BANK OF CANADA
BANQUE DU CANADA

Working Paper/Document de travail
2014-3

Search-for-Yield in Canadian Fixed-Income Mutual Funds and Monetary Policy

by Sermin Gungor and Jesus Sierra

Bank of Canada Working Paper 2014-3

January 2014

Search-for-Yield in Canadian Fixed-Income Mutual Funds and Monetary Policy

by

Sermin Gungor and Jesus Sierra

Financial Markets Department
Bank of Canada
Ottawa, Ontario, Canada K1A 0G9
sgungor@bankofcanada.ca
sier@bankofcanada.ca

ISSN 1701-9397

Acknowledgements

We would like to thank Sami Alpanda, Antonio Diez de los Rios, Scott Hendry, Jonathan Witmer, Jun Yang, Bo Young Chang and seminar participants at the Bank of Canada for helpful comments and useful conversations. All remaining errors and omissions are our own.

Abstract

This paper investigates the effects of monetary policy on the risk-taking behavior of fixed-income mutual funds in Canada. We consider different measures of the stance of monetary policy and investigate *active* variation in mutual funds' risk exposure in response to monetary policy. We find evidence in support of a systematic link between monetary conditions and intertemporal variation in the risk-taking behavior of mutual funds. Specifically, following an expansionary monetary shift, funds actively increase default-risk exposure (i.e., search-for-yield). This is particularly evident in the post-crisis period where interest rates were kept low for a prolonged period of time.

JEL classification: G23, E52
Bank classification: Financial institutions; Transmission of monetary policy

Résumé

Les auteurs étudient l'incidence de la politique monétaire sur le comportement de prise de risques affiché au Canada par les gestionnaires de fonds communs de placement à revenu fixe. Pour ce faire, ils emploient différentes mesures de l'orientation de la politique monétaire et analysent les changements que ces gestionnaires apportent *activement* au niveau d'exposition au risque de leurs fonds, en fonction de la politique monétaire mise en œuvre. Les résultats qu'ils obtiennent plaident pour l'existence d'un lien de nature systématique entre les conditions monétaires et la variation dans le temps du comportement de prise de risques adopté par les gestionnaires de fonds. En l'occurrence, lorsque la politique monétaire se fait expansionniste, ces derniers accroissent activement le niveau d'exposition au risque de défaillance des fonds qu'ils gèrent (c'est-à-dire qu'ils se tournent vers des placements plus risqués en vue d'obtenir de meilleurs rendements). Ce phénomène est particulièrement manifeste en ce qui concerne la période qui a suivi la crise récente, dans laquelle les taux d'intérêt ont été maintenus à de bas niveaux pendant longtemps.

Classification JEL : G23, E52
Classification de la Banque : Institutions financières; Transmission de la politique monétaire

1 Introduction

The historically low interest rates following the financial crisis of 2007-2009 pushed the yields on debt of major advanced economies to unprecedented lows. Such low yields spurred investors to search for investment opportunities that offered some extra return. The result was an increased appetite for risky assets, called *search-for-yield* (Rajan (2006), Borio and Zhu (2012)). Most of the available evidence on increased risk taking by financial intermediaries as a response to low interest rates has concentrated on how commercial banks relax their lending standards when policy rates are low. However, as originally noted by Rajan (2006) and later stressed by Stein (2013), the search-for-yield phenomenon encompasses not just commercial banks, but also other financial intermediaries involved in the pricing of credit. In this paper, we contribute to the literature by investigating whether another important type of financial intermediary in credit markets, fixed-income mutual funds, alters its risk exposure in response to a change in the level of short-term interest rates.

Fixed-income mutual funds are key intermediaries in credit markets. As part of their portfolios, they hold a variety of public and private debt issues, ranging from provincial bonds to asset-backed and mortgage-backed securities.[1] Mutual funds in general are also an important part of retirement portfolios of Canadian households,[2] and therefore changes in their portfolio value can affect consumption and labor supply via wealth effects. Although (long-term) fixed-income mutual funds do not usually figure in discussions of the shadow banking system, such as money market mutual funds, in reality they provide funding to shadow banks by purchasing securitization-related bonds (Pozsar et al. (2012)), and they often provide liquidity in the repo market.[3] Therefore, a change in their risk appetite can have an impact on the availability of credit, both overnight and term. Finally, mutual funds' assets are supported by *demandable equity*, which means that investors have the right to redeem their shares upon demand and, depending on the portfolio composition of the fund, clustered redemptions could lead to fire sales of assets (Stein (2013)). For all of these reasons, a study of how these managed portfolios change their exposure to risk is a useful step toward a more comprehensive understanding of how the actions of central banks affect the economy.

This paper relates to two strands of the literature on the risk-taking behavior of financial

[1]Based on holdings data as of 31 March 2013, Canadian-domiciled fixed-income mutual funds held 7.18% of the outstanding Canada Mortgage Bonds; 7.39% of outstanding credit card asset-backed securities (ABS); 23.13% of auto ABS; 3.88% of Government of Canada bonds; 7 and 4%, respectively, of provincial and municipal bonds; and 15.98% of outstanding corporate debt.

[2]The Investment Funds Institute of Canada estimates that "mutual funds and mutual fund wraps account for 30% of Canadians' financial wealth."

[3]Based on holdings data as of 31 March 2013, fixed-income mutual funds in Canada held CAD 1.03 billion in overnight cash/repurchase agreements.

intermediaries. The first strand has studied the so-called risk-taking channel of monetary policy in the commercial banking sector.[4] These studies show that the risk tolerance of banks increases when risk-free rates remain low for extended periods of time, manifesting itself as a change in a bank's portfolio composition from less-risky to more-risky borrowers (Paligorova and Sierra (2012)).

The second strand has studied mutual fund risk taking as a function of the compensation incentives faced by fund managers (see, for example, Sirri and Tufano (1998), Brown et al. (1996), Chevalier and Ellison (1997), Ferreira et al. (2012)). The central empirical finding on the determinants of individual mutual fund flows is that these flows are, on average, convex in past performance, meaning that funds that outperform their benchmarks in a given period experience significant inflows in the next period, while funds that underperform do not suffer an equally proportional outflow. This empirical finding, together with the usual practice of compensating fund managers with a fixed percentage of assets under management, implies that managers can increase their compensation if they outperform their benchmarks, while not suffering commensurate cuts in their earnings if they instead underperform. This paper contributes to both strands of the literature by studying how fixed-income mutual funds change their exposure to risk as a function of an aggregate variable, the level of safe interest rates. A similar analysis can be extended to study risk taking in equity funds.

Aggregate forces can be important determinants of fund flows, and their interaction with the compensation incentives can influence managers' portfolio decisions. Specifically, in times of adverse economic conditions and greater uncertainties, fixed-income funds benefit from flight-to-safety due to the perceived safety of bonds.[5] On the other hand, it is exactly at these adverse times that central banks conduct expansionary policies to lower rates and provide liquidity. Higher inflows combined with lower yields can push funds to take increasing risks in their search for higher yields to deliver a steady stream of income and maintain/increase assets under management. Since mutual fund managers' compensation depends on inflows, they may have an incentive to tilt their portfolios toward riskier but higher-yielding assets in order to generate such extra income. Furthermore, mutual funds have yield-oriented clientele, such as retirees or pensioners, who finance part of their living expenses with the income they receive from their investments. As low interest rates on safe assets decrease this source of

[4]For a non-exhaustive list, see Rajan (2006), Borio and Zhu (2012), Jimenez et al. (2008), Gambacorta (2009), Ioannidou et al. (2010), Adrian and Shin (2010), Paligorova and Santos (2012).

[5]Morningstar's 2013 Fund Research publication states that, since 2009, investors poured more than CAD 56.7 billion into bond funds, while the outflow in equity funds reached CAD 45.9 billion over the same period. The same trend was observed worldwide, where only in 2012 the inflow into bond mutual funds reached \$535.2 billion (the equivalent of 95% of all net inflows into mutual funds that year), and \$124.7 billion flowed out of equity funds.

income, any investment vehicle that offers comparatively higher interest income becomes extremely valuable and thus can be expected to attract inflows, giving a further incentive to managers to search-for-yield. Due to their incentive structures as well as their prominent role in credit markets, fixed-income mutual funds are a natural place to look for search-for-yield behavior.

Our empirical approach employs a two-step estimation procedure. In the first step, we estimate the time-varying risk exposures (betas) of individual funds through rolling regressions of a two-factor model, where the two systematic risk factors are interest rate risk (TERM) and default risk (DEF), as in Fama and French (1993). Similarly, we estimate the same model for the benchmark portfolio[6] to obtain its time-varying betas as a measure of passive changes in the risk exposures. Second, using the fund beta estimates from the first step as dependent variables, we test whether funds' intertemporal *active* risk exposures (fund risk exposures in excess of the passive risk exposure of the benchmark portfolio) vary with the stance of monetary policy. Our definitions of active management and passive management follow Cremers and Petajisto (2009): passive management of a portfolio is simply replicating the return on an index by buying and holding all, or almost all, assets within an index at the official proportions; any deviation from passive management is considered to be active management. Given the widely known empirical fact that the risk exposure of passive portfolios can also vary over time (Ang and Kristensen (2012)), we isolate a manager's intentional changes in the portfolio as a response to interest rates by disentangling the 'active' and 'passive' risk components. Our approach resembles the methodology employed by Ferson and Schadt (1996), which accommodates time-varying risk exposures by allowing factor loadings to change on a month-to-month basis as functions of observable conditioning variables. We abstain from the conditional model to distinguish active from passive risk taking.

