Unemployment Risk and the Conditional Ex-Ante Equity Premium

Klaus E. Buhr

Department of Accounting and Finance, Unitec Institute of Technology, New Zealand

Nonthipoth Buranavityawut*

Department of Accounting and Finance, Unitec Institute of Technology, New Zealand

Mark C. Freeman

Bradford University School of Management, United Kingdom

*Corresponding author

Unitec Institute of Technology

Department of Accounting and Finance

Private Bag 92025, Victoria St West

Auckland, New Zealand

Phone: +64 9 8154321 ext 7098

Fax: +64 9 8154563

E-mail: nburanavityawut@unitec.ac.nz

Abstract

We estimate unemployment risk's contribution to the ex-ante equity premium. This is achieved by taking a closed form conditional asset pricing model that incorporates unemployment risk, and calibrating it using Markov-GARCH and Dynamic Conditional Correlation methods. While potential job loss is not, on average, a major contributor to the equity premium, this relationship is highly time-varying and often of economic significance. Our results strongly suggest that there was a structural downward shift in this effect in the mid 1940s in the UK. Given the potential severity of the current economic crisis, there is the possibility that this will be reversed, causing larger than expected declines in equity market values.

Unemployment Risk and the Conditional Ex-Ante Equity Premium

1. Introduction

In this paper we examine whether unemployment risk is a significant contributing factor to the ex-ante equity premium and whether this effect is time-varying. This is achieved by taking a closed-form conditional asset pricing model that explicitly allows for job loss and calibrating it using Markov-GARCH and Dynamic Conditional Correlation methods. Using monthly data for the UK for the period from 1888-2000, we construct 60-month rolling average estimates of the prevailing equity premium. We show that, on average, job loss does not appear to be an important factor influencing the equity premium. However, particularly in the first half of the century, the effects were often of economic significance and there appears to have been a structural downward shift in this effect after the mid 1940s. There is also some evidence of a structural rise in the 1970 followed by another structural decline in the 1980s. If the forecast sharp rise in unemployment during the current credit crisis materialises, then we deduce that this would be associated with an economically significant increase in the equity premium. In turn, this would result in us predicting more severe equity market declines than would be inferred from many alternative asset pricing model.

The motivation for this paper is that it is well established that the observed average excess returns to broad stock market indices over Treasury bill rates has been too high when compared to theoretical estimates (for example, Mehra and Prescott, 1985, 2003; Mehra, 2008). One of the potential explanations for this "equity premium puzzle" is that the standard model assumes that the market is complete and so does not take into account the effect of idiosyncratic risk. By incorporating uninsurable risk, the theoretical equity premium normally increases and becomes closer to the observed average excess return.

There has been much debate over which source of uninsurable risk has the greatest impact on asset prices. Suggestions have included labour income risk (for example, Heaton and Lucas, 1996), housing assets (Fratantoni, 2001) and proprietary income (Heaton and Lucas, 2000). Amongst the most important sources of labour income risk is unemployment, and this has, therefore, been linked with a potential resolution to the equity premium puzzle. Matthews and Benjamin (1992) and Constantinides (2002) suggest that job loss can be a major uninsurable income shock of households. Not only does job loss have persistent implications on household income, but unemployment compensation is also inadequate. Benito (2006) finds that job loss risk represents a major source of income uncertainty for most households. Seligman and Wenger (2006) find that unemployment risk can amplify the stock market investment loss of stockholding workers. Workers can lose both from being unemployed and from the stock market decline if unemployment occurs at the time of a falling equity market. In addition, during times of high unemployment risk, the impact of unemployment risk is not only on individuals who become unemployed, but also on those who remain employed. This is because high unemployment risk signals a lower sense of financial security, which raises concerns among those employed that their employment status is in jeopardy.

While a clear theoretical link has been established between unemployment risk and the equity premium puzzle, to our knowledge, no previous detailed empirical work has examined this relationship. We conjecture that there are two problems that have prevented previous researchers from filling this gap, both of which we largely overcome in this paper. First, deriving closed-form theoretical asset pricing models in the presence of heterogeneous agents is extremely complex. Here, we use the extension of Constantinides and Duffie (1996) by Freeman (2002), which explicitly incorporates unemployment-type events. While this is a stylised model, like Constantinides and Duffie (1996), it does make the estimation of the equity premium reasonably straightforward even in the presence of uninsurable risk. Second, we know of no reliable source of monthly unemployment data for the US prior to 1929, making it impossible to calibrate such a model on US data over the sample period of Mehra and Prescott (2003). We overcome this problem by reverting to UK data, which is available on a monthly basis back to 1888. A further advantage of using UK data is that unemployment is a more persistent phenomenon in the UK than the US.

As we calibrate our conditional asset pricing model using Markov-GARCH and Dynamic Conditional Correlation methods, we are able to construct 60-month rolling estimates of the impact of unemployment risk on the ex-ante equity premium. This also allows us to contribute to the current debate on what is the appropriate forward looking equity premium. At present, there is no agreement on this topic. For example, in the 2008 update of Welch (2000) the 5%-95% percentile estimates of the 30-year arithmetic equity premium are 3.0% to 8.6%. Our results strongly suggest that there was a structural weakening of the relationship between the ex-ante equity premium and unemployment risk in the mid-1940s. This places our estimate of the current equity premium towards the bottom of Welch's range as of 2000, but would expect our estimates to increase during the current credit crisis.

2. Unemployment and the Equity Premium

Mehra and Prescott (1985) document that the historically observed excess returns appeared to be much greater than the ex-ante equity premium that a rational investor would require. One of the potential explanations for the equity premium puzzle is that the standard model assumes that the market is complete, so it does not take into account uninsurable risk. Mankiw (1986) illustrates that, in the absence of certain contingent claims markets, individuals are subject to different risks and, as a result, they are heterogeneous, which in turn influences equilibrium asset prices and normally increases the theoretical equity premium. A number of studies have investigated the implications of various sources of uninsurable risk on asset prices and portfolio choice. Among those sources, labour income uncertainty is one of the most studied background risks. This is because it is a pervasive economy-wide risk that is difficult to diversify. Jagannathan and Wang (1993) and Jagannathan, Kubota and Takehara (1998) report that labour income constitutes a relatively large percentage of the total national income in the US, Japan and many European countries. Matthews and Benjamin (1992) and Constantinides (2002) suggest that job loss can be considered as a major uninsurable income shock. Seligman and Wenger (2006) find that unemployment risk can amplify the stock market investment loss of stockholding workers.

Although the literature discussed above is based on the US, it can be argued that unemployment shocks are more of a UK rather than a US problem. The evidence in support of this argument can be found in Layard, Nickell and Jackman (1991). Based on their study, they find that unemployment duration is much higher in the UK than that in the US. Layard et al. (1991, p. 228) show that the average uncompleted length of spell among unemployed workers is 21 months for Britain (males) and 3 months for the US. Another example is that the average completed length of spell for entrants to unemployment is 7 months for Britain (males) and 2 months for the US. Also, it is the case that the exit rate, which they define as the proportion of unemployed people leaving unemployment within the next three-month, tends to be lower the longer the unemployment duration. From those findings, Layard et al. (1991, p. 39), therefore, state "In the USA there are very few long-term unemployed. But in Britain, where there are many, the exit rates are much lower for the long-term unemployed..." Hence, it can be said that the unemployment incidence in the UK is more severe than that in the US. As a consequence, unemployment is more likely to explain the high observed average excess return to stocks in the UK than in the US.

One can argue that unemployment risk should not be considered uninsurable because unemployment insurance does exist. In the UK, the 1911 National Insurance Act was launched as the first contributory system of insurance against unemployment for the British working classes. The intention of the Act was to accumulate reserves in good times in order to make payments in bad times. However, literature documents that unemployment compensation is inadequate. Matthews and Benjamin (1992) find that the compensation can cover only a portion of the workforce. In addition, high inflation, especially during and immediately after the World War I, eroded real compensation benefits significantly. Hence, the receipt of unemployment benefits became an unattractive alternative relative to work. Dynarski and Sheffrin (1987) assert that complete and private unemployment insurance cannot exist because of the following reasons. First, ideal insurance schemes cannot be perfectly designed as long as moral hazard and adverse selection exist. Second, some macroeconomic shocks to the economy make the private sector unable to provide contemporaneous self-insurance. In addition, Seligman and Wenger (2006) point out that the risk-coping strategies such as self-insurance and institutionalised insurance are not entirely sufficient in the context of unemployment. There are two reasons for this assertion. First, workers have firm-specific and general skills that are costly to acquire, thus this makes diversification expensive. Second, insuring against income loss as a result of unemployment is only partially accomplished through the unemployment insurance system. Based on these counterarguments, it is justified to consider unemployment risk as at least partially uninsurable.

Two methods of measuring unemployment risk are discussed in previous empirical research. The first is to use econometric models. Dunn (1998), Carroll, Dynan and Krane (2003) and Seligman and Wenger (2006) use the probit model to estimate the probability of job loss. This probability indicates the likelihood that a currently employed individual will become unemployed in the next period. Boyd, Hu and Jagannathan (2005) estimate unemployment shocks by subtracting unemployment forecast – obtained from an Autoregressive Moving Average (ARMA) model with exogenous variables – from actual unemployment. Chetty and Szeidl (2007) conduct an event study to examine the response of households' consumption following the unemployment shock by using PSID data. They measure the magnitude of the unemployment shock in each year based on the amount of reduction in wage income due to an unemployment spell of the households' head in that particular year. The second method of measuring unemployment risk is to use a survey. Lusardi (1998) obtains a measure of subjective job insecurity by conducting a survey of self-reported unemployment expectations on a sample of workers in the US. Benito (2006) estimates unemployment risk from the survey that asks each employed individual questions such as "In the next twelve months, how likely do you think it is that you will become unemployed?". There is no general consensus on how to measure unemployment risk. In this paper we develop a new measure of the average probability of an employed individual becoming unemployed that is based only on macroeconomic unemployment data and estimates of the duration of unemployment. This allows us to make this estimate over our very long time series.

Constantinides and Duffie (1996) show that, in situations where idiosyncratic shocks are persistent and the conditional variance of idiosyncratic shocks vary inversely with stock returns, such risk can have the potential to explain the equity premium puzzle. However, this model cannot accommodate unemployment risk. In order to resolve this problem, Freeman (2002) extends the model of Constantinides and Duffie (1996) to incorporate unemployment-type risk. However, the theoretical specification proposed by Freeman (2002) is still difficult to examine empirically, so we use a simplified version of his model in the next section.

