_0=0 given a_1=0 in the regression above, again using ad-hoc critical values).

Unit root test statistics on wages and assets were calculated using an appropriate number of lags to control for residual autocorrelation, starting with the assumption of both intercept and trend in the data generating process. Table 5.1 reports the results using Phillips Perron test statistics with four lags (as it is recommended with quarterly data), which appears to be more appropriate because residuals from several of the ADF regressions display signs of heteroskedasticity. Most conclusions do not change employing ADF test statistics, but, when differences arise, these are reported in the

[12] In general the Augmented Dickey Fuller test (Dickey and Fuller, 1979) is more powerful with autoregressive processes or processes with negative moving average components, while Phillips Perron (Phillips and Perron, 1988) is recommended for processes containing positive moving average terms (see Enders, 1995). The main difference between the two is that Phillips-Perron allows less restrictive assumption on the distribution of the error terms

[13] If γ=0 the series in level would be characterised by a unit root process (x_t=x_{t-1}+ε_t), which implies non-stationarity

text. Table 5.2 reports the results obtained with the annual data set (Phillips Perron tests with 1-year lag).

5.2 *Earnings and wages*

The monthly and quarterly data sets are considered first. The reported results are based on variables in logs and the quarterly data set but no significant differences arise using monthly data. Nominal salaries increased faster during the 1970s (an average of over 3.5% on a quarter to quarter basis compared to just above 2% in the overall sample), slowly in the 1980s and 1990s and were almost flat in the 1960s. In fact a regression of quarterly salary growth on 4 time dummy variables representing each decade explain almost 40% of the variance of nominal salary growth[14].

From the ADF regression we concluded that the data generating process was not likely to include an intercept or deterministic time trend component, while the Phillips Perron test suggested a time trend might be present. In this case we believe Phillips Perron is definitely more appropriate because residuals from the ADF equation appear clearly heteroskedastic.

It is not straightforward to determine the order of integration of the data-generating process of nominal wages because the Augmented Dickey Fuller tests cannot reject the hypothesis of a second unit root for both wage indices, although the test statistic is not far from the rejection area. Phillips-Perron on the other hand suggest the process is I(1), or difference stationary[15]. The long-term data set should provide more clues as to whether the nominal salary returns series can be considered stationary, although the rate of growth has been so remarkably different across sub-periods after the 1960s.

The picture is different when we consider real salaries because there is no evidence of a significantly different rate of growth across sub-periods. We can also confirm that inflation and not real salaries caused nominal wage growth to differ significantly across decades. If we regress inflation on the 4 period dummies, we obtain in fact highly significant coefficients and a R-squared of 35%.

With real salaries on the other hand, there is some evidence of a possible structural change occurring shortly after the 1975 oil crisis when real salary growth has been negative for a number of quarters. If we perform Perron (1989) regression on UK real wages (the total compensation series) using the fourth quarter of 1975 as a breakpoint we find some support for the structural change

[14] Similar results were obtained with the AEI and the wage index from national accounts

[15] This means that the series constructing by taking first differences of the nominal salary index (the log return series given that the original index is in log terms) is stationary

hypothesis. The conclusion would be that the oil shock has led to a permanent change in the intercept and slope of the real wages process. The intercept appears to have gone up after 1975 while the slope seems to have decreased. However, both coefficient and slope in the Perron regression turns out to be insignificant if we use the real average earnings index and therefore we are unable draw a definite conclusion.

Aside from the issue of structural breaks, the Phillips Perron test cannot reject the hypothesis of stationarity around a time trend[16] for AEI salaries, while the ADF test reaches the opposite conclusion. The two tests however agree in strongly rejecting the hypothesis of an I(2) process[17] for both real indices. Similarly to what happens with nominal data, the Phillips Perron test cannot reject, under the hypothesis of a unit root, the null hypothesis of no linear trend for the AEI index. Moreover with real data also the null hypothesis of no intercept is rejected at 95% confidence level, but even in this case, this happens only with the average earnings index. The conclusion appears to be that the presence of a deterministic linear trend component in the salary process may depend on how salaries are actually measured. This is because an index arising from survey data is likely to be characterised by a smoother path than national accounts wages.

With the annual data set, we have one nominal salary series and two alternative real salary indices constructed with different price deflators. Interestingly, visual inspection of the return series suggests all processes are stationary, although some unusual periods can be identified. With nominal salaries, growth was higher than normal in the 1920s, later followed by a severe contraction, and in the 1970s. With real wages, volatility was higher in the 19th century[18] and the same pattern of nominal salaries was recorded in the 1920s (but not in the 1970s where higher than normal growth was caused by inflation).

Unit root tests confirm this view suggesting wages both in nominal and real terms are difference stationary. The hypothesis of non-stationarity of the return series is however strongly rejected in the case of real earnings, while the conclusion is less clear-cut for money wages, probably to the influence of the 1970s. The hypothesis of no linear time trend cannot be rejected for the three series but test statistics are close to the rejection area for real wages deflated with the

[16] At 90% confidence level, while it is very close to the 95% threshold

[17] I(2) would imply that not only the series in levels but also the first difference (or log return/growth) series is non-stationary

[18] This can be due to the unreliability of price data, the problem is in fact more severe with Global Financial Data as opposed to Mitchell (1994) deflator

Global Financial Data price series. Finally, all processes are likely to contain an intercept because the model with no intercept would imply explosive dynamics.

5.3 *Equities*

Turning our attention to financial assets we start examining the properties of the equity total return series. The hypothesis of a different rate of growth across sub-periods is not supported by the data for both nominal and real returns, although a dummy variable, which takes the value of one in 1975, is highly significant for both nominal and real returns. The average nominal return was in fact 23% higher in 1975 than in the rest of the sample and the real return 16% higher. Interestingly, also a dummy for the period 2001:1-2002:3 is significant indicating that average nominal returns after the millennium were 10% lower than in the rest of the sample (and real returns almost 9% lower).

The 1970s apparently did not bring a structural break but a one-off temporary shock to the mean of the process following the first oil shock (and quite evidently also to its variance given the large movements recorded in 1974-75). Conversely a structural break may have occurred in 2000 but, although it appears to be more persistent than in the 1970s, it is difficult to say whether it is temporary or permanent.

Both unit root tests suggest nominal and real processes are I(1) and they have no significant deterministic trend component. It must be highlighted that this does not imply stock prices are a random walk with no stationary component, it simply says the predictable component of equity prices, if exists, is something more complex than a simple linear trend.

With the annual long-term data set the nominal equity returns are characterised by rising volatility, especially after the 1920s with a peak in the mid-1970s where very large swings were recorded. Phillips-Perron test statistics are likely to be a better guidance here because the error terms in unit root regression are likely to be heteroskedastic. Unit root tests suggest the equity process is likely to be difference stationary and to contain an intercept but no linear trend component.

Real equity returns are characterised by rising volatility like the nominal process and unit root tests achieve similar results.

5.4 *Bonds*

For nominal bond returns a dummy for the period 1980:1-1990:4 is positive and significant (1.2% higher average returns) indicating a possible structural break in the early 1980s with the start of a process of transition towards lower interest rates. For real bond returns a dummy for the 1970s is also negative and significant due to inflationary pressures depressing real bond returns.

Both unit root tests suggest nominal bonds are difference stationary, while less clear-cut is the conclusion on the presence of a linear trend and an intercept component. The hypothesis of no trend and no intercept cannot be rejected using the ADF and Phillips Perron critical values but the test statistic are very close to the rejection area, especially for the linear trend hypothesis (2.69 against 2.79 critical value at 95% using Phillips Perron). In addition to that, it appears unwise to assume no intercept and no trend because this would lead to an explosive process (γ>0 and statistically significant).

Real bonds are found to be difference stationary and with a possible linear trend component, as the test statistics for the hypothesis a_1=0 are higher than 2 both with the ADF and Phillips Perron test (although the likelihood of a linear trend is slightly weaker than with nominal data)

In the annual data set nominal bonds appear to be difference stationary as well, although the ADF test (but not Phillips Perron) cannot reject the hypothesis of a second unit root. This is because the process has a relatively long memory, as it can be observed in the 1970s with a sustained period of high interest rates and since the 1980s with a decreasing trend fuelled by low inflation. Conversely, no doubts arise for real bonds as both ADF and Phillips Perron strongly reject the hypothesis of a second unit root. The tests with annual data for both nominal and real bonds suggest the processes are likely to contain an intercept but do not provide much support for the linear trend component hypothesis.

Finally, with index-linked[19] bonds, although monthly data are employed, although it is hard to draw meaningful conclusions from only 20 years of data, unit root tests suggest the process is difference stationary and the linear trend component is significant.

5.5 *Other asset classes*

We consider here bills, house prices and US equities using only the 1963-2002 quarterly data set (monthly for US equities). For nominal bills the regression on dummy variables representing each decade explain 40% of return variation across the sample. For real bills the conclusion is similar to real bonds indicating that high inflation in the 1970s has had a negative impact on real returns.

[19] These results are not reported in the tables

The nominal process appears to be characterised by a constant and no time trend while the real process appears to have a significant positive linear trend component[20]. Both unit root tests cannot reject the presence of a second unit root for the nominal bill index, while for the real bill Index the I(2) hypothesis is rejected by Phillips-Perron but not ADF.

With nominal house prices a regression of index returns on decades' dummies explain about 18% of the total variation and mean returns are significantly higher in the 1970s and the 1980s than in the 1990s and the 1960s. However this could be misleading unless we take a closer look at a simple historical graph. Nominal house prices experienced in fact three clear 'boom and bust' cycles: early 1970s, late 1970s and late 1980s. For real house prices a similar regression yield non-significant coefficients, while dummies representing peaks and lows of the three cycles are highly significant.

The ADF test cannot reject the hypothesis of a second unit root for nominal house prices, while Phillips Perron, which is more appropriate because of heteroskedasticity in the residuals, would suggest nominal house prices are I(1) and display a significant intercept but not linear trend component.

With real house prices, the hypothesis of a unit root is rejected at 10% confidence level by the ADF test (but not by Phillips Perron) suggesting the hypothesis property investments have been a reasonable inflation hedge. Cardinale (2003) explores this issue in more detail. Both the ADF and Phillips Perron tests suggest real returns are stationary and the ADF equation in levels suggests the presence of an intercept and a time trend, while Phillips Perron reaches the opposite conclusion.

US equity returns were significantly correlated with the UK stock market (48% unhedged and 58% hedged) and the correlation increases if the sub-sample 1980-2002 is considered (52% unhedged and 71% hedged). Diversification benefits of unhedged equities arose because the pound has displayed a tendency to depreciate against the dollar since the 1960s (and the trend accelerated in the 1980s when US equity returns were very high).

Unit root tests on the S&P 500 nominal and real total return series (hedged and unhedged) display a broadly similar picture to the one depicted for UK equities: difference stationarity and no significant linear trend component. With US equities even the presence of an intercept in the data generating process is rejected by both the ADF and the Phillips Perron tests.

[20] We used Philips Perron critical values because a White test for heteroskedasticity (White, 1980) suggests the mean of the residuals is different across periods (higher in the 1970s due to higher volatility)

6 Pairwise bivariate relationships and multivariate cointegration models

In this section we focus our attention on four asset classes in which UK pension funds can invest: bonds, bills, UK equities, property and US equities. The objective is to ascertain whether there is evidence of long-run equilibrium relationships between domestic wages and a range of assets, applying the cointegration theory framework discussed in Section 4. This is clearly a relevant question when it comes to assess long-term investment strategies for a final salary pension scheme, whose active members' liabilities rise in line with wages. This will be explored considering both variables in real and nominal terms.

We will also briefly focus on the relationship between domestic salary and domestic equity returns and look more closely at the long-run equilibrium between wages and productivity, the stock market and productivity, wages and corporate profits and finally wages and dividends.

Intuitively, when two series are cointegrated, there should be a meaningful association between values in one sequence and values in the second sequence, which could be approximated by a regression line. There would be random but not systematic deviations from the estimated regression line. Figures 9-12 display scatter plots with band regression lines to visualise the historical relationship between salaries and a range of assets both in nominal and real terms. The graphs can provide a snapshot to judge the extent to which there might be a long-term equilibrium between assets and salaries. The graphs are presented in levels but no significant differences arise if the log scale is used.

The graphs suggest in most cases the presence of meaningful positive relationships between salary and asset indices, although not surprisingly there appears to be substantial departures from long-term trends in some occasions. For instance, the equity market downturn after 2000 represents the most substantial exception to an overall positive relationship between nominal salaries and nominal equities. This is evident both in the 1963-2002 and in the 1850-2001 data sets. However this can also be read as a correction towards a more realistic long-term relationship following a marked rise in the steepness of the estimated regression line after the end of the 1980s. A similar pattern can be detected in the house prices/wages scatter plot, which displays an evident hump-shaped behaviour in the years of the property market crash of the end 1980s. Even in the relationship between nominal bonds and nominal salaries there have been deviations from an overall positive correlation: the late 1990s, as it appears evident from the 1963-2002 graph, and the 25 years between WWI and WWII, as shown in the annual sample graph.

In real terms the picture is even more complex to read. There is indeed some evidence of positive relationships between salaries and assets but, using the 1963-2002 data set, the 1970s represents a problem for all estimated relationships, the property crash of the late 1980s is a further problem for house prices and the equity market downturn after 2000 is an additional problem for equities. It appears that no single asset class is a perfect hedge for salary increases. To investigate this further we plotted a similar graph using index-linked bonds, which should be a better indicator of real returns than ex post inflation-adjusted nominal returns[21]. It is difficult to draw conclusions from such a limited sample but it appears that after 1998 it was increasingly difficult to detect a positive relationship. With an index-linked bond price index the regression line is flat after 1998 and this may be due to depressed real yields because of distortions in the market (i.e. flight to safe assets by pension funds and insurance companies).

Interestingly, using the long-term annual data set there appears to be a more evident positive long-term relationship between real equities and real salaries rather than between real bonds and real salaries, although this is challenged by recent market movements. With real bonds there was in fact little evidence of a positive relationship between the 1920s and 1950s, because the real bond total return index has been almost flat (except in the early 1920s when real bond returns were negative).

The evidence presented in the previous section showed that almost all series considered appeared to be difference stationary, except perhaps the nominal bill index for which the hypothesis of an I(2) process could not be rejected. Cointegration tests usually assume that variables are integrated of the same order. We use two main cointegration tests here: the Engle Granger (1987) methodology and the Johansen (1988) estimator.

Engle and Granger (1987) rely on a two-step procedure involving estimating the long-run equilibrium relationship by ordinary least square and then perform unit root tests on the residuals[22]. If the residuals appear to be 'white noise'[23], the hypothesis of cointegration cannot be rejected and an error correction model can then be estimated. Two-steps procedures have been criticised because they are sensitive to the variable chosen for the normalisation, an essentially arbitrary assumption.

Johansen (1988) introduced a new procedure based on full information maximum likelihood to estimate the linear space spanned by the cointegrating vectors. Unlike residual based tests (e. g.

[21] See for instance Arak and Kreicher (1985)

[22] Unit root test statistics must however by compared with critical values computed by Engle and Granger (1987) and not with the ordinary Dickey Fuller table

[23] 'White noise' is a stationary process with no autocorrelation

the Engle and Granger methodology), the procedure was also found to be reasonably robust to deviations from classical regression assumptions, such as residual autocorrelation, non-Gaussian errors, ARCH effects or unconditionally leptokurtic innovations (see Gonzalo, 1994). The Johansen test requires assumptions on whether the data-generating process contains an intercept and a linear time trend or drift. In reality however the data-generating process is not known and therefore results should be analysed both with and without a time trend, as suggested by Diebold, Gardeazabal and Yilmaz (1994).

The Johansen procedure hinges on a multivariate extension of the Dickey Fuller test, as it focuses on the rank of the matrix π, constructed as the difference between the matrix of parameters and the identity matrix. Under the null hypothesis of no cointegration the rank of π will be equal to zero and the procedure can be repeated to test for the presence of more than one cointegrating relationship if at least three variables are included in the system. The Johansen test procedure yields two statistics to test for the hypothesis of no correlation: the trace and the maximum eigenvalue statistics. The two typically give the same answer but careful attention must be made where conflicting results arise.

We now test for the presence of a long-run equilibrium relationship for UK salaries (using both the AEI and the national accounts measure) and five asset classes: bonds, UK equities, bills, house prices and US equities.

6.1 *Salaries and bonds*

This is a very important relationship to focus on in order to judge the claim bonds are a good hedge for final salary schemes active members' liabilities. Cross-correlogram analysis of differenced series did not provide a clear answer, although it found some evidence, especially in the quarterly data set, of a relationship between bond returns and past salary increases. However, differencing a series and considering only the return (or growth) process means losing a lot of information present in the original data. Cointegration is designed to address this issue by investigating the stationarity of linear combinations of the series in levels, whilst attempting to explain the return or growth process in terms of an adjustment towards a long-run equilibrium.

As shown in Tables 6.1 and 6.5 the Johansen procedure provides strong evidence of a cointegration relationship between real bonds and both salary indices in real terms if quarterly or monthly data sets are considered. If the sample is broken into two, the null of no cointegration cannot be rejected for the sub-period 1964-1979, while the opposite is true for the subsequent period (1980-2002). The sample is divided up non-symmetrically because, as discussed in Section

4, there is evidence of a possible structural break at the end of the 1970s. In fact a dummy for the 1970s was found to be negative and significant for real bond returns, while a dummy for the 1980s was positive and significant for nominal bond returns.

The results are highly sensitive to the trend assumption, as outlined in Table 6.4, which shows that the hypothesis of cointegration is rejected if a linear time trend is allowed for. In Section 5 we showed that it was unclear whether the real salary process contained a deterministic trend component. The no trend hypothesis could in fact be rejected with the Phillips Perron test with the average earnings index but not using national accounts real wages.

The long-term relationship linking real bonds to real wages estimated with monthly data (see Table 6.7) and assuming no trend suggests a long-term positive co-movement, and this could be interpreted as the result of a common factor affecting the labour and the bond market, possibly the real interest rate or the bond maturity premium. There could be in other words a pattern of dependency between the rate of intertemporal substitution affecting consumption/savings decision and the rate of substitution between work and leisure. Moreover, LM restriction tests[24] suggest that the hypothesis of a one-to-one relationship cannot be rejected for the overall sample, while for the sub-sample 1980-2002 the coefficient estimate is significantly higher than 1.

Quarterly data are however more problematic because coefficient estimates are not significant at conventional level, even assuming no trend, unless the relationship is re-estimated using only the sub-sample 1980-2002. This suggests that either a structural change has occurred in the relationship during the 1970s or the true relationship cannot be inferred due to small sample bias. In particular, if a dummy equal to 1 from the fourth quarter of 1975 onwards and a dummy for the decade of the 1970s are added to the cointegrating regression, a positive and significant relationship between real bond and real national accounts wages can be estimated.

In general the inclusion of a dummy variable in multivariate regression models (and this applies to the error correction model of a cointegrating relationship as well), measures the impact on the dependent variable of a switch between categories defined by the dummy variable, controlling for all the other predictors. The inclusion of dummy variables, whose presence help make residuals from the long-run relationship visually more stationary, highlights the potential impact of extraordinary events on long-run relationships. If extraordinary events are simply one-off factors, cointegration does not break down; however if they occur with random periodicity, cointegration

[24] LM restriction tests are standard econometric tests of coefficient restrictions. In this case the null hypothesis is that the coefficient associated to the bond index in the cointegrating vector is equal to one

analysis would have to be extended with Markov switching model (see for instance Miles and Timmermann, 1999). This however would lie outside the scope of this paper.

The instability of the estimated long run relationship can also be inferred by the adjustment coefficients in the error correction models, which represent the sensitivity of endogenous variables to the disequilibrium from the long run relationship (a positive coefficient means that when the system is above the equilibrium the endogenous variable will tend to rise). These are positive for both real salaries and real bonds and the real bonds coefficient is also higher in magnitude (see Table 6.5 for quarterly data, similar results arise with monthly data). This suggests that when real bonds are higher than what would be implied by the long-term equilibrium, real salaries tend to rise to restore the equilibrium, but at the same time real bonds tend to continue rising driving the system away from the equilibrium.

For nominal bonds, the results suggest a long-term one-to-one relationship with nominal salaries and this is a more robust finding as it does not depend on trend assumptions (the graph in Figure 9 confirms this). This is clear however if the entire sample is considered, while mixed evidence arises (even a negative relationship with monthly data between 1980 and 2002) if the sample is broken down in two. And even with the full sample the relationship may be subject to potential instability because the two adjustment coefficients are both negative (see Table 6.5) indicating that positive deviations from the equilibrium (nominal bonds higher than the equilibrium level) cause bond prices but also salaries to fall. However, in absolute terms, the magnitude of the speed of adjustment coefficient for the bond market is substantially higher (at least with quarterly data), indicating a better probability of restoring the equilibrium than in the case of the variables in real terms.

Interestingly, residual based tests would achieve a different conclusion on cointegration, both with nominal and real data, as shown in Table 6.2. With the Engle Granger two-steps methodology, the hypothesis of no long-term relationship cannot be rejected, although with the sub-sample 1980-2002 and the national accounts series, both ADF and Phillips Perron test statistics are closer to the rejection area. Failure of residual based tests to reject non-stationarity could be interpreted as evidence of a long memory process, in which the effect of a shock declines at a slow rate and deviations from equilibrium are likely to persist for a reasonably long time.[25]

[25] This was highlighted by Baillie and Bollerslev (1994), who argued that fractional cointegration, which allows for a long memory adjustment, could be an interpretation of these conflicting results. Fractional cointegration postulates that the residual series is neither I(0) nor I(1) but I(d) with d<1 and is characterised by slow adjustment and long-term cycles. A full test of the hypothesis however would require a greater number of observations to detect cycles

With the long-term annual data set (see Table 6.9-6.11) the results are broadly consistent, but there is stronger evidence of cointegration with the variables in nominal terms, consistently with visual inspection of the data (Figures 10 and 11). Cointegration tests using the full 1850-2002 sample suggest the presence of a common stochastic trend between salaries and bonds both in real and nominal terms. Estimated coefficients are higher than 1 but the hypothesis of a long-run one-to-one relationship cannot be rejected with variables in nominal term. The evidence is weaker if the 19th century is removed from the sample, because with 20th century data the Johansen test would suggest the hypothesis of a common stochastic trend only between nominal bonds and nominal wages. With annual data (both nominal and real) adjustment coefficients (Table 6.5) are negative indicating a potential source of instability in the systems (wages falling instead of rising to catch up with the equilibrium), although the bond market coefficients are greater in magnitude.

Figures 13 and 14 plot the long-run equilibrium estimated respectively in real and nominal terms with annual data and no linear trend (but adding a dummy variable for 1924 onwards to remove the effect of the unusual behaviour of real bond and real wages in the 1920s).