For an equally-weighted portfolio of fixed-income funds over the January 2000 - January 2013 period, we show that a decline in the interest rate is associated with a statistically and economically significant increase in the "active" component of the systematic default-risk exposure, whereas the active component of interest rate risk exposure does not show a clear relation to monetary conditions. Moreover, an analysis of the pre-crisis and post-crisis subperiods reveals that the higher default-risk exposure in response to monetary policy is especially evident in the post-crisis subperiod. This is not surprising given that interest rates were kept at historically low levels for an extended period of time in the aftermath of the

[6] All mutual funds declare a benchmark against which their performance is evaluated. Among several different benchmarks that are used by Canadian fixed-income funds, the DEX Universe Bond Index is the most popular. The DEX Index is a broad measure of the Canadian investment-grade fixed-income market. The data provider PC-Bond Analytics indicates that, as of 31 December 2010, the Universe Index consisted of 1,103 securities, with a total market value of approximately $1.031 trillion.

2007-2009 financial crisis, and spurred investors to search-for-yield. For an average fund, the default-risk exposure in the post-crisis period, relative to its unconditional average over the same period, increases by up to 50% as a response to a one standard deviation decrease in the risk-free rate.

As Roll (1977) and Ang et al. (2010) point out, portfolio groupings can result in a loss of information about the cross-sectional behavior of the constituents. To avoid this problem and utilize as much information as possible, we also conduct our analysis using individual funds as test assets. We find that the individual fund results agree with those from the portfolio analysis, confirming that most funds increase systematic default-risk exposure in response to a decline in the short-term interest rates.

Overall, our empirical evidence supports the contention that recognizing the relevance of monetary policy is important for understanding the risk-taking incentives of financial intermediaries. Our findings suggest that, for fixed-income funds in Canada, the decision of managers to invest in assets with higher exposure to systematic default risk (which pay higher yields) is conditional on the prevailing monetary policy stance. These findings represent a step forward in understanding the different links between central bank actions and the incentives for financial intermediaries.

The rest of the paper proceeds as follows. In section 2, we discuss the methodology and provide details on the dynamic asset-pricing model, as well as the dynamics of the risk-exposure structure of mutual funds. Section 3 describes the data and the variables used in the analysis. Section 4 performs a subperiod analysis for an equally-weighted portfolio of all funds, which represents an average fund. It also reports our empirical findings at the individual fund level. Finally, section 5 concludes.

2 Methodology

This section explains the methodology employed to evaluate mutual fund active risk-taking behavior in response to monetary conditions. Before going into further detail, one might ask whether there is any evidence of time-varying risk in mutual fund returns. Using the volatility of the average return as an initial measure of aggregate risk, Figure 1 shows that the risk of fixed-income mutual funds varies considerably over time.[7] This figure makes no attempt to link the source of volatility variation to monetary policy. However, there appears to be an especially large spike in the fund risk following the collapse of Lehman Brothers. Indeed, mutual funds exhibited low risk after the onset of the crisis, until Lehman Brothers filed for

[7]The volatility of the average return is computed using monthly data over 24-month rolling windows.

bankruptcy in September 2008. Following the Lehman bankruptcy, the increased volatility in the financial markets and the stream of economic events within and outside Canada may have contributed to the increased return volatility in the sample of funds considered. For instance, in the United States, the Federal Reserve started slashing the federal funds rate and the discount rate to provide liquidity to financial markets. By October 2008, the fed funds rate and discount rate were lowered to 1% and 1.75%, respectively. Also in October 2008, the Bank of Canada started to reduce its target for the overnight rate, reaching 0.25% in April 2009, at which time the Bank announced a conditional commitment to keep the main policy rate as low as 0.25% for over a year. The preliminary evidence indicates that risk in fixed-income mutual fund returns is not constant over time. Its association with monetary conditions will be explored in the following sections.

2.1 Multifactor model

The empirical analysis focuses on excess returns on fixed-income mutual funds. Our choice of asset-pricing model follows Fama and French (1993) (FF), who show that two variables, the returns on portfolios exposed to default and interest rate risk, dominate the common variation in government and corporate passive bond portfolio returns. The FF model identifies two standard risk factors used in studies of corporate bond returns (e.g., Gebhardt et al. (2005), Lin et al. (2011)), allows for an intuitive and parsimonious description of bond returns, and provides a reasonable description of the returns on our sample of mutual funds (see section 4.1).

We assume that the excess returns on fixed-income funds are generated by the following two-factor model:

$$R_{p,t} = \alpha_{p,t} + \beta_{p,t}^{def} DEF_t + \beta_{p,t}^{term} TERM_t + \varepsilon_{p,t}, \tag{1}$$

where $R_{p,t}$ is the excess return of fund p in month t. DEF, the proxy for default risk, is defined as the difference between the return on a value-weighted portfolio of long-term corporate bonds and the long-term government bond return. TERM, the proxy for interest rate risk, is the difference between the returns of a portfolio of long-term government bonds and the 1-month Treasury bill rate. $\alpha_{p,t}$ is the pricing error; $\beta_{p,t}^{def}$ and $\beta_{p,t}^{term}$ measure the fund risk exposure, which may vary over time; and $\varepsilon_{p,t}$ is a zero mean error term uncorrelated with the risk factors.

If funds dynamically adjust their portfolio holdings in response to changes in the economy, then estimates from a constant coefficient model will generally be systematically biased (Ferson and Schadt (1996)). To account for the time variation in risk exposure, we estimate the dynamic model in (1) for each fund with rolling estimation windows, which yields time-varying exposure estimates. The analysis is conducted using a 24-month rolling window on monthly data over the January 2000 - January 2013 period, allowing us to capture the conditional commitment period of the Bank of Canada over April 2009 - May 2010. During this period, the overnight rate was committed to be kept at 0.25%. It was gradually increased afterwards to reach 1% in September 2010.

Figures 2a and 2b show the average DEF and TERM risk exposures of all funds in our sample and the risk exposure of a passive portfolio, namely the DEX Universe Bond Index (a broad measure of the Canadian investment-grade fixed-income market). The vertical lines indicate the periods of high and low interest rates. We identify January 2002 - December 2005 and July 2008 - January 2013 as a low interest rate environment and January 2006 - June 2008 as a high interest rate environment. Both figures display large variations in the interest rate and default-risk exposures over time, suggesting that a dynamic model is required. Moreover, they reveal that funds' default and interest rate risk exposures increase substantially during periods of low rates, while the trend is reversed when the rate is high. Although for the majority of the sample period the trends of the systematic risk exposure of an average fund and the DEX index are analogous, occasional sharp deviations are observable. For example, the default-risk exposure of fixed-income mutual funds displays an upward deviation from that of the DEX index during the low interest rate environment of July 2008 - January 2013, indicating an increase in the default-risk tolerance of an average fixed-income mutual fund.

2.2 Active and passive risk exposures

In principle, variation over time in mutual fund risk-exposures $\beta_{p,t}^{def}$ and $\beta_{p,t}^{term}$ can derive from two sources: an *active* change due to the fund manager's actions; and a *passive* change that can be due to variation in the risk exposures of underlying assets (Ang and Kristensen (2012)), or in the relative weights that each asset has in the portfolio. As a stated objective in most fund prospectuses, managers attempt to outperform a self-declared benchmark. They can do this by either buying underpriced bonds or loading on systematic factors that they think will outperform general market movements. However, they also do not want to significantly underperform their benchmark, because this can lead to outflows and/or termination (Chevalier and Ellison (1999)). The only way a manager can outperform a benchmark is by holding a portfolio that is different than the benchmark, but at the same time this raises the probability of underperformance and termination, so in practice some managers end up holding a portfolio that is similar to the benchmark and deviate only slightly from it. We follow Cremers and Petajisto (2009) and define passive management as a strategy of holding all the securities in the benchmark index in the official proportions, and define active management as any deviation from passive management.

The main objective of this paper is to test whether fund managers *actively* expose themselves to more risk when interest rates are low, because this would provide evidence that the actions of the central bank have an effect on managers' attitudes toward risk, manifested as a change in the risk exposures of the portfolios they manage. However, since variation over time in betas can derive from factors unrelated to managers' actions, we control for the passive component of variation over time in mutual fund betas, and test whether the remaining variation is related to monetary policy.

Let the weight of security $k = 1, ..., K$ in portfolio p be denoted as $\omega^p_{k,t-1}$, where the subscript $t-1$ indicates that portfolio weights are assumed to depend on lagged information variables, denoted by Z_{t-1}, such that $\omega^p_{k,t-1} = F(Z_{t-1})$. Also, let the excess return on individual security k be denoted as $R_{k,t}$. Then, the excess return of portfolio p, $R_{p,t}$, can be written as $R_{p,t} = \sum_{k=1}^{K} \omega^p_{k,t-1} R_{k,t}$. From the definition of portfolio return, and given the linearity of the covariance operator, portfolio p's beta or risk exposure to systematic risk factor f_t can be written as

$$\beta^f_{p,t} = \sum_{k=1}^{K} \omega^p_{k,t-1} \beta^f_{k,t}. \tag{2}$$

Equation (2) states that any portfolio's beta is a weighted average of individual security betas. If we denote as $\omega^b_{k,t-1}$ the weight of security k in a benchmark portfolio b, then it follows that the risk exposure of the benchmark index b can likewise be written as

$$\beta^f_{b,t} = \sum_{k=1}^{K} \omega^b_{k,t-1} \beta^f_{k,t}. \tag{3}$$

Then, equations (2) and (3) imply that portfolio p's risk exposure can be decomposed into two parts, as follows:

$$\beta^f_{p,t} = \beta^f_{b,t} + \sum_{k=1}^{K} (\omega^p_{k,t-1} - \omega^b_{k,t-1}) \beta^f_{k,t}. \tag{4}$$

The first term on the right-hand side of equation (4) is the beta of the passive benchmark index and, as such, its variation over time is unrelated to active management. The second term represents the potential deviation from passive management that a manager might choose to undertake.