With regards to current estimates of the ex-ante equity premium for the US, Welch (2000), with updates in 2001 and 2008, provides summary statistics from surveys of a large number of academics in the area. The mean estimate of the 30-year arithmetic equity premium in 2008 was 5.7%. By contrast, Graham and Harvey (2006) has reported that the average US equity premium for the next ten years relative to a ten-year US Treasury bond obtained from the November 2005 survey of the 325 US Chief Financial Officers (CFOs) was 2.39%. Dimson, Marsh and Staunton (2006, p. 5) describe that the arithmetic mean equity premium suggested by many key finance textbooks are on average at 8.5%, while their own estimate based on the arithmetic average of data from 1900 to 2005 is 6.14% for the UK. Also, Gregory (2007) suggests that the ex-ante arithmetic mean equity premium for the UK may be as low as 2%. Our results will suggest that estimates at the lower end of this range may be more justifiable than ones at the top.

3. The Theoretical Model

3.1 The Asset Pricing Model

Let $E_t\left(r_{t+1}\right)-r_{ft}$ be the equity premium, where $E_t\left(r_{t+1}\right)$ denotes the expected return on the market and r_{ft} denotes the risk-free rate of return. In the Constantinides and Duffie (1996) model, the equity premium can be approximated by

$$E_t\left(r_{t+1}\right)-r_{ft} \quad \approx \quad \gamma \operatorname{cov}_t\left(\Delta c_{t+1}, r_{t+1}\right)-\frac{1}{2}\gamma(\gamma+1)\operatorname{cov}_t\left(y_{t+1}^2, r_{t+1}\right) \tag{1}$$

The first term on the right-hand-side of Equation (1) is the product of the coefficient of relative risk aversion, γ, and the conditional covariance between consumption growth, Δc_{t+1}, and the return on risky asset, r_{t+1}. These variables determine the magnitude of the equity premium in the standard complete market scenario. The effect of the idiosyncratic risk on the magnitude of the equity premium is captured in the second term on the right-hand-side of the equation. The parameter y_{t+1}^2 determines the degree of income heterogeneity in the economy at that time. The larger the negative covariance between the magnitude of idiosyncratic risk and stock returns, the larger is the equity premium. Note that Constantinides and Duffie's (1996) model assumes that the idiosyncratic shocks to personal income are normally distributed. However, certain types of personal income risk, such as unemployment risk, do not take this form. Hence, Freeman (2002) extends the Constantinides and Duffie (1996) model within a Markov setting to situations where investors face rare, but severe, shocks to their labour income. He argues that if no trade is an equilibrium, knowing the pricing kernel of any individual is sufficient to determine equilibrium expected returns.

Consider any two periods of a multi-period economy. There are N finite states of the economy. The notation i and j is used to denote the macroeconomic states of the economy prevailing at time t and $t+1$, respectively. $Q(i,j)$ denotes the fixed probability that the economy moves from state i to state j between times t and $t+1$. $E^Q[\bullet], \operatorname{cov}^Q[\bullet,\bullet]$ denote expectations and covariances with respect to this probability.

In this paper, we use a simplified version of Freeman's (2002) generic model.[1] There are two employment states; $v = 2$, where state 1 denotes employment and state 2 denotes unemployment. It is assumed that all individuals who are in the same employment states at both times t and $t+1$ have the same change in income. Finally, we assume that the change in income from moving from employment state 1 to employment state $v^* \in \{1,2\}$ is independent of the prevailing macroeconomic conditions at times t and $t+1$. As unemployment is economically disadvantageous, it is required that $J(1,1) > J(1,2)$; the change in income if an individual remains employed is greater than if he or she becomes unemployed.

In this case, even though investors are heterogeneous in outcome, they all apply an identical pricing kernel to asset pricing problems. From equation (6) of Freeman (2002, p. 4), we can price assets as if there were a representative agent, whose pricing kernel, $E(i,j)$, on moving from state i to state j is:

$$E(i,j) \quad = \quad e^{-\rho} \sum_{v^*=1}^{2} P_{1v^*}(i,j) J^{-\gamma}(1,v^*) \qquad (2)$$

Where ρ is the time preference coefficient and $P_{1v^*}(i,j)$ is the probability of any individual moving from employment state 1 to employment state $v^* \in \{1,2\}$ conditional on the macroeconomic state going from i to j.

It can be shown that under these assumptions, the ex-ante equity premium is given by

$$E^{Q}\left[R(i,j) - R_f(i)\right] \quad = \quad \Gamma \operatorname{cov}^{Q}\left[P_{12}(i,j), R(i,j)\right] \qquad (3)$$

Where $R(i,j)$ is the gross market return when the macroeconomic state moves from i to j between t and $t+1$, $R_f(i)$ is the gross one-period risk-free rate at time t and $\Gamma = -e^{-\rho} R_f(i)\left[J^{-\gamma}(1,2) - J^{-\gamma}(1,1)\right]$.[2] This is analogous to equation (1), except here all idiosyncratic risk is of "jump" form while in equation (1) it is all of "diffusion" form.

[1] This paper follows Freeman's (2002) notations.
[2] Strictly speaking, Γ is a function of i as the risk-free rate depends on today's state i. However, the observed variability of real Treasury bill rate is very low. Therefore, this paper assumes that Γ is a constant and it doesn't depend on the state at time t.

From Equation (3), the ex-ante equity premium is predominantly driven by three variables. The first is the relative magnitude of the change in income on becoming unemployed, $J(1,2)$, to the change in income on remaining employed, $J(1,1)$. The second is the risk aversion of the agent, γ. The third is the covariance between the probability of a currently employed person becoming unemployed, $P_{12}(i,j)$, and the market return, $R(i,j)$. As $J^{-\gamma}(1,2) - J^{-\gamma}(1,1) > 0$ for $\gamma > 0$, the covariance term in Equation (3) must be negative in order to ensure that the ex-ante equity premium is positive. The ex-ante equity premium should be large and positive if stock returns co-vary negatively with the probability of becoming unemployed.

The equity premium could be estimated from Equation (3) by calibrating the Markov process directly. In this paper, though, we prefer a time-series empirical approach, and therefore base our estimates of the conditional equity premium on the time-series equivalent of Equation (3):

$$E_t\left[R_m(t+1) - R_f(t+1)\right] \quad = \quad \Gamma \operatorname{cov}_t\left[P_{12}(t+1), R_m(t+1)\right] \tag{4}$$

$P_{12}(t+1)$ is the probability of an employed individual becoming unemployed over the time interval $[t, t+1]$ and $R_m(t+1)$, $R_f(t+1)$ are the realised gross market return and gross risk-free rate over the same time interval.

3.2 Estimating the probability of becoming unemployed

Define $LF = EM_t + UN_t$, where LF is the total number of members of the labour force in an economy, which is assumed to remain unchanged through time, EM_t and UN_t denote the number of those who are employed and unemployed at time t, respectively. The unemployment and employment rates can be defined respectively as $U_t = UN_t / LF$ and $E_t = EM_t / LF$. The unemployment rate today is $U_t = 1 - E_t$. Note that the change in employment status of individuals from being employed to becoming unemployed may result from two circumstances. The first is the circulation of labour between employed and unemployed workers, which can occur even when the unemployment rate remains unchanged and the number of jobs available in the market is constant. The second is the change in economic conditions that affects the supply side of the economy, which leads to the change in the number of jobs available in the market and, in turn, causes a change in the unemployment rate. In order to solve the empirical difficulty in modelling the probability of becoming unemployed, it is necessary to introduce an intermediate pseudo-time-period into the model.

Let $t+1-\delta$ be a pseudo time period, where δ is an instant of time. In the time interval from t to $t+1-\delta$, the unemployment rate remains unchanged at U_t. Workers in this time interval can become unemployed because of the circulation of labour alone. There are constant probabilities $\pi_{11},\pi_{12},\pi_{21},\pi_{22}$ that workers will remain employed, become unemployed, become employed and remain unemployed, respectively, even though the unemployment rate hasn't changed. The unemployment rate at $t+1-\delta$, which is restricted to equal U_t, arises from $U_t = \pi_{22}U_t + \pi_{12}(1-U_t)$. Rearranging gives:

$$\pi_{12} \quad = \quad (1-\pi_{22})\frac{U_t}{1-U_t} \tag{5}$$

In the time interval from $t+1-\delta$ to $t+1$, the unemployment rate can change from U_t to U_{t+1} due to the change in the macroeconomic condition. Note that the change in the employment status during this time interval must come only from the movement in unemployment rate. In this time instant, there are four constant probabilities $\phi_{11},\phi_{12},\phi_{21},\phi_{22}$ that workers will, respectively, remain employed, become unemployed, become employed and remain unemployed as a result of the change in the unemployment rate. In this case, employed workers can only become unemployed if the aggregate rate of unemployment rises. Similarly, unemployed workers can only become employed if the unemployment rate falls. The probability of becoming unemployed during this time interval is given by

$$\phi_{12} \quad = \quad \max\left[\frac{U_{t+1}-U_t}{1-U_t},0\right] \tag{6}$$

while the probability of becoming reemployed is $\phi_{21} = \max\left[\frac{U_t-U_{t+1}}{U_t},0\right]$, so that

$$\phi_{22} \quad = \quad 1-\max\left[\frac{U_t - U_{t+1}}{U_t}, 0\right] \tag{7}$$

To obtain the probability of becoming unemployed from time t to $t+1$, the evolution of the employment status in the two time intervals must be considered together. Figure 1 illustrates that workers can change their employment status from being employed at time t to become unemployed at $t+1$ through two paths.

[Insert Figure 1 around here]

Given this, $P_{12}(t+1) = \pi_{11}\phi_{12} + \pi_{12}\phi_{22}$. Noting that $\pi_{11} = 1-\pi_{12}$ and substituting in from Equations (5), (6) and (7):

$$\begin{aligned} P_{12}(t+1) \quad = \quad & \max\left[\frac{U_{t+1} - U_t}{1-U_t}, 0\right] + (1-\pi_{22})\frac{U_t}{1-U_t} \times \\ & \left(1-\max\left[\frac{U_t - U_{t+1}}{U_t}, 0\right] - \max\left[\frac{U_{t+1} - U_t}{1-U_t}, 0\right]\right) \end{aligned} \tag{8}$$

The probability of becoming unemployed between time t and $t+1$ only then depends on the aggregate unemployment rates at the two times and the probability of remaining unemployed under labour circulation, π_{22}. This variable can be estimated from the duration of unemployment at times when the aggregate rate of unemployment is stable. The expected duration of unemployment in such situations is $\sum_{t=1}^{\infty} t\pi_{22}^{t-1}(1-\pi_{22}) = 1/(1-\pi_{22})$. In our baseline calibrations, we set $\pi_{22} = 0.5$, implying an average duration of unemployment of two months under stable macroeconomic conditions. This estimate is a lower estimate for the UK given the empirical evidence cited in Section 2. Our results do not appear to be particularly sensitive to this choice of value and additional results are available on request.