In real terms, although cointegration is not rejected by the two tests, the evidence does not appear conclusive, mainly because of the 1920s and the 1970s, when real bonds were far below the level implied by the cointegrating relationship. However, whilst in the 1920s the equilibrium was quickly restored, the system has stayed below the equilibrium for over 50 years after WWII shedding doubts on the existence of a true adjustment process. Only at the end of the 1990s the equilibrium was finally restored after a period of falling inflationary expectations and rising real ex post bond returns.

In nominal terms the relationship appears to be characterised by a very slow adjustment process, because salaries have been below their long-run equilibrium with bonds throughout the post WWII years and only after the 1980s there has been a gradual convergence through higher bond returns.

Ideally one would want to estimate a cointegrating relationship between salaries and index-linked bonds but, as outlined in the previous section, reliable estimates with the index-linked series are not possible because no index-linked bonds were available before the 1980s. However the index-linked return profile can be proxied by inflation and a cointegrating model can be estimated with the RPI price index series. This is essentially a test of whether a stable long-run equilibrium exists between nominal wages and prices.

Figure 15 provide a visual illustration of the deviations from the long-run relationship between prices and national accounts wages (1963-2002 quarterly sample) and between prices and the historical wage index (1850-2001 annual sample). The estimated cointegrating equation is also reported, although at conventional levels the null hypothesis of no cointegration cannot be rejected due to the length of the adjustment process.

With the quarterly sample the system appears to be above the equilibrium position until the beginning of the 1970s, while the reverse was true after the mid-1970s when anti-inflationary policies aimed at cutting the transmission mechanism between wage and price inflation were implemented. The long-run equilibrium however does not seem to be stable, and not surprisingly the estimated adjustment coefficient of the salary error correction model is positive suggesting salaries displayed a tendency to increase when they were already above the equilibrium level. With the annual sample the relationship appear even less stationary by visual inspection (even if a dummy for 1924 onwards is included), with the exception of the period between 1900 and 1950. In fact, wages have been below the long-run equilibrium before 1900, while after 1950 they have been above it, indicating a possible structural change (e.g. higher productivity leading salaries to grow faster than in the past relative to prices).

6.2 *Salaries and UK equities*

We are testing here the hypothesis of a common stochastic factor driving the labour market and the stock market. This could be interpreted as productivity, which would exercise an upward pressure on real salaries while at the same time raising the expectations for future real dividend growth. Tables 6.12-6.19 show a similar picture to what discussed concerning the bond market. With the entire sample, there is evidence of cointegration between real equity and both indices of wages, but, as with bonds, the results depend on the linear trend assumption. With equities however the appropriate Dickey Fuller test statistics more clearly suggested the absence of a linear trend component in the data generating process, as outlined in Section 5.

Similar to what happened with bonds, results are different across sub-samples with rejection of cointegration in the period 1964-1979 and strong evidence of cointegration in the period 1980-2002. Interestingly, with national accounts real wages and quarterly data, the hypothesis of one cointegrating vector in the sub-sample 1980-2002 cannot be rejected even if we allow for the presence of a deterministic time trend.

Estimated cointegrating relationships, both with the quarterly and monthly data set, suggest the coefficients in the equilibrium relationship between real equity and real salary are higher than 1.

This can be interpreted as evidence of companies' ability to generate real revenues, which in the long-run have exceeded their labour costs in real terms.

Error-correction models suggest that adjustment to the long-run equilibrium is generally performed through higher salary growth rather than a fall in the stock market given that the adjustment coefficients associated with real equities are not significant (see Table 6.16). In a model in which both the stock market and wages are moved by underlying productivity changes, this pattern could be interpreted if equities adjust more swiftly to productivity rises, while real salaries tend to catch up with a lag, because workers' aspirations adjust slowly, as discussed by Ball and Moffitt (2001). However, in Barsky (1989) framework, this long-run relationship may break up if the substitution effect prevails, leading individuals to consume more, save less and shun the equity market when growth prospects are higher. This could explain a slow adjustment to a long-run equilibrium position in which the stock market reflects its underlying fundamentals.

With nominal equities and nominal salaries, cointegration evidence is stronger and the trend assumption is irrelevant, as it happens with bonds. However, the relationship is negative in the 1964-1979 sub-sample and strongly positive after the 1980s suggesting that in a higher inflation environment such as the 1970s salary growth still keeps up with prices while the equity market suffers. There is a whole literature discussing the fact that equities are a poor inflation hedge when inflation is high, highlighting that inflation causes capital market inefficiencies negatively affecting stocks (see Barnes, Boyd and Smith, 1999). Another hypothesis is that the stock market irrationally discount real dividends at nominal interest rates undervaluing stocks when inflation is high and overvaluing them when inflation is low (Modigliani and Cohn, 1979).

In real terms, on the other hand, there was no significant relationship between equities and salaries in the 1964-1979 sub-sample, and this could be explained if real salary growth was not aligned with productivity, but rather helped fuel inflationary pressures (Engle, 1982 explored a similar hypothesis with 1958-1977 data in his model of UK inflation in which lagged real wage helps predict inflation).

Finally, even with nominal and real equities, the Engle Granger methodology achieves a different conclusion (no cointegration), providing support for the hypothesis of slow adjustment towards the equilibrium (see the discussion in 6.1). Even here this might suggest fractional integration for the residuals, that is a slowly decaying stationary process rather than a pure random walk.

With the long-term annual data set, cointegration tests depict a similar picture, as they reject the null of no cointegration for equities and salaries both in nominal and real terms. Interestingly, using the full 1850-2002 sample the evidence is even stronger because the result does not even depend on trend assumptions. In other words, even assuming that equities and salaries shared a common deterministic trend, cointegration between the detrended series could still not be rejected. The shape of the cointegrating vector in the real term equation confirms a positive association with a coefficient significantly higher than 1. With nominal variables on the other hand, the magnitude of the coefficient is higher than 2, unlike with quarterly and monthly data whereby the hypothesis of a long-run one-to-one relationship could not be rejected. Finally, unlike with quarterly and monthly data, the adjustment coefficients suggest the equilibrium is restored mainly by the stock market, both using nominal and real variables (Table 6.16).

Figures 16 and 17 show the equilibrium errors from the long-run cointegrating relationship estimated respectively in real and nominal terms with annual data and no linear trend (but inserting a dummy from 1975 onwards to remove the effects of the extreme swings recorded in that year). The adjustment pattern appears particularly evident in the late 1940s, when real equities were below their long-run position, given that the stock market experienced a healthy growth in the subsequent years, and in the early 1970s, when real equities were above the level implied by the cointegrating relationship, just before a fall in the stock market index. After the 1980s real equities have been above the level implied by the cointegrating relationship, even during the recent bear market.

In nominal terms Figure 17 shows that equities were above the equilibrium between the 1930s and the 1970s as adjustments were not large enough in magnitude to restore the equilibrium until the 1970s. Since the 1970s the adjustment was quicker in real terms, while in nominal terms the equilibrium was restored only in the mid 1990s.

6.3 *Salaries, profits and productivity*

The discussion in Section 2 suggests that introducing measures of productivity could help better understand the relationship between the labour market and the stock market. Corporate profits in the short-run can grow because of higher prices, lower wages or higher output per worker. This may justify the observed negative relationship between wages and stock prices described by Bottazzi, Pesenti and van Wincoop (1996). However we argued that to consistently achieve higher profits while wages decrease or remain flat appears unlikely to be sustainable in the long-run.

One should therefore expect to observe a positive long-run relationship between productivity and both real salaries and real profits. This would imply that productivity shocks would affect both the labour and the stock market, and this could help explain a positive relationship between stock prices and wages (assuming that the substitution effect in Barsky, 1989 framework prevails, leading individuals to buy more equities if profits are expected to grow faster in the future).

In the short-term however profits and wages growth may not be aligned and, as a consequence, a positive relationship between them may not necessarily hold when short-term considerations play a substantial role in moving equity prices and influencing company salary policies. In this framework, when real salaries increase faster than productivity, like in the early 1970s in the UK, they are likely to contribute fuelling inflationary pressures, which will eventually be curbed through restrictive monetary and fiscal policies leading to a contraction in real salary growth. Conversely, with competitive labour and products markets, higher productivity is unlikely to result in faster growth in company profits without an increase in real salaries.

Table 6.23 shows estimated bivariate cointegrating relationships between real profits and productivity and, separately, between real salaries and productivity. With the full 1964-2002 sample, a one-to-one cointegrating relationship appears to exist between both national accounts and AEI real salaries with productivity measured by output per job (similar results would arise using output per filled job). This conclusion holds even if the sample is broken in two indicating that equilibrium positions are restored in less than 20 years through lower real salary growth. This is precisely what occurred in the 1970s when real salary had grown faster than productivity until 1975, but this was compensated by negative real salary growth between 1975 and 1979.

At the same time, a positive cointegrating relationship cannot be ruled out for real profits (measured as either national accounts profits or earnings per share of listed companies, the latter being more comparable to a per capita measure such as output per job) and also real dividends with productivity, but the coefficients are not significant with the full sample because of the pattern observed between the 1960s and 1970s. In fact, if the estimation is performed using the 1964-1979 sub-sample the null of no-cointegration cannot be rejected at 95% level and the estimated regression coefficients are negative. This might suggest that the increase in oil prices and inflation had significant real effects, depressing profits even in the absence of a significant productivity slowdown.

Furthermore, if dividends are considered the null of no cointegration cannot be rejected even in the 1980-2002 sub-sample due to the well documented international trend towards lower dividend yields (see Campbell and Shiller, 2001 for a review and Cardinale, 2002 for UK evidence). In essence, the lack of a relationship between dividend growth and productivity after the 1980s may just be due to a temporary departure from a long-run relationship or may be due to the fact dividends underestimate total shareholder compensation which includes share buyback programmes.

6.4 *Other asset classes*

We examine here whether the hypothesis of a common stochastic trend with UK salaries holds only for domestic bonds and domestic equities or whether similar patterns arise with other asset classes in which defined benefit pension funds might invest.

First of all, we consider short-term money market instruments (Table 6.26-6.28), because we might expect a similar pattern to that observed for nominal bonds, if we believe a long-run relationship between short and long-term bonds holds, consistently with the expectation hypothesis of the term structure of interest rates[26].

With variables in real terms, the picture is broadly similar to the one depicted for bonds. There appears to be evidence of a long-term positive co-movement but the relationship does not hold if the sample is broken down in two (if the estimate is carried out between 1964 and 1979 there is virtually no relationship while after the 1980s there is a positive association with an estimated coefficient not significantly higher than 1). Moreover, cointegration test results are sensitive to trend assumptions and both adjustment coefficients are positive suggesting instability.

With variables in nominal terms, the evidence of cointegration is slightly stronger, suggesting a one-to-one long-run relationship. The adjustment coefficient is negative and significant for bills while it is insignificant for AEI salaries and negative and significant for total compensation wages. However, a closer look at the result suggests a single full cycle in the period considered: a trend away from the equilibrium until 1980, as salaries grew more than bills, and a gradual move towards the equilibrium since then, within a lower inflation environment.

Secondly, we investigate whether UK property is correlated with wages (Tables 6.29-6.31). There may be a mortgage financing argument here, as most house purchases in the UK are financed through mortgages based on multiples of individual wages. One could argue that house prices in the long-run cannot grow too fast with respect to wages and when house prices are too high economic forces would restore the long-term equilibrium.

The results are encouraging in real terms as cointegration cannot be ruled out for the entire 1963-2002 sample and the two sub-samples in the hypothesis of no linear trend (except for the total

[26] See Campbell and Shiller (1986) for a first empirical test of the cointegrating relationship between long-term bonds and bills in the US

compensation index after 1980). Interestingly, unlike with bonds, equities and bills, the evidence of cointegration is even stronger (it holds independently of trend assumptions) for the period 1964-1979, indicating that in the high inflationary environment of the 1970s real properties have been the best hedge for real salary growth. Coefficient estimates are broadly consistent across sub-samples and the adjustment coefficients suggest that the equilibrium is more likely to be restored through a higher growth in wages rather than through a property crash[27]. Interestingly, the opposite conclusion on adjustment arises if a dummy variable equal to 1 from 1989 onwards, which is significant in the error correction equations, is included in the model. This implies that the property crash in the early 1990s has induced a one-time change in the mean of the house prices series. Figure 18, which plots the equilibrium errors for the model with a dummy for 1989 onwards, suggests that, after the property crash of the early 1990s, house prices have been below their equilibrium level with wages, although the gap has recently been narrowing due to strong property returns.

In nominal terms, however, the evidence on cointegration is strong only if the sub-sample 1964-1979 is considered, unless a dummy for 1989 onwards is added as an exogenous variable. If a dummy is added to the regression the results suggest a one-to-one relationship in the long-run, although with a slow adjustment in the 1990s.

Finally, with unhedged US equities (Tables 6.32-6.34) coefficients estimates are broadly in line with those estimated with UK equities, given the significant positive correlation between UK equities and US equities, but the evidence of a cointegrating relationship with UK wages is weaker, both in nominal and real terms. This is consistent with the argument of Baxter and Jermann (1995) on the international diversification puzzle. They posited that individuals should invest more abroad than in domestic assets because they have an implicit holding in a domestic assets (human capital) whose returns (salaries) depend on domestic factors. Final salary pension schemes have however a different perspective on this, as their liabilities are tied to domestic salary growth.

6.5 *Towards a multivariate long-run model*

After assessing pairwise relationship between UK salaries and a range of asset classes, this section focuses on combining asset classes together to compare alternative long-term multivariate models of wages. The estimations were carried out using quarterly data between 1963 and 2002.

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[27] The adjustment coefficient is positive for real wages (or real AEI) and not significantly different from zero for real property

Table 6.35 compares results of alternative models to capture the long-term behaviour of UK real salaries, starting from bivariate models (Model 1, 1b and 2) to alternative multivariate specifications (Model 3-6b). No dummy variables were added to these specifications and the estimation is carried out using 4 lags in the error correction model and with the assumption of no linear trend.

Interestingly, in all models the hypothesis of no long-term equilibrium is rejected using the Johansen procedure while the hypothesis of at most one cointegrating relationship cannot be rejected. This would suggest the presence of a single stochastic trend driving real salaries and up to five assets, showing a high degree of interdepedence between the UK labour market, the economy (i.e. exchange rate) and asset markets. Moreover, a closer examination of the long-run equilibrium coefficients suggests:

1) A positive and significant long-term relationship between real salaries and the stock market in real terms, independently of what other assets are considered at the same time. Coefficient estimates range from 0.29 to 0.53, depending on the menu of assets and the index used to measure wages.

2) A negative and significant relationship between real salaries and real bonds, when the real equity index is controlled for, in all Models, except Model 5. Coefficient estimates vary, ranging from -0.45 to -0.67, again depending on the menu of assets and the index used to measure wages. This pattern suggests the real interest rate may be negatively associated with productivity and real wages[28], through lower labour demand from companies facing higher borrowing costs, perhaps combined with higher supply from impatient individuals requiring a high compensation to defer their consumption. Model 5 provides support for this interpretation as the relationship turns out to be negative for short-term bills while it is insignificant for bonds. However in a bivariate model the coefficient linking real salaries and real bills is not significantly different from zero (Model 1b) suggesting the presence of a pro-cyclical component in real bills, which is removed by controlling for the stock market.

3) A positive relationship between real salaries and the housing market in real terms, which is only significant if real bills are included in the model (Model 5).

4) A residual positive, but not always significant, relationship between US unhedged equities deflated using UK prices and real salaries when the positive impact of the domestic stock market is controlled for. This would suggest a linkage between domestic wages and international factors, such

[28] In the long run one would expect to find a positive association between real interest rates and total returns on bonds even if in the short terms bond returns fall when interest rates rise

as the exchange rate or the performance of US companies, which employ a large number of people through direct subsidiaries in the UK.

Figure 19 shows the equilibrium errors arising from the key models presented in Table 6.35. The graphs enable to compare simple bivariate models with more complex multivariate structure by looking at the extent to which a stable adjustment pattern can be inferred.

Bivariate relationships linking real wages and either real bonds or real equities (Model 1 and Model 2) do not provide convincing evidence of a stable adjustment pattern, although the null hypothesis of no cointegration is rejected in both cases. Firstly, real bonds and real equities have been below the estimated cointegrating relationship throughout the entire sample (except real equities in the mid 1970s). The latter could be addressed by including one-off exogenous factors to control for the impact of unusual market movements (e.g. a dummy variable for 1975) or by estimating the models with a linear trend. More fundamentally, the forces which should bring the systems back to the equilibrium through the error correction models do not seem particularly strong, especially from the 1980s onwards. In fact, since the 1980s, both systems have either stayed far from the equilibrium or departed from it (the bond model). From Table 6.36, which displays error correction adjustment coefficients from all estimated models[29], it appears evident that the adjustment in Model 1 and Model 2 was carried out through lower salary growth (negative coefficient) rather than higher bond or stock returns (coefficients for bonds and equities are either negative or not significant). This is in line with the results of the previous section.

Multivariate models (in particular Models 5 and 6) provide a more interesting picture. The magnitude of the errors is lower in absolute terms and there is a more evident pattern of adjustment to restore the equilibrium. As shown in the last graph of Figure 6, with a dummy variable for the year 1975 and one from 1975 onwards included as exogenous factors, the adjustment pattern is even more evident, because the unusual behaviour of both labour and financial markets in the mid 1970s may have pushed upwards the estimated long-run equilibrium position of Models 5 and 6. However, because adding ad-hoc dummy variables to control for unusual events is essentially an arbitrary exercise, it was deemed preferable to focus on models estimated with no exogenous dummy variables.

[29] The coefficients reported in the table are derived from each endogenous variable error correction model and represent the sensitivity of each endogenous variable to the disequilibrium between the level of the salary index and its equilibrium position. Because the nature of the dynamic adjustment process is not the main focus of this paper (Clements and Galvão, 2004 is a reference on the latter), coefficients referred to autoregressive factors in the error correction models are not reported

From Model 5 or 6, one could read the dynamic relationship between wages and asset markets as follows: when wages are above the equilibrium the adjustment can take place in three alternative ways: 1) one or more of the variables with a positive long-term relationship with salaries (e.g. equities) grow faster than real wages, 2) one or more of the variables with a negative long-term relationship with salaries (e.g. bonds) grow slower than real wages, 3) real salaries fall. Clearly, the reverse would be true when wages are below their long-term equilibrium.

To illustrate this, from the graph referred to Model 6, one could argue real wages were very high in the mid 1970s compared to their long-run position (possibly because they did not reflect underlying productivity gains but only union-sponsored pay deals). Subsequently, a fall in average earnings between 1975 and 1979 has helped the system restore the equilibrium. Real salaries were then below their long-run level in the early 1980s, perhaps due to strong equity returns, but a combination of factors (higher real salary growth and bond returns, together with the housing market crash) have once again brought the system back towards the equilibrium in the early 1990s. Finally, slow real salary growth at the beginning of the 1990s and strong equity returns after 1995 have pushed the system again below the long-run equilibrium, leaving it to the prolonged bear market after 2000 to correct the imbalance.

However, from Table 6.36 only the adjustment coefficient for real salaries, real bonds and US equities are significant at 95% level. In fact, between the 1980s and 2000 the system has been below its equilibrium and one could argue rising bond returns, with a shift in inflationary expectations, have played a major role in restoring the equilibrium position.

If Models 5 and 6 appear to have the best long-run properties, in the short-term models should be assessed using the error correction equation for wages and considering their ability to predict movements in the wage index. Models can also be compared the predicting power of the wage error-correction models using different asset classes. The dependent variable here is wage growth (equal to first difference as we are working in logs) and the independent variables are the distance from the long-run equilibrium and lagged asset returns (up to lag 4). All of the eight wage growth models considered in Table 6.37 are significant (F test rejects the hypothesis all coefficient are equal to zero). The R-square, which captures the proportion of salary growth explained by the model (and the adjusted R-square, which penalises the introduction of unnecessary variables), is highest in Model 5 which includes bills alongside bonds (40.8%), while all of the other models have an R-square in the range of 20%. Interestingly, Model 1 has a higher R-square than Model 2, confirming that in the short-term (but not in the long-term as discussed before) bonds have a higher correlation with wages than equities.

Finally, because cointegration analysis involves the estimation of error correction models for all variables in the system, it is important to consider the correlation between the residuals from each error-correction equation and the response implied by the error correction models to innovations (random shocks to one of the variables in the system), which are transmitted from an equation to another through the correlation structure between residuals.

Table 6.38 shows residual correlation implied by Model 6 and Figure 20 shows impulse-response functions to a real salary innovation, implied by a standard Cholesky decomposition of the covariance matrix of residuals. The portion of real salary variance unaccounted for by the model appears to be correlated with the unexpected component of real house prices, perhaps indicating an additional common factor between the property market and wages not captured by the model. Arguably, the interrelation between labour demand pressures and property supply constraints could be candidates for the missing variable in this context. Table 6.38 also shows basic residual-based mis-specification tests carried out on the error correction model using Model 6. Autocorrelation and non-normalities are rejected at conventional levels, while there are signs of heteroskedasticity due to the pattern observed in the 1970s, which led to a concentration of high residuals in absolute value.

Figure 20 shows the response implied by the cointegrating model (Model 6) to a positive unexpected shock to real wages. Following Sims (1980), residuals are orthogonalised using the Cholesky decomposition to enable identification (common components are attributed to the variable that comes first in the Cholesky ordering). Interestingly, while the immediate response of the stock market is negative, in the longer term there is a positive co-movement between the two series (after 3 years both real salaries and real equities are above the level they would have reached in the absence of a shock). Conversely, real bonds and real house prices display a short-term positive correlation with real salaries. This provides further support to the hypothesis of a long-term (but not a short-term) link between real salaries and the stock market. Given that impulse-response results depend on the Cholesky order, the exercise was performed using alternative orders as a sensitivity test and no significantly different results emerged.

As a second step, the multivariate analysis was replicated in nominal terms, with the inclusion of the price index in some specifications to proxy the return profile of index-linked bonds. The hypothesis of a common stochastic trend cannot be rejected at 95% level for a range of alternative models, although the evidence is more mixed than in the case of variables in real terms (Table 6.39). With Models 5 and 6b, which include the price index, the hypothesis of at most one cointegrating relationship is rejected by the Trace Statistic (but not the Maximum Eigenvalue) test, suggesting the possibility of a more complex pattern with two long-run equilibrium relationships: one between bonds and wages and the other between equities and prices.

Figure 21 shows the estimated relationships and the equilibrium errors for the key bivariate and multivariate models. In all cases the pattern of the 1970s has influenced the results, even more evidently than with variables in real terms. Model 3 would suggest an insignificant relationship between wages and the stock market, but the equilibrium errors do not suggest it is a stable equilibrium. On the other hand, model 5b, which is estimated with a linear trend because the nominal AEI process is likely to be characterised by a trend (see the discussion in Section 5.2), suggests a positive and significant relationship with the stock market. Finally Model 6b, whose equilibrium errors are visually more stationary, suggests a close association between prices and wages, but a significant residual relationship between wages and US unhedged equities.