A few implications of equation (4) are worth noting. For an index fund[8] or a fund that does not deviate from the benchmark, a regression of $\beta^f_{p,t}$ on $\beta^f_{b,t}$ should have an estimated intercept close to zero, a slope coefficient close to 1 and an R^2 statistic close to 1, and additional regressors should have no explanatory power. In addition, if, for some or all securities, the weights are different than those of the benchmark, but not time-varying (i.e., $\omega^p_{k,t-1} - \omega^b_{k,t-1} = \Delta\bar{\omega}_k \neq 0$), then the manager has chosen to actively change the composition of the portfolio, and any time variation in the difference $\beta^f_{p,t} - \beta^f_{b,t}$ cannot be attributed to a dynamic trading strategy, but rather to time variation in $\beta^f_{k,t}$. In this last case, if the time variation in $\beta^f_{k,t}$ is correlated with an information variable Z_{t-1}, then a regression of $\beta^f_{p,t}$ on $\beta^f_{b,t}$ and Z_{t-1} might find a significant coefficient on Z_{t-1}, while the manager is clearly not dynamically changing the portfolio as a function of Z_{t-1}.

Taking into account these possibilities, in the next section we explain our empirical strategy to study active changes in risk exposure as a function of the stance of monetary policy.

2.3 Empirical model

To test whether the active component of mutual fund risk exposure is related to the aggregate variables Z_{t-1}, we specify $\beta^{def}_{p,t}$ and $\beta^{term}_{p,t}$ as linear functions of the risk exposure of a passive benchmark portfolio and observable macroeconomic variables, as follows:

$$\beta^{def}_{p,t} = \beta^{def}_p + \delta^{def}_p \beta^{def}_{b,t} + \gamma^{def}_p Z_{t-1} + e_{p,t}., \tag{5}$$

$$\beta^{term}_{p,t} = \beta^{term}_p + \delta^{term}_p \beta^{term}_{b,t} + \gamma^{term}_p Z_{t-1} + e_{p,t}. \tag{6}$$

We estimate (5) and (6) using ordinary least squares (OLS). Z_{t-1} is a vector of observable variables known at time $t-1$, and the time-varying risk exposure of a passive index portfolio ($\beta^f_{b,t}$) is estimated using a rolling window on monthly data. Note that the constant beta model is nested in the above specification where $\delta_p = \gamma_p = 0$, or $\gamma_p = 0$ and $\beta^f_{b,t} = \beta^f_b$.

We recognize that the special case of a non-zero but time-invariant $\Delta\bar{\omega}_k$ alluded to in the discussion of equation (4) is a potential drawback to our approach to detecting active risk-taking behavior. However, in such a special case, the OLS coefficient on Z_{t-1} represents the change in $\beta^f_{p,t}$ to be brought about by a change in Z_{t-1} that is *uncorrelated* with $\beta^f_{b,t}$ but correlated with the individual security risk-exposure $\beta^f_{k,t}$. For example, if Z_{t-1} is the overnight rate, then this would mean that there is a component of the overnight rate that is somehow not correlated with the risk exposure of the representative benchmark index

[8]Index funds are funds that replicate as closely as possible the return of a passive index.

$\beta^f_{b,t}$, but correlated with the risk exposure of such an individual security. Since we only use macro variables in Z_{t-1}, we find this possibility remote: we would expect that the effect of an aggregate variable on the risk exposure of a class of securities to a given risk factor (say, the duration risk of AAA-rated corporate bonds) should be captured by the risk exposure of a diversified, value-weighted, passive portfolio such as the benchmark. However, in order to determine exactly which percentage of the variation in $\beta^f_{p,t}$ derives from the difference $\omega^p_{k,t-1} - \omega^b_{k,t-1}$, and which from the individual security beta variation $\beta^f_{k,t}$, data on security-level holdings are needed. We leave this task for future work, but note that, using data on hedge fund holdings, Patton and Ramadorai (2013) provide a similar decomposition and show that the dominant force behind time-varying risk exposure is indeed portfolio weight variation.

2.3.1 Information variables

The vector of observable information variables $Z_{t-1} = [mp, ts, ds]'_{t-1}$ consists of a monetary policy variable (mp), which is the central variable of interest, and two control variables: term spread (ts) and default spread (ds). In our empirical tests, we employ four different proxies for the stance of monetary policy: (i) residuals from an interest rate policy rule, (ii) ex-post real interest rate, (iii) level of short rate, and (iv) first principal component of a cross-section of government bond yields (see section 3.3 for details).

Note that the term and default spreads are different than the term and default-risk factors employed in equation (1). The term spread captures the slope of the yield curve measured as the yield spread between long-term and short-term government bonds. The default spread captures the state of business conditions and is measured as the yield spread between BBB- and AA-rated long-term corporate bonds. The evidence suggests that the term spread predicts economic activity (see, for example, Estrella and Hardouvelis (1991) and Fama and Bliss (1987)). Moreover, Fama and French (1989) identify both term spread and default spread as good predictors for the excess returns on passive portfolios of bonds and stocks.

3 Data Description

3.1 Mutual funds

The mutual fund data at the share class level are obtained from Morningstar, Inc. The data consist of monthly investor returns (net-of-fees), assets under management, latest declared fund benchmark and category[9] for all fixed-income mutual funds domiciled in Canada for the period January 2000 to January 2013 (157 months). The initial sample includes 1,441 share classes in 408 funds.

We apply a series of filters to arrive at our final sample. We exclude share classes that are denominated in U.S. dollars or that have an international/global orientation,[10] that attempt to replicate the performance of a particular benchmark as their investment objective (index funds), or that specialize in high-yield debt. We also remove funds that have less-than-continuous 157 monthly return observations, in order to have the longest possible time series to estimate the rolling risk exposures with some precision. Finally, we restrict the sample to funds for which a self-declared benchmark identity is available.[11] Although most funds report a benchmark, we concentrate on funds that declare the DEX Universe Bond Index as their benchmark, because the DEX is the most widely used and broadest benchmark to measure performance in the Canadian fixed-income market,[12] and it can be considered the "market portfolio" of investment-grade bonds in Canada. After these filters, our final sample consists of 41 funds with complete return data, where fund-level returns are the equal-weighted returns across share classes.[13] Blake et al. (1993) use a similar-sized sample to study U.S. bond mutual fund performance during the 1979-1988 period.

Note that our sample may suffer from survivorship bias because we include only funds that have complete return histories over the 2000-2013 period. Potentially, this may cause an upward bias in the distribution of pricing errors. However, since our main focus is not to measure performance but rather to analyze intertemporal variation in risk exposures, this is less of a concern.

Panel A of Table 1 reports the coverage of our sample within the universe of fixed-income mutual funds. Although the number of funds and total assets under management vary over time, our sample of 41 funds encompasses, on average, 38% of the dollar value of assets under management in Canadian-domiciled fixed-income funds in the Morningstar database, and about 26% of the number of funds, with a higher coverage in the beginning of the sample. The summary statistics in Panel B show that the average fund in our sample has CAD 453 million in total net assets and a 0.23% return in excess of the risk-free rate.

---

[9]The categories are: Canadian Fixed Income, Inflation-Protected, Long-Term Fixed Income, Short-Term Fixed Income, Global Fixed Income and High Yield Fixed Income.

[10]These investments would be affected by the monetary policies of other countries.

[11]For 12.76% of the fund-month observations, benchmark information is missing.

[12]In our sample, 46% of the fund-month observations record the DEX Universe as the benchmark. For more information on the DEX indices, see http://www.canadianbondindices.com/Debt_Market_indices.asp.

[13]We use this approach instead of value-weighting the returns because some share classes (including the share class with the highest total net assets) contain missing net assets in some months.

3.2 Systematic risk factors

The return data on the DEX Universe Bond Index are provided by PC-Bond Analytics via DataStream. The systematic risk factors TERM and DEF are constructed using long-term government and corporate total return indices (TRI), also from PC-Bond Analytics via DataStream, and the 1-month T-bill rate from the Bank of Canada. At the end of each month, the corresponding TRI returns are computed as the natural logarithm of the change in the value of the index. The TERM factor returns are then computed as the difference between the return on the long-term government bond index and the lagged 1-month T-bill rate. To obtain the DEF factor, the return of the long-term government index is subtracted from the long-term corporate TRI. Finally, the excess return of the DEX Index is computed by subtracting the lagged 1-month T-bill rate.

3.3 Monetary policy stance and control variables

Given the uncertainty about the "correct" empirical measure of the stance of monetary policy (Bernanke and Mihov (1998)), we estimate equations (5) and (6) using four alternative proxies. Of these, two are relative monetary policy indicators and the remaining two are absolute indicators. The relative indicators are the residual from an interest rate policy rule and the ex-post real interest rate. The policy rule residual is obtained from an OLS regression of the 3-month T-bill rate on 12-month core CPI inflation and a measure of the output gap. The ex-post real rate is obtained by subtracting contemporaneous realized 12-month inflation from the lagged 3-month T-bill rate, similar to Mishkin (1981). The two absolute proxies of the level of interest rates are the 1-month T-bill rate and the first principal component from a cross-section of government bond yields. The first principal component is obtained from a cross-section that includes the 1-, 3- and 6-month Treasury bill rates, and the 1-, 2-, 3-, 5-, 7- and 10-year yields.

Most of our relative proxies for the stance of monetary policy have been used in previous studies of the risk-taking channel of monetary policy, such as Gambacorta (2009), Bekaert et al. (2013), Jimenez et al. (2008) and Ioannidou et al. (2010). We use the 1-month T-bill rate instead of the overnight rate because the former tracks the level of the latter, but displays more variability, which is important for our tests that try to measure covariation between interest rates and risk exposures. We consider the first principal component because it measures the general level of interest rates and represents the best proxy for the variable that causes changes in portfolio returns via the duration-risk channel.