In Figure 2, we present values of $P_{12}(t+1)$ for various values of U_t, U_{t+1}. The nonlinearities in unemployment risk are clear. For fixed U_{t+1}, $P_{12}(t+1)$ is decreasing in U_t while $U_t < U_{t+1}$, but after this point it becomes increasing in U_t, giving a "V shape" when plotted univariately against U_t. By contrast, for fixed U_t, $P_{12}(t+1)$ is monotonic increasing in U_{t+1}. However, the slope becomes steeper when $U_{t+1} > U_t$, giving a "kink" shape to the graph. This stresses the importance of placing the correct functional form for unemployment risk, $P_{12}(t+1)$ estimated from Equation (8), into Equation (4) to provide estimates of the conditional ex-ante equity premium.

[Insert Figure 2 around here]

4. The Econometric Models

By allowing for conditionality in the covariance structure, Equation (4) allows for estimates of a time-varying ex-ante equity premium. This conditionality is incorporated through two approaches. First, the conditional correlation between those two variables can be estimated using the Dynamic Conditional Correlation (DCC) model of Engle (2002). The univariate conditional volatility of stock returns and $P_{12}(t+1)$ are estimated using Gray' (1996) Regime-Switching GARCH-type model. The model was developed initially for modelling the short-term interest rate. The justifications for applying such model to the unemployment series are as follows. First, Bean (1994), Phelps (1994), Nickell (1997 and 1998) and Blanchard (1999) have emphasised the role of the real interest rates as one of the input prices in determining unemployment. Caporale and Gil-Alana (2002) find that there is a long term relationship between interest rates and unemployment. Second, both interest rate and unemployment exhibit mean reversion behaviour. Third, the unconditional distributions of changes in the interest rate and the unemployment rate are leptokurtic. Engle (1982) shows that a possible causes of the leptokurtosis in the unconditional distribution can be attributed to conditional heteroskedasticity. Fourth, Gray (1996) shows that interest rate should be fitted using nonlinearity model. Finally, Johnes (1999) finds that nonlinearities are present in the UK unemployment data. Bianchi and Zoega (1998) also find that the behaviour of unemployment in major OECD countries is stationary around an infrequently changing mean.

4.1. The Markov Regime-Switching GARCH Model

The GRS model of Gray (1996) is described as follows. Let y_t be the return at time t modelled as a constant plus a disturbance term such that

$$y_t = \mu(s_t) + e_t(s_t) \tag{9}$$

and

$$e_t(s_t) \mid \psi_{t-1} = h_t(s_t) \times z_t \tag{10}$$

where $(s_t) = \{1,2\}$ is a notation denoting an unobserved state variable at time t that follows a first-order, two-state Markov process; $e_t(s_t)$ is a state-dependent residual term; z_t is a standard normal random variable; and $h_t(s_t)$ is a state-dependent, conditional standard deviation of y_t. In Gray's (1996) model, the conditional volatility is assumed to follow a GARCH(1,1) process

$$h_t^2(s_t) = \gamma(s_t) + \alpha(s_t)e_{t-1}^2 \times \beta(s_t)h_{t-1}^2 \tag{11}$$

where $\gamma(s_t)$, $\alpha(s_t)$ and $\beta(s_t)$ are state dependent coefficients.

For the appropriate ARCH-GARCH specification, this paper will also use the root mean square error (RMSE), the mean absolute error (MAE), the mean absolute percentage error (MAPE), the Theil Inequality Co-efficient and the modified Diebold and Mariano test statistic of Harvey, Leybourne et al., (1999) (DM-LS) to test for the appropriate ARCH-GARCH process. These measurement models do not provide a formal statistical indication of whether one model is significantly better than another is. So this paper uses forecast-encompassing tests. Diebold and Mariano (1995) proposed several methods to test the null hypothesis of equal forecast accuracy. The modified version of the Diebold and Mariano (1995) test proposed by Harvey, Leybourne et al., (1999) uses the squared prediction errors to make pair-wise comparisons of different models, with its statistics being adjusted for the presence of ARCH forecast errors. Their underlying assumptions about the forecasting error distributions are non-Gaussian, nonzero, serially correlated and contemporaneously correlated.

If two forecasts have produced two sets of forecasting errors (ε_{1t} and ε_{2t}, $t = 1,2,...,n$) then the forecasting accuracy could be tested by using the following null hypothesis.

$$E(\varepsilon_{1t} - \varepsilon_{2t}) = 0 \qquad \text{and} \qquad d_t = \varepsilon_{1t} - \varepsilon \tag{12}$$

and the observed sample mean can be calculated as:

$$\bar{d} \quad = \quad \sum_{t=1}^{n} \frac{d_t}{n} \tag{13}$$

In Harvey, Leybourne et al., (1997), the series d_t is likely to be autocorrelated. It is also assumed that all autocorrelations of order h, or higher, of the d_t are zero for h-steps ahead forecasts and the variance of d_t will be asymptotically. They pointed out the Diebold-Mariano test could be oversized in the case of a two-steps-ahead prediction (h=2), getting worse when h increases.

The test compares a statistic (S) to a critical value drawn from a student t-distribution with (n-1) degrees of freedom (n is the number of independent point forecasts in the out-of-sample period), where S is given by:

$$S \quad = \quad \left[\frac{n+1+2h+n-1h(h-1)}{n}\right]^{-\frac{1}{2}} \hat{V}(\bar{d})^{-\frac{1}{2}} \bar{d} \tag{14}$$

$$\bar{d} = \frac{1}{n}\sum_{t=1}^{n} d_t \; ; \;\; d_t = g(e_{1,t}) - g(e_{2,t}); \;\; \hat{V}(\bar{d}) = n^{-1}\left(\hat{\gamma}_0 + 2\sum_{i=1}^{h-1} \hat{\gamma}_i\right); \tag{15}$$

where $\hat{\gamma}_i$ is the estimated auto-covariance of the series of squared prediction errors and h is the forecast horizon being considered.

In Gray's (1996) model, the variances are defined as

$$h_t^2 \quad = \quad p_{1t}(\mu_1^2 + h_{t,1}^2) + (1 - p_{1t})(\mu_2^2 + h_{t,2}^2) - [p_{1t}\mu_1 + (1 - p_{1t})\mu_2]^2 \tag{16}$$

where $p_{1t} = \Pr\left(s_t = 1 \mid \psi_{t-1}\right)$ is the regime probability of being in state 1 given all information up to time $t-1$. The residual is given by

$$e_t = y_t - \left[p_{1t}\mu_1 + (1 - p_{1t})\mu_2\right] \tag{17}$$

Gray (1995, 1996) derives a nonlinear recursive expression of the regime probability as follows:

$$p_{1t} = P\left[\frac{g_{1t-1}p_{1t-1}}{g_{1t-1}p_{1t-1} + g_{2t-1}\left(1 - p_{1t-1}\right)}\right] + (1-Q)\left[\frac{g_{2t-1}\left(1 - p_{1t-1}\right)}{g_{1t-1}p_{1t-1} + g_{2t-1}\left(1 - p_{1t-1}\right)}\right] \tag{18}$$

where the transition probabilities that the regime 1 and 2 at time $t-1$ followed by regime 1 and 2 at time t are given, respectively as $P = \Pr\left[s_t = 1 \mid s_{t-1} = 1\right]$ and $Q = \Pr\left[s_t = 2 \mid s_{t-1} = 2\right]$.

The conditional probability density function is

$$g_{it} = \frac{1}{\sqrt{2\pi}h_{it}}\exp\left\{\frac{-\left(y_t - \mu_{it}\right)^2}{2h_{it}^2}\right\} \quad ; \quad i = 1,2 \tag{19}$$

The parameters $\theta = \{P, Q, \mu(s_t), \gamma(s_t), \alpha(s_t), \beta(s_t)\}$ for $s_t = 1,2$ can be estimated by maximising the following log-likelihood function with respect to θ:

$$L(\theta) = \sum_{t=1}^{T}\log\left\{p_{1t}\frac{1}{\sqrt{2\pi}h_{t,1}}\exp\left[\frac{-\left(y_t - \mu_1\right)^2}{2h_{t,1}^2}\right] + (1 - p_{1t})\frac{1}{\sqrt{2\pi}h_{t,2}}\exp\left[\frac{-\left(y_t - \mu_2\right)^2}{2h_{t,2}^2}\right]\right\} \tag{20}$$

4.2. The DCC Model

Define $y_t \mid I_{t-1} \sim N(0, H_t)$, where y_t is the $k \times 1$ vector of demeaned variable values conditional on information, denoted I_{t-1}, available at time $t-1$; y_t is assumed to be conditionally multivariate normal; H_t is the conditional covariance matrix and

$$H_t \quad = \quad D_t R_t D_t \tag{21}$$

where R_t is the $k \times k$ time-varying correlation matrix and D_t is a $k \times k$ diagonal matrix of conditional standardised residuals, ε_t.

Engle (2002) shows that the likelihood function of the DCC estimator may be written as:

$$L \quad = \quad -0.5\sum_{t=1}^{T}\left(k\log(2\pi) + 2\log(|D_t|) + \log(|R_t|) + \varepsilon_t' R_t^{-1}\varepsilon_t\right) \tag{22}$$

The correlation estimator can be written as:

$$R_t \quad = \quad (1-\alpha-\beta)\overline{R} + \alpha\varepsilon_{t-1}\varepsilon_{t-1}' + \beta R_{t-1} \tag{23}$$

where $\overline{R}$ is the unconditional correlation matrix of ε_{t-1}.

5. Empirical Estimates of the Input Variables into the Asset Pricing Model

5.1. The Data

The data used in this paper are as follows: (i) three series of UK unemployment rates, hereafter called “UNNEW”, “UNOLD” and “UNTAB”, (ii) the FTSE All-Share total return index and (iii) the UK Consumer Price Index (CPI). All data are at monthly frequency. The period of study is from January 1888 to December 2000. Most of the data are from Global Financial Data (GFD) website (http://www.globalfindata.com), except for UNTAB in which only part of it comes from this source. The detailed information about the sources for, and problems associated with, the UNTAB series are available upon request. For more details on how UNOLD and UNNEW were calculated, see Global Financial Data (2005). Note that all unemployment rates used in this paper were seasonally adjusted.

