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The rise of shadow banking and the hidden benefits of diversification

L'essor des activités bancaires non traditionnelles et les bénéfices cachés de la diversification

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The rise of shadow banking and the hidden benefits of diversification

Abstract

The diversification benefits associated with banks off-balance-sheet activities (OBS), and particularly non-traditional activities, is a question much debated in the literature. These activities, related to the emergence of shadow banking, greatly contribute to the volatility of bank operating revenues, but their impact on accounting returns is less clear (Stiroh and Rumble 2006). In this paper, we use a Canadian dataset to revisit the risk-return trade-off associated with banks OBS activities and study the evolution of the endogeneity of banks decision to expand their market-oriented business lines. Consistent with the changing mix of noninterest income OBS activities generate, we identify a structural break in 1997 which coincides with an increased impact of endogeneity on banks returns, and which also leads to an increased return on assets (*ROA*) and a surge in banking risk. We trace the sources of the greater volatility of noninterest income to a tighter co-integrating relationship between noninterest income and stock market indices after 1997. Introducing a new, robust estimation method based on a modification of the Hausman procedure, we find that neglecting endogeneity greatly underestimates the positive impact of shadow banking on bank accounting returns, even when the subprime crisis is considered. Our main results suggest that the influence of market-based activities on the risk-return trade-off might be larger than what was previously thought.

JEL classification: C32; G20; G21.
Keywords: Noninterest income; Hausman test; Structural break; Shadow Banking; Endogeneity: Macro-prudential analysis.

L’essor des activités bancaires non traditionnelles et les bénéfices cachés de la diversification

Résumé

L’ampleur des gains de diversification associés aux activités bancaires hors-bilan (OBS), et particulièrement aux activités non traditionnelles, est une question fort débattue dans la littérature. Ces activités, qui sont issues de l’émergence du système bancaire parallèle (*shadow banking*), contribuent fortement à la volatilité des revenus d’opération des banques, mais leur impact sur les rendements comptables est moins clairement établi (Stiroh and Rumble 2006). Dans cet article, nous réexaminons l’arbitrage rendement-risque relié aux activités OBS de manière à étudier l’évolution de l’endogénéité de la décision des banques de développer leurs opérations orientées vers les marchés financiers. En conformité avec le changement dans la composition des revenus autres que d’intérêt que les activités OBS occasionnent, nous identifions un changement structurel en 1997 qui coïncide avec l’impact accru de l’endogénéité sur les rendements bancaires, et qui se traduit également par une augmentation du rendement sur les actifs *(ROA)* et un gonflement de la volatilité des revenus. Nous associons ce sursaut de volatilité au chapitre des revenus autres que d’intérêt à une relation de cointégration plus étroite entre les revenus autres que d’intérêt et les indices boursiers après 1997. En nous basant sur une nouvelle version robuste de la régression artificielle d’Hausman, nous trouvons que la méthode des moindres carrés ordinaires, qui fait abstraction de l’endogénéité, donne lieu à une sous-estimation importante de l’impact positif du *shadow banking* sur les rendements bancaires, et ce même après la prise en compte de la crise 2007-2009. Nos principaux résultats suggèrent que l’influence, sur les rendements, des activités bancaires centrées sur les marchés financiers pourrait être beaucoup plus importante qu’on ne le pensait jusqu’ici.

Classification JEL: C32; G20; G21.
Mots-clefs: Revenu autre que d’intérêt; test d’Hausman ; Changement structurel; *Shadow banking*; Endogénéité ; Analyse macroprudentielle.