Interestingly, adjustment coefficients in nominal models provide a mixed answer on the role of bonds in the adjustment process. Whilst in Model 5 rising bond returns restore the equilibrium when wages are above it, the reverse is true in Model 6b, in which the adjustment coefficient for bonds is negative.

In nominal terms the residual component of salary growth is negatively correlated with US unhedged equities, perhaps through an exchange rate effect (salary growth may slow down as the UK currency weakens). Finally, impulse-response analysis (Figure 22) is consistent with the pattern observed in real terms: although there is a short-term negative response of the stock market to salary innovations, there is evidence of a positive longer-term association. The latter suggests that, although the behaviour of nominal salaries cannot be easily modelled through a single cointegrating relationship, a meaningful link with the stock market may indeed exist.

6.6 *A longer-term monthly sample*

This section replicates the analysis using a longer term sample (1920-2003), which is less influenced by the unusual behaviour of the 1970s. The key conclusions of Section 5 on trends and unit roots are confirmed with the longer sample, in particular the presence of a linear trend in the nominal and real average earnings process. In addition to this, there is also evidence of a linear trend in consumer prices, which suggests models in nominal terms are better estimated if a linear trend is allowed for. Table 6.43 displays the results and Figure 23 shows equilibrium errors from selected models.

In real terms, the analysis of longer-term monthly data confirms the hypothesis of a long-run equilibrium with the stock market, while the results are more mixed for real bonds (first graph in

Figure 23) due to lengthy adjustment process and some evidence of a structural break around the 1950s. On the other hand, as far as the stock market is concerned, the second graph in Figure 23 shows real salaries above the cointegrating relationship in the 1940s and 1950s as well as in the mid-1970s, leading to an adjustment process which predominantly relied on stronger equity returns. The reverse was true in the late 1990s, when the system was below the level implied by the cointegrating relationship just before the stock market downturn.

In nominal terms, there is significant evidence of a long-term relationship between nominal salaries and consumer prices, as shown in the third graph of Figure 23. Interestingly, although adjustment cycles are long (e.g. 20 years between the 1950s and the 1970s), the evidence is here stronger than in the case of the quarterly and annual sample of Section 6.1. Ultimately, the existence of a stable equilibrium depends on whether the system moves because of secular changes due to stronger underlying productivity growth or because of cyclical salary growth fuelling inflationary pressures.

The fourth graph in Figure 23 shows a multivariate model with salary growth explained by prices and equity returns (as well as bonds whose coefficient is however not significant), which depicts a very similar picture to that of the real salary-real equity equilibrium (Model 5). Finally, a model to explain nominal salaries using nominal and index-linked bonds was also fitted, but, although a time span of just over 20 years is not enough from which to draw definite conclusions, there is little evidence of a stable relationship, as salaries have been far below the equilibrium in the early 1980s and above it since the mid-1990s (Model 6).

6.7 *Implications of results for pension fund asset allocation*

The results showed so far have displayed evidence of a long-term relationship between wages and a range of asset classes, including equities. The Appendix relates the results to minimum risk asset allocation by means of a simple model in which pension funds incorporate the long-term equilibrium relationship in a simple asset liability modelling exercise. This assumes non wage-related liabilities can be hedged by an appropriately structured bond portfolio, while the wage-related component is akin to a zero coupon bond maturing at retirement.

The empirical implementation is closer to an ABO definition of liability, but the results are compatible with either an ABO or a PBO definition, given that whether the current or the projected wage is the variable of interest should not have a material impact on the estimated long-run relationships. This is because with an ABO (Accrued Benefit Obligation) definition the liability is re-calculated every time an actuarial valuation is performed using the current wage of active members, while in a PBO framework the liability is calculated each time using the projected wage at retirement of active members (the latter usually derived as a deterministic function of the current wage). An extension to this model could look at the relationship between current wages and lagged asset prices to proxy a PBO setting.

The results suggest the long-run hedging portfolios for real active members' liabilities will have an equity share (including international equities) different from zero (in the multivariate estimated models the coefficients range from just below 30% to around 50%). There are however some important caveats to bear in mind when it comes to assessing the full implications for the overall risk management of a pension fund.

Firstly, liabilities can be easily broken down into a real and a nominal component if the volatility of inflation is low and there are stable relationships between prices and wages. Historically, the relationship between prices and wages as well as between prices and financial markets has been significantly more complex. In fact, whilst even with variables in nominal terms there is significant support for a bivariate long-run equilibrium with the stock market, results are more mixed if translated into multivariate models. Although with the longer-term monthly sample there is some evidence of a single stochastic trend shaping the relationship between salaries, equities, bonds and prices, the quarterly sample suggests there might be more than one cointegrating vector. The relationship between each asset class and consumer prices may in fact be characterised by different properties in terms of both the equilibrium and the adjustment process. Ahmed and Cardinale (2004) review the vast literature in this area (Fama and Schwert, 1977b and Boudoukh and Richardson, 1993 are two of the most important contributions) and explore the relationship between equities and prices in the four largest economies. Although they find some evidence of a dynamic relationship, no definite conclusion is reached on the existence of a stable adjustment pattern. Similar conclusions are reached by Cardinale (2003) on the relationship between property returns and inflation.

Secondly, cointegrating equilibriums, similarly to other regression-based models, are estimated with linear models. The true underlying relationship may be non-linear, perhaps due to asymmetric behaviour for positive vs. negative shocks. An extension to this paper may apply the framework proposed by Granger and Yoon (2002) to deal with non-linearities. Furthermore, cointegrating relationships may not be stable over time, as highlighted in several occasions in the paper, and therefore historical equilibrium relationships should not be the only inputs to pension funds risk budgeting exercises. However, equilibrium relationships should still represent a key reference point, from which any departure would have to be justified with a reasonable explanation of why historical relationships no longer hold in a given economic environment.

Thirdly, an asset allocation which tracks long-term equilibrium relationships leads to short-term volatility, because the system dynamically adjust to the long-term position and, as the results showed, adjustment lags can be significantly long. An alternative strategy would entail dynamically exploiting the deviations from the long-run equilibrium position, using the information contained in the error correction model (e.g. adjustment coefficients). As anticipated in Section 1, the ultimate choice depends on rebalancing costs to follow the long term equilibrium strategy and the level of resources available within the pension fund, given that exploiting temporary deviations from the equilibrium is riskier and require more careful and frequent monitoring.

Fourthly, the paper considers the aggregate relationships between salaries and asset markets, implicitly assuming a high degree of correlation between firm-wide salary increases and the aggregate index. Pension funds should assess whether the age structure of the workforce within the company is significantly different from the aggregate earnings index or whether the age distribution of earnings is influenced by firmwide policies or sector patterns.

Fifthly, the paper used a portfolio of 10-year nominal bonds in the estimation of cointegrating equilibriums in the quarterly and annual sample, while, bonds with longer durations and consols are employed in the longer term monthly sample (Section 6.6). Results may change using different maturities, although this is not expected to significantly alter the key implications.

Finally, the paper concentrates on active members' liabilities and therefore the above mentioned findings (positive equity share) relate to the portion of the total liability referring to active members. A portfolio close to 100% fixed income may therefore be justified as a minimum risk position in mature schemes with high proportion of pensioners and deferred members.

7 Conclusion

If active members' liabilities are defined on an ongoing basis, risk management for defined benefit pension schemes must take into serious consideration the relationship between salary increases and the return profile of alternative assets in which pension funds may invest.

This paper investigated whether a long-run equilibrium relationship exists between the labour market and asset markets using UK historical data. Consistently with economic theory, the paper finds evidence of a long-run link not only between salaries and bonds but also with other assets such as domestic and international equities. Long-run movements in real salaries can be captured by a multivariate equilibrium model which includes bonds, equities and property. In the short-term however the evidence of correlation is more mixed, because, whilst there is evidence of a dynamic adjustment process towards the equilibrium, salaries have departed from the estimated long-term relationships for relatively long periods. This also translates into a significant long-run correlation between nominal salaries and nominal assets including equities, although the evidence is less conclusive as to whether the equilibrium between wages, prices and asset markets can be characterised by a single long-run equation. A more complex model could therefore incorporate the dynamic relationship between each asset class and consumer prices.

From the point of view of defined benefit pension funds, the implication of these results challenges the view according to which salary-linked liabilities can be perfectly hedged with fixed income instruments, as claimed by several influential papers in the actuarial literature (Wilkie, 1995 or Smith, 1998). The key contribution of the paper is to show that no asset has historically been a perfect hedge for salary-linked liabilities and the best historical hedge was a composite portfolio, which included not only fixed income instruments, but also domestic equities, foreign equities and property. Although index linked bonds were not in the menu of assets because of lack of historical data, consumer prices have not been historically a perfect hedge for salaries and indeed long-run coefficients associated with other assets including equities remained significant in multivariate models with the Retail Price Index among the predictors of wages. The Appendix reviews how long-run equilibrium relationships can be translated into minimum risk asset allocations.

It is clear, however, that long-run relationships will not be the only factor driving minimum risk asset allocations for defined benefit pension funds. Rebalancing costs and the extent to which salaries and assets deviate from the long-run equilibrium path in the short- and medium-term should also be taken into account. If deviations are substantial and persistent and rebalancing costs are high, minimum risk asset allocations will likely to look significantly different from those consistent with the long-run equilibrium. Moreover, salary-linked liabilities are only a fraction of defined benefit pension fund total liabilities and therefore an overall minimum risk portfolio will have also to take into account how significant is the proportion of salary-linked liabilities as a proportion of total pension liabilities.

Whilst a breakdown of defined benefit liabilities is not disclosed in company accounts, the ratio of current service cost (referred to active members) over interest cost on pension scheme liabilities (referred to all members) can be employed as a proxy. Figure 24 display average values of the ratio by broad sectors using FRS 17 disclosures for the accounting year 2002 collected by Watson Wyatt for companies in the FTSE 350 share index (the overall average within the sample is just above 60% and the median is slightly below 55%)[30]. The conclusions of the paper will be of greater importance for relatively young schemes (e.g. in sectors such as IT & Telecom or Retail where the ratio is close or above 1) with a high proportion of active members as opposed to traditional sectors (e.g. General Industrials) were schemes are on average more mature.

[30] The first graphs considers only UK defined benefit schemes while the second includes non-UK schemes of UK companies' overseas subsidiaries

8 Appendix: Cointegration and Pension Liabilities

This appendix illustrates the relationship between cointegrating relationships between wages and various asset classes indices. The level of accrued pension liability for an individual active member is proportional to the current level of the real wage index times an appropriately structured bond or bond portfolio[31]:

$$L \propto WB \tag{8.1}$$

where W is the wage and B is the bond. Taking logarithms of Eq. (8.1), we have:

$$\log L = C + \log W + \log B \tag{8.2}$$

or:

$$l = C + w + b \tag{8.3}$$

where $w \equiv \log(W)$ and $b \equiv \log(B)$

8.1 *Cointegrating Vectors*

We separately have a cointegrating relationship:

$$w = \delta + \alpha b + \sum_{k=1}^{N} \beta_k x_k \tag{8.4}$$

where the x_k are the level of various indices (e.g., equity index) which could be invested in, all expressed in either nominal or real terms. Under the assumption of cointegration, the difference between the level of the wage index and the cointegrating function in Eq. (8.4) is stationary. It must be noted that the difference equation does not have to be white noise and it will be typically an autocorrelated process with relatively short memory. Therefore, although (8.4) will not hold at every point in time, it will hold in the long-run.

Substituting Eq. (8.4) into Eq. (8.3) we have:

$$l = \kappa + (\alpha + 1)b + \sum_{k=1}^{N} \beta_k x_k \tag{8.5}$$

where $\kappa = C + \delta$. Suppose that β_k =0, $k = 1...N$ then clearly a bond matched portfolio is the best match for the liability in the long-run. Whether this is so in practice is an empirical question. Cointegration allows tests of the joint null hypothesis: $\alpha = 0$, β_k =0, $k = 1...N$.

8.1.1 Minimum Risk Asset Allocation

The next step from Eq. (8.5) is to determine the minimum risk asset allocation.

[31] In the case of the UK the bond portfolio will be closer to a real bond portfolio because pensions in payment are indexed to prices (albeit with a 5% cap)

Our assets are:

$$A = \sum_{k=1}^{N} Q_k X_k + Q_b B + M \qquad (8.6)$$

where Q are quantities (number of shares, bonds, properties etc.) and M is cash. We want to choose the Q and M to minimize the difference with liabilities. To do this we take logarithms so that:

$$a = \log\left[\sum_{k=1}^{N} Q_k X_k + Q_b B + M\right] \qquad (8.7)$$

Hence:

$$a - l = \log\left[\sum_{k=1}^{N} Q_k X_k + Q_b B + M\right] - \left[\kappa + (\alpha+1)b + \sum_{k=1}^{N} \beta_k x_k\right] \qquad (8.8)$$

which can then be used to calculate sensitivities to levels of the indices and therefore to compute minimum risk asset allocations.

8.2 *Minimum Risk Portfolio*

To determine the minimum risk asset allocations, we compute sensitivities with respect to the levels of bond prices and asset market indices in Eq. (8.8). We find:

$$\frac{\partial(a-l)}{\partial x_k} = \frac{Q_k e^{x_k}}{A} - \beta_k \qquad (8.9)$$

$$\frac{\partial(a-l)}{\partial b} = \frac{Q_b e^{b}}{A} - (\alpha+1) \qquad (8.10)$$

We choose asset allocations to set these sensitivities to zero, thereby achieving a minimum level of risk. Therefore, the minimum risk long-run asset allocations are a fraction $(\alpha+1)$ in bonds, β_k in asset k and the remainder (which can be negative) in cash.

8.2.1 Example

We consider as an example the following cointegrating relationship:

$$w = -4.501 + 0.483\, ftse - 0.568b \qquad (8.11)$$

then in Eq. (8.5) we have:

$$l = C - 4.501 + (-0.568+1)b + 0.483\, ftse \qquad (8.12)$$

or:

$$l = \kappa + (0.422)b + 0.483\, ftse \qquad (8.14)$$

Consider now the minimum risk portfolio. Using Eq. (8.8), we have:

$$a - l = \log\left[Q_k e^{ftse} + Q_b e^b + M\right] - \left[\kappa + (0.422)b + 0.483\,ftse\right] \qquad (8.15)$$

The sensitivity of this imbalance to b is:

$$\frac{\partial(a-l)}{\partial b} = \frac{Q_b e^b}{A} - 0.422 \qquad (8.16)$$

and with respect to $ftse$ is:

$$\frac{\partial(a-l)}{\partial ftse} = \frac{Q_{fts} e^{ftse}}{A} - 0.483 \qquad (8.16)$$

Using Eq. (8.15), we have that for sensitivity to be zero, the share of bonds in total assets is 42.2%. Similarly, the share of equity in the minimum risk long-run portfolio is 48.3%. The remainder of the portfolio is held in cash.

8.3 *Cointegration and Career-Average Pension Schemes*

For career-average schemes, the analysis is similar. The liability for a career-average scheme is:

$$L_N \propto B\left[\sum_{k=1}^{N} C_{ik} \frac{W_R}{W_k}\right] \qquad (8.17)$$

where B is a bond, C_{ik} is the wage of individual i in period k and W_k is the value of the wage index and R is the projected retirement date. Taking logarithms of Eq. (8.17) we have:

$$l = b + w + \log\left[\sum_{k=1}^{N} \frac{C_{ik}}{W_k}\right] \qquad (8.18)$$

where the last term is an idiosyncratic term capturing ratio of an individual's wages to the wage index. With the exception of the idiosyncratic term, the form of Eq. (1.1) is equivalent to our earlier analysis. Given we expect no relationship between an aggregation of idiosyncratic terms and asset allocations, the results above about minimum-risk portfolios follow.

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Tables

Table 4.1 Nominal quarterly log returns (quarterly sample): basic statistics

	BILL	BOND	FTSE	HOUSE	AEI	WAGES	SP500
Mean	2.08%	2.36%	2.95%	2.33%	2.13%	2.21%	2.80%
Median	1.86%	2.26%	4.65%	2.08%	1.77%	1.97%	3.19%
Maximum	3.98%	19.57%	59.01%	11.96%	8.00%	10.47%	23.72%
Minimum	0.92%	-6.33%	-32.00%	-4.71%	-1.65%	-0.61%	-40.51%
Std. Dev.	0.78%	3.24%	10.39%	2.69%	1.48%	1.47%	9.61%
Obs.	158	158	158	158	158	158	158

Table 4.2 Real quarterly log returns (quarterly sample): basic statistics

	BILL	BOND	FTSE	HOUSE	AEI	WAGES	SP500
Mean	0.47%	0.75%	1.32%	0.71%	0.50%	0.59%	1.18%
Median	0.68%	0.99%	2.69%	0.81%	0.55%	0.67%	1.75%
Maximum	2.78%	12.57%	49.70%	9.93%	5.61%	7.12%	22.51%
Minimum	-6.93%	-9.31%	-32.59%	-5.95%	-4.42%	-7.08%	-41.38%
Std. Dev.	1.29%	3.46%	10.13%	2.75%	1.42%	1.47%	9.62%
Obs.	158	158	158	158	158	158	158

Table 4.3 Real and nominal log annual returns (annual sample): basic statistics

	BOND	REAL BOND	FTSE	REAL FTSE	WAGES	REAL WAGES (2)	REAL WAGES (1)
Mean	4.77%	2.18%	7.20%	4.62%	3.87%	1.29%	1.31%
Median	3.78%	1.80%	5.47%	3.65%	3.54%	1.34%	1.23%
Maximum	33.47%	49.85%	92.47%	70.16%	25.13%	19.49%	19.49%
Minimum	-21.20%	-41.14%	-72.89%	-90.22%	-25.66%	-20.66%	-18.09%
Std. Dev.	6.76%	9.69%	15.31%	16.17%	6.33%	5.34%	4.25%
Obs.	151	151	151	151	151	151	151

Table 4.4 Correlation between nominal log returns (quarterly sample)

	SP500	BILL	BOND	FTSE	HOUSE	AEI	WAGES
SP500	1	4.28%	4.35%	62.81%	-5.19%	-4.34%	-12.33%
BILL	4.28%	1	15.18%	4.23%	-13.00	41.05	33.52%
BOND	4.35%	15.18%	1	51.79%	-19.64%	-4.03%	8.85%
FTSE	62.81%	4.23%	51.79%	1	-5.43%	-5.64%	4.50%
HOUSE	-5.19%	-13.00%	-19.64%	-5.43%	1	21.01%	25.56%
AEI	-4.34	41.05%	-4.03%	-5.64%	21.01%	1	73.32%
WAGES	-12.33%	33.52%	8.85%	4.50%	25.56%	73.32%	1

Table 4.5 Correlation between real log returns (quarterly sample)

	SP500	BILL	BOND	FTSE	HOUSE	AEI	WAGES
SP500	1	8.72%	7.62%	63.41%	-3.31%	-4.28%	-11.45%
BILL	8.72%	1	42.79%	7.22%	15.36%	51.09%	48.99%
BOND	7.62%	42.79%	1	49.39%	-4.09%	14.99%	27.33%
FTSE	63.41%	7.22%	49.39%	1	-4.18	-6.65	3.9%
HOUSE	-3.31	15.36%	-4.09%	-4.18%	1	26.36%	31.75%
AEI	-4.28	51.09%	14.99%	-6.65	26.36%	1	72.06%
WAGES	-11.45%	48.99%	27.33%	3.9%	31.75%	72.06%	1

Table 4.6 Correlation between nominal and real log returns (annual sample)

	BOND	REAL BOND	FTSE	REAL FTSE	WAGES	REAL WAGES (2)	REAL WAGES (1)
BOND	1	72.27%	45.99%	46.01%	1.84%	5.05%	9.62%
REAL BOND	72.27%	1	24.73%	52.96%	-43.39%	9.9%	37.97%
FTSE	45.99%	24.73%	1	91.49%	10.54%	2.93	2.78%
REAL FTSE	46.01%	52.96%	91.49%	1	-16.55%	5.79%	20.79%
WAGES	1.84%	-43.39%	10.54%	-16.55%	1	37.95%	38.09%
REAL WAGES (2)	5.05%	9.9%	2.93	5.79%	37.95%	1	54.04%
REAL WAGES (1)	9.62%	37.97%	2.78%	20.79%	38.09%	54.04%	1

Table 4.7 Autocorrelation of nominal and real returns (quarterly sample, partial autocorrelation at different lags). * Within two standard errors bound

Series	Lag 1	Lag 2	Lag 3	Lag 4	Lag 5	Lag 6	Lag 7	Lag 8	Lag 9
Nominal AEI	0.58*	0.20*	0.22*	0.20*	-0.09	0.02	0.06	0.17*	0.04
Nominal wages	0.64*	0.39	-0.09	-0.08	0.25*	0.06	0.01	-0.01	0.13
Real AEI	0.04	-0.29*	-0.10	0.14	-0.14	-0.101	0.01	0.18*	0.04
Real wages	-0.08	0.12	-0.18*	0.15*	0.02	-0.08	0.07	0.17*	-0.15
Nominal bonds	-0.01	0.04	-0.03	0.10	-0.11	0.06	0.10	0.15*	-0.15*
Real bonds	-0.02	0.18*	-0.03	0.16*	-0.05	0.13	0.08	0.13	-0.184*
Nominal equities	0.10	-0.05	0.07	-0.12	-0.08	0.01	-0.01	0.02	0.04
Real equities	0.10	-0.02	0.07	-0.10	-0.06	0.03	-0.02	0.03	0.03
Nominal house prices	0.62*	0.064	0.20*	0.12	-0.14	-0.22*	0.17*	-0.03	-0.07
Real house prices	0.60*	0.20*	0.06	0.04	-0.17*	-0.20*	0.09	0.00	-0.05
Nominal bills	0.94*	-0.24*	-0.02	0.01	0.01	0..02	-0.06	0.03	-0.06
Real bills	0.39*	0.29*	0.07	0.55*	-0.20*	0.02	0.14	0.24*	-0.07

Table 5.1 Phillips Perron unit root tests, quarterly data-set, 4 lags, critical values at 95% level