The Augmented Dickey Fuller (ADF) test results (not reported) show that the null hypothesis of unit root for the variables $P_{12}(t+1)$ and π_{12} cannot be rejected at the 5% significance level. However, the same hypothesis can be rejected after those two series were expressed in first-difference. This implies that we can write the unemployment risk variable as:

$$P_{12}(t+1) \quad = \quad P_{12}(t) + \upsilon_{t+1} \tag{24}$$

For some non-integrated random variable $\upsilon_{t+1} = \Delta P_{12}(t+1)$. By substituting this into Equation (4) and noting that $P_{12}(t)$ is non-stochastic at time t:

$$E_t\left[R_m - R_f\right] \quad = \quad \Gamma \operatorname{cov}_t\left[\Delta P_{12}(t+1), R_m(t+1)\right] \tag{25}$$

From an empirical perspective, we also consider three other possible non-integrated measures of unemployment risk in place of the theoretically justified variable $\Delta P_{12}(t+1)$. Doing so will help determine whether the results are sensitive to different unemployment risk measures. The three measures are $\Delta U_{t+1} = (U_{t+1} - U_t)/U_t$, $\Delta u_{t+1} = \ln(U_{t+1}) - \ln(U_t)$ and $\Delta \pi_{12,t}$, which is defined in an obvious way from Equation (5). Notice that ΔU_{t+1} is a relative, rather than absolute, change in unemployment as this helps keep the variable scale independent.

Descriptive statistics of the data used in this study are reported in Table 1 to 4.

[Insert Tables 1- 4 around here]

From these statistics, it can be seen that the properties of two of the variables, $\Delta P_{12}(t+1)$ and $\Delta \pi_{12,t}$, are highly similar. Therefore, we would expect to see similar estimates of the equity premium from these two variables. By contrast, ΔU_{t+1} and Δu_{t+1} have much higher standard deviations and therefore, a priori, we would expect higher estimates of the equity premium from these measures of "unemployment risk".

5.2. The Conditional Volatility Results

Figures 1 to 3 plot the conditional standard deviations, obtained from the Markov regime-switching GARCH model, of Δu_t, ΔU_t, $\Delta \pi_{12,t}$ and $\Delta P_{12}(t+1)$ for UNOLD, UNNEW and UNTAB unemployment rate series, respectively. It is apparent from those figures that the magnitudes of unemployment risks in the first half of the Twentieth century were higher than those in the latter. Unsurprisingly, all measures of unemployment risk during the First World War of 1914-1918 (WWI) and the Second World War of 1939-1945 (WWII) were (artificially) low because of the war demands for labour. From these graphs it is clear, though, that if unemployment risk were uninsurable and it does affect the demand for stocks, then, ceteris paribus, the equity premium should be higher during the interwar period than those in the pre-WWI and the post-WWII periods

[Insert Figures 3-5 around here]

The next input in the estimation of the ex-ante equity premium is the conditional volatility of stock returns. Bollerslev (1986) suggests that the conditional variances of stock returns can be estimated by fitting the $\text{GARCH}(1,1)$ model to the stock returns data.

[Insert Figure 6 around here]

Figure 6 plots the conditional standard deviations of real stock returns obtained from the $\text{GARCH}(1,1)$ model. The noticeable spikes of the conditional standard deviations of stock return shown in the figure represent the effects of the following events in history on stock returns: (i) the breakout of the First World War in 1914, (ii) the Great Depression of 1929, (iii) the breakout of the Second World War in 1939, (iv) the Oil shock of 1976 and (v) Black Monday of 1987. It can be seen that the UK stock market risk had increased over time during the previous century. This finding is not consistent with the case of US in which the stock market risk had been found to be declining. For empirical works on the reduction of the US stock market risk, see, for example, Pastor and Stambaugh (2001) and Kim, Morley and Nelson (2005). The results show that while the UK stock market risk, as illustrated in Figure

6, had increased over the Twentieth century, unemployment risks, as illustrated in Figures 3 to 5, has declined.

5.3. The Conditional Correlation Results

Figure 7 to 9 plot the conditional correlations between real stock return and each of the four measures of unemployment risk across three unemployment rate series. From those figures, the conditional correlations for all cases are mostly negative, which is consistent with the argument that unemployment risk increases the equity premium. Negative correlation between unemployment risk and stock return destabilises the consumption smoothing pattern of individuals. For stocks to be attractive they must, therefore, offer higher expected returns. Specifically, negative correlation between unemployment risk and stock return contributes to the positive ex-ante equity premium. The more negative of the correlation, the higher is the positive ex-ante equity premium. From Figure 7 to 9, the magnitudes of the correlation are the most negative around 1918, which was the end of the First World War period. Also, in Figures 3 to 5, the magnitudes of conditional standard deviations of unemployment risk are also higher during the aftermath of the First World War relative to other periods. It can therefore be hypothesised that the ex-ante equity premium was largest around the end of the First World War.

[Insert Figures 7-9 around here]

5.4. Estimating Γ

In order to estimate the ex-ante equity premium from Equation (25), it is necessary to calibrate the variable $\Gamma = -e^{-\rho} R_f\left(i\right)\left[J^{-\gamma}\left(1,2\right) - J^{-\gamma}\left(1,1\right)\right]$. Table 5 reports a range of possible values for this variable.

[Insert Table 5 around here]

The second and the third columns of Table 5 present the range of parameter values deemed reasonable in the financial economics literature, although this is largely subjective. For the risk aversion coefficient, γ, and the rate of time preference coefficient, ρ, values are taken from empirical research. For example, estimates of γ obtained from Mankiw (1985) and Summers (1981) are 3. More recently, Chetty (2003) uses labour supply data and estimates γ to be about 1. However, Barsky et al., (1997) and Metrick (1995) use experimental data to obtain a gamma value and it is in the range of 0 to 15. Dreze (1987) uses insurance contracts data and Coutant (2000) uses option prices data to estimate gamma. They both obtain estimates of gamma range from 1 to 11. Wada (2007) uses a risk aversion coefficient that ranges from 0 to 10 in his analysis. Hence, the proposed range for the risk aversion coefficient used in this paper is from 1 to 15. For the magnitude of the ex-ante equity premium as reported below, this paper uses $\gamma = 6$ in the estimation.

The rate of time preference, ρ, as shown in the second and third column of Table 5, varies from 0 to 0.14. This range is based on the following empirical evidence. Moore and Viscusi (1990) estimate the discount rate to be around 2%. However, Lawrence (1991) estimates the discount rate of median-income households to be around 4% to 13%. Carroll and Samwick (1997) report point estimates for the discount rate ranging from 5% to 14%. However, Gourinchas and Parker (2001) report point estimates of 4.0% to 4.5%. For literature review on the time preference, see Frederick, Loewenstein and O' Donoghue (2002). This paper uses $\rho = 0.02$ in estimating the ex-ante equity premium as reported below.

For the average change in income of those who remain employed given that they were employed in the previous period, $J(1,1)$, this paper arbitrarily determines its range as 0.98 to 1.075. This means that for those who remain employed, their income can reduce by 2% at most or increase by 7.5% at most. The ex-ante equity premium presented below was estimated based on $J(1,1) = 1.04$. On the other hand, the average change in income of those who become unemployed given that they were employed in the previous period, $J(1,2)$, is from 0.4 to 0.8. The ex-ante equity premium presented below was estimated based on $J(1,2) = 0.6$.

We estimate the real risk-free rate from the fundamental theorem of asset pricing; $R_f = 1/E^Q[E(i,j)]$, where the pricing kernel is given in Equation (2). For the calibration of Γ, it is assumed that this variable is constant across time and so we take fixed estimates of each of the input parameters. For the expected probability of remaining employed, $E^{Q}\left[P_{11}(i,j)\right]$, this paper assumes that it ranges from 90% to 99.5 with a baseline estimate of $E^{Q}\left[P_{11}(i,j)\right]=0.99$. Since $E^{Q}\left[P_{12}(i,j)\right]=1-E^{Q}\left[P_{11}(i,j)\right]$, this results in the probability of becoming unemployed equals to 0.01. Carroll, Dynan and Krane (2003) estimate the probability of job loss and they obtain such probability equal to 0.02. This value is relatively close to the probability of becoming unemployed used in this paper.

By using the parameter values presented in the last column of Table 5, the implied gross risk-free rate is $R_f(i)=1.0134$. This is close to standard empirical estimates. As a result, our calibrations are based on $\Gamma=-20.7105$. It should be noted that the value of Γ is particularly sensitive to γ and $J(1,2)$.

6. The Ex-Ante Equity Premium

6.1 Estimates of the Unconditional Ex-Ante Equity Premium

Table 6 reports the average annual ex-ante equity premia estimated from Δu_t, ΔU_t, $\Delta\pi_{12,t}$ and $\Delta P_{12,t}(i,j)$ for three series of unemployment rate as estimated from Equation (25). The variation in the equity premia is due to the difference in the unemployment risk measures used in the estimation. In the first column of the table, the average annual ex-ante equity premia estimated based on Δu_t ranges from 3.01% for UNNEW series of unemployment rate to 4.55% for UNTAB series. The average annual ex-ante equity premia estimated from ΔU_t, reported in the second column of the table, ranges from 4.75% to 6.44%. These are broadly consistent with standard estimates of the unconditional equity premium for the UK. For example, Vivian (2007) obtains estimates, based on the dividend model, of an equity premium of 4.40% over the 1901-2004 period. Empirical estimates by Dimson, Marsh and Staunton (2006) find that the arithmetic mean annual equity premium relative to bills for the UK over the 1900-2005 period was 6.14%.

[Insert Table 8 around here]

By contrast, the third and last columns of the table, based on unemployment risk measures $\Delta\pi_{12,t}$ and $\Delta P_{12,t}\left(i,j\right)$ respectively, present estimates of the equity premium that are an order of magnitude lower. Estimates, in this case, are in the range of 0.21% to 0.35%. As stressed in Section 5, this finding is unsurprising given the unconditional standard deviations of the different employment risk series.

Of course, the results that are most theoretically robust are those driven by the risk variable $\Delta P_{12,t}\left(i,j\right)$. These imply that unemployment risk alone is unlikely to resolve the equity premium puzzle. It is, though, important to note that estimates of the equity premium documented throughout this section reflect the incremental effects of incorporating unemployment risk into the asset pricing model. This is in addition to any contribution that aggregate consumption uncertainty plays in determining the equity premium in a standard Mehra and Prescott (1985) type model.

6.2 Estimates of the Conditional Equity Premium

Figure 10 to 13 plot the 60-month moving averages of the ex-ante equity premium estimated from $P_{12}(t+1)$, Δu_t, ΔU_t and $\Delta\pi_{12,t}$ respectively. The magnitudes of the estimated equity premia are expressed on an annual percentage basis. In each figure, the premia obtained from all three unemployment rate series are presented together.