1. Introduction

Banks off-balance-sheet (OBS) activities, and in particular securitization, have fuelled the last lending boom, enabling banks to increase their operational funding. This eventually led to a standard liquidity crisis driven by maturity mismatch (Farhi and Tirole 2009, Gorton and Metrick 2009). At the core of the problem is the recent change in the banking landscape, which now, thanks to deregulation, comprises the whole leveraged financial system, including market based banking[1]. This new type of banking, which Adrian and Shin (2009, 2010) call shadow banking, presents a considerable challenge to central banks and regulators. In the context of this new banking era, it becomes crucial for central banks to fully understand the behaviour of OBS activities. What we know so far is that the increase in banks non-traditional activities has had a significant influence on banks risk-return trade-off (DeYoung and Roland 2001, Estrella 2001, Acharya 2002, Clark and Siems 2002, Stiroh 2004, Stiroh and Rumble 2006, Baele et al. 2007, De Jonghe 2009, Nijskens and Wagner 2011). International evidence suggests that it triggered a substantial increase in the volatility of banks net operating revenue growth, without a clear increase in returns (Stiroh 2004, Baele et al. 2007, De Jonghe 2009, Calmès and Liu 2009, Nijskens and Wagner 2011). The Canadian experience also suggests that the contribution to banks income of the revenues generated from market-oriented activities, i.e. noninterest income, rapidly became a key, procyclical determinant of banks profits after 1997 (Calmès and Théoret 2010). However, to our knowledge, the literature does not provide any rigorous evidence about the *evolution* of the endogeneity of the share of noninterest income (*snonin*) during the transition period the banking business

[1] See Shin (2009) for a detailed explanation.

underwent. The aim of this paper is to fill this gap and check whether the change that occurred in the banking system, namely the rise of shadow banking, is associated with a change in the degree of OBS endogeneity. Indeed, if banks non-traditional activities are better integrated to the banking business, we should expect more regularity in the relationship between *snonin* and *ROA* . To study this question, we introduce a new approach based on a modified version of the Hausman test specifically designed to gauge the changes in the endogeneity of banks decision to expand their market-based activities.

Another motivation for adopting this alternative approach comes from the fact that treating endogeneity too casually can leave spurious correlations between *snonin* and unobservables not accounted for in banks returns equations (Stiroh 2004). In particular, the remaining non-orthogonality of *snonin* with the innovation in the returns equations can cause serious biases in the parameters estimates and yield misleading results. To treat endogeneity, Stiroh and Rumble (2006), and Baele et al. (2007) introduce fixed effects or lagged explanatory variables in their panel regressions. Some authors also introduce various control variables and other techniques to deal with this issue (Fluck and Lynch 1999, Chevalier 2000, Lamont and Polk 2001, Maksimovic and Phillips 2002, Graham et al. 2002, Villalonga 2004, Laeven and Levine 2007). However, this kind of approach does not completely alleviate the problem. In particular, it is generally not suited to investigate the changes in the relative contribution of noninterest income to banks profits. To our knowledge, the study of Goddard et al. (2008) is the only one using conventional instrumental variables to treat the endogeneity of *snonin*. We depart from their method however, introducing a new technique, an *h* test based on an artificial regression equivalent to a two-stage least squares (TSLS) procedure, but modified to gauge endogeneity biases. The key advantage of this procedure is that it provides a direct measure of the changing biases in the endogenous variable coefficient associated with noninterest income, while also delivering robust instruments built with the higher moments of the explanatory variables (Fuller 1987, Lewbel 1997, Racicot and Théoret 2008, Meng et al. 2011).

We apply this framework to a Canadian dataset to study the emergence of shadow banking and assess its positive impact on the aggregate banking risk-return trade-off over the whole sample, which runs from the first fiscal quarter of 1988 to the second fiscal quarter of 2010[2]. Consistent with the findings of European studies (Baele et al. 2007, Lepetit et al. 2008, Busch and Kick 2009) we detect an improvement of the risk-return trade-off, OBS activities leading to greater returns on assets and equity after 1997. More importantly, we find that this change is associated with a structural break which coincides with a sharp increase in the volatility of banks net operating revenues growth and in the ratio of noninterest income. As evidenced by financial flows accounts, the magnitude of banks financial flows jumps after 1997, providing evidence that banks were entering into a new risk regime.