Series	Test $\gamma=0$ (PP statistic)	Critical value	Test $a_1=0$ given $\gamma=0$	Critical value	Test $a_0=0$ given $\gamma=0$	Critical value
Nominal AEI	0.93	-3.44	-3.10	2.79	1.33	3.11
Nominal wages	0.37	-3.44	-2.30	2.79	-0.95	3.11
Real AEI	-3.36	-3.44	3.31	2.79	3.53	3.11
Real wages	-2.64	-3.44	2.61	2.79	2.49	3.11
Nominal bonds	-2.57	-3.44	2.69	2.79	2.89	3.11
Real bonds	-0.89	-3.44	2.28	2.79	0.76	3.11
Nominal equities	-2.19	-3.44	1.84	2.79	2.36	3.11
Real equities	-2.10	-3.44	1.82	2.79	2.01	3.11
Nominal house prices	-1.09	-3.44	0.27	2.79	1.85	3.11
Real house prices	-1.95	-3.44	0.98	2.79	0.90	3.11
Nominal bills	-0.85	-3.44	0.03	2.79	0.70	3.11
Real bills	-0.53	-3.44	3.04	2.79	0.18	3.11
Nominal (unhedged) US equities	-2.27	-3.44	2.06	2.79	2.60	3.11
Real (unhedged) US equities	-1.66	-3.44	1.52	2.79	1.61	3.11

Table 5.2 Phillips Perron unit root tests, annual data-set, 1 lag, critical values at 95% level

Series	Test $\gamma=0$ (PP statistic)	Critical value	Test $a_1=0$ given $\gamma=0$	Critical value	Test $a_0=0$ given $\gamma=0$	Critical value
Nominal wages	-0.17	-3.44	2.07	2.79	-0.57	3.11
Real wages (1)	-3.21	-3.44	3.26	2.79	3.11	3.11
Real wages (2)	-1.90	-3.44	1.95	2.79	1.83	3.11
Nominal bonds	1.66	-3.44	-0.17	2.79	-1.53	3.11
Real bonds	-1.78	-3.44	1.59	2.79	1.66	3.11
Nominal equities	0.02	-3.44	0.96	2.79	0.16	3.11
Real equities	-1.67	-3.44	1.84	2.79	0.80	3.11

Table 6.1 Bonds and salaries: **Johansen cointegration tests (4 lags, intercept, no linear trend) Quarterly data 1963-2002**

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Bond	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	21.83**	28.38**	12.53	17.18	19.67*	23.32*
Real Wages	27.96**	33.12**	11.43	16.95	22.18**	31.62**
Bond						
Salary SA	26.39**	32.18**	12.22	18.41	20.75**	27.06**
Wages	37.42**	43.46**	19.40*	28.00**	22.51**	30.69**

Table 6.2 Bonds and salaries: Engle and Granger cointegration tests (4 lags, intercept, no linear trend) Quarterly data 1963-2002

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Bond	**ADF**	**Phillips Perron**	**ADF**	**Phillips Perron**	**ADF**	**Phillips Perron**
Real Salary SA	-1.19	-1.19	-1.42	-1.58	-1.72	-1.96
Real Wages	-1.48	-1.46	-1.67	-2.17	-2.66	-2.59
Bond						
Salary SA	-0.78	-0.67	-2.35	-2.32	-1.74	-1.90
Wages	-1.10	-0.99	-2.54	-2.21	-2.41	-2.12

Table 6.3 Bonds and salaries cointegrating relationships: long-run coefficient estimates (4 lags, intercept, no linear trend) Quarterly data 1963-2002

^=Not significant at 95% level

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Bond	**α**	**β**	**α**	**β**	**α**	**β**
Real Salary SA	-4.57	2.15^	5.68	0.14^	-11.90	4.13
Real Wages	-4.70	0.88^	2.52^	0.23^	-23.56	2.57
Bond						
Salary SA	2.19	1.08	2.33	0.77	-21.97	5.20
Wages	4.90	1.05	-3.51	0.85	6.33	0.22^

Table 6.4 Bonds and salaries: Johansen cointegration tests (4 lags, intercept, linear trend) Quarterly data 1963-2002

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Bond	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	7.09	7.54	5.30	9.95	4.81	5.54
Real Wages	7.11	11.68	5.57	8.72	9.78	9.79
Bond						
Salary SA	14.54*	19.02*	10.38	13.54	9.55	15.19
Wages	19.02**	23.71**	13.28	19.84*	11.94	18.93*

Table 6.5 Bonds and salaries: error correction models adjustment coefficients (4 lags, intercept, no linear trend) Quarterly data 1963-2002

^=Not significant at 95% level . Dependent variable in the cointegrating model in bold

	1963:1-2002:3	
Real Bond	**D(Real Bond)**	**D(Real Salary SA/Wage)**
Real Salary SA	0.006	0.004
Real Wages	0.021	0.009
Bond	**D(Bond)**	**D(Salary SA/Wage)**
Salary SA	-0.040	-0.009
Wages	-0.044	-0.012
	1850-2001	
Real Bond	**D(Real Bond)**	**D(Real Wage)**
	-0.025	-0.013
Bond	**D(Bond)**	**D(Wage)**
	-0.037	-0.011

Table 6.6 Bonds and salaries: Johansen cointegration tests (4 lags, intercept, no linear trend) Monthly data 1963-2002

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:12		1963:1-1979:12		1980:1-2002:09	
Real Bond	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	20.16**	26.51**	7.45	11.33	26.86**	31.55**
Bond						
Salary SA	83.81**	92.56**	41.09**	45.69**	48.26**	53.25**

Table 6.7 Bonds and salaries: Engle and Granger cointegration tests (4 lags, intercept, no linear trend) Monthly data 1963-2002

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:12		1963:1-1979:12		1980:1-2002:09	
Real Bond	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	6.48	7.81	5.39	7.99	6.96	7.40
Bond						
Salary SA	34.97**	43.55**	20.15**	23.08**	22.64**	26.55**

Table 6.8 Bonds and salaries cointegrating relationships: long-run coefficient estimates (4 lags, intercept, no linear trend) Monthly data 1963-2002
^=Not significant at 95% level

	1963:1-2002:12		1963:1-1979:12		1980:1-2002:09	
Real Bond	**α**	**β**	**α**	**β**	**α**	**β**
Real Salary SA	2.85	0.35	8.88	-0.96	-10.11	3.78
Bond						
Salary SA	2.39	1.01	2.28	0.84	22.16	-2.52

Table 6.9 Bonds and salaries: Johansen cointegration tests (1 lag, intercept, no linear trend) Annual data 1850-2001

***=significant at 95% level, **=significant at 99% level**

	1850-2001		1900-2001		1932-2001	
Real Bond	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Wage 2	20.28**	24.54*	15.21	18.83	15.46	16.65
Bond						
Wage	66.07**	72.74**	45.98**	51.14**	43.54**	45.92**

Table 6.10 Bonds and salaries: Johansen cointegration tests (1 lag, intercept, linear trend) Annual data 1850-2001

***=significant at 95% level, **=significant at 99% level**

	1850-2001		1900-2001		1932-2001	
Real Bond	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Wage (2)	7.57	7.74	6.90	7.14	2.04	2.08
Bond						
Wage	39.45**	45.20**	28.07**	30.94**	28.58**	28.75**

Table 6.11 Bonds and salaries cointegrating relationships: long-run coefficient estimates (1 lag, intercept, no linear trend) Annual data 1850-2001

^=Not significant at 95% level

	1850-2001		1900-2001		1932-2001	
Real Bond	**α**	**β**	**α**	**β**	**α**	**β**
Real Wage (2)	1.55	1.76	2.06^	1.66	0.86^	2.11
Bond						
Wage	4.75	1.57	4.96	1.45	6.22	1.05

Table 6.12 Equities and salaries: Johansen cointegration tests (4 lags, intercept, no linear trend) Quarterly data 1963-2002. Dependent variables in bold.

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Equity	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	21.08**	24.77**	10.34	15.45	21.83**	25.09**
Real Wages	25.64**	28.52**	12.87	18.92	27.09**	36.68**
Equity						
Salary SA	17.43*	22.39*	9.37	15.84	23.39**	30.20**
Wages	24.67**	29.67**	16.26*	24.37*	26.62**	32.15**

Table 6.13 Equities and salaries: Engle and Granger cointegration tests (4 lags, intercept, no linear trend) Quarterly data 1963-2002. Dependent variables in bold

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Equity	**ADF**	**Phillips Perron**	**ADF**	**Phillips Perron**	**ADF**	**Phillips Perron**
Real Salary SA	-2.13	-2.24	-2.34	-2.23	-1.22	-1.50
Real Wages	-2.05	-2.07	-2.38	-2.26	-1.10	-1.60
Equity						
Salary SA	-1.76	-1.82	-2.42	-2.31	-1.51	-1.72
Wages	-1.82	-1.85	-2.43	-2.28	-1.07	-1.32

Table 6.14 Equities and salaries cointegrating relationships: long-run coefficient estimates (4 lags, intercept, no linear trend) Quarterly data 1963-2002. Dependent variables in bold

^=Not significant at 95% level

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Equity	**α**	**β**	**α**	**β**	**α**	**β**
Real Salary SA	-13.69	4.63	4.31	0.77^	-21.71	6.71
Real Wages	-34.13	3.60	4.25^	0.83^	-50.83	5.17
Equity						
Salary SA	2.15	1.28	2.59^	-1.07^	-7.73	3.14
Wages	-5.79	1.23	19.26	-2.32	-23.88	2.67

Table 6.15 Equities and salaries: Johansen cointegration tests (4 lags, intercept, linear trend) Quarterly data 1963-2002. Dependent variables in bold

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Equity	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	7.58	7.92	8.00	10.10	9.08	12.32
Real Wages	11.56	11.67	7.80	9.74	20.33**	28.36**
Equity						
Salary SA	11.33	14.08	6.61	10.71	19.17**	25.99**
Wages	12.08	14.72	9.00	16.33*	19.32**	24.14**

Table 6.16 Equities and salaries: error correction models adjustment coefficients (4 lags, intercept, no linear trend). Quarterly data 1963-2002 and annual data 1850-2001

^=Not significant at 95% level . Dependent variable in the cointegrating model in bold

	1963:1-2002:3	
Real Equity	**D(Real Equity)**	**D(Real Salary SA/Wage)**
Real Salary SA	0.003^	0.006
Real Wages	0.012^	0.009
Equity	**D(Equity)**	**D(Salary SA/Wage)**
Salary SA	-0.041	-0.006
Wages	-0.031	-0.006
	1850-2001	
Real Equity	**D(Real Equity)**	**D(Wage)**
Real Wages	-0.082	-0.010
Equity	**D(Equity)**	**D(Wage)**
Wages	-0.043	-0.009

Table 6.17 Equities and salaries: Johansen cointegration tests (4 lags, intercept, no linear trend) Monthly data 1963-2002. Dependent variables in bold

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:12		1963:1-1979:12		1980:1-2002:09	
Real Equity	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	22.71**	27.19**	9.46	13.08	28.60**	32.71**
Equity						
Salary SA	72.27**	76.27**	9.37	15.84	23.39**	30.20**

Table 6.18 Equities and salaries: Engle and Granger cointegration tests (4 lags, intercept, no linear trend) Monthly data 1963-2002. Dependent variables in bold

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:12		1963:1-1979:12		1980:1-2002:09	
Real Equity	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	8.03	8.17	7.61	9.06	9.14	13.24
Equity						
Salary SA	34.96**	37.39**	14.16*	19.68*	25.27**	27.87**

Table 6.19 Equities and salaries cointegrating relationships: long-run coefficient estimates (4 lags, intercept, no linear trend) Monthly data 1963-2002. Dependent variables in bold

^=Not significant at 95% level

	1963:1-2002:12		1963:1-1979:12		1980:1-2002:09	
Real Equity	**α**	**β**	**α**	**β**	**α**	**β**
Real Salary SA	-12.71	4.43	1.96^	0.76^	-21.18	6.60
Equity						
Salary SA	2.11	1.28	1.26^	-1.83	-11.75	3.89

Table 6.20 Equities and salaries: Johansen cointegration tests (1 lag, intercept, no linear trend) Annual data 1850-2001. Dependent variables in bold

***=significant at 95% level, **=significant at 99% level**

	1850-2001		1900-2001		1932-2001	
Real Equity	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Wage (2)	23.11**	36.34**	17.81*	27.61**	15.59	24.50*
Equity						
Wage	37.66**	49.73**	23.76**	33.50**	20.02*	28.01**

Table 6.21 Equities and salaries: Johansen cointegration tests (1 lag, intercept, linear trend) Annual data 1850-2001. Dependent variables in bold

***=significant at 95% level, **=significant at 99% level**

	1850-2001		1900-2001		1932-2001	
Real Equity	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Wage (2)	16.33*	17.28*	13.77	14.96	9.08	9.39
Equity						
Wage	17.90*	27.73**	13.71	18.05*	12.94	13.16

Table 6.22 Equities and salaries cointegrating relationships: long-run coefficient estimates (1 lag, intercept, no linear trend) Annual data 1850-2001. Dependent variables in bold

^=Not significant at 95% level

	1850-2001		1900-2001		1932-2001	
Real Equity	**α**	**β**	**α**	**β**	**α**	**β**
Real Wage 2	-20.97	3.82	-21.22	3.81	-16.47^	2.70^
Equity						
Wage	-6.67	2.29	-6.23	2.05	-4.85	1.67

Table 6.23 Profits, dividends, wages and productivity cointegrating relationships: Johansen cointegration tests (4 lags, intercept, no linear trend) Quarterly data 1963-2002. Dependent variables in bold. EPS is earnings per share

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real profits	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Output per job	21.24**	31.13**	11.65	15.94	17.56**	25.09**
Real EPS						
Real Wages	21.17**	32.40**	23.78*	23.35*	10.98	15.80
Real dividend						
Output per job	21.82**	24.15*	10.91	18.27	13.69	15.93
Real salary SA						
Output per job	30.74**	42.43**	18.54*	26.02**	14.02	18.06
Real Wages						
Output per job	39.55**	48.85**	25.89**	32.93**	20.61**	31.40**

Table 6.24 Profits, dividends, wages and productivity cointegrating relationships: long-run coefficient estimates (4 lags, intercept, no linear trend) Quarterly data 1963-2002. Dependent variables in bold. EPS is earnings per share

^=Not significant at 95% level

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real profits	**α**	**β**	**α**	**β**	**α**	**β**
Output per job	14.33	0.51^	12.84	-0.65^	5.08	1.15
Real EPS						
Real Wages	-2.66	0.59	-0.47^	0.39^	2.25	0.15
Real dividend						
Output per job	4.79^	1.59^	7.04	-0.94	-13.94^	3.65
Real salary SA						
Output per job	0.20^	0.89	0.44^	0.82	0.95^	0.76
Real Wages						
Output per job	6.76	0.97	6.64	0.99	5.44	1.27

Table 6.25 Profits, dividends, wages and productivity cointegrating relationships: Johansen cointegration tests (4 lags, intercept, linear trend) Quarterly data 1963-2002. Dependent variables in bold. EPS is earnings per share

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real profits	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Output per job	9.90	11.81	4.53	8.68	10.96	11.25
Real EPS						
Real Wages	13.54	13.78	8.43	9.34	9.54	10.59
Real dividend						
Output per job	4.11	4.48	8.11	10.81	3.01	4.29
Real salary SA						
Output per job	14.86*	16.36*	12.29	14.37	4.38	5.10
Real Wages						
Output per job	17.38*	18.69*	15.94*	18.08*	14.16	14.17*

Table 6.26 Treasury Bills and salaries cointegrating relationships: Johansen cointegration tests (4 lags, intercept, no linear trend) Quarterly data 1963-2002. Dependent variables in bold

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Bills	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	22.48**	26.15**	10.83	17.23	21.16**	27.38**
Real Wages	22.05**	26.07**	13.53	20.04*	15.86*	20.18*
Bill						
Salary SA	18.22*	23.38*	10.96	16.46	17.49**	25.38**
Wages	17.91*	22.61*	11.07	14.66	7.92	11.41

Table 6.27 Treasury Bills and salaries cointegrating relationships: long-run coefficient estimates (4 lags, intercept, no linear trend) Quarterly data 1963-2002.
Dependent variables in bold

^=Not significant at 95% level

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Bills	**α**	**β**	**α**	**β**	**α**	**β**
Real Salary SA	-12.15^	4.34	9.27	-0.55^	0.83^	1.75
Real Wages	-3.66^	0.97	12.37^	-0.50^	-6.14^	1.29
Bill						
Salary SA	4.36	0.97	4.43	1.00	1.93	1.39
Wages	-1.84	0.92	-2.24	0.98	25.98^	-1.18^

Table 6.28 Treasury Bills and salaries cointegrating relationships: Johansen cointegration tests (4 lags, intercept, linear trend) Quarterly data 1963-2002. Dependent variables in bold

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Bills	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	5.43	6.12	6.84	8.67	8.39	12.33
Real Wages	6.58	6.83	6.75	7.36	7.37	8.14
Bill						
Salary SA	13.04	17.38*	7.58	13.08	17.45*	22.92**
Wages	12.68*	16.74*	7.81	11.25	5.09	6.65

Table 6.29 House prices and salaries cointegrating relationships: Johansen cointegration tests (4 lags, intercept, no linear trend) Quarterly data 1963-2002. Dependent variables in bold

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Property	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	19.21*	30.13**	17.84*	24.73**	20.62**	24.20*
Real Wages	16.98*	28.03**	19.56*	27.48**	6.60	10.96
Property						
Salary SA	13.29	21.20*	22.57**	28.49**	12.93	15.62
Wages	14.88	26.46**	26.91**	37.08**	24.36**	27.10**

Table 6.30 House prices and salaries cointegrating relationships: long-run coefficient estimates (4 lags, intercept, no linear trend) Quarterly data 1963-2002.
Dependent variables in bold

^=Not significant at 95% level

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Property	**α**	**β**	**α**	**β**	**α**	**β**
Real Salary SA	-1.13	1.46	-1.08	1.59	-2.41	1.76
Real Wages	-11.67	1.34	-9.70	1.37	177.5	-17.6
Property						
Salary SA	0.89	1.04	0.99	1.00	2.33	6.76
Wages	-7.92	1.14	-5.32	0.93	-11.87	1.41

Table 6.31 House prices and salaries cointegrating relationships: Johansen cointegration tests (4 lags, intercept, linear trend) Quarterly data 1963-2002. Dependent variables in bold

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real Property	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	11.97	12.74	17.78*	19.47*	10.24	10.41
Real Wages	11.08	11.11	18.63**	19.25*	4.36	4.54
Property						
Salary SA	13.23	15.54*	22.34**	25.64**	5.11	5.99
Wages	12.13	14.73	26.20**	32.20**	6.39	6.85

Table 6.32 US equities (unhedged) and salaries cointegrating relationships: Johansen cointegration tests (4 lags, intercept, no linear trend) Quarterly data 1963-2002. Dependent variables in bold

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real US Equities	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	20.03*	23.25*	8.55	14.24	18.92*	22.56*
Real Wages	22.57**	25.09**	9.76	16.78	19.42**	25.56**
US Equities						
Salary SA	14.82	18.58	9.94	16.64	9.30	14.83
Wages	20.12*	23.73*	11.48	19.82	14.69	18.89

Table 6.33 US equities (unhedged) and salaries cointegrating relationships: long-run coefficient estimates (4 lags, intercept, no linear trend) Quarterly data 1963-2002. Dependent variables in bold

^=Not significant at 95% level

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real US Equities	**α**	**β**	**α**	**β**	**α**	**β**
Real Salary SA	-12.58	3.96	4.24	-0.01^	-23.90	6.90
Real Wages	-25.51	2.72	-322.10^	29.79^	-45.73	4.61
US Equities						
Salary SA	2.68	1.03	1.63	0.49	-18.44	4.89
Wages	-3.32^	0.97	14.19^	-1.76^	-27.24	2.82

Table 6.34 US equities (unhedged) and salaries cointegrating relationships: Johansen cointegration tests (4 lags, intercept, linear trend) Quarterly data 1963-2002
Dependent variables in bold

***=significant at 95% level, **=significant at 99% level**

	1963:1-2002:3		1963:1-1979:4		1980:1-2002:3	
Real US Equities	**Max Eigenvalue**	**Trace**	**Max Eigenv.**	**Trace**	**Max Eigenv.**	**Trace**
Real Salary SA	5.37	5.41	6.29	8.99	7.12	9.31
Real Wages	9.74	9.77	7.11	9.04	17.08*	18.85*
US Equities						
Salary SA	6.74	9.69	8.95	11.02	5.74	11.16
Wages	7.18	10.14	8.57	12.38	6.34	10.36

Table 6.35 MULTIVARIATE MODELS: long-run coefficient estimates with standard errors in parenthesis and Johansen cointegration tests. Estimation with quarterly 1963-2002 data 4 lags, intercept, no linear trend. Variables in real log terms

***=significant at 95% level, **=significant at 99% level**

Model		**Max Eigenv.**	**Trace**
Model 1	RS = 2.13(1.72) + 0.46(0.31)RB	21.83**	28.38**
Model 1b	RS = 2.80(2.84) + 0.23(0.38)RBI	22.48**	26.15**
Model 2	RS = 2.96(0.26) + 0.22(0.04)RE	21.08**	24.77**
Model 3	RS = 4.50(0.54) - 0.57(0.17)RB + 0.48(0.08)RE	32.83**	47.13**
Model 4	RS = 3.32(0.94) - 0.45(0.17)RB + 0.40(0.1)RE + 0.19(0.13)RH	39.35**	66.15**
Model 5	RS = 4.03(0.86) + 0.27(0.21)RB - 0.53(0.20)RBI + 0.32(0.06)RE + 0.25(0.06)RH + 0.12(0.06)RUSE	43.62*	119.19**
Model 6	RS = 3.53(1.05) - 0.53(0.24)RB + 0.38(0.09)RE + 0.21(0.12)RH + 0.06(0.12)RUSE	40.22**	82.95*
Model 6b	RW = 10.25(1.56) - 0.67(0.36)RB + 0.29(0.14)RE + 0.25(0.17)RH + 0.32(0.17)RUSE	46.02**	82.14*

RS= Real Salary (AEI index, seasonally adjusted)
RW= Real Wages (National accounts)
RB= Real 10-year Government Bond Total Return Index
RE= Real FTSE All Share Total Return index
RH= Real Nationwide House Price Index
RBI= Real Bill Index
RUSE= Real S&P 500 Index (in pound sterling)

Note: with models 3-6b the hypothesis of at most one cointegrating relationship hypothesis cannot be rejected (no statistical support for more than one cointegrating relationship)

Table 6.36 MULTIVARIATE MODELS: error-correction models adjustment coefficients deriving from cointegrating relationships in Table 6.35 (standard errors in parenthesis). Variables in real log terms