Figure 3 shows the four monetary policy indicators applied in the empirical analysis. Although there are significant differences in levels, the figure reveals a close relationship in their intertemporal fluctuations. The rates unanimously started to decline toward the end of 2007 and the downward trend accelerated around the collapse of Lehman Brothers in September 2008. In the aftermath of the Lehman collapse, nominal rates were kept low, where the 1-month Treasury bill rate ranged around 0.25% and 1%, and ex-post real rates went down to negative levels.

The data for the control variables, the term and default spreads, are from the Bank of Canada and PC-Bond Analytics via DataStream. The term spread is computed as the difference between the yield of a 10-year government bond minus the yield of a 2-year government bond; the default spread is the difference between the redemption yield of a value-weighted portfolio of medium-term BBB-rated corporate bonds minus the redemption yield of a value-weighted portfolio of medium-term AA-rated corporate bonds.[14] The summary statistics for these variables are reported in Panel C of Table 1.

4 Results

This section reports our main results in three parts. First, in section 4.1, we provide evidence on the validity of the two-factor pricing model by analyzing the pricing errors and conclude that (1) adequately describes fund returns. Section 4.2 reports the results from estimating equations (5) and (6) for an equally-weighted portfolio of 41 funds. Finally, section 4.3 describes the relation between fund risk exposures and monetary policy, using individual funds as the test assets.

4.1 Fund-level pricing errors

Table 2 shows the cross-sectional distribution of pricing errors (alphas) and beta estimates along with their t-statistics from estimation of an unconditional version of equation (1), where the model parameters are constant over time. As can be seen, the average loadings on TERM and DEF factors are positive and significantly different than zero, indicating that funds in our sample take interest rate and default risk to generate excess returns.

The results in the following sections are based on the two-factor pricing model (1), hence it is desirable to test the validity of this model before moving forward. A linear factor pricing model can be evaluated by testing the model-imposed constraints that the corresponding pricing errors should be zero. A correct pricing model captures the relation between the expected risk premium on individual funds and systematic risk, resulting in zero pricing

[14]We use medium-term portfolios instead of long-term portfolios because the medium-term maturity segment (between 5 and 10 years of time-to-maturity) includes the largest share of volume outstanding. See http://www.canadianbondindices.com/Debt_Market_indices.asp.

errors. However, our mutual fund return data are net of management fees, while the factors do not include such costs. As a result, mutual fund risk-adjusted returns tend to be negative rather than zero (Fama and French (2010)). The distribution of alpha estimates in Table 2 indicates that only three funds (7.3% of the sample) have significantly positive pricing errors. The average alpha across all funds is 0.01, with an average t-statistic of 0.29, and the two-factor model explains 82% of the variation in excess returns of a typical fund.

We interpret the evidence in Table 2 as consistent with the notion that the two-factor model provides a reasonable description of the average risk-return relationship for fixed-income funds in our sample. In the next sections, we investigate how funds' risk exposures vary over time as a response to variation in interest rates, assuming (1) is an adequate description of the main risks that funds take in order to generate abnormal returns.

4.2 Portfolio-level analysis

In this section we investigate how an equally-weighted portfolio of funds changes its active risk exposures as a function of monetary policy indicators. An equally-weighted portfolio of funds can be considered a representative fund and can thus help to concisely illustrate the average changes in risk exposures of the funds in our sample.

4.2.1 Unconditional risk exposures

Table 3 reports parameter estimates from an unconditional version of the two-factor model (1) for an equally-weighted portfolio of all funds in our sample. We observe that, over the whole sample, the average fund has higher loadings on the TERM factor than on the DEF factor. Over the pre-crisis period, the average fund seems to load more on TERM risk when compared to the whole sample, but it is not significantly exposed to DEF risk. However, in the post-crisis sample, the average fund decreased its exposure to TERM risk, when compared to the pre-crisis sample, and in turn increased its exposure to higher-yielding securities, as evidenced by the higher loading on the DEF factor. This finding suggests that the average fund increased its portfolio allocation to higher-yielding bonds precisely at a time when the central bank kept the level of interest rates at historical lows. Thus, the subsample results show that the average fund modified its risk exposure over time when macroeconomic conditions changed. In the next subsection, we analyze whether this modification is intentional and how it relates to monetary policy.

4.2.2 Active component of risk exposures

In this section, we study how the active risk exposure of an equally-weighted portfolio of funds changes over time as a function of the stance of monetary policy. As explained in section 2.3, we do this in two steps. First, we estimate by OLS the risk exposures of the equally-weighted portfolio and the benchmark index using 24-month rolling windows. Then, in the second step, for each risk factor, we regress the portfolio's exposure on the benchmark exposure, monetary policy proxy and control variables. Tables 4 and 5 report the results from estimating equations (5) and (6), respectively. In both tables, Panel A provides results for the whole sample, while panels B and C provide the estimates from the pre- and post-crisis periods, respectively.

Table 4 analyzes how the average fund actively adjusts its default-risk exposure. It shows that, for the whole sample, the average fixed-income fund increased its exposure to default risk when the stance of monetary policy was more accommodative; that is, when interest rates went down. As can be seen from the first row of Panel A, the coefficient on the policy variable is negative across all proxies of the policy stance. This finding has the interpretation that the average fund takes on more default risk, that is, holds more bonds that pay higher yields, when the level of interest rates decreases. In addition, Panels B and C reveal that the negative response to policy is stronger in the post-crisis period, since all of the coefficients in that period are negative and strongly statistically significant, while for the pre-crisis period, the stance of monetary policy has no statistically significant relation to risk exposures for two of the proxies (the Taylor-rule residual and the ex-post real rate), and even records a positive coefficient for the remaining two proxies (the 1-month T-bill rate and the first principal component). Therefore, the subsample results reveal that it was precisely in the period of historically low interest rates, the post-crisis period, when funds actively increased their exposure to default risk.

Table 5 reports the results for interest rate risk exposure. Here, a different conclusion emerges. Over the whole sample, the average fund appears to decrease its active exposure to interest rate risk when rates go down, although the statistical significance of the relationship is less strong than in the case of default-risk exposure. In contrast to the case of default risk, the sensitivity of the active component of interest rate risk exposure to policy rates seems to be fairly stable across subperiods. However, the important difference is that, whereas in the pre-crisis period, across all monetary policy proxies there was an average sensitivity of around 0.68 to the benchmark's risk exposure (second row, Panel B), in the post-crisis period such sensitivity increased to about 1.03 (second row, Panel C). This suggests that in the post-crisis period, the average fund might have aligned the duration of its portfolio more closely with that of the benchmark to avoid suffering capital losses greater than those of the benchmark when rates start to increase.[15] This is intuitive given the desire of managers not to underperform their benchmark, and given the fact that the interest rate risk factor (TERM) has a higher volatility than the default-risk factor (see Table 1).

In summary, the results of this section suggest that the average fund actively increases its exposure to default risk when the level of interest rates goes down, while decreasing its exposure to interest rate risk. While these results inform about the direction of the effect, they are silent on their economic significance. We gauge the economic significance of the estimated changes in risk exposures due to a change in the stance of policy in section 4.3.2. In the next section, we first verify that the relatively stronger impact of policy on the active component of default-risk exposures is present at the fund level.

4.3 Fund-level analysis

In this section, we evaluate the effect of monetary policy on the active component of risk exposures using individual funds instead of portfolios. The rationale for doing this is that, as Roll (1977) and Ang et al. (2010) point out, creating portfolios ignores important information. Specifically, funds with different alphas and betas can cancel out within a portfolio, leading to an incorrect inference about the significance of the model coefficients. To investigate whether the effects we document at the portfolio level are still present at the fund level, we use individual funds as test assets to investigate active changes in risk exposures. We follow the same two-step procedure as in section 4.2.2, first obtaining rolling risk exposures for each fund, and then regressing these on the benchmark's risk exposure, a monetary policy indicator, and control variables.

4.3.1 Active component of DEF and TERM risk exposures

Table 6 summarizes the cross-sectional distributions of coefficients and t-statistics resulting from estimation of equation (5) at the fund level, which relates individual fund DEF risk exposures to benchmark index default-risk exposures, a policy stance variable and two additional control variables. The model is estimated for each of the proxies of monetary policy explained in section 3.3. On average, the estimated coefficients are negative and the tails of the t-statistics distributions are thicker than normal. The heteroskedasticity and serial-correlation robust generalized method of moments (GMM) test for joint significance strongly rejects the hypothesis that funds do not actively change their risk exposures when policy rates vary. Across all proxies for the policy stance, at least 27 out of 41 funds have a negative

[15]This observation is probably another dimension of active management. However, we do not pursue this idea further in this paper.

coefficient on the policy variable, and for the case of the policy rule residual (the column labelled "taylor"), 29 funds, or 70% of the sample, record a negative sign on the policy variable. We find that the default-risk exposure of an average fund is 0.01 to 0.02 higher when the policy stance is more accommodative. Economically, these numbers correspond to 1.5% higher default betas when interest rates decrease by 10 basis points.[16]

In terms of statistical significance, we find that across all proxies for policy, at least 21 funds had a statistically significant negative coefficient, with 23 funds for the case of the ex-post real rate. These numbers suggest that approximately 50% of the funds in our sample had a statistically significant negative response of their active default-risk exposure to increases in policy rates. On the other hand, across all proxies for the stance of policy, out of the (at most) 14 funds that *increased* their active exposure to default risk when rates increased, only 7 of them (17% of the sample) had a statistically significant positive coefficient. We interpret this fund-level evidence as being consistent with the notion that, on average, most funds increased their active exposure to default risk when the level of safe interest rates declined.