[Insert Figures 10-13 around here]

The low average level of the average equity premium, as reported in Table 6, is clear from this figure. However, it is also clear that the equity premium effect is highly time-varying, with an apparent regime shift in the mid 1940s. This is consistent with an effect for the US documented by Kim et al. (2005). However, the explanatory factor for the effect is different. It was shown in Figure 6 that market volatility actually increased after WWII in the UK, while Kim et al.'s (2005) argument for a declining equity premium is derived from lower market risk in the second half of the century. Here the declining equity premium results from lower unemployment risk in the aftermath of WWII. This finding also contrasts with studies such as Vivian (2007), who argues that the equity premium is quite constant over the two halves of the century in the UK: 4.40% over the 1901-2004 period, 4.22% in the 1901-1950 sub-period and 4.58% for the 1951-2004 sub-period. Similarly, Gregory (2007) reports the arithmetic average ex-ante equity premia obtained from the dividend growth model of 4.67% over the 1925-2006 period, 5.13% in the 1925-1950 sub-period and 4.46% for the 1951-2006 sub-period. There is also evidence that the equity premium increased again in the 1970s, before declining further in the mid 1980s in some, but not all series.

In addition, while, on average, unemployment risk appears to have little influence on the equity premium, at certain times the effects are certainly of economic significance – over 1% in some years on the baseline parameterisation for Γ. This implies that, while unemployment is not a factor that significantly influences the equity premium during good times, in recessions it is an important contributory factor.

Figures 12-13, while different in levels to Figure 10, as expected from Table 6, provide strong support for the conclusion that the influence of unemployment on the ex-ante equity premium is highly time-varying. The shapes of all graphs are also broadly consistent and demonstrate a structural decrease in the equity premium in the mid 1940s and again in the mid 1980s preceded by an offsetting rise in the 1970s. Figure 11, based on Δu_t, suggests that the reduction in the cost of capital occurred in the mid 1920s but is otherwise consistent with the other exhibits.

7. Conclusion

In this paper we have shown that while, in good times, the potential risk of unemployment has little effect on the ex-ante equity premium, in severe downturns the relationship is much stronger. As we enter into a credit crisis, when unemployment rates are expected to increase dramatically, this strongly suggests that we are facing a sharp increase in the equity premium. For example, if the cost of equity capital in economic booms is 4.5% (1.5% risk-free + 3% equity premium) and the expected real future growth rate of dividends in perpetuity is 2%, then the Static Gordon Growth model would suggest an equilibrium price/dividend ratio of 40. In a recession, a standard analysis might allow for a decline in the dividend growth rate to 1.5%, reducing the price/dividend ratio to 33.3, or a decline in market value of 17%. However, this paper would suggest that it may also be appropriate to simultaneously increase the cost of equity capital by at least 1%, giving a predicted price/dividend ratio now of 25, or a 38% decline in the market. By also factoring in the additional impact that is likely to arise from increasing aggregate consumption volatility within a complete market model, this may well be an underestimate.

As commentators draw parallels with the 1920s, the potential impact of rising unemployment (and more general macroeconomic) fears on equilibrium asset pricing is now a pressing area for further investigation. We have used an asset pricing model that explicitly incorporates such risks and have estimated “unemployment risk” in a robust way that allows for estimation using only macroeconomic employment data and information on the average duration of unemployment. We conclude that the relationship between the equity premium and unemployment is likely to be substantially different in the next 10 years than it has been over the last half century.

References

Barsky, R. B., F. T. Juster, M. S. Kimball and M. D. Shapiro (1997) Preference Parameters and Behavioral Heterogeneity: An Experimental Approach in the Health and Retirement Study, *Quarterly Journal of Economics*, 112:2, 537-579

Bautista, C. C. (2003) Interest Rate-Exchange Rate Dynamics in The Philippines: A DCC Analysis, Applied Econometrics Letters, 10, 107-111

Bean, C. (1994) European Unemployment: A Survey, *Journal of Economic Literature*, 32, 573-619

Benito, A. (2006) Does Job Insecurity Affect Household Consumption?, *Oxford Economic Papers*, 58:1, 157-181

Bianchi, M. and G. Zoega (1998) Unemployment Persistence: Does the Size of the Shock Matter?, *Journal of Applied Econometrics*, 13, 283-304

Blanchard, O. J. (1999) Revisiting European Unemployment: Unemployment, Capital Accumulation and Factor Prices, *NBER Working Paper* no. 6566

Bollerslev, T. (1986) Generalized Autoregressive Conditional Heteroscedasticity, *Journal of Econometrics*, 31, 307-327

Bollerslev, T (1990) Modelling the Coherence in Short-Run Nominal Exchange Rates: A Multivariate Generalized Arch Model, *Review of Economics and Statistics*, 72, 498–505

Boyd, J. H., J. Hu and R. Jagannathan (2005) The Stock Market's Reaction to Unemployment News: Why Bad News is Usually Good for Stocks, *Journal of Finance*, 60:2, 649-672

Buhr, K. E., Li, X., and Rose, L. C. (2008). Forecasting Stock Market Volatility and Assessing Information Content Present in S&P/ASX 200 Index Option Prices in Australia. *Journal of Accounting and Finance* (7), 15-30.

Caporale, G. M. and L. A. Gil-Alana (2002) Unemployment and Input Prices: A Fractional Cointegration Approach, *Applied Economics Letters*, 9:6, 347-351

Carroll, C. D., K. E. Dynan and S. D. Krane (2003) Unemployment Risk and Precautionary Wealth: Evidence from Households' Balance Sheets, *Review of Economics and Statistics*, 85:3, 586-604

Carroll, C. D. and A. Samwick (1997) The Nature of Precautionary Wealth, *Journal of Monetary Economy*, 40, 41-71

Chetty, R. (2003) A New Method of Estimating Risk Aversion, *NBER Working Paper*, 9988

Chetty, R. and A. Szeidl (2007) Consumption Commitments and Risk Preferences, *Quarterly Journal of Economics*, 122:2, 831-877

Constantinides, G. M. (2002) Rational Asset Prices, *Journal of Finance*, 57:4, 1567-1591

Constantinides, G. M. and D. Duffie (1996) Asset Pricing with Heterogeneous Consumers, *Journal of Political Economy*, 104:2, 219-240

Coutant, S. (2000) Time-Varying Implied Risk Aversion in Options Prices Using Hermite Polynomials, *Preceedings of the Chicago Board of Trade 13th Annual European Futures Symposium*, Glasgow, Scotland

Diebold, F. X., and Mariano, R. S. (1995). Comparing Predictive Accuracy. *Journal of Business and Economic Statistics, 13*, 253-263.

Dimson, E., P. Marsh and M. Staunton (2006) *The Worldwide Equity Premium: A Smaller Puzzle*, working paper, London Business School, http://ssrn.com/abstract=891620

Dreze, J. H. (1987) Decision Theory with Moral Hazard and State-Dependent Preferences, In Dreze, J. H., editor, *Essays on Economic Decisions Under Uncertainty*, Chapter 2, 23-89, Cambridge University Press, Cambridge

Dunn, W. E. (1998) Unemployment Risk, Precautionary Savings, and Durable Goods Purchase Decisions, *Finance and Economics Discussion Series 1998-49*, Washington: Board of Governors of the Federal Reserve System

Dynarski, M. and S. M. Sheffrin (1987) Consumption and Unemployment, *Quarterly Journal of Economics*, 102: 2, 411-428

Engle, R. F. (1982) Autoregressive Conditional Heteroscedasticity with Estimates of the Variance of United Kingdom Inflation, *Econometrica*, 50, 987-1007

Engle, R. F. (2002) Dynamic Conditional Correlation: A Simple Class of Multivariate Generalized Autoregressive Conditional Heteroskedasticity Models, *Journal of Business and Economic Statistics*, 20:3, 339-350

Fratantoni, M. C. (2001) Homeownership, committed expenditure risk and the stockholder puzzle, *Oxford Economic Papers,* 53, 241-259.

Frederick, S., G. Loewenstein and T. O'Donoghue (2002) Time Discounting and Time Preference: A Critical Review, *Journal of Economic Literature*, 40:2, 351-401

Freeman, M. C. (2002) Asset Pricing with Jump/Diffusion Permanent Income Shocks, *Economics Letters*, 77:1, 1-8

Global Financial Data (2005) *GFD Encyclopedia of Global Financial Markets*, 10th Edition, Global Financial Data, available at Global Financial Data website, http://www.globalfindata.com/index.php3?action=gfd_encyclopedia

Gourinchas, P. O. and J. Parker (2001) The Empirical Importance of Precautionary Saving, *American Economics Review*, 91:2, 406-412

Graham, J. R. and C. R. Harvey (2006) The Equity Risk Premium in January 2006: Evidence from the Global CFO Outlook Survey, *Working Paper*, Duke University

Gray, S. F. (1995) An Analysis of Conditional Regime-Switching Models, *Working Paper*, Duke University

Gray, S. F. (1996) Modeling the Conditional Distribution of Interest Rates as a Regime-Switching Process, *Journal of Financial Economics*, 42, 27-62

Gregory, A. (2007) How Low is the UK Equity Risk Premium?, *Xfi Centre for Finance and Investment Working Paper*, University of Exeter

Harvey, D. I., Leybourne, S. J., and Newbold, P. (1997). Testing the Equality of Prediction Mean Square Errors. *International Journal of Forecasting, 13*, 273-281.

Harvey, D. I., Leybourne, S. J., and Newbold, P. (1999). Forecast Evaluation Tests in the Presence of ARCH. *Journal of Forecasting, 18*, 343-445.

Heaton, J. and Lucas, D. (1996) Evaluating the Effects of Incomplete Markets on Risk Sharing and Asset Pricing, *Journal of Political Economy,* 104, 443-87.

Heaton, J. and Lucas, D. (2000) Portfolio Choice and Asset Prices: the Importance of Entrepreneurial Risk, *Journal of Finance,* 55, 1163-1198.