	D(Real salary/wages)	D(Real Bond)	D(Real Equity)	D(Real house)	D(Real SP500)	D(Real Bill)
Model 1	-0.01(0.00)	-0.01(0.00)				
Model 1b	-0.01(0.00)					-0.00(0.00)
Model 2	-0.03(0.01)		-0.01(0.05)			
Model 3	-0.04(0.01)	-0.07(0.02)	0.03(0.07)			
Model 4	-0.05(0.01)	-0.10(0.03)	0.03(0.09)	0.01(0.02)		
Model 5	-0.12(0.02)	-0.04(0.05)	0.01(0.02)	-0.03(0.03)	0.04(0.16)	-0.07(0.01)
Model 6	-0.05(0.01)	-0.10(0.03)	0.05(0.09)	0.01(0.02)	0.16(0.08)	
Model 6b	-0.04(0.01)	-0.08(0.02)	0.05(0.06)	-0.01(0.01)	0.14(0.05)	

Table 6.37 MULTIVARIATE MODELS: salary error-correction model key statistics

	R-squared	R-squared (adjusted)	F test
Model 1	24.7%	20.5%	5.94
Model 1b	26.1%	22.1%	6.42
Model 2	17.5%	12.9%	3.83
Model 3	28.5%	22.4%	4.69
Model 4	29.5%	21.3%	3.59
Model 5	40.8%	29.8%	3.70
Model 6	31.6%	21.3%	3.07
Model 6b	30.0%	19.4%	2.84

Table 6.38 MULTIVARIATE MODELS: residual correlation and residual specification tests (model 6)

	REAL SALARYSA	REAL FTSE	REAL HOUSE	REAL BOND	REAL SP500POUND
REALSALARYSA	1.00	-0.04	0.24	0.06	0.01
REALFTSE	-0.04	1.00	-0.01	0.43	0.62
REALHOUSE	0.24	-0.01	1.00	0.04	-0.12
REALBOND	0.06	0.43	0.04	1.00	0.00
REALSP500POUND	0.01	0.62	-0.12	0.00	1.00

Test	Statistic	P-value
Autocorrelation (LM test)	23.29 (Lag 5)	0.5608
Heteroskedasticity (White Test)	793.45	0.0000
Normality (Jarque Bera)	17.50	0.0640

Table 6.39 MULTIVARIATE MODELS: long-run coefficient estimates with standard errors in parenthesis and Johansen cointegration tests. Estimation with quarterly 1963-2002 data 4 lags, intercept, no linear trend. Variables in nominal log terms

***=significant at 95% level, **=significant at 99% level**

Model		Max Eigenv.	Trace
Model 1	S = -2.02(0.29) + 0.92(0.05)B	26.39**	32.18**
Model 1b	S = -4.51(0.28) + 1.03(0.04)BI	18.22*	23.38*
Model 2	S = -1.67(0.44) + 0.78(0.06)E	17.43*	22.39*
Model 3	S = -1.93(0.70) + 0.81(0.52)B + 0.08(0.37)E	29.74**	47.60**
Model 4	S = -4.34(1.01) + 0.33(0.49)B - 0.18(0.16)E + 0.91(0.44)BI	34.45**	63.61**
Model 5	S = -2.47(0.62) + 0.98(0.51)B – 0.21(0.32)E + 0.33(0.23)RPI	35.54**	76.83**
Model 5b	(allowing for linear trend) S = -0.94 - 0.08(0.14)B + 0.28(0.09)E + 0.81(0.06)RPI	22.74	49.55*
Model 6	S = 1.01(1.09) - 2.88(1.01)B + 0.34(0.53)E - 0.12(0.61)H + 1.12(0.48)USE + 2.87(0.86)RPI	35.03	108.43*
Model 6b	W = 5.48(0.13) + 0.06(0.12)B + 0.01(0.06)E + 0.13(0.07)H + 0.11(0.05)USE + 0.85(0.10)RPI	45.85*	126.18**

S= Nominal Salary (AEI index, seasonally adjusted)
W= Nominal Wages (National accounts)
B= Nominal 10-year Government Bond total return Index
E= Nominal FTSE All Share total return index
H= Nominal Nationwide House Price index
BI= Nominal Bill Index
USE= Nominal S&P 500 Index in pound sterling
RPI= Retail Price Index

Note: with model 5 the hypothesis of at most one cointegrating relationship hypothesis is rejected at 99% confidence level only by Trace Statistic test, with model 6b the same hypothesis is rejected at 95% confidence level again only by Trace Statistic test

Table 6.40 MULTIVARIATE MODELS: error-correction models adjustment coefficients deriving from cointegrating relationships in Table 6.39 (standard errors in parenthesis). Variables in nominal log terms

	D(Salary/ Wages)	D(Bond)	D(Equity)	D(RPI)	D(SP500)	D(Bill)
Model 1	0.01(0.00)	0.04 (0.01)				
Model 1b	0.01(0.01)					0.01(0.00)
Model 2	0.01(0.00)		0.05(0.01)			
Model 3	0.01(0.00)	0.04(0.01)	0.01(0.03)			
Model 4	0.01(0.01)	-0.01(0.02)	-0.02(0.08)			0.01(0.00)
Model 5	0.01(0.00)	0.04(0.01)	-0.00(0.04)	0.00(0.00)		
Model 5b	-0.03(0.02)	0.13(0.05)	0.32(0.17)	0.04(0.01)		
Model 6	-0.00(0.00)	-0.03(0.01)	0.01(0.03)	0.00(0.00)	0.06(0.02)	
Model 6b	-0.06(0.02)	-0.22(0.05)	0.01(0.19)	0.05(0.02)	0.31(0.18)	

Table 6.41 MULTIVARIATE MODELS: nominal salary error-correction model key statistics

	R-squared	R-squared (adjusted)	F test
Model 1	43.9%	40.8%	14.20
Model 1b	46.7%	43.8%	15.88
Model 2	45.6%	42.6%	15.22
Model 3	45.7%	41.1%	9.90
Model 4	49.4%	43.5%	8.37
Model 5	55.0%	49.8%	10.50
Model 5b	56.9%	51.4%	10.55
Model 6	59.0%	51.4%	7.74
Model 6b	67.7%	61.7%	11.28

Table 6.42 MULTIVARIATE MODELS: residual correlation (model 6b).

	WAGES	FTSE	BOND	HOUSE	SP500 POUND	RPI
WAGES	1.00	-0.03	-0.05	0.08	-0.21	0.05
FTSE	-0.03	1.00	0.44	0.02	0.66	0.11
BOND	-0.05	0.44	1.00	-0.06	0.04	0.11
HOUSE	0.08	0.02	-0.06	1.00	-0.10	0.09
SP500POUND	-0.21	0.66	0.04	-0.10	1.00	0.07
RPI	0.05	0.11	0.11	0.09	0.07	1.00

Table 6.43 MULTIVARIATE MODELS: long-run coefficient estimates with standard errors in parenthesis and Johansen cointegration tests. Estimation with monthly 1920-2003 data, 12 lags. Variables in nominal or real log terms

***=significant at 95% level, **=significant at 99% level**

Model		Max Eigenv.	Trace
Model 1	RS = 2.00(0.67) + 0.04(0.15)RB with linear trend RS = 2.75 + 0.16(0.07)RB	25.26** 3.82	27.36** 4.58
Model 1b	RS = 1.97(0.27) + 0.20(0.05)RBI	7.31	11.09
Model 2	RS = 2.28(0.08) + 0.27(0.02)RE with linear trend RS = 2.21 + 0.27(0.01)RE	18.43* 17.31*	32.62** 17.33*
Model 3	S = 2.49(0.06) + 0.28(0.02)RPI with linear trend S = 2.60 + 0.28(0.02)RPI	20.22** 17.11*	27.88** 17.13*
Model 4	RS = 2.26(0.10) – 0.04(0.07)RB + 0.26(0.06)RE	55.53**	31.94**
Model 5	with linear trend S = -1.17 - 0.00(0.03)B + 0.24(0.03)E + 0.77(0.06)RPI	58.87**	28.38*
Model 6	with linear trend (1981-2003) S = -2.37 – 1.64(0.62)B + 2.61(0.86)IL	18.26	33.09*

S= Nominal Salary (historical average earnings indices)
B= Nominal Long-Term Government Bond Total Return Index
E= Nominal FTSE All Share total return index
BI= Nominal Bill Index
RPI= Retail Price Index
RS= Real Salary (historical average earnings indices)
RB= Real Long-Term Government Bond Total Return Index
RE= Real FTSE All Share Total Return index
RBI= Real Bill Index
IL= Index-Linked Total Return Index

Note: with models 4 the hypothesis of at most one cointegrating relationship hypothesis is rejected at 95% confidence level by Trace Statistic and Maximum Eigenvalue Test

Figure 1 National accounts wages and AEI in nominal and real terms (quarterly data)

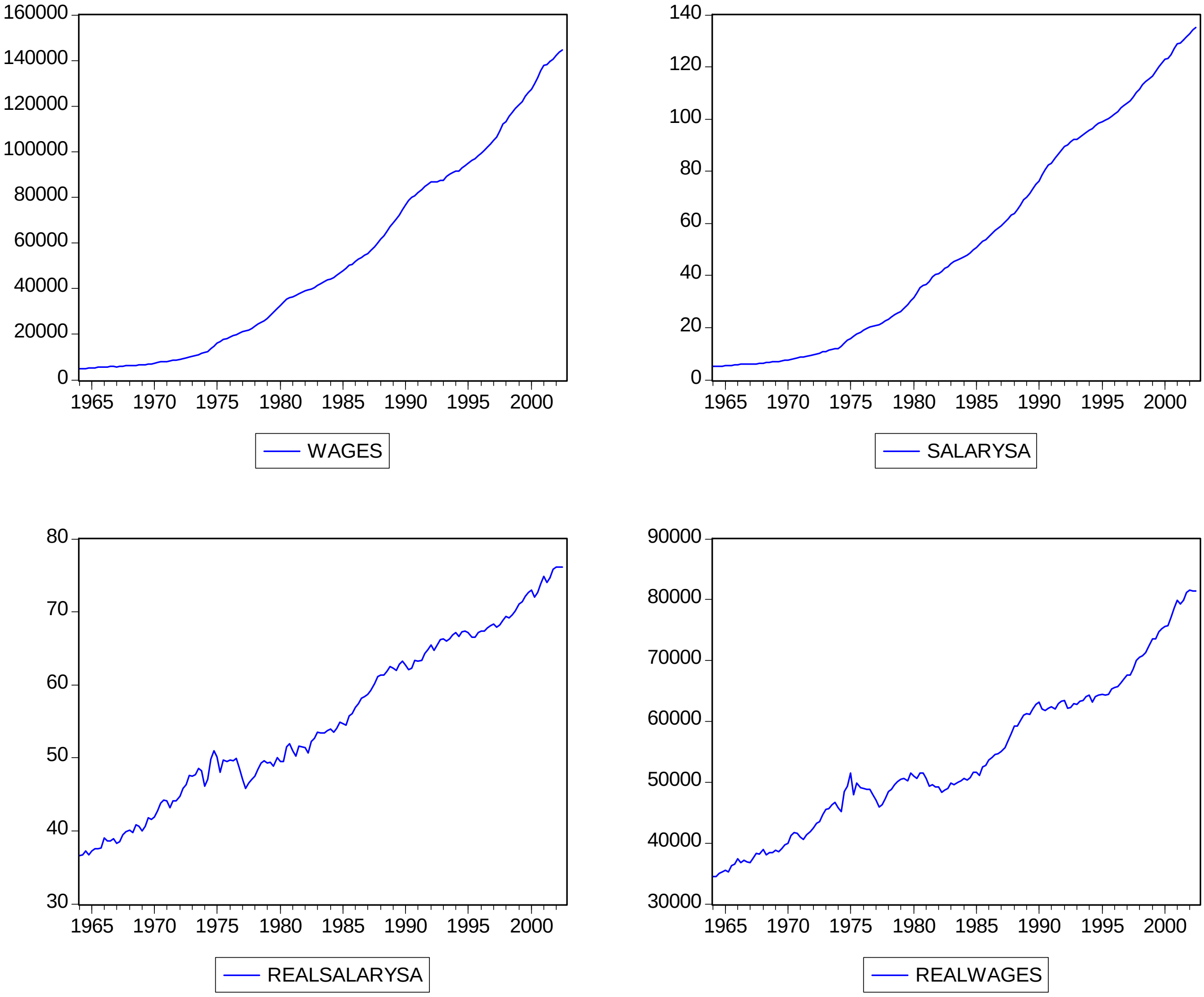


Figure 2 Nominal assets (quarterly data)

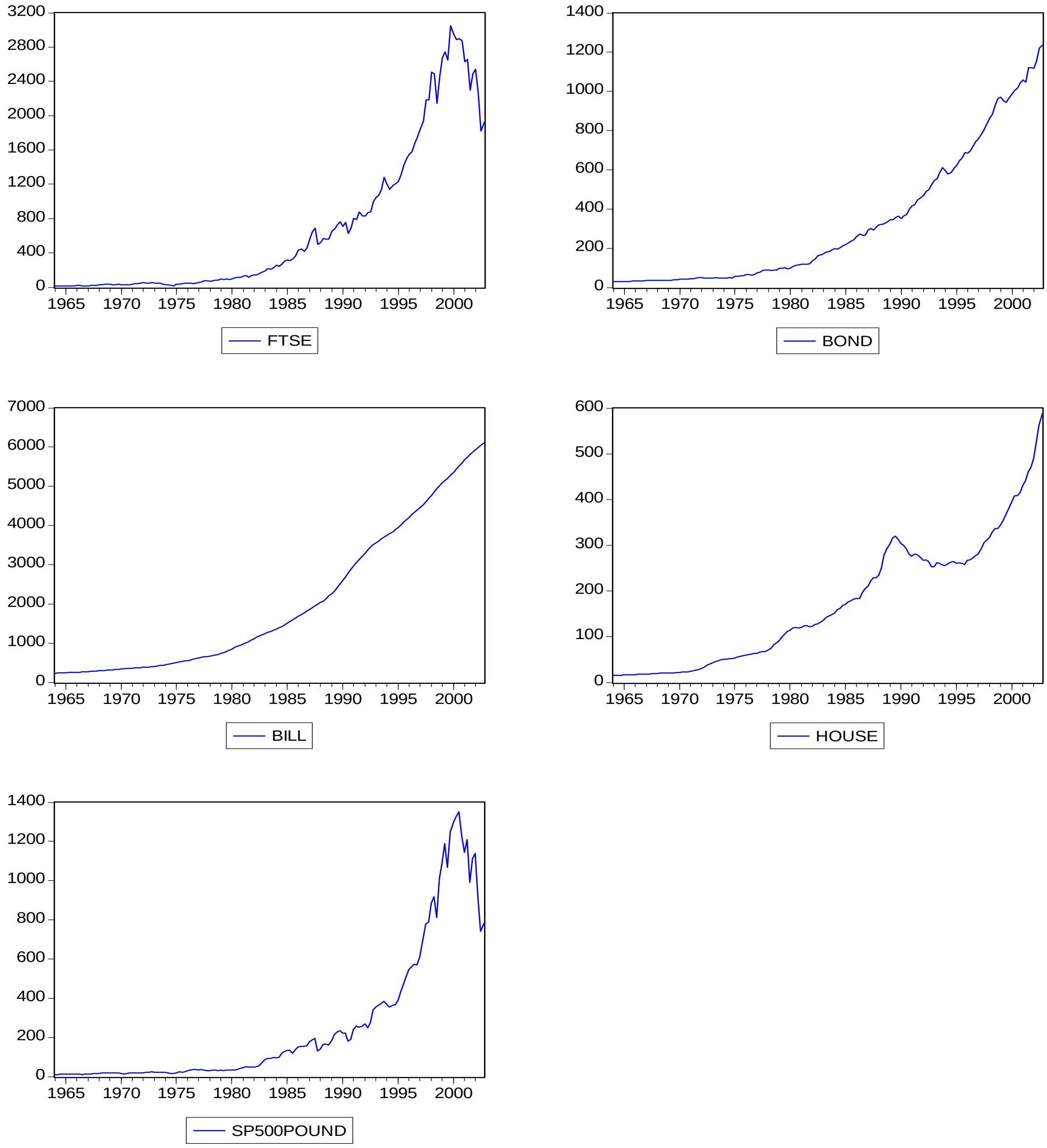


Figure 3 Real assets (quarterly data)

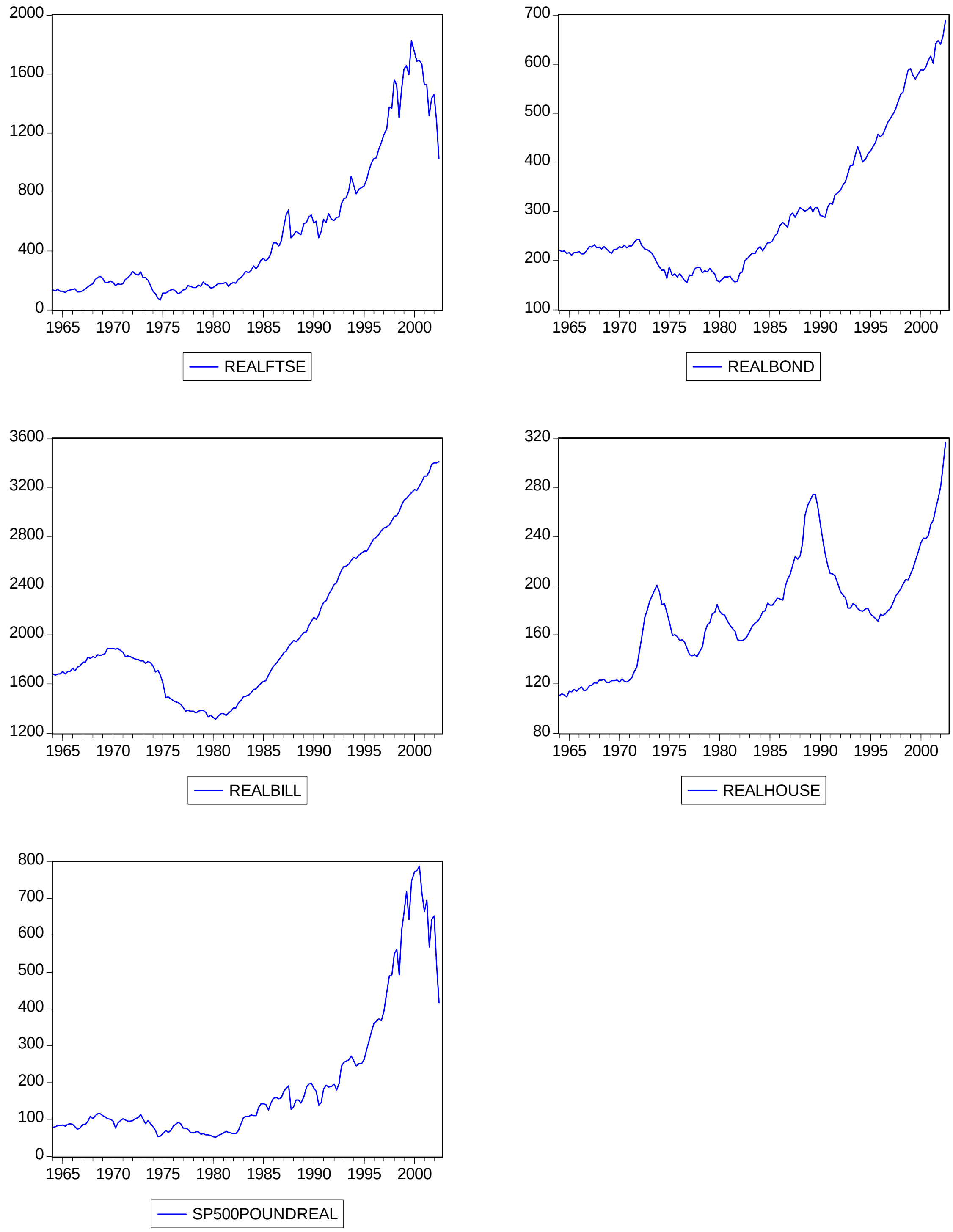


Figure 4 Nominal assets and wages (annual data set, 1850-1949)

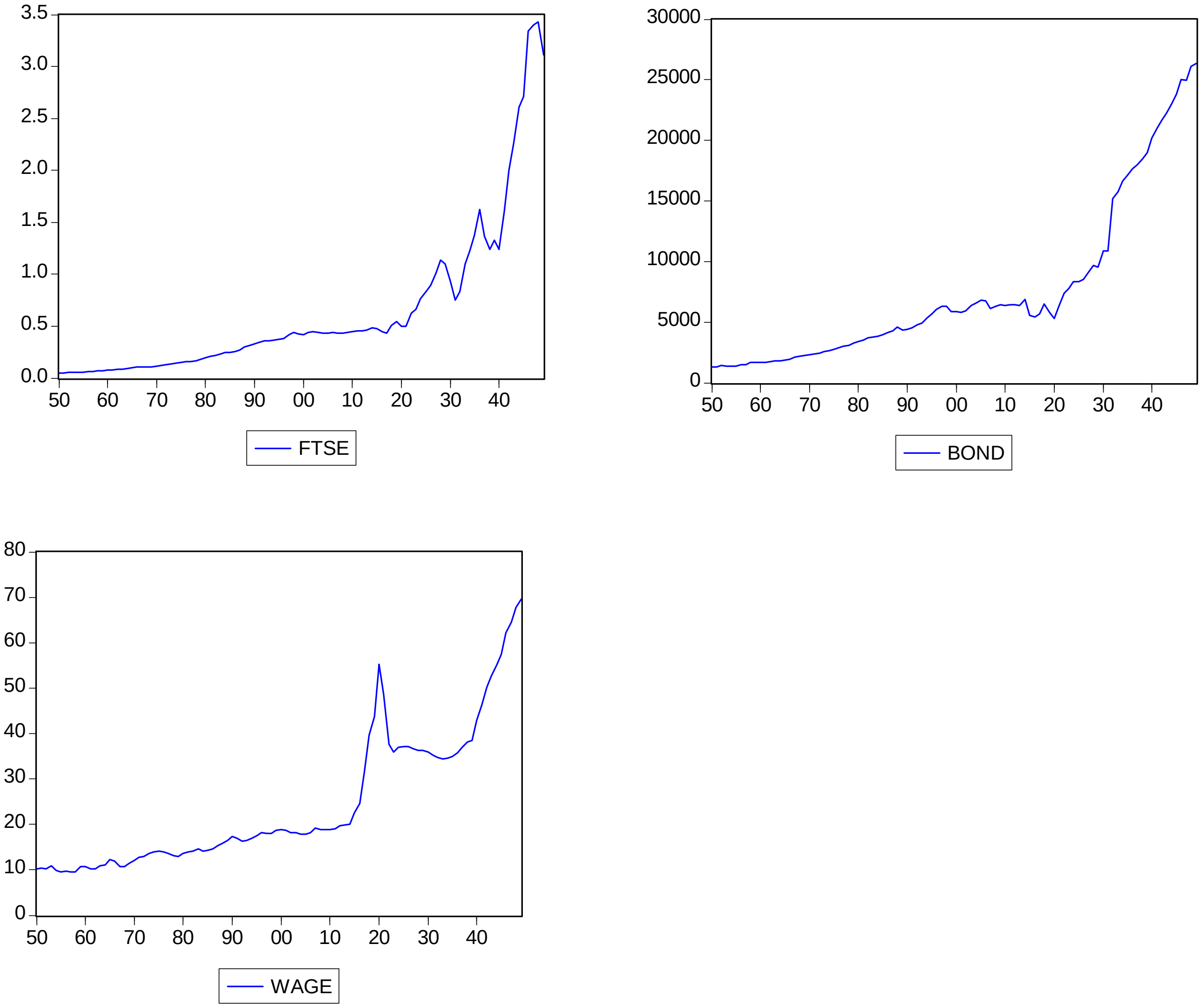


Figure 5 Nominal assets and wages (annual data set, 1950-2001)