A somewhat different picture emerges when we analyze active TERM (or interest rate risk) exposures in Table 7. The average estimated coefficients on monetary policy proxies are positive, with thicker-than-normal tails of t-statistics distributions. The GMM test for joint significance rejects the hypothesis that funds do not actively change their risk exposures when policy rates vary, which is due to the fact that across all policy proxies, at least 20 funds had a statistically significant non-zero (positive or negative) coefficient on the policy stance variable. In general, the pattern of signs of coefficients is reversed when compared to the case of DEF exposures, since now the majority of funds have a positive coefficient on the policy stance variables. The main difference with that case, however, is that out of the coefficients that are positive, across all measures of policy no more than 16 funds (39% of the sample) have a statistically significant coefficient. Thus, the results suggest that although the most common action is to decrease active exposure to interest rate risk when policy rates are low, the effect is somewhat less strong when compared to the case of default-risk exposure. It is important to emphasize here that the effects we study are above and beyond the impact that the term or default spreads have on interest rate or default-risk exposures, since those variables are also included in the regressions as control variables.

To summarize, the individual fund results suggest that while the level of policy rates appears to influence active interest rate risk exposures in a moderate way, most of the individual funds significantly increase their exposure to default risk when interest rates decrease.

[16]The point estimate for the average DEF risk exposure of funds, from Table 6, is 0.11. This suggests that a 10 basis point decrease in the 1-month T-bill rate would translate into an (-0.016)*(-0.1)/0.11 = 0.0145 or 1.45% higher DEF risk exposure.

This confirms the findings at the portfolio level in section 4.2.2.

4.3.2 Economic significance of changes in risk exposures

The above portfolio- and fund-level results provide evidence that a more accommodative stance of monetary policy is statistically significantly associated with an increase in default-risk exposures. While statistical significance is an important criterion for evaluating the strength of the relationship between two variables, economic importance should not be overlooked. To gauge the economic significance of the association between the fund risk exposures and monetary conditions, Table 8 shows the relative change in the exposure to each risk factor implied by the parameter estimates in Tables 3, 4 and 5. The table reports the percentage change in betas, with respect to the average exposure, when the policy stance variable increases by one standard deviation. In Panel A, we find that although interest rate and default-risk exposures change in opposite directions, the percentage change in default-risk exposure is much higher. Indeed, on average across all stances of policy, default-risk exposure increases approximately 19% when interest rates decrease, while interest rate risk decreases about 1.67%. Importantly, the subsample estimates in Panels B and C reveal that it is especially in the post-crisis period when the increase in default-risk exposure becomes more acute. Overall, the results suggest that the changes in default-risk exposure brought about by low interest rates are non-trivial, and point to an increase in portfolio risk following periods of low interest rates, given the magnitudes of the sample volatilities of the DEF and TERM factors.

5 Conclusions

This paper provides evidence that a subset of the universe of fixed-income mutual funds in Canada increases their exposure to default risk when the level of safe interest rates decreases. We use a linear factor pricing model to derive estimates of the time-varying risk exposures of our sample of funds, as well as their self-declared benchmark index. Then, we linearly relate the funds' risk exposures to the benchmark index exposure and to macroeconomic variables. We find that for most funds, the component of the fund's exposure to default risk that is not correlated with the benchmark index exposure, which we refer to as the active component, increases when the level of interest rates declines. On the other hand, we find no strong relationship between a fund's active exposure to interest rate risk and the level of short-term interest rates. In addition, when we similarly study the active risk exposure of an equally-weighted portfolio of all funds, we find statistically significant evidence that the active component of default-risk exposure increases when the stance of monetary policy is accommodative. Importantly, subsample analysis reveals that this effect derives mainly from the post-crisis period of historically low interest rates. In terms of economic significance, the increase in default-risk exposure brought about by a one standard deviation decrease in the policy rate is on average 19%, when compared to the full-sample average exposure estimates. On the other hand, we find that the active exposure to interest rate risk decreases when interest rates are lowered, but the effect is economically insignificant and close to zero. We interpret our findings as evidence that the stance of monetary policy can influence the risk-taking behavior of non-bank financial intermediaries.

References

Adrian, T. and H. Shin (2010). Liquidity and leverage. *Journal of Financial Intermediation 19*, 418–437.

Ang, A. and D. Kristensen (2012). Testing conditional factor models. *Journal of Financial Economics 106*(1), 132 – 156.

Ang, A., J. Liu, and K. Schwarz (2010). Using stocks or portfolios in tests of factor models. Working Paper, Columbia University.

Bekaert, G., M. Hoerova, and M. L. Duca (2013). Risk, uncertainty and monetary policy. *Journal of Monetary Economics 60*(7), 771 – 788.

Bernanke, B. S. and I. Mihov (1998). Measuring monetary policy. *The Quarterly Journal of Economics 113*(3), 869–902.

Blake, C. R., E. J. Elton, and M. J. Gruber (1993). The performance of bond mutual funds. *The Journal of Business 66*(3), pp. 371–403.

Borio, C. and H. Zhu (2012). Capital regulation, risk-taking and monetary policy: A missing link in the transmission mechanism? *Journal of Financial Stability 8*(4), 236 – 251.

Brown, K. C., W. V. Harlow, and L. T. Starks (1996). Of tournaments and temptations: An analysis of managerial incentives in the mutual fund industry. *The Journal of Finance 51*(1), 85–110.

Chevalier, J. and G. Ellison (1997). Risk taking by mutual funds as a response to incentives. *The Journal of Political Economy 105*, 1167–1200.

Chevalier, J. and G. Ellison (1999). Career concerns of mutual fund managers. *The Quarterly Journal of Economics 114*(2), 389–432.

Cremers, K. J. M. and A. Petajisto (2009). How active is your fund manager? A new measure that predicts performance. *Review of Financial Studies 22*(9), 3329–3365.

Estrella, A. and G. Hardouvelis (1991). The term structure as a predictor of real economic activity. *The Journal of Finance 46*, 555–576.

Fama, E. and R. Bliss (1987). The information in long-maturity forward rates. *American Economic Review 77*, 680–692.

Fama, E. and K. French (1989). Business conditions and expected returns on stocks and bonds. *Journal of Financial Economics 25*, 23–49.

Fama, E. and K. French (1993). Common risk factors in the returns on stocks and bonds. *Journal of Financial Economics 33*, 3–56.

Fama, E. and K. French (2010). Luck versus skill in the cross-section of mutual fund returns. *The Journal of Finance 65*(5), 1915–1947.

Ferreira, M. A., A. Keswani, A. F. Miguel, and S. B. Ramos (2012). The flow-performance relationship around the world. *Journal of Banking and Finance 36*(6), 1759 – 1780.

Ferson, W. and R. Schadt (1996). Measuring fund strategy and performance in changing economic conditions. *The Journal of Finance 51*, 425–461.

Gambacorta, L. (2009). Monetary policy and the risk-taking channel. *BIS Quarterly Review (December)*, 43 – 53.

Gebhardt, W. R., S. Hvidkjaer, and B. Swaminathan (2005). The cross-section of expected corporate bond returns: Betas or characteristics? *Journal of Financial Economics 75*(1), 85 – 114.

Ioannidou, V., S. Ongena, and J. Peydró (2010). Monetary policy, risk-taking and pricing: Evidence from a quasi-natural experiment. CentER Discussion Paper Series No. 2009-31S.

Jimenez, G., S. Ongena, J. Peydró, and J. Saurina (2008). Hazardous times for monetary policy: What do twenty-three million bank loans say about the effects of monetary policy on credit risk-taking? Banco de España Working Paper No. 833.

Lin, H., J. Wang, and C. Wu (2011). Liquidity risk and expected corporate bond returns. *Journal of Financial Economics 99*(3), 628 – 650.

Mishkin, F. (1981). The real interest rate: An empirical investigation. *Carnegie-Rochester Conference Series on Public Policy 15*, 151–200.

Newey, W. and K. West (1987). A simple, positive semidefinite, heteroskedasticity and autocorrelation consistent covariance matrix. *Econometrica 55*, 703–708.

Paligorova, T. and J. Santos (2012). When is it less costly for risky firms to borrow? Evidence from the bank risk-taking channel of monetary policy. Bank of Canada Working Paper No. 2012-10.

Paligorova, T. and J. Sierra (2012). Monetary policy and the risk-taking channel: Insights from the lending behaviour of banks. *Bank of Canada Review* (Autumn), 23–30.

Patton, A. and T. Ramadorai (2013). On the high-frequency dynamics of hedge fund risk exposures. *The Journal of Finance 2*, 597–635.

Pozsar, Z., T. Adrian, A. Ashcraft, and H. Boesky (2012). Shadow banking. *FRBNY Staff Report No. 458*.

Rajan, R. G. (2006). Has finance made the world riskier? *European Financial Management 12*(4), 499–533.

Roll, R. (1977). A critique of the asset pricing theory tests; part i: On past and potential testability of the theory. *Journal of Financial Economics 4*, 129–176.

Sirri, E. R. and P. Tufano (1998). Costly search and mutual fund flows. *The Journal of Finance 53*(5), 1589–1622.

Stein, J. (2013). Overheating in credit markets: Origins, measurement, and policy responses. In *Restoring Household Financial Stability after the Great Recession: Why Household Balance Sheets Matter*. Remarks.