Jagannathan, R., K. Kubota and H. Takehara (1998) Relationship between Labor-Income Risk and Average Return: Empirical Evidence from the Japanese Stock Market, *Journal of Business*, 71:3, 319-347

Jagannathan, R. and Z. Wang (1993) The CAPM is Alive and Well, *Staff Report*, No. 165, Federal Reserve Bank of Minneapolis

Johnes, G. (1999) Forecasting Unemployment, *Applied Economics Letters*, 6, 605-607

Kearney, C. and V. Poti (2003) DCC-GARCH Modelling of Market and Firm-Level Correlation Dynamics in the Dow Jones Eurostoxx50 Index, *Paper Submitted to the European Finance Association Conference*, Edinburgh, August 2003

Kim, C. J., J. C. Morley and C. R. Nelson (2005) The Structural Break in the Equity Premium, *Journal of Business and Economic Statistics*, 23:2, 181-191

Lawrence, E. (1991) Poverty and the Rate of Time Preference: Evidence from Panel Data, *Journal of Political Economy*, 119, 54-77

Layard, R., S. Nickell and R. Jackman (1991) *Unemployment: Macroeconomic Performance and the Labour Market*, Oxford University Press

Lusardi, A. (1998) On the Importance of the Precautionary Saving Motive, *American Economic Review*, Papers and Proceedings, 88:2, 449-453

Mankiw, N. G. (1985) Consumer Durables and the Real Interest Rate, *Review of Economics and Statistics*, 67:3, 353-362

Mankiw, N. G. (1986) The Equity Premium and the Concentration of Aggregate Shocks, *Journal of Financial Economics*, 17, 211-219

Matthews, K. And D. Benjamin (1992) *US and UK Unemployment between the Wars: A Doleful Story*, Institute of Economic Affairs

Mehra, R. (2008) *Handbook of the Equity Risk Premium*, Elseveier

Mehra, R. and E. C. Prescott (1985) The Equity Premium: A Puzzle, *Journal of Monetary Economics*, 15, 145-161

Mehra, R. and E. C. Prescott (2003) The Equity Premium in Retrospect, in G.M. Constantinides, M. Harris and R. Stulz, eds: *Handbook of the Economics of Finance, Volume 1B: Financial Markets and Asset Pricing,* North Holland.

Metrick, A. (1995) A Natural Experiment in "Jeopardy!", *American Economic Review*, 85:1, 240-253

Moore, M. J. and W. K. Viscusi (1990) Discounting Environmental Health Risks: New Evidence and Policy Implications, *Journal of Environmental Economics and Management*, 18, S51-S62

Nickell, S. J. (1997) Unemployment and Labour Market Rigidities: Europe versus North America, *Journal of Economic Perspectives*, 11, 55-74

Nickell, S. J. (1998) Unemployment: Questions and Some Answers, *Economic Journal*, 108, 802-816

Pastor, L., and R. F. Stambaugh (2001) The Equity Premium and Structural Breaks, *Journal of Finance*, 56, 1207-1239

Phelps, E. S. (1994) *Structural Slumps: The Modern Equilibrium Theory of Unemployment, Interest and Assets*, Harvard University Press, Cambridge, MA

Seligman, J. S. and J. B. Wenger (2006) Asynchronous Risk: Retirement Savings, Equity Markets, and Unemployment, *Journal of Pension Economics and Finance*, 5:3, 237-255

Summers, L. H. (1981) Capital Taxation and Accumulation in a Life Cycle Growth Model, *American Economic Review*, 71:4, 533-544

Vivian, A. (2007) The UK Equity Premium: 1901-2004, *Journal of Business Finance and Accounting*, 34:9-10, 1496-1527

Wada, K. (2007) The Knightian Uncertainty and the Risk Premium and the Risk Free Rate Puzzles in Japan and the U.S., *Economics Letters*, 95:3, 386-393

Welch, I. (2000) Views of Financial Economists on the Equity Premium and on Professional Controversies, *Journal of Business*, 73:4, 501-537

Table 1
Descriptive Statistics for Real Stock Returns (1888:02-2000:12)

Observations	1,355
Minimum	-0.3112
Maximum	0.4052
Mean	0.0039***
Std.Devn.	0.0420
Skewness	-0.33***
Excess Kurtosis	11.28***
Jarque-Bera	7,212.39***

Notes: The null hypotheses for the sample mean, skewness, excess kurtosis and Jarque-Bera normality tests are $mean = 0$, $sk = 0$, $ku = 0$ and $JB = 0$, respectively. The asterisks *** indicates statistically significant at 1%.

Table 2
Descriptive Statistics for Four Measures of UNOLD Series (1888:03-2000:12)

	Δu_t	ΔU_t	$\Delta \pi_{12,t}$	$P_{12}(t+1)$
Observations	1,354	1,354	1,354	1,354
Minimum	-0.6295	-0.4671	-0.1133	-0.0780
Maximum	1.4856	3.4174	0.0746	0.0616
Mean	-1.35×10^{-4}	0.0089	-0.08×10^{-4}	-0.07×10^{-4}
Std.Devn.	0.1266	0.1645	0.0087	0.0067
Skewness	1.93***	8.93***	-1.93***	-0.92***
Excess Kurtosis	24.69***	154.67***	42.78***	46.68***
Jarque-Bera	35,217.43***	1,367,598***	104,071.34***	123,135.40***

Notes: The null hypotheses for the sample mean, skewness, excess kurtosis and Jarque-Bera normality tests are $mean = 0$, $sk = 0$, $ku = 0$ and $JB = 0$, respectively. The asterisks *** indicates statistically significant at 1%.

Table 3
Descriptive Statistics for Four Measures of UNNEW Series (1888:03-2000:12)

	Δu_t	ΔU_t	$\Delta \pi_{12,t}$	$P_{12}(t+1)$
Observations	1,354	1,354	1,354	1,354
Minimum	-0.6344	-0.4698	-0.0518	-0.0543
Maximum	1.4844	3.4123	0.0652	0.0611
Mean	-1.35×10^{-4}	0.0085	-0.08×10^{-4}	-0.07×10^{-4}
Std.Devn.	0.1236	0.1626	0.0062	0.0049
Skewness	2.17***	9.24***	1.04***	0.88***
Excess Kurtosis	26.80***	161.49***	33.52***	58.90***
Jarque-Bera	41,571.28***	1,490,602***	63,622.18***	195,861.52***

Notes: The null hypotheses for the sample mean, skewness, excess kurtosis and Jarque-Bera normality tests are $mean = 0$, $sk = 0$, $ku = 0$ and $JB = 0$, respectively. The asterisks *** indicates statistically significant at 1%.

Table 4

Descriptive Statistics for Four Measures of UNTAB Series (1888:03-2000:12)

	Δu_t	ΔU_t	$\Delta \pi_{12,t}$	$P_{12}(t+1)$
Observations	1,354	1,354	1,354	1,354
Minimum	-0.5202	-0.4056	-0.0438	-0.0545
Maximum	0.7965	1.2177	0.0656	0.0612
Mean	-4.23×10^{-4}	0.0045	-0.13×10^{-4}	-0.06×10^{-4}
Std.Devn.	0.0965	0.1075	0.0054	0.0042
Skewness	1.22***	3.93***	1.38***	1.20***
Excess Kurtosis	16.11***	36.99***	34.77***	71.06***
Jarque-Bera	14,973.99***	80,668.63***	68,631.91***	285,225.59***

Notes: The null hypotheses for the sample mean, skewness, excess kurtosis and Jarque-Bera normality tests are $mean = 0$, $sk = 0$, $ku = 0$ and $JB = 0$, respectively. The asterisks *** indicates statistically significant at 1%.

Table 5

Proposed Parameter Values and Values used in the Calibration of Γ

	Minimum	Maximum	Values Used
γ	1	15	6
ρ	0	0.14	0.02
$J(1,1)$	0.98	1.075	1.04
$J(1,2)$	0.4	0.8	0.6
$E^Q\left[P_{11}(i,j)\right]$	0.9	0.995	0.99

where

$$\Gamma = -e^{-\rho} R_f(i)\left[J^{-\gamma}(1,2) - J^{-\gamma}(1,1)\right]$$

$$E^Q\left[P_{12}(i,j)\right] = 1 - E^Q\left[P_{11}(i,j)\right],$$

$$R_f(i) = \frac{1}{P_f} = \frac{1}{E^Q\left[E(i,j)\right]} \text{ and}$$

$$E^Q\left[E(i,j)\right] = e^{-\rho} J^{-\gamma}(1,1) E^Q\left[P_{11}(i,j)\right] + e^{-\rho} J^{-\gamma}(1,2) E^Q\left[P_{12}(i,j)\right]$$

Table 6

Average Annual Ex-Ante Equity Premia in the Incomplete Market Framework Estimated from Four Unemployment Risk Measures (1888:5-2000:12)

Panel A: UNOLD			
Δu_t	ΔU_t	$\Delta \pi_{12,t}$	$P_{12}(t+1)$
3.98%	6.44%	0.29%	0.35%
Panel B: UNNEW			
Δu_t	ΔU_t	$\Delta \pi_{12,t}$	$P_{12}(t+1)$
3.01%	4.75%	0.22%	0.21%
Panel C: UNTAB			
Δu_t	ΔU_t	$\Delta \pi_{12,t}$	$P_{12}(t+1)$
4.55%	6.42%	0.28%	0.24%

Notes: The table reports the average annual ex-ante equity premia estimated from four risk measures (in columns) across three series of unemployment rate (in rows). The equity premium, as defined in Equation (25), is

$$E_t\left[R_m - R_f\right] = \Gamma \operatorname{cov}_t\left[\Delta P_{12}(t+1), R_m(t+1)\right]$$

The magnitudes of the equity premium presented in the table use $\Gamma = -20.7105$. The difference among four columns comes from the difference in unemployment risk measures used in the calculation. The average equity premia in the first column was calculated based on the following unemployment risk measure.

$$\Delta u_t = \ln\left(\frac{U_t}{U_{t-1}}\right)$$

The average equity premia in the second column was calculated using

$$\Delta U_t = \frac{U_t - U_{t-1}}{U_{t-1}}$$

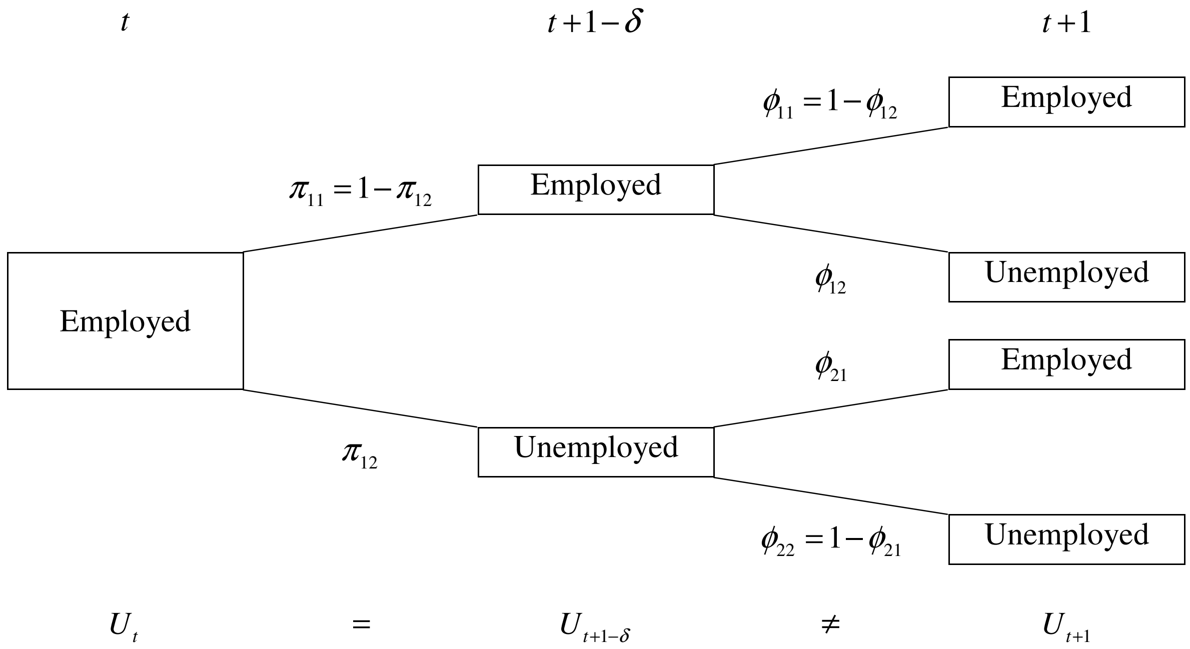
In the third column, the average equity premia were calculated from the change in the probability of becoming unemployed when the unemployment rate has not changed (labour circulation) specified as follows:

$$\pi_{12} = \frac{U_t - \pi_{22} U_t}{1 - U_t}$$

The fourth column represents average equity premia calculated using the theoretically justified measure of unemployment risk, $\Delta P_{12}(t+1)$ as derived from Equation (8)