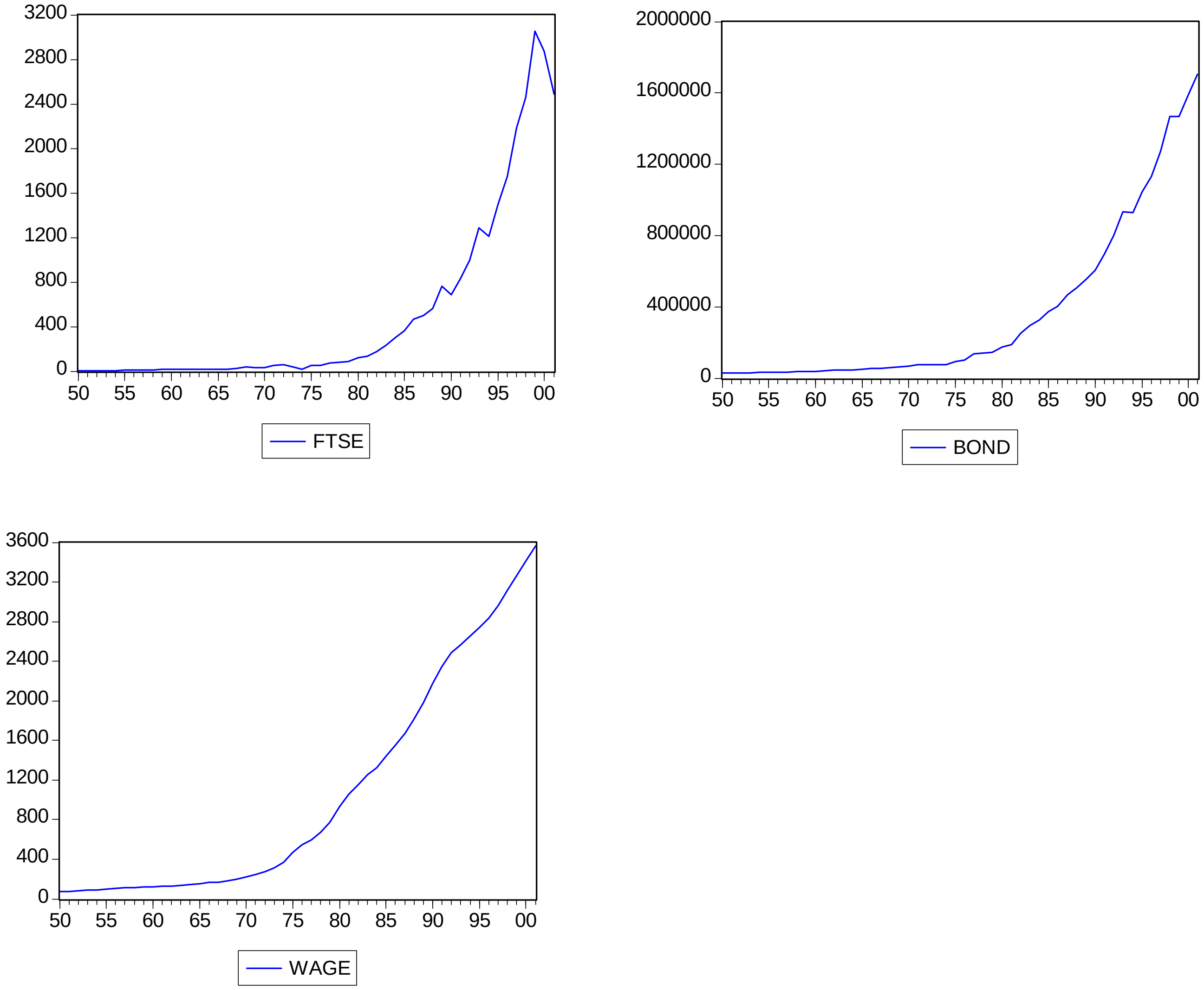


Figure 6 Real assets and real wages (annual data set, 1850-2001)

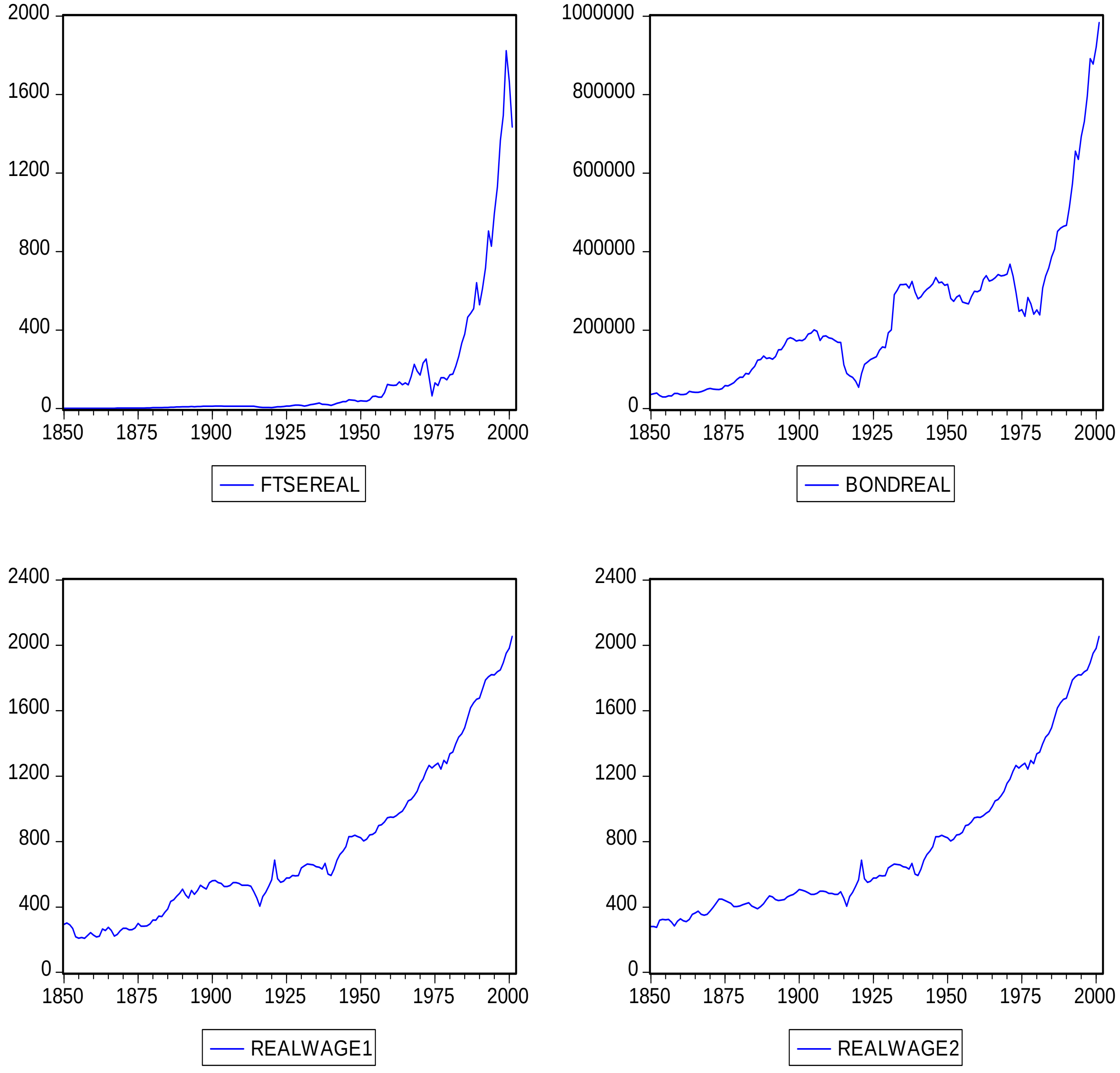


Figure 7 Cross correlograms of nominal asset returns with nominal salary growth (1963-2002)

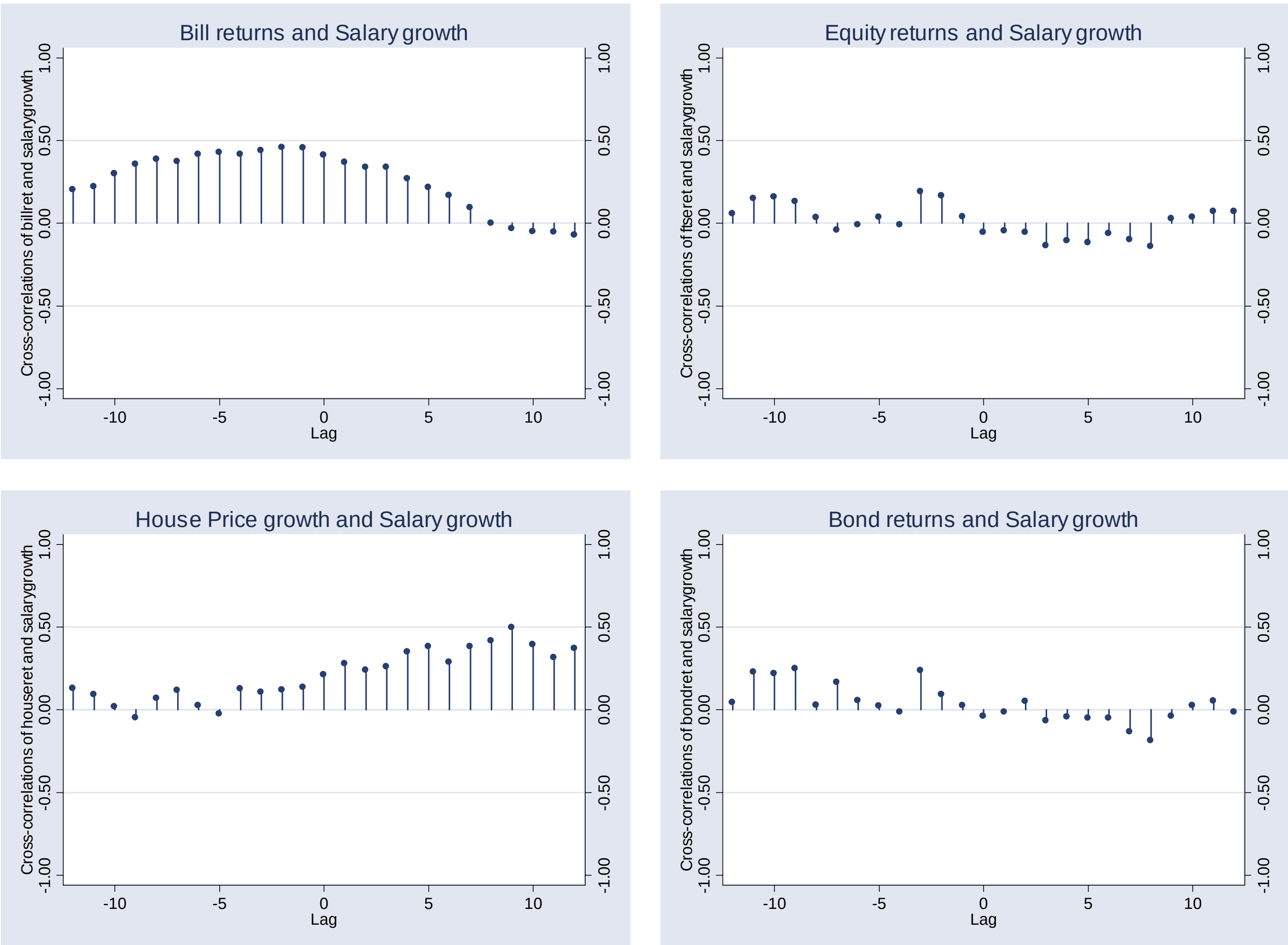


Figure 8 Cross correlograms of real asset returns and real salary growth (1963-2002)

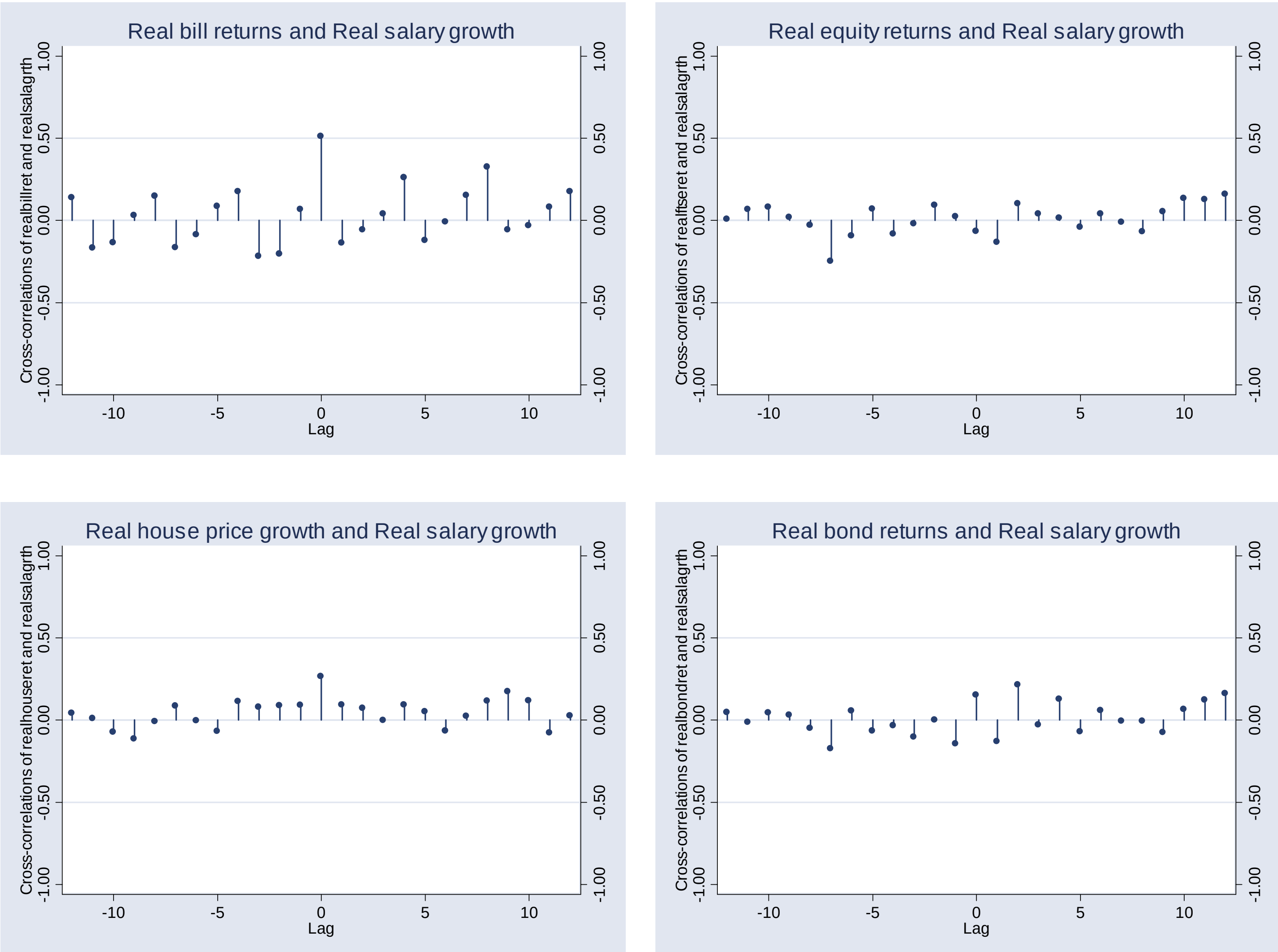


Figure 9 Scatter plots of nominal asset indices and nominal salary index (1963-2002)

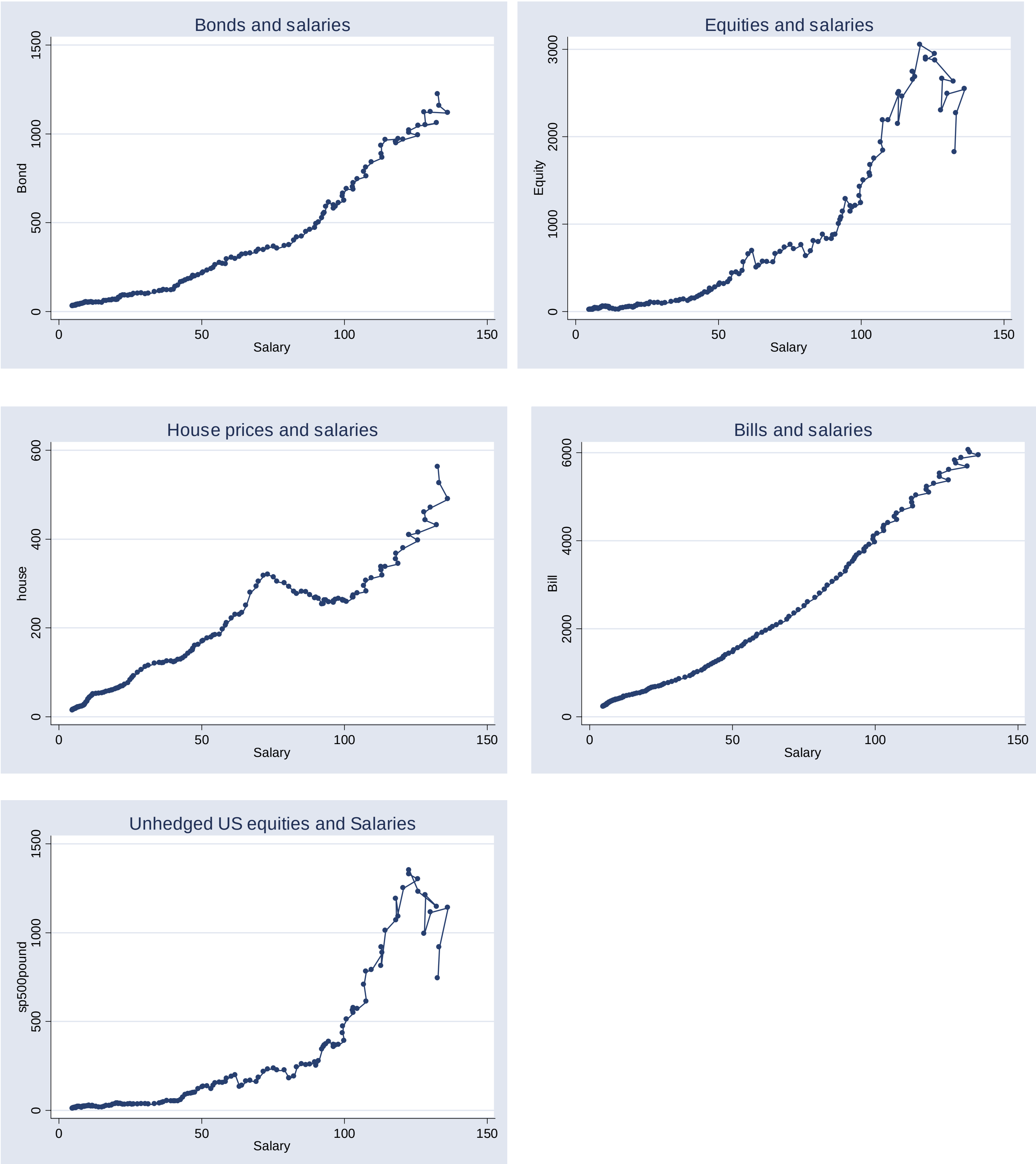


Figure 10 Scatter plots of real asset indices and real salary index (1963-2002)