Table 1: **Descriptive statistics**

Panel A: Mutual fund sample coverage

	2000	2005	2010	2012	average 2000-2012
Assets under management (CAD billion)					
universe	16.0	32.1	62.5	92.1	
sample	5.7	13.6	19.0	33.1	
coverage (%)	36%	42%	30%	36%	38%
Number of funds					
universe	110	158	224	274	
sample	41	41	41	41	
coverage (%)	37%	26%	18%	15%	26%

Panel B: Descriptive statistics of mutual fund sample

	mean	stdev	min	max
Total net assets (CAD million)	453	655	0.23	4330
Return	0.43	0.95	-3.43	3.99
Excess return	0.23	0.95	-3.61	3.73
β_p^{term}	0.46	0.10	0.08	0.56
β_p^{def}	0.11	0.07	-0.03	0.30
alpha	0.01	0.04	-0.09	0.08

Panel C: Descriptive statistics of policy proxies, pricing factors and conditioning variables

	mean	stdev	min	max
1-month Tbill	2.44	1.54	0.12	5.49
Taylor Residual	-1.44	1.73	-4.47	3.35
Real rate	0.78	1.60	-1.88	4.59
PC1	-0.89	2.83	-5.09	4.89
TERM	0.46	1.90	-4.07	5.76
DEF	0.05	1.20	-5.32	6.90
DEX	0.32	1.01	-2.16	3.00
term spread	1.04	0.79	-0.41	3.05
def spread	0.89	0.46	0.31	2.17

This table presents descriptive statistics for our sample of funds, pricing factors, policy proxies and control variables. Panel A presents the number of funds and assets under management of all fixed-income funds in Canada included in the Morningstar Direct database, together with the same figure for the sample used in the paper. Panel B presents descriptive statistics for the mutual funds included in our sample: total net assets is reported in CAD millions; return is the investor return and the excess return is the investor return in excess of the lagged 1-month T-bill rate; the β_p^{term} and β_p^{def} are the full-sample two-factor model regression coefficients on the pricing factors; and *alpha* is the intercept in the factor model regressions. Panel C presents descriptive statistics for: the pricing factors (TERM and DEF) and benchmark index returns (DEX), which are described in section 3.2; and the monetary policy proxies and control variables (term and default spreads), which are described in section 3.3. The data are sampled at a monthly frequency, from the period January 2000-January 2013.

Table 2: **Cross-sectional distribution of two-factor model coefficient estimates: Individual funds**

	Coefficient estimates			t-statistics			adj. R^2
	alpha	TERM	DEF	alpha	TERM	DEF	
Minimum	-0.09	0.08	-0.03	-3.04	5.32	-1.09	0.21
Average	0.01	0.46	0.10	0.29	23.33	2.95	0.82
Maximum	0.08	0.56	0.30	2.34	31.33	8.11	0.90
No. of funds	41						
No. of fund with < 0 alpha	16						
No. of fund with > 0 alpha	25						
No. of funds with < 0 alpha significant at 5%	3						
No. of funds with > 0 alpha significant at 5%	3						

This table presents the cross-sectional distribution of estimated coefficients and t-statistics for each of the 41 funds from the unconditional two-factor model $R_{p,t} = \alpha_p + \beta_p^{def} DEF_t + \beta_p^{term} TERM_t + \varepsilon_{p,t}$, where the model intercept and the factor loadings are constant over time. The last two rows show the number of funds with alpha estimates significantly different than zero at the 5% level.

Table 3: **Two-factor coefficient estimates: single equally-weighted portfolio**

	alpha	TERM	DEF	N	adj. R^2
All sample: Jan 2002 - Jan 2013					
	-0.000	**0.462**	**0.107**	133	0.88
	(-0.01)	(25.68)	(3.60)		
Pre-crisis: Jan 2002 - Dec 2006					
	-0.069	**0.502**	0.079	60	0.93
	(-1.93)	(27.24)	(1.52)		
Post-crisis: Jan 2007 - Jan 2013					
	0.053	**0.433**	**0.105**	73	0.86
	(1.22)	(17.36)	(3.57)		

This table presents the estimated coefficients and t-statistics from an unconditional version of the two-factor pricing model in (1). The test asset is an equally-weighted portfolio of all funds, and the full sample period covers January 2002 - January 2013 (N = 133 time-series observations). The pre-crisis and post-crisis subperiods are January 2002 - December 2006, and January 2007 - January 2013, respectively. The t-statistics in parentheses are computed using Newey-West standard errors with six lags. The bold entries indicate significance at least at the 5% level.

Table 4: **Explaining β^{def}: single equally-weighted portfolio**

	1-month Tbill Rate	Taylor Residual	Ex-post Real Interest Rate	PC1
Panel A. All sample: Jan 2002 - Jan 2013				
monetary policy$_{t-1}$	**-0.016**	**-0.013**	**-0.017**	**-0.008**
	(-3.50)	(-3.47)	(-3.95)	(-3.59)
β_t^{dex}	1.022	1.005	1.002	1.024
	(18.50)	(20.33)	(19.98)	(18.60)
term$_{t-1}$	-0.013	-0.006	-0.012	-0.004
	(-2.16)	(-0.75)	(-2.14)	(-0.71)
def$_{t-1}$	0.044	0.020	0.029	0.044
	(2.74)	(1.65)	(2.41)	(2.73)
constant	0.053	0.007	0.039	-0.003
	(2.76)	(0.30)	(2.23)	(-0.11)
N	133	133	133	133
Adj. R^2	97.00%	96.72%	97.07%	97.01%
Panel B. Pre-crisis: Jan 20002 - Dec 2006				
monetary policy$_{t-1}$	**0.022**	-0.001	0.008	**0.009**
	(2.67)	(-0.63)	(1.51)	(2.25)
β_t^{dex}	0.891	0.913	0.909	0.888
	(31.50)	(30.34)	(34.99)	(26.94)
term$_{t-1}$	-0.009	-0.027	-0.022	-0.022
	(-1.31)	(-3.95)	(-5.17)	(-4.65)
def$_{t-1}$	-0.001	0.013	0.016	0.000
	(-0.06)	(1.14)	(1.88)	(0.02)
constant	0.007	0.072	0.060	0.084
	(0.29)	(5.42)	(6.27)	(5.25)
N	60	60	60	60
Adj. R^2	99.37%	99.26%	99.28%	99.34%
Panel C. Post-crisis: Jan 2007 - Jan 2013				
monetary policy$_{t-1}$	**-0.041**	**-0.017**	**-0.031**	**-0.019**
	(-10.17)	(-4.56)	(-6.97)	(-9.83)
β_t^{dex}	1.120	0.972	0.970	1.102
	(17.00)	(18.18)	(16.94)	(15.49)
term$_{t-1}$	-0.030	0.000	-0.021	-0.009
	(-5.63)	(0.01)	(-2.71)	(-2.36)
def$_{t-1}$	-0.064	0.043	-0.020	-0.060
	(-2.85)	(1.09)	(-1.47)	(-2.56)
constant	0.171	-0.024	0.077	0.030
	(7.82)	(-1.18)	(4.45)	(2.39)
N	73	73	73	73
Adj. R^2	95.32%	90.94%	93.39%	95.06%

This table presents estimated coefficients and t-statistics of the equally-weighted portfolio from the following regression: $\beta_{p,t}^{def} = \beta_p^{def} + \delta\beta_{b,t}^{def} + \gamma_{p,1}mp_{t-1} + \gamma_{p,2}term_{t-1} + \gamma_{p,3}def_{t-1} + e_{p,t}$, where $\beta_{p,t}^{def}$ is exposure to default risk, $\beta_{b,t}^{def}$ is the benchmark index exposure to default risk, $term_{t-1}$ is the term spread and def_{t-1} is the default spread. Panel A reports the results for the full sample over January 2002 - January 2013. Panels B and C show the estimates from the pre-crisis and post-crisis subperiods, respectively. Four different variables are used to proxy the policy stance: columns 2-5 report the results for the 1-month T-bill rate, the residual from a Taylor-rule OLS regression, the ex-post real interest rate and the first principal component from a cross-section of zero-coupon government rates, respectively. The numbers in parentheses are the t-statistics calculated using Newey and West (1987) standard errors with six lags. The bold entries for the monetary policy proxy indicate significance at least at the 5% level.

Table 5: **Explaining β^{term}: single equally-weighted portfolio**

	1-month Tbill Rate	Taylor Residual	Ex-post Real Interest Rate	PC1
Panel A. All sample: Jan 2002 - Jan 2013				
monetary policy$_{t-1}$	0.006	0.003	**0.006**	0.003
	(1.92)	(1.54)	(2.13)	(1.87)
β_t^{dex}	0.766	0.816	0.786	0.767
	(11.51)	(12.06)	(13.36)	(11.31)
term$_{t-1}$	0.009	0.005	0.008	0.006
	(2.39)	(1.45)	(2.28)	(2.12)
def$_{t-1}$	-0.009	-0.004	-0.006	-0.009
	(-2.48)	(-1.02)	(-1.50)	(-2.44)
constant	0.060	0.052	0.059	0.081
	(1.93)	(1.39)	(1.93)	(2.21)
N	133	133	133	133
Adj. R^2	92.42%	90.83%	92.06%	92.35%
Panel B. Pre-crisis: Jan 2002 - Dec 2006				
monetary policy$_{t-1}$	**0.006**	**-0.003**	**-0.003**	**0.002**
	(3.06)	(-5.15)	(-4.11)	(2.96)
β_t^{dex}	0.736	0.642	0.630	0.711
	(22.79)	(23.18)	(21.46)	(28.45)
term$_{t-1}$	0.005	0.001	0.000	0.002
	(3.70)	(1.05)	(0.14)	(2.18)
def$_{t-1}$	-0.004	-0.008	-0.006	-0.004
	(-2.32)	(-6.44)	(-5.95)	(-2.56)
constant	0.073	0.147	0.160	0.109
	(3.21)	(8.76)	(8.77)	(7.27)
N	60	60	60	60
Adj. R^2	95.55%	96.05%	95.12%	95.28%
Panel C. Post-crisis: Jan 2007 - Jan 2013				
monetary policy$_{t-1}$	**0.010**	0.003	**0.009**	**0.005**
	(4.09)	(1.34)	(4.37)	(4.02)
β_t^{dex}	0.999	1.075	1.026	0.998
	(19.67)	(12.20)	(18.38)	(19.30)
term$_{t-1}$	0.020	0.011	0.018	0.015
	(5.57)	(4.44)	(6.14)	(5.96)
def$_{t-1}$	-0.008	-0.039	-0.009	-0.007
	(-0.67)	(-2.14)	(-0.78)	(-0.51)
constant	-0.069	-0.056	-0.064	-0.035
	(-2.74)	(-1.42)	(-2.40)	(-1.46)
N	73	73	73	73
Adj. R^2	96.22%	94.19%	96.18%	96.34%