Figure 1

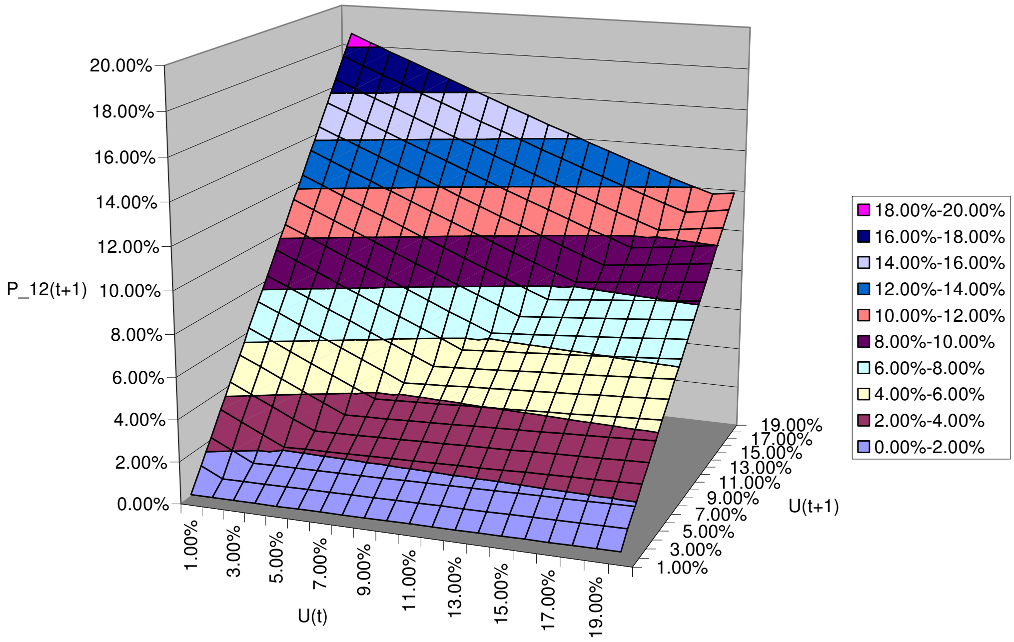
The Evolution of Employment Status of Individuals across Time



This figure illustrates how the probability of becoming unemployed due to circulation effects (π_{12}) and changes in macroeconomic conditions (ϕ_{12}) combine to determine the overall probability of becoming unemployed between times t and $t+1$, $P_{12}(i, j)$.

Figure 2

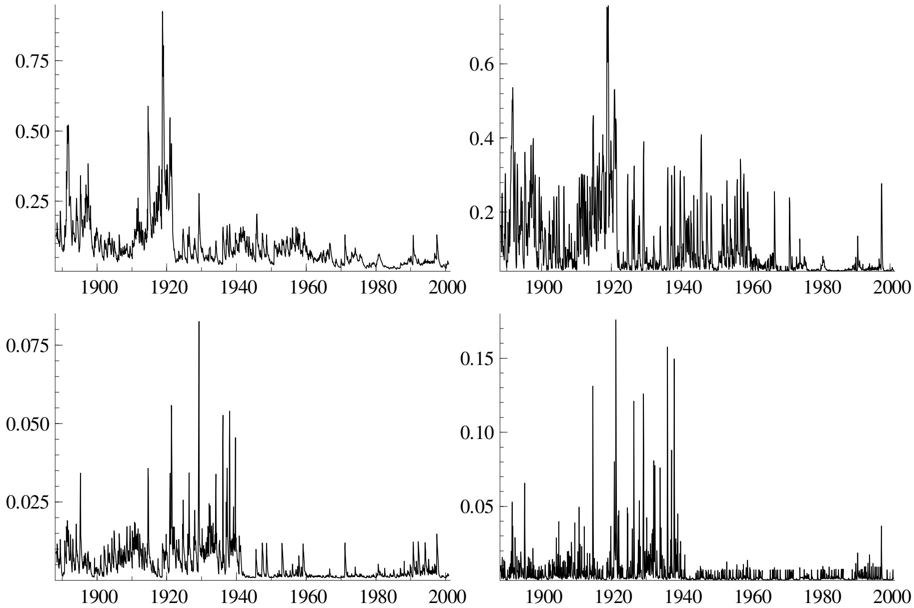
The probability of becoming unemployed



This presents estimates of the main unemployment risk variable, $P_{12}(t+1)$, for different values of unemployment at the start and end of the period, U_t, U_{t+1}, with $\pi_{22} = 0.5$

Figure 3

Conditional Standard Deviations of UNOLD Series (1888:05-2000:12)



For this exhibit, Figures 4-5 and Figures 7-9, the graphs should be read in the following order:

Δu_t	ΔU_t
$\Delta \pi_{12,t}$	$\Delta P_{12}(t+1)$

So, for example, ΔU_t is at the upper right of all figures

Figure 4

Conditional Standard Deviations of UNNEW Series (1888:05-2000:12)

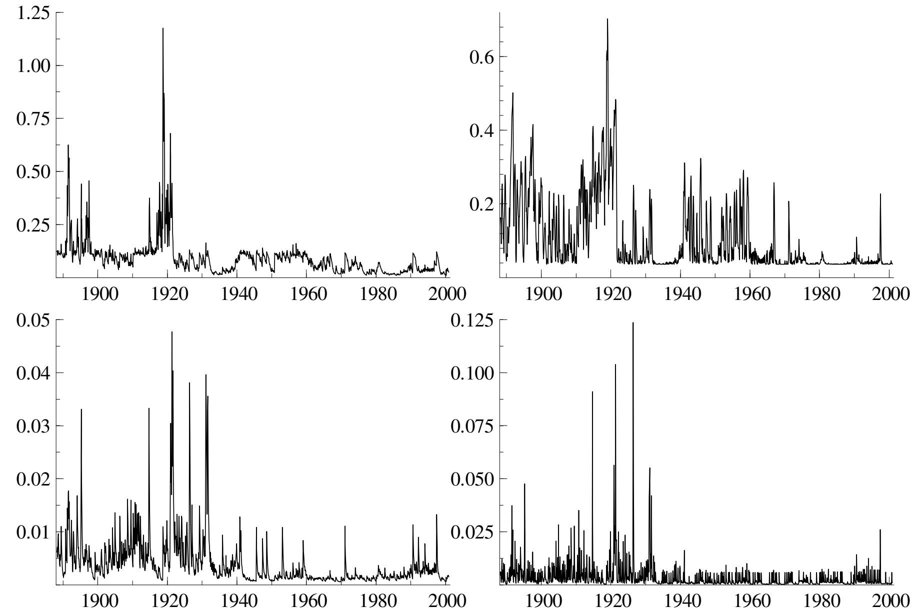


Figure 5

Conditional Standard Deviations of UNTAB Series (1888:05-2000:12)

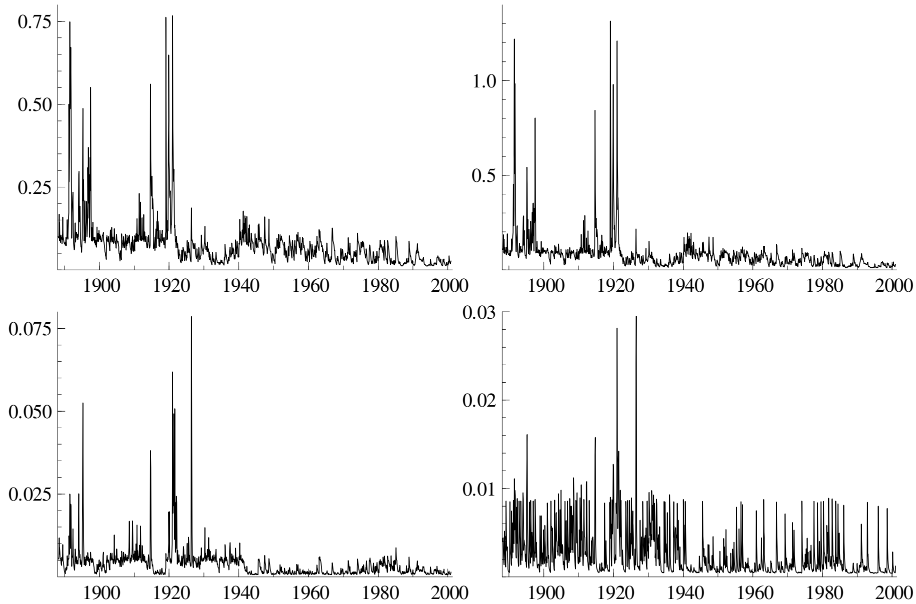


Figure 6

Conditional Standard Deviations of Real Stock Returns (1888:05-2000:12)