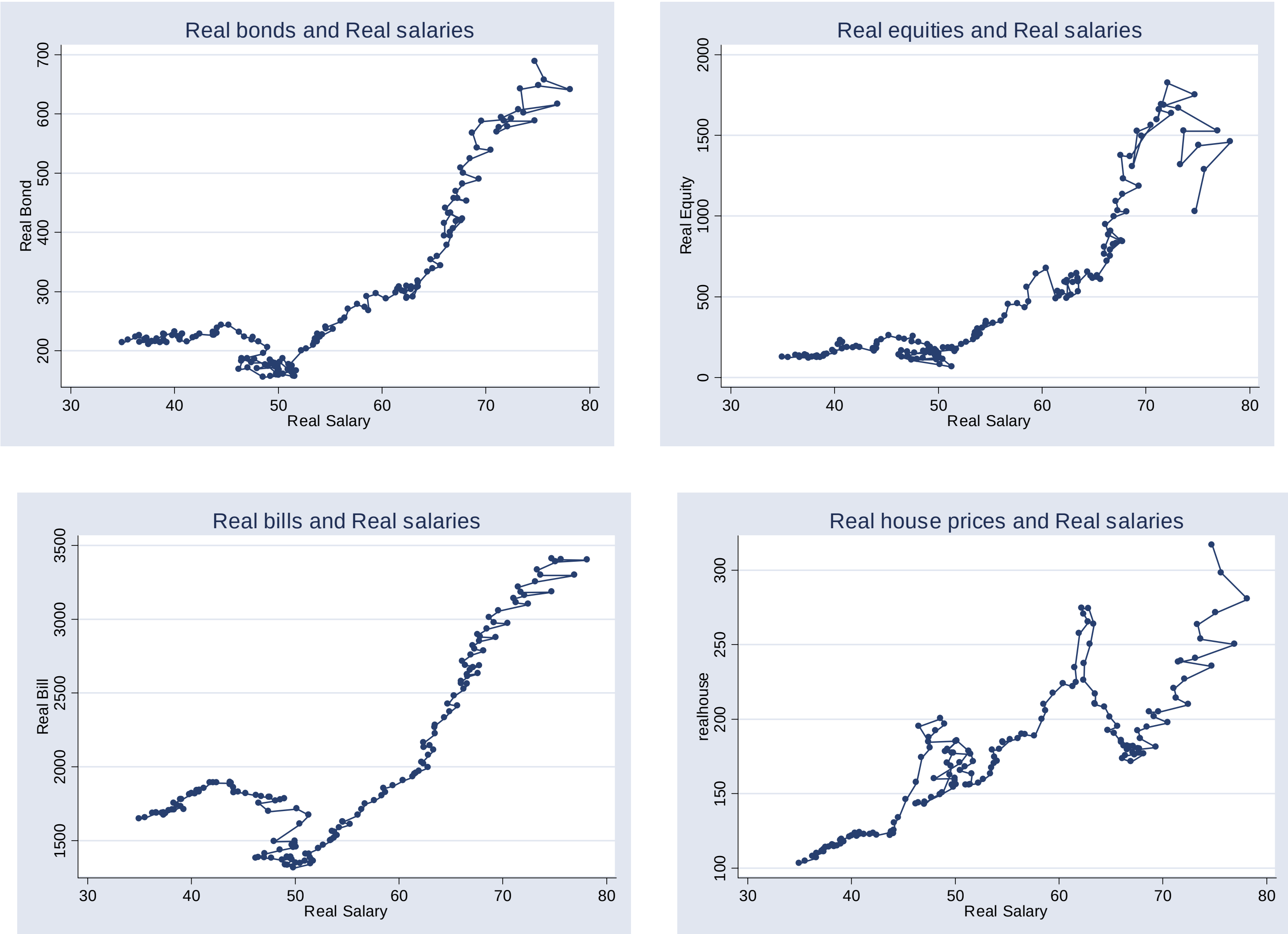
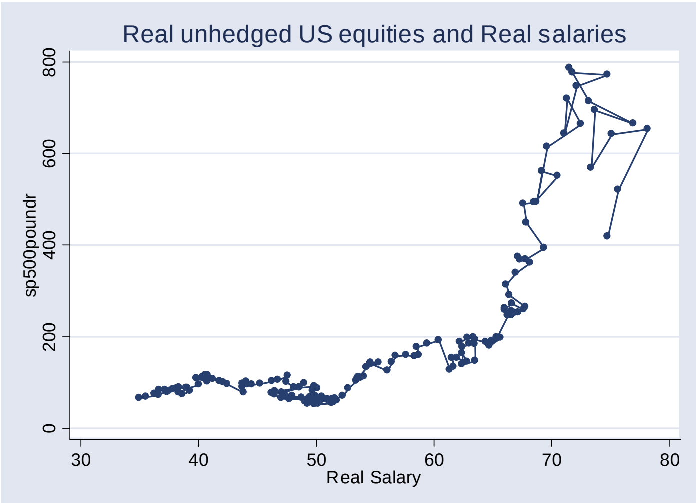
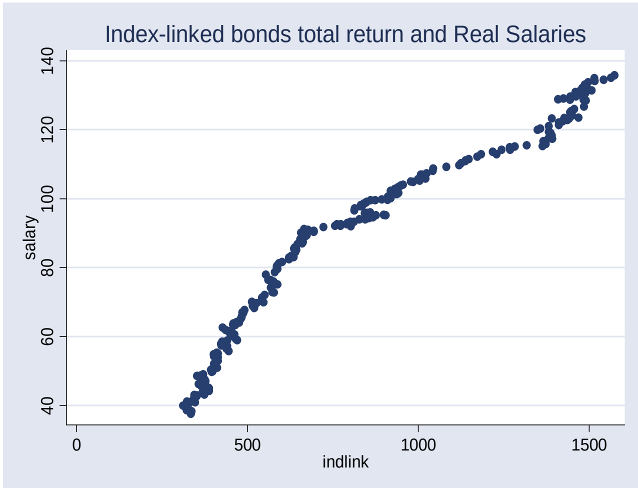


Figure 10 (continues) Scatter plots of real asset indices and real salary index (1963-2002)





Note: index-linked total return against real salaries is computed on monthly data since 1981

Figure 11 Scatter plots of nominal asset indices and nominal salary index (1850-2001)

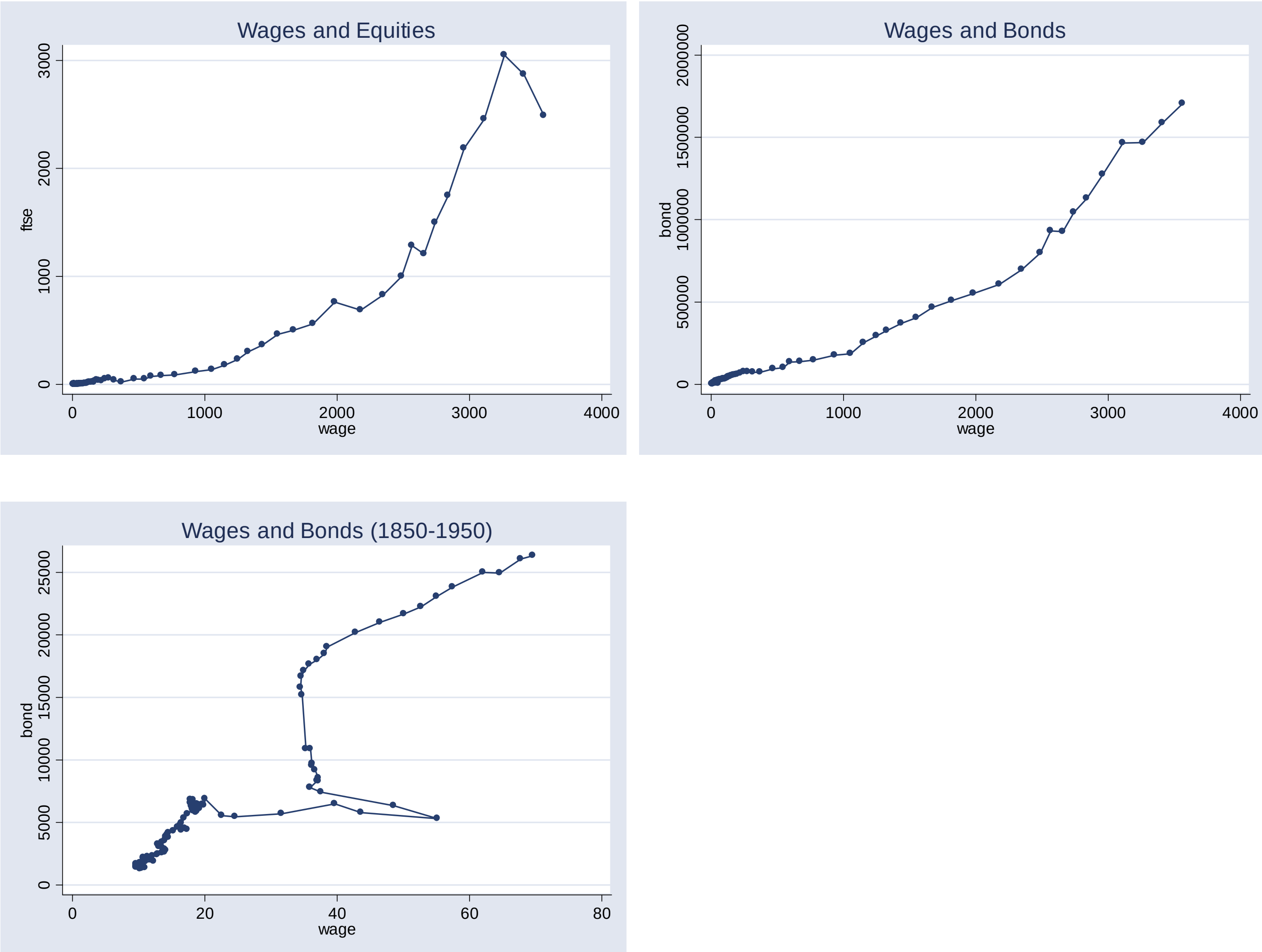


Figure 12 Scatter plots of real asset indices and real salary index (1850-2001)

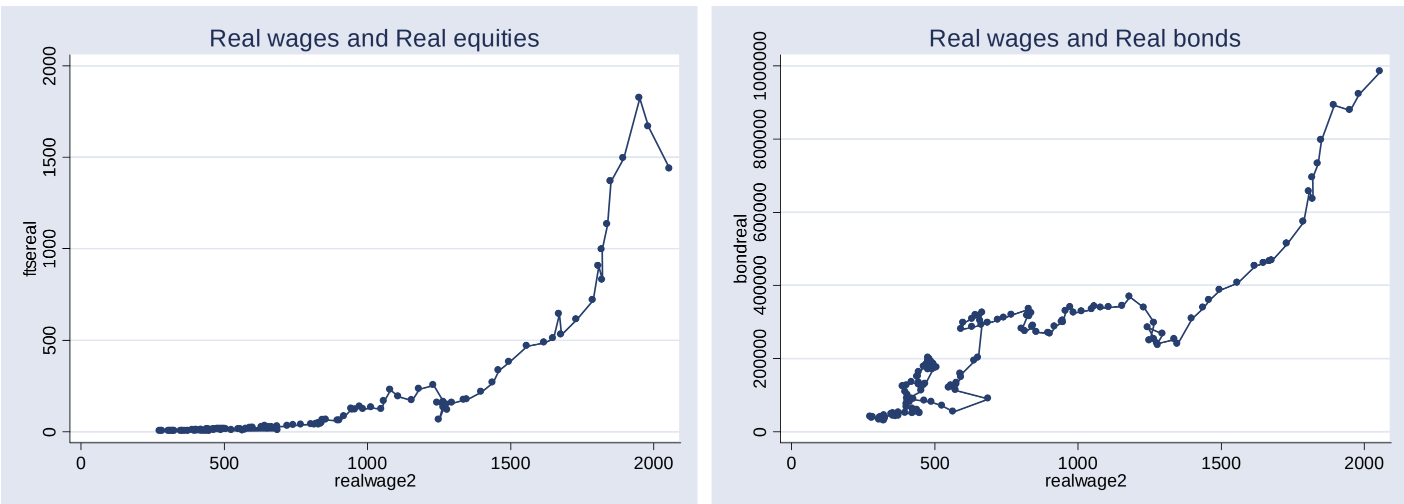
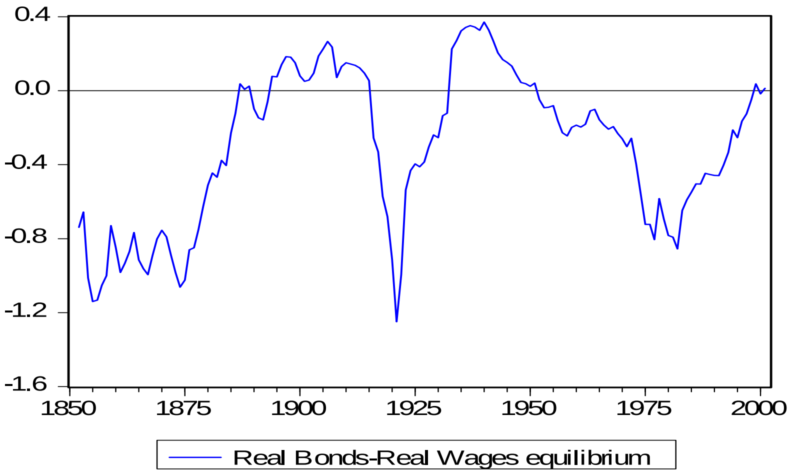
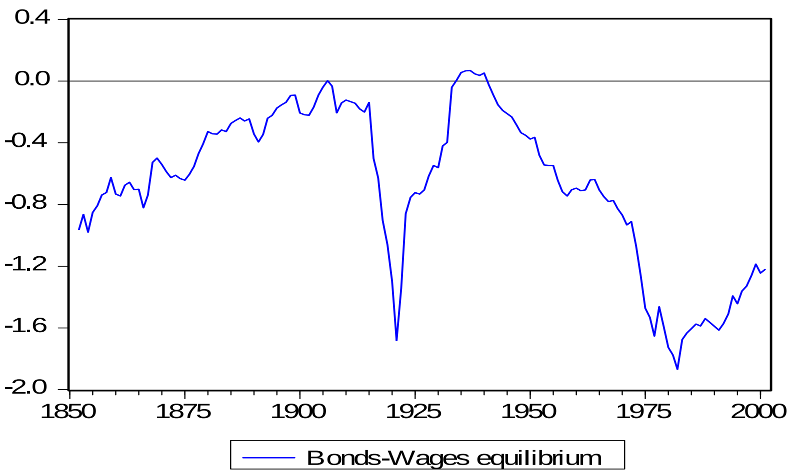


Figure 13 Long-term cointegrating equilibrium and equilibrium errors between real bonds and real wages



The estimated model (1850-2001 annual data, 1 lag, no linear trend, dummy for 1924 onwards) is
RB = 4.23(3.01) + 1.25(0.50)RW

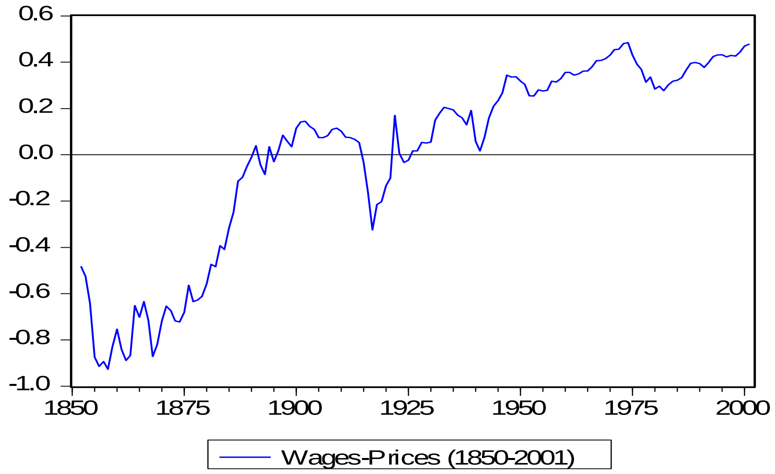
Figure 14 Long-term cointegrating equilibrium and equilibrium errors between nominal bonds and nominal wages



The estimated model (1850-2001 annual data, 1 lag, no linear trend, dummy for 1924 onwards) is
RB = 5.16(0.20) + 1.27(0.07)RW

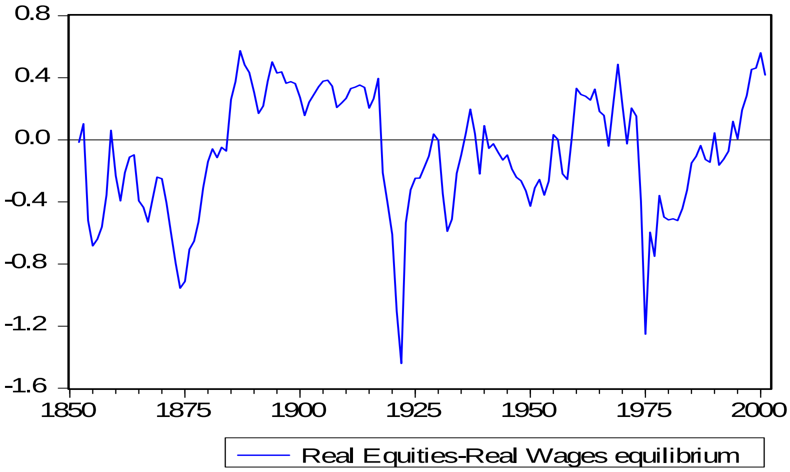
Figure 15 Long-term cointegrating equilibrium and equilibrium errors between nominal wages and prices

The estimated model (1963-2002 quarterly data, 4 lags, linear trend) is
W = 4.65 + 1.39(0.07)RPI



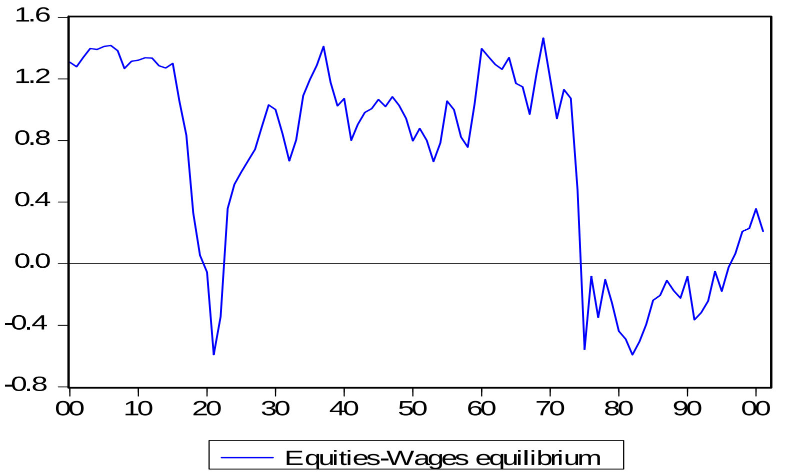
The estimated model (1850-2001, 1 lag, linear trend, dummy for 1924 onwards) is
W = 1.30 + 1.23(0.18)RPI

Figure 16 Long-term cointegrating equilibrium and equilibrium errors between real equities and real wages



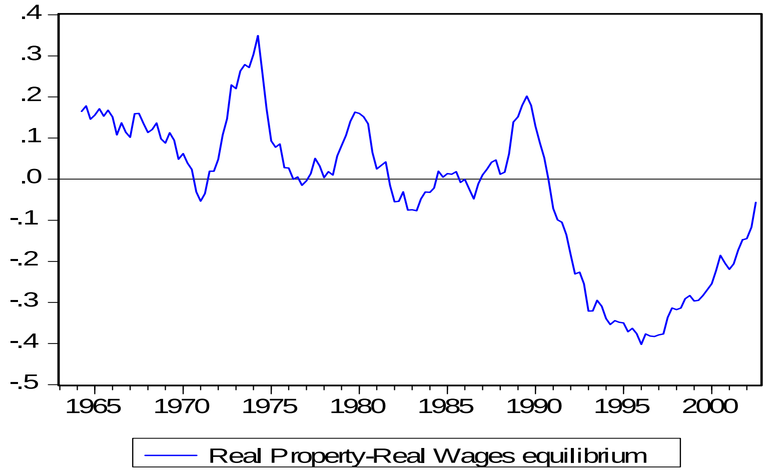
The estimated model (1850-2001 annual data, 1 lag, no linear trend, dummy for 1975 onwards) is RE = -18.83(1.79) + 3.40(0.28)RW

Figure 17 Long-term cointegrating equilibrium and equilibrium errors between nominal equities and nominal wages



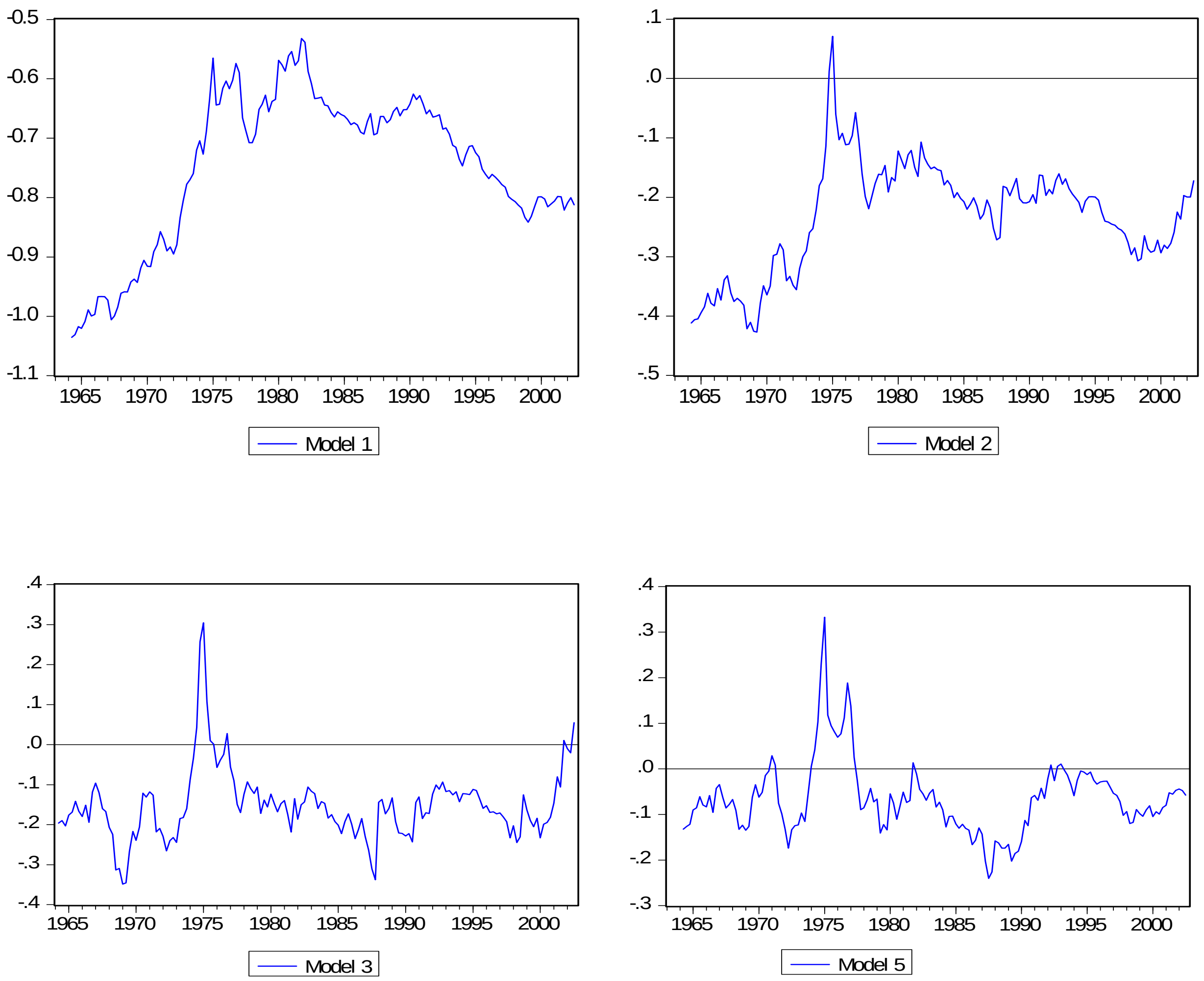
The estimated model (1850-2001 annual data, 1 lag, no linear trend, dummy for 1975 onwards) is E = 7.74(1.25) + 1.90(0.31)W

Figure 18 Long-term cointegrating equilibrium and equilibrium errors between real property and real wages



The estimated model (1963-2002 quarterly data, 4 lags, no linear trend, dummy for 1989 onwards) is RH = -1.44(0.88) + 1.66(0.23)RS

Figure 19 Multivariate long-term cointegrating equilibriums (1963-2002) and equilibrium errors (Table 6.35, variables in real log terms)