This table presents estimated coefficients and t-statistics of the equally-weighted portfolio from the following regression: $\beta_{p,t}^{term} = \beta_p^{term} + \delta\beta_{b,t}^{term} + \gamma_{p,1} mp_{t-1} + \gamma_{p,2} term_{t-1} + \gamma_{p,3} def_{t-1} + e_{p,t}$, where $\beta_{p,t}^{term}$ is the exposure to interest rate risk, $\beta_{b,t}^{term}$ is the benchmark index exposure to default risk, $term_{t-1}$ is the term spread and def_{t-1} is the default spread. Panel A reports the results for the full sample over January 2002 - January 2013. Panels B and C show the estimates from the pre-crisis and post-crisis subperiods, respectively. Four different variables are used to proxy the policy stance: columns 2-5 report the results for the 1-month T-bill rate, the residual from a Taylor-rule OLS regression, the ex-post real interest rate and the first principal component from a cross-section of zero-coupon government rates, respectively. The numbers in parentheses are the t-statistics calculated using Newey and West (1987) standard errors with six lags. The bold entries for the monetary policy proxy indicate significance at least at the 5% level.

Table 6: **The cross-sectional distribution of t-statistics for the monetary policy indicators: Default-risk exposure in individual funds**

	tb1m		taylor		real		pc1	
	coeff	t-stat	coeff	t-stat	coeff	t-stat	coeff	t-stat
Minimum	-0.097	-11.87	-0.058	-8.74	-0.090	-8.65	-0.046	-11.63
Average	-0.016	-2.11	-0.013	-2.02	-0.017	-2.16	-0.008	-2.15
Maximum	0.064	4.25	0.051	3.78	0.063	4.01	0.031	4.39
GMM	21814		2638		10316		24172	
p-value	0.00		0.00		0.00		0.00	
No. and % of funds								
t-stat < -2.58		19		15		18		19
		46.34%		36.59%		43.90%		46.34%
-2.58 < t-stat < -1.96		2		5		2		2
		4.88%		12.20%		4.88%		4.88%
-1.96 < t-stat < -1.65		1		1		3		1
		2.44%		2.44%		7.32%		2.44%
-1.65 < t-stat < 0		5		9		5		5
		12.20%		21.95%		12.20%		12.20%
0 < t-stat < 1.65		7		8		7		8
		17.07%		19.51%		17.07%		19.51%
1.65 < t-stat < 1.96		2		1		1		1
		4.88%		2.44%		2.44%		2.44%
1.96 < t-stat < 2.58		0		0		2		0
		0.00%		0.00%		4.88%		0.00%
2.58 < t-stat		5		2		3		5
		12.20%		4.88%		7.32%		12.20%
Total no. of funds		41		41		41		41
No. of significantly < 0 funds		22		21		23		22
No. of significantly > 0 funds		7		3		6		6

This table presents the cross-sectional distribution of fund-level estimated coefficients and t-statistics on the monetary policy indicator mp_{t-1} from the following regression: $\beta^{def}_{p,t} = \beta^{def}_{p} + \delta\beta^{def}_{b,t} + \gamma_{p,1}mp_{t-1} + \gamma_{p,2}term_{t-1} + \gamma_{p,3}def_{t-1} + e_{p,t}$, where $\beta^{def}_{p,t}$ is fund p's exposure to default risk, $\beta^{def}_{b,t}$ is the benchmark index exposure to default risk, $term_{t-1}$ is the term spread and def_{t-1} is the default spread. It also shows the GMM test statistics and bootstrapped p-values for the null hypothesis that the coefficients on the policy variable are jointly equal to zero across all funds. Under the null hypothesis, the policy indicator does not explain time variation in the active component of mutual fund risk exposures. Four different variables are used to proxy the policy stance: in columns 2-3, the policy measure employed is the 1-month T-bill rate; in columns 4-5, the residual from a Taylor-rule OLS regression; in columns 6-7, the ex-post real interest rate; and in columns 8-9, the first principal component from a cross-section of zero-coupon government rates. Under the null hypothesis, the GMM test statistic is distributed as a chi-square random variable with N degrees of freedom. For the individual fund regressions, the t-statistics are computed using Newey and West (1987) standard errors with six lags.

Table 7: **The cross-sectional distribution of t-statistics for the monetary policy indicators: Interest rate risk exposure in individual funds**

	tb1m		taylor		real		pc1	
	coeff	t-stat	coeff	t-stat	coeff	t-stat	coeff	t-stat
Minimum	-0.020	-5.65	-0.016	-3.43	-0.019	-4.06	-0.010	-6.50
Average	0.006	0.88	0.003	0.60	0.005	0.98	0.003	0.81
Maximum	0.047	8.69	0.027	4.31	0.042	6.52	0.023	9.20
GMM	56844		3687		15512		47352	
p-value	0.00		0.00		0.00		0.00	
No. and % of funds								
t-stat < -2.58		6		2		5		7
		14.63%		4.88%		12.20%		17.07%
-2.58 < t-stat < -1.96		4		2		1		3
		9.76%		4.88%		2.44%		7.32%
-1.96 < t-stat < -1.65		0		1		1		0
		0.00%		2.44%		2.44%		0.00%
-1.65 < t-stat < 0		6		9		9		6
		14.63%		21.95%		21.95%		14.63%
0 < t-stat < 1.65		8		11		8		8
		19.51%		26.83%		19.51%		19.51%
1.65 < t-stat < 1.96		2		3		3		1
		4.88%		7.32%		7.32%		2.44%
1.96 < t-stat < 2.58		3		4		2		4
		7.32%		9.76%		4.88%		9.76%
2.58 < t-stat		11		8		11		11
		26.83%		19.51%		26.83%		26.83%
Total no. of funds		40		40		40		40
No. of significantly < 0 funds		10		5		7		10
No. of significantly > 0 funds		16		15		16		16

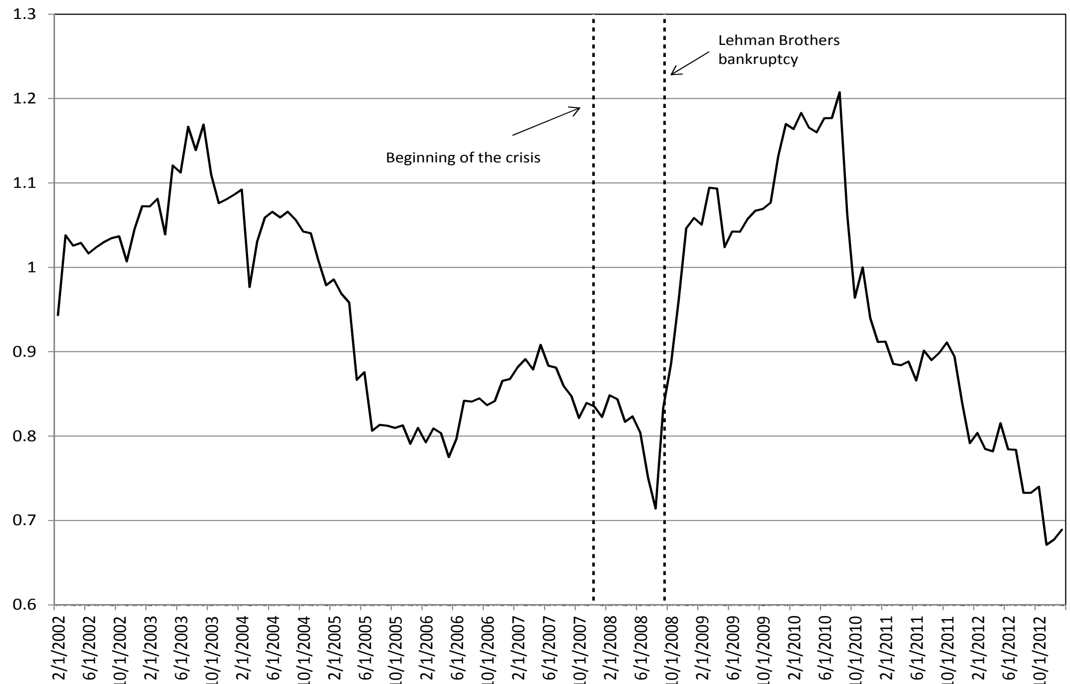
This table presents the cross-sectional distribution of fund-level estimated coefficients and t-statistics on the monetary policy indicator mp_{t-1} from the following regression: $\beta_{p,t}^{term} = \beta_{p}^{term} + \delta\beta_{b,t}^{term} + \gamma_{p,1}mp_{t-1} + \gamma_{p,2}term_{t-1} + \gamma_{p,3}def_{t-1} + e_{p,t}$, where $\beta_{p,t}^{term}$ is fund p's exposure to interest rate level risk, $\beta_{b,t}^{term}$ is the benchmark index exposure to interest rate level risk, $term_{t-1}$ is the term spread and def_{t-1} is the default spread. It also shows the GMM test statistics and bootstrapped p-values for the null hypothesis that the coefficients on the policy variable are jointly equal to zero across all funds. Under the null hypothesis, the policy indicator does not explain time variation in the active component of mutual fund risk exposures. Four different variables are used to proxy the policy stance: in columns 2-3, the policy measure employed is the 1-month T-bill rate; in columns 4-5, the residual from a Taylor-rule OLS regression; in columns 6-7, the ex-post real interest rate; and in columns 8-9, the first principal component from a cross-section of zero-coupon government rates. Under the null hypothesis, the GMM test statistic is distributed as a chi-square random variable with N degrees of freedom. For the individual fund regressions, the t-statistics are computed using Newey and West (1987) standard errors with six lags.