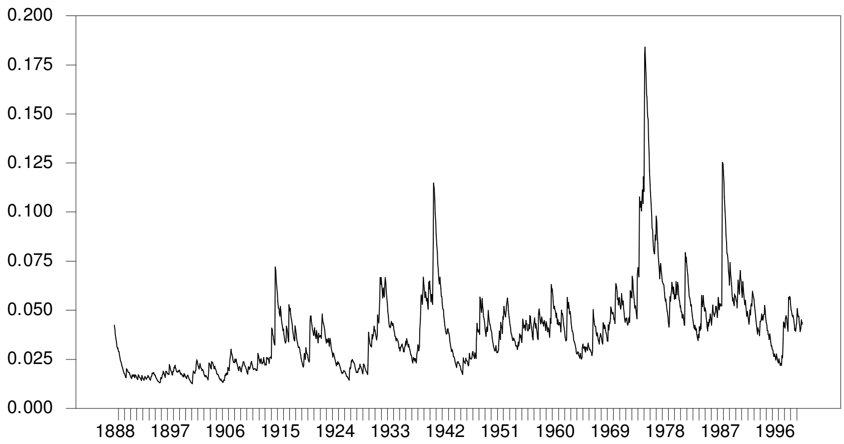


Figure 7

Conditional Correlations between Four Measures of UNOLD Series and Real Stock Returns

(1888:05-2000:12)

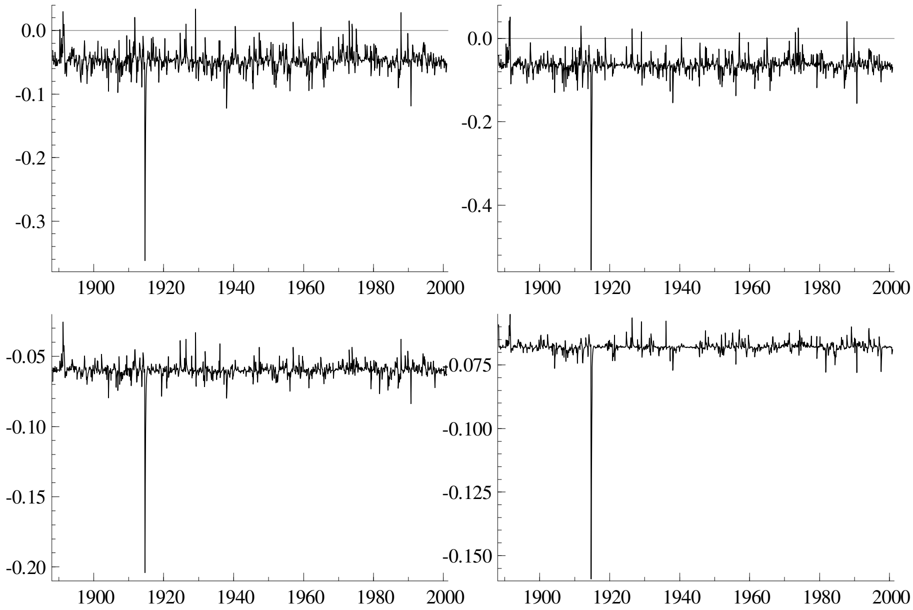


Figure 8

Conditional Correlations between Four Measures of UNNEW Series and Real Stock Returns

(1888:05-2000:12)

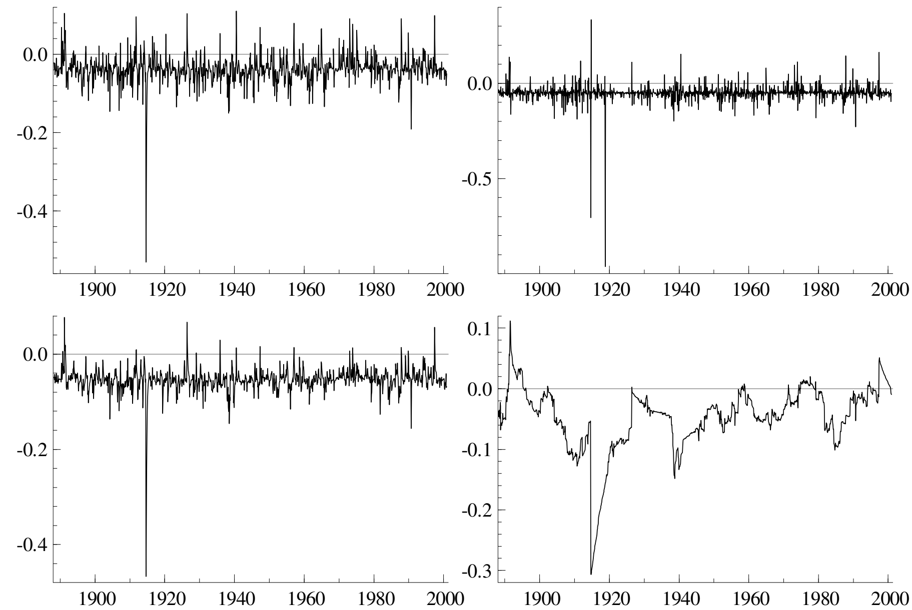


Figure 9

Conditional Correlations between Four Measures of UNTAB Series and Real Stock Returns

(1888:05-2000:12)

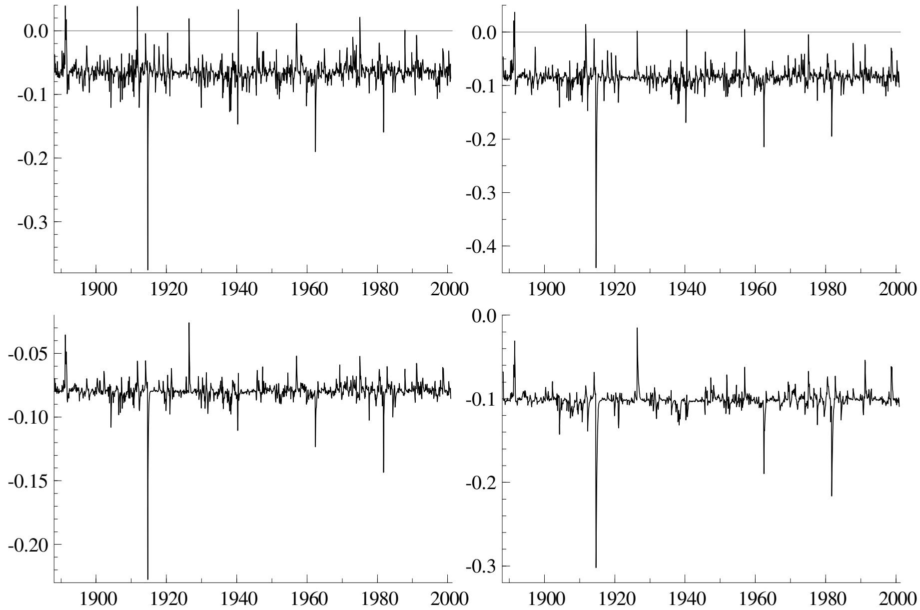


Figure 10

60-Month Moving Average of the Equity Premia Based on $\Delta P_{12}(t+1)$ (1893:5-2000:12)

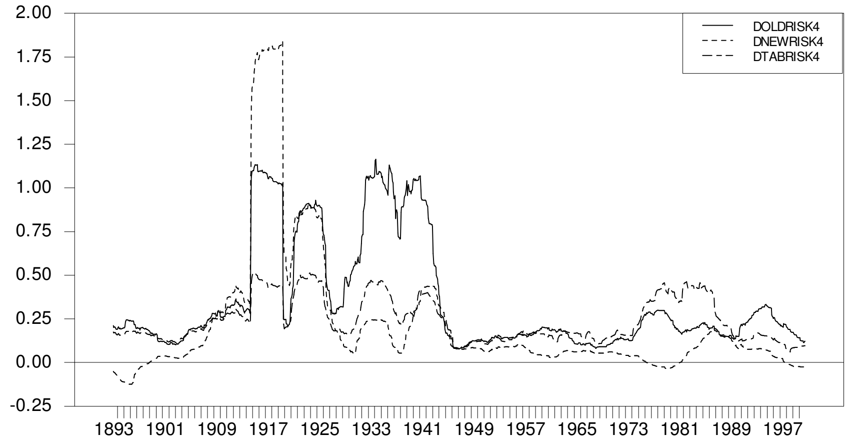


Figure 11

60-Month Moving Average of the Equity Premia Based on Δu_t (1893:5-2000:12)

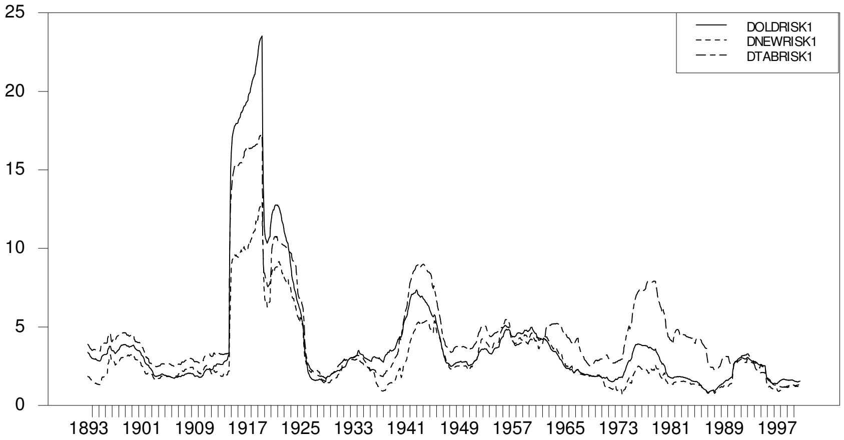


Figure 12

60-Month Moving Average of the Equity Premia Based on ΔU_t (1893:5-2000:12)

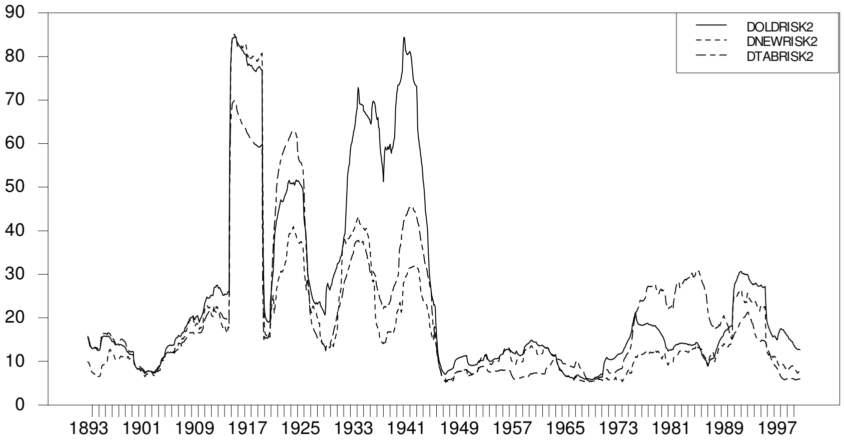


Figure 13

60-Month Moving Average of the Equity Premia Based on $\Delta \pi_{12,t}$ (1893:5-2000:12)