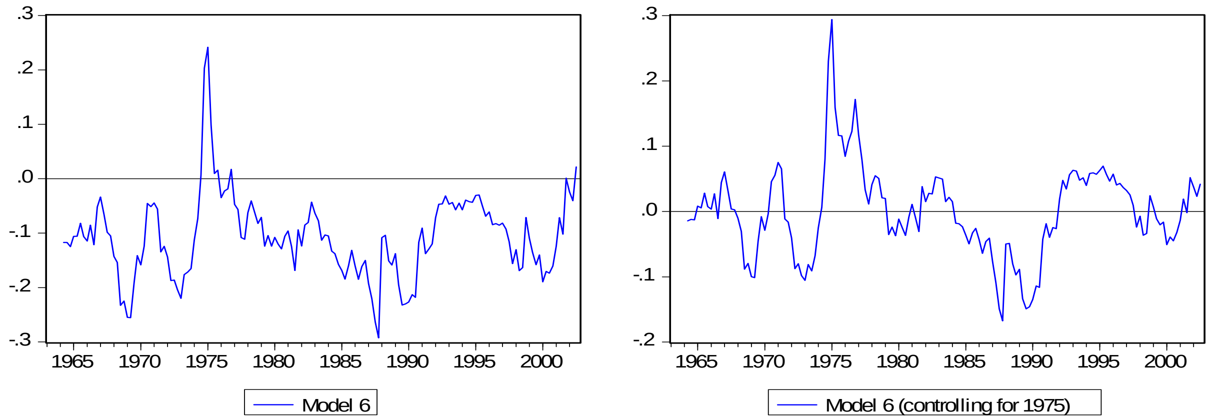
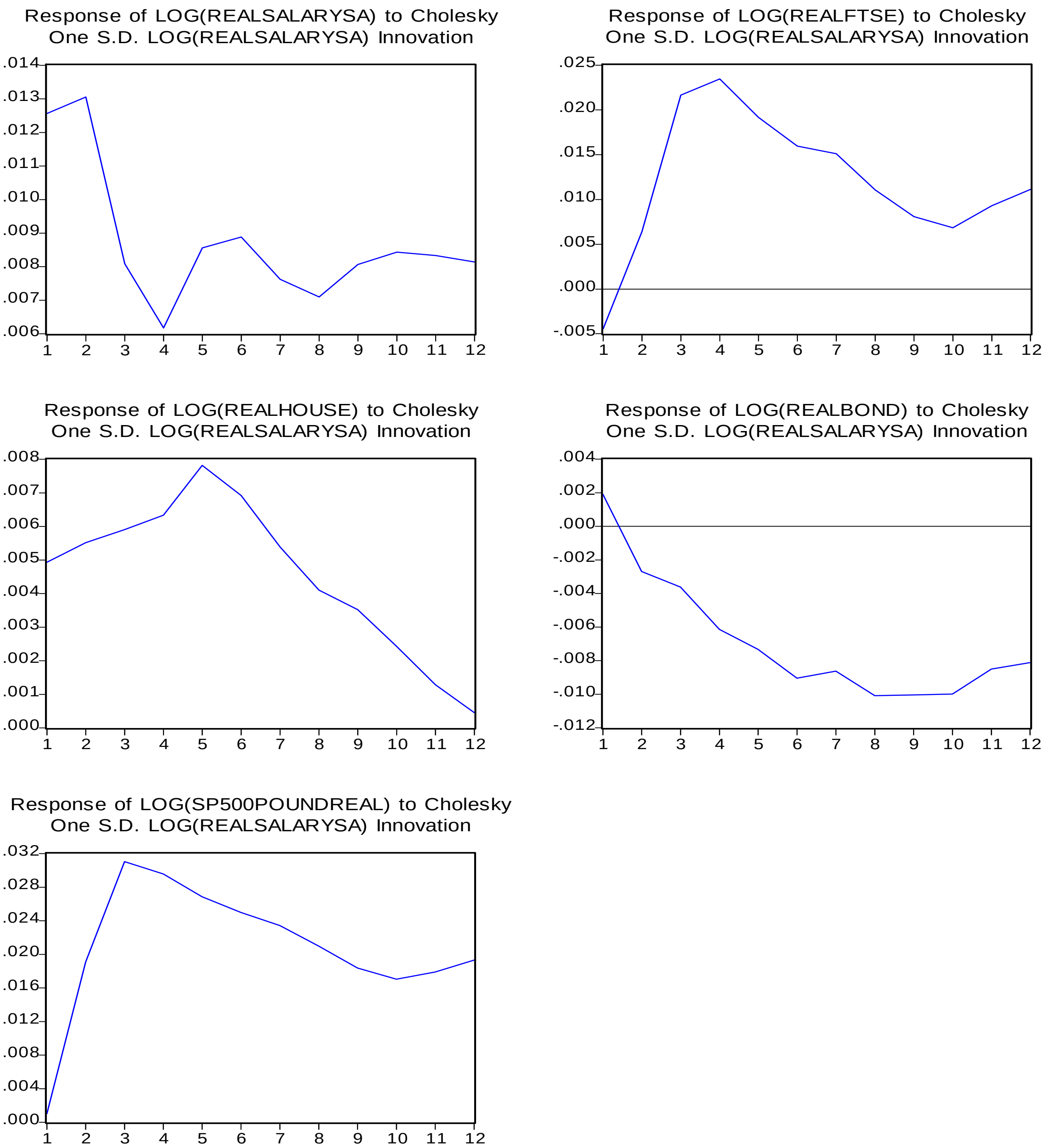


Figure 20 Impulse-response functions from Cholesky decomposition arising from real salary innovation (Table 6.35, model 6, variables in real log terms)



Note: Cholesky ordering is: REAL SALARY-REAL FTSE-REAL HOUSE-REAL BOND-SP500POUNDREAL

Figure 21 Multivariate long-term cointegrating equilibriums (1963-2002) and equilibrium errors (Table 6.39, variables in nominal log terms)

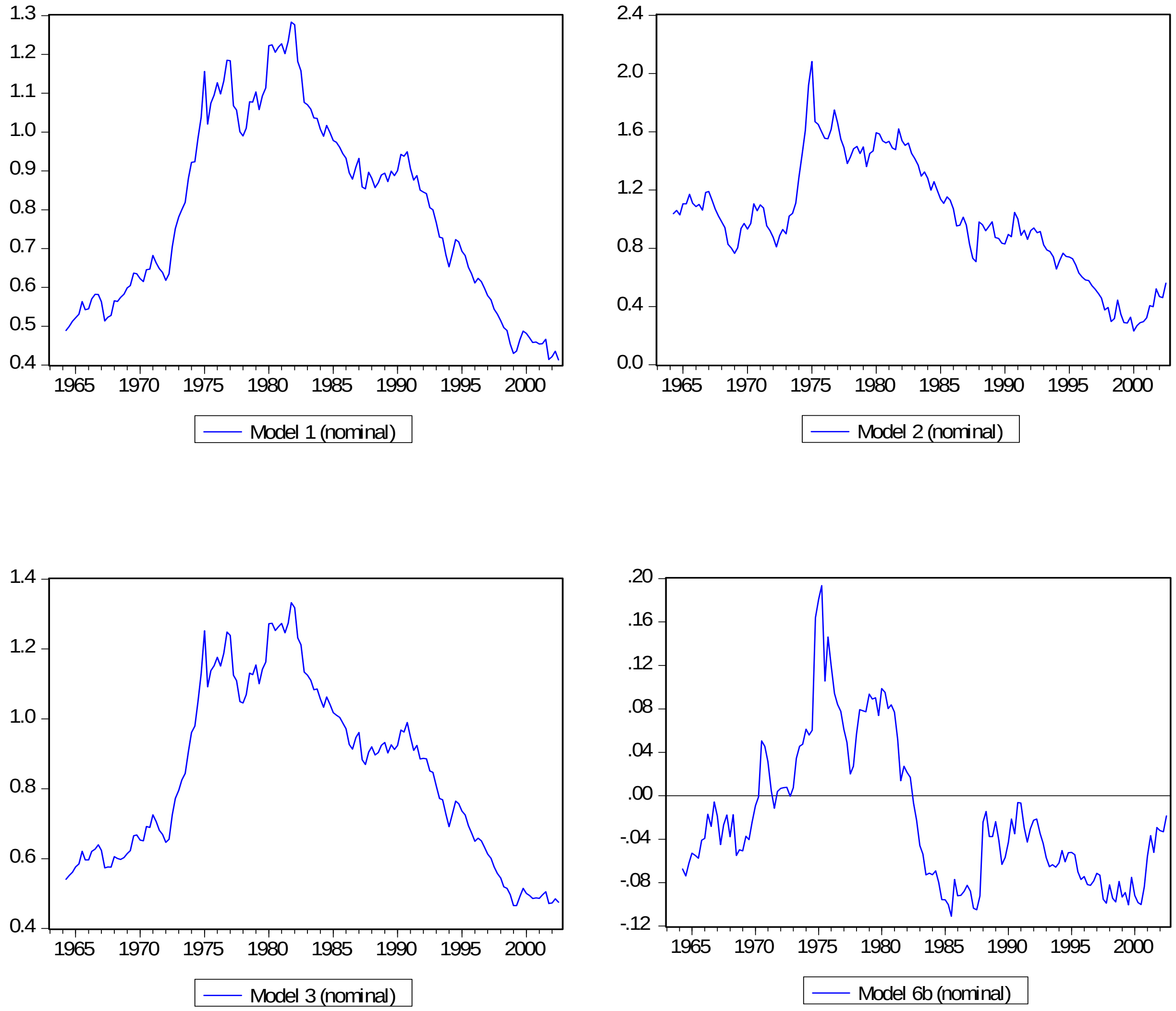
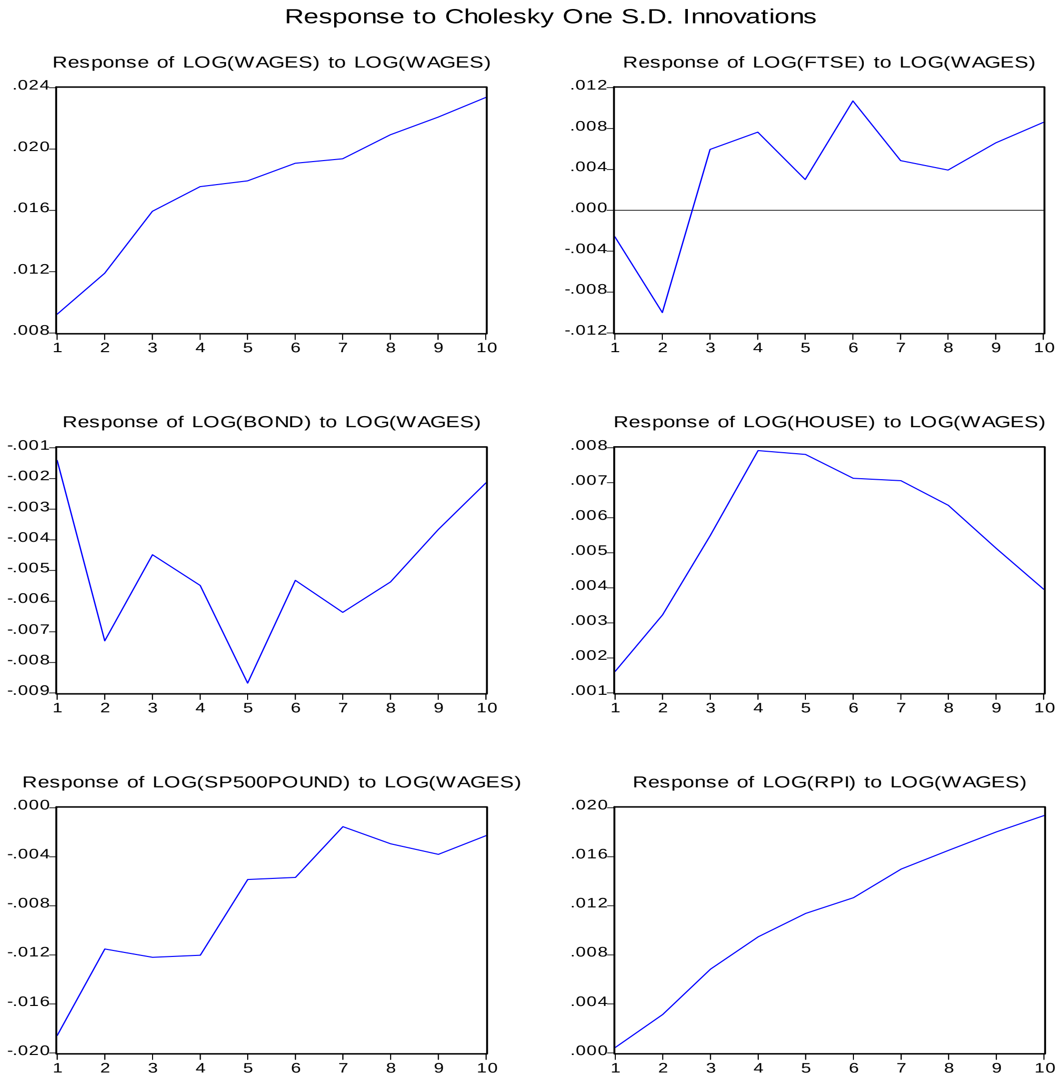


Figure 22 Impulse-response functions from Cholesky decomposition arising from nominal salary innovation (Table 6.39, model 6b, variables in nominal log terms)



Note: Cholesky ordering is: WAGES-FTSE-BOND-HOUSE-SP500POUND-RPI

Figure 23 Multivariate long-term cointegrating equilibriums (1920-2003) and equilibrium errors (Table 6.43, variables in nominal or real log terms)

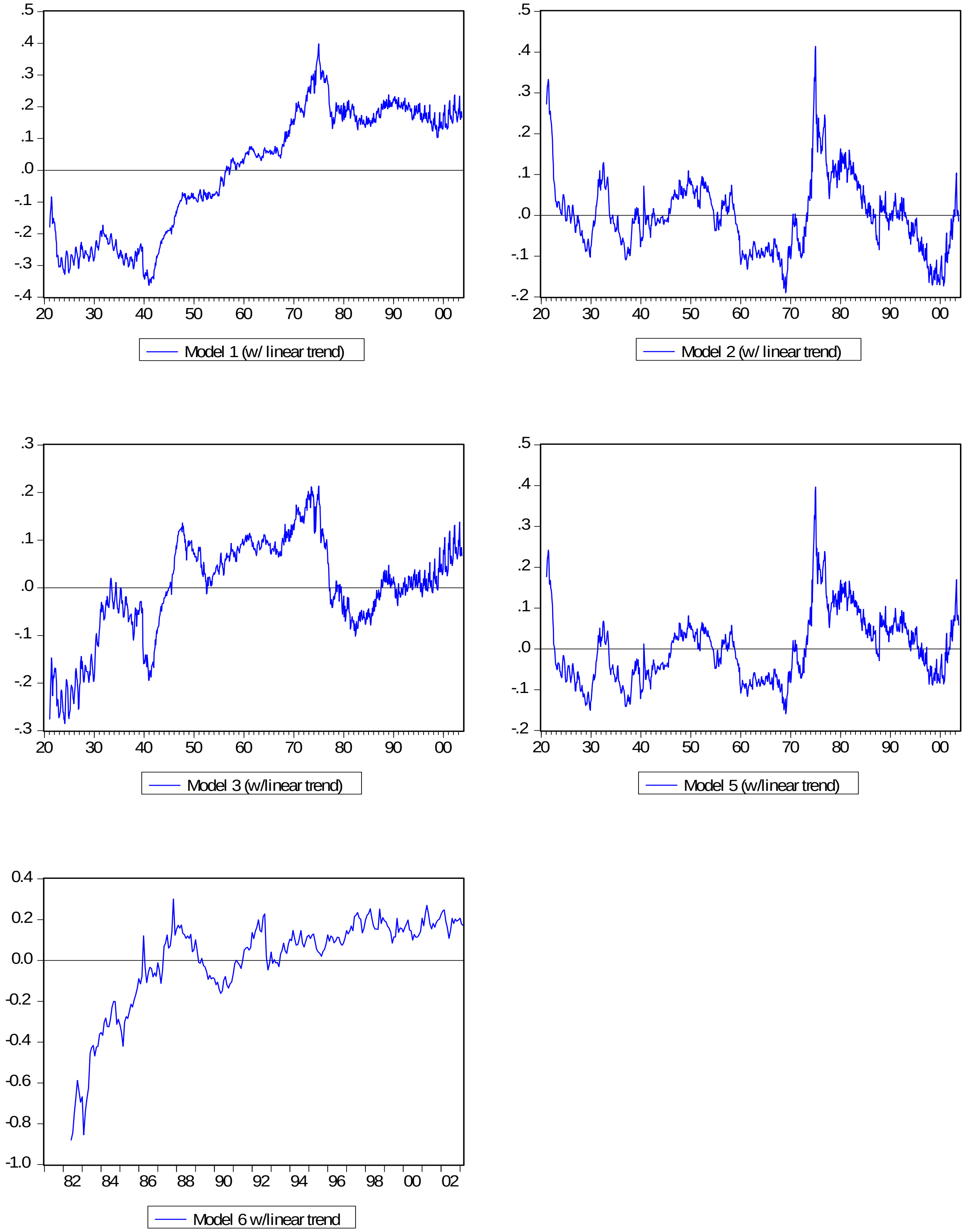
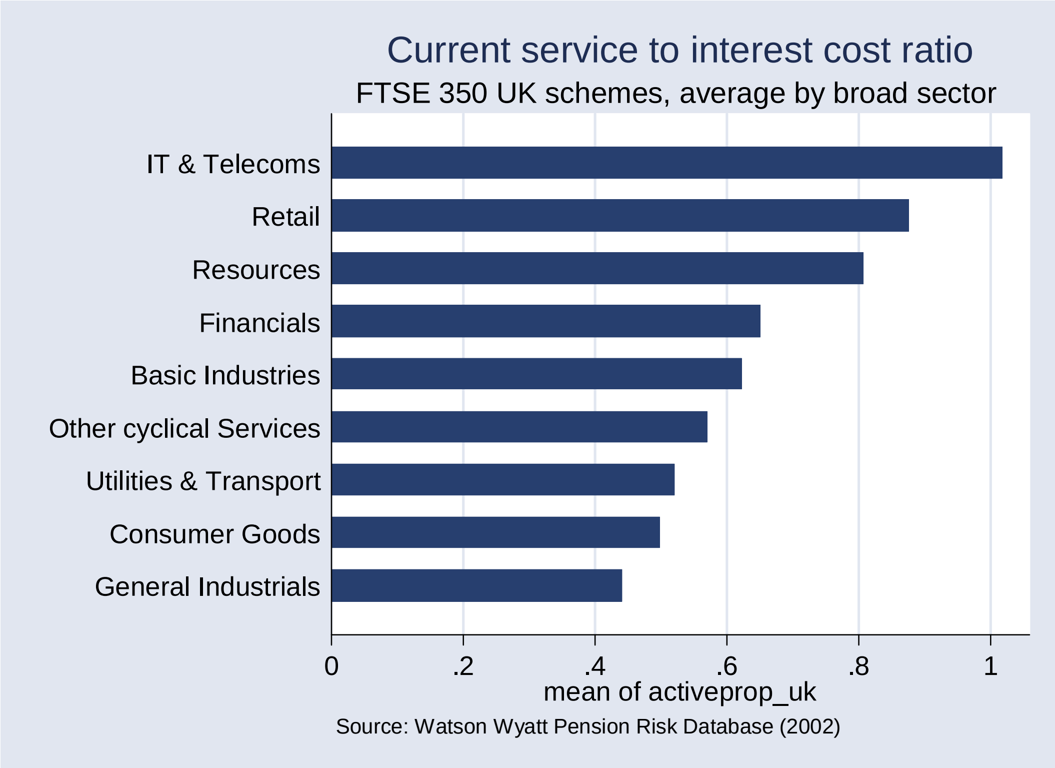
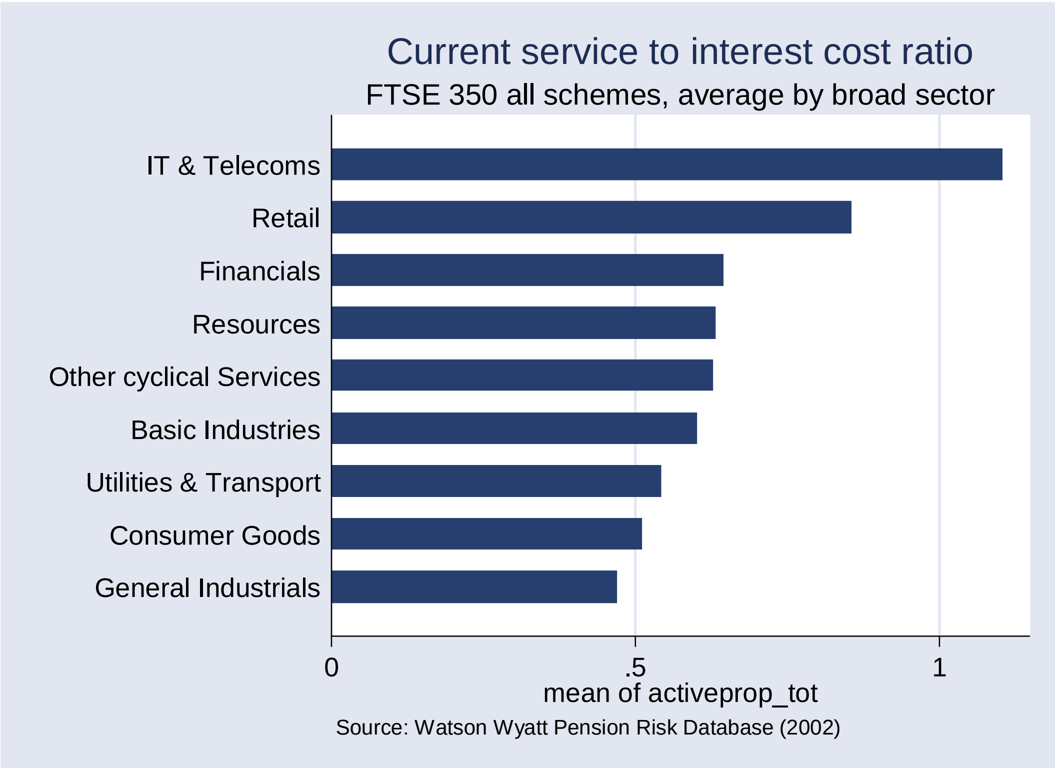


Figure 24 Average maturity of UK defined benefit pension schemes (current service cost over interest on pension liabilities reported in FRS 17 disclosures)


Current service to interest cost ratio
FTSE 350 UK schemes, average by broad sector
IT & Telecoms
Retail
Resources
Financials
Basic Industries
Other cyclical Services
Utilities & Transport
Consumer Goods
General Industrials
0
.2
.4
.6
.8
1
mean of activeprop_uk
Source: Watson Wyatt Pension Risk Database (2002)


Current service to interest cost ratio
FTSE 350 all schemes, average by broad sector
IT & Telecoms
Retail
Financials
Resources
Other cyclical Services
Basic Industries
Utilities & Transport
Consumer Goods
General Industrials
0
.5
1
mean of activeprop_tot
Source: Watson Wyatt Pension Risk Database (2002)