Table 8: **Economic significance: single equally-weighted portfolio**

	1-month Tbill Rate	Taylor Residual	Real Interest Rate	PC1
Panel A. All sample: Jan 2002 - Jan 2013				
Std. Dev.	1.28	1.30	1.23	2.26
$\Delta\beta^{def}$	-23.20%	-11.83%	-22.38%	-20.57%
$\Delta\beta^{term}$	2.77%	0.85%	1.61%	1.48%
Panel B. Pre-crisis: Jan 2002 - Dec 2006				
Std. Dev.	0.66	1.08	0.70	0.83
$\Delta\beta^{def}$	18.10%	-1.35%	7.04%	9.39%
$\Delta\beta^{term}$	1.32%	-0.65%	-0.42%	0.33%
Panel C. Post-crisis: Jan 2007 - Jan 2013				
Std. Dev.	1.38	1.35	1.39	2.29
$\Delta\beta^{def}$	-50.18%	-24.58%	-37.99%	-41.56%
$\Delta\beta^{term}$	3.21%	0.94%	2.92%	2.66%

This table shows the percentage change in β^{term} and β^{def} due to one standard deviation (Std. Dev.) change in the monetary policy indicator.

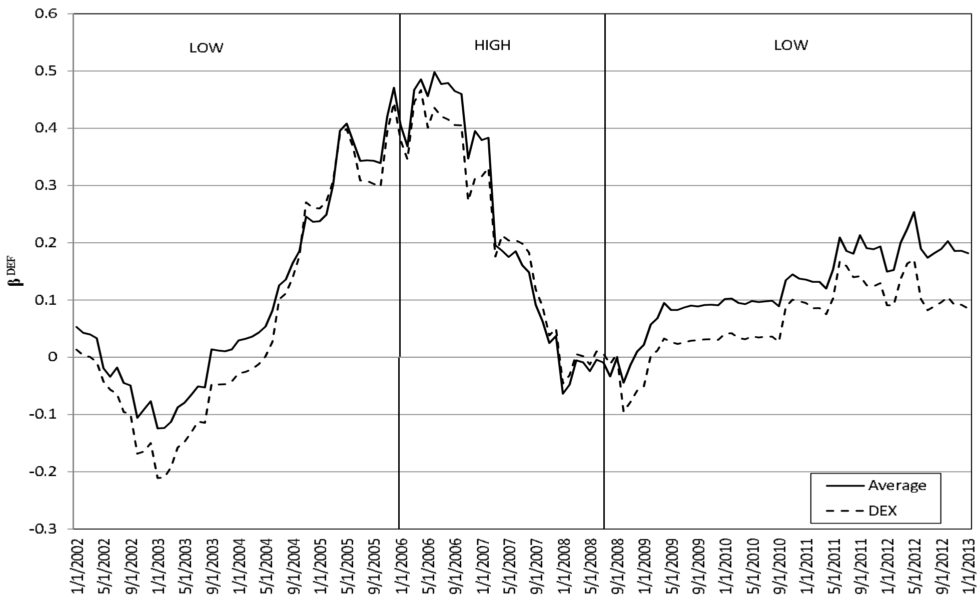
Figure 1: Average Fund Return Volatility



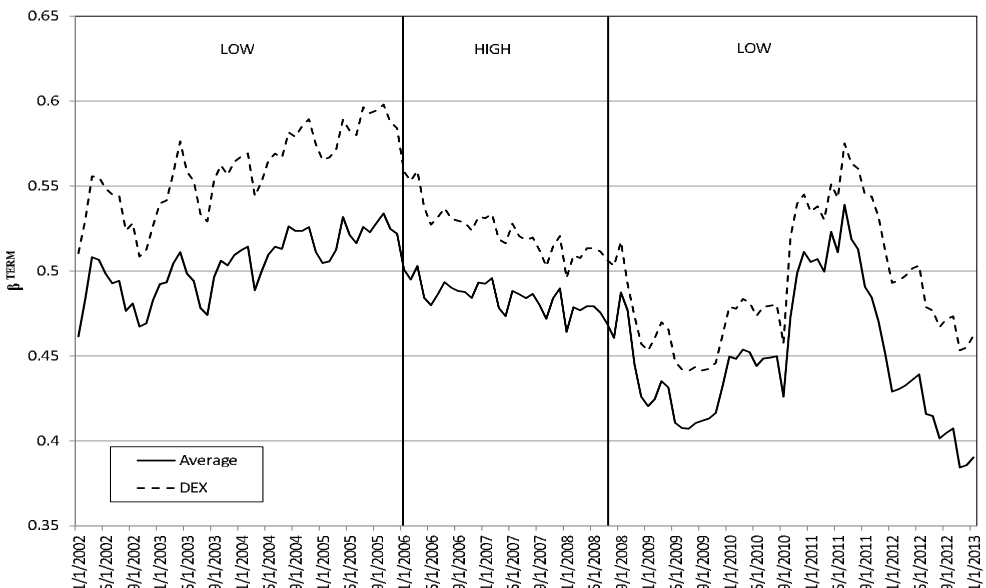
This figure shows the average volatility of fund returns in the sample. The beginning of the crisis in December 2007 and the collapse of Lehman Brothers in September 2008 are marked by dashed lines.

Figure 2: Time-varying risk exposures

(a) DEF Exposures

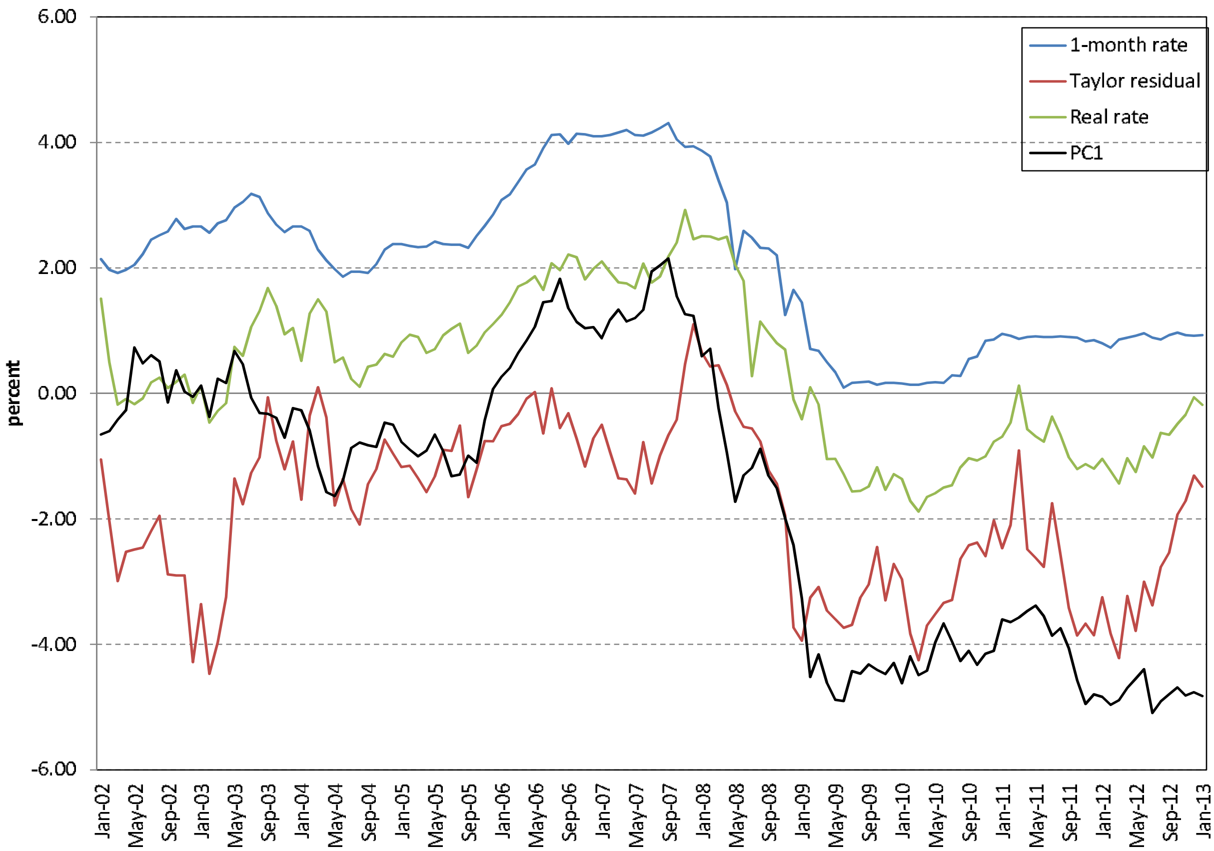


(b) TERM Exposures



This figure shows the rolling betas for TERM and DEF factors using a 24-month rolling window over the January 2002 - January 2013 period. The solid line indicates the fluctuations in the average risk exposure of all funds in the sample. The dashed line shows the variation in the risk exposure of the DEX Universe Bond Index, which is a broad measure of the Canadian investment-grade fixed-income market, used as a benchmark to evaluate the performance of all the funds in our sample. The LOW and HIGH interest rate periods are separated by vertical lines, where LOW rate periods are January 2002 to December 2005 and July 2008 - January 2013. The HIGH rate period is January 2006 to June 2008.

Figure 3: Monetary Policy Proxies



This figure shows the employed monetary policy proxies: 1-month T-bill rate, real rate, Taylor rule residual, and the first principal component of a cross-section of yields, which are described in section 3.3.