The year 1997 is a plausible break since it is around this date that banks modified the mix of their OBS activities in Canada, giving a much greater weight to their market-based operations, like trading and capital markets operations, which boosted the volatility of their operating income. Accompanying this volatility spike, our main results suggest that endogeneity, which was a minor concern before 1997, *increases* substantially thereafter, a fact generally overlooked in the literature. What we find is that the link between

[2] Note that the involvement of Canadian banks in OBS activities was quite restricted before 1987, banks being not allowed to get involved in investment banking until this date. For example, before 1987, Canadian banks reported very low commissions.

ROA and *snonin* becomes stronger and much more significant after the structural break, and that this result is greatly reinforced when properly accounting for endogeneity. Compared to TSLS, ordinary least-squares (OLS) estimations substantially understate the sensitivity of *ROA* to *snonin* during the second subperiod, 1997-2010. Overall, the new evidence we gather suggests a marked increase in the endogeneity of noninterest income, which strongly supports the idea that the bank regime shift might be deeper than previously thought.

This paper is organized as follows. In the next section, we present the data and some basic stylized facts about the behaviour of noninterest income. Then, in section 3, we expose the banks returns model and the modified Hausman method we introduce to monitor the change in the endogeneity of noninterest income. The fourth section details our results and various robustness checks, and the last section concludes with some straightforward policy implications.

2. The change in noninterest income

2.1. The data

The sample we use runs from the first fiscal quarter of 1988 to the second fiscal quarter of 2010. In total, we consider eight banks, and quarterly data for about twenty two years, so that, aggregating, we have around ninety observations, a number of observations reasonable to perform standard time series regressions.

In the study, we use aggregate data of the whole Canadian banking system. Data come from the Canadian Bankers Association, the Office of the Superintendent of Financial Institutions, the Bank of Canada and CANSIM. The sample comprises the eight major Canadian domestic banks, which, taken together, account for 90 percent of the Canadian banking business. All of them are chartered banks, i.e. commercial banks regulated by the Canadian Bank Act, running a broad range of activities, from loan business to brokerage, investment dealing, fiduciary services, insurance and securitization.

Given the high degree of concentration of the Canadian banking sector, the banks are generally well funded, with extremely low probability of bankruptcy. Considering the small number of banks in the sample, we obviously need to focus our analysis on aggregate data in order to get robust regression results. Indeed, with panel data regressions, we would need more observations to ensure reliable findings.

Note that a specificity of the Canadian shadow banking is that it is much concentrated in, and controlled by, the traditional banking sector, and, therefore, not divided between commercial banks and security dealers (e.g., investment banks). In other words, this homogenous dataset offers the key advantage of being easy to work with. Compared to the US or the European banking sectors, the Canadian banking sector might appear quite small to draw any meaningful inference about the emergence of a new banking environment (i.e. shadow banking). However, our methodological choice, based on aggregate time series, comprising 90 observations, and very parsimonious models, is more than enough to derive robust results.

2.2. The evolution of the noninterest income series

Insert Figure 1 about here

Figure 1 presents the evolution of the performance of the Canadian banking system during the whole sample period. First note that banks returns, as measured by the return on assets (*ROA*) and the return on equity (*ROE*), share a very close relationship[3]. The Hodrick-Prescott trends indicate that these two returns measures tend to move upward since the beginning of the 1990s. This movement can be explained by the downward trend in the loan loss provisions ratio, but it might also relate to the better integration of banks traditional and OBS activities. Relatedly, Figure 2 illustrates the growth in the banks noninterest income share (in net operating revenues). By 2000, noninterest income accounted for 57% of net operating revenue, up from only 25% in 1988. This ratio seems to have stabilized thereafter, as the new banking business lines matured. The ratio somewhat increased after the high tech bubble burst, culminating at 60% in the first quarter of 2006, but decreased again thereafter, particularly during the recent credit crisis.

More importantly note that the fluctuations of *snonin* are much larger after 1997. In particular, *snonin* becomes increasingly sensitive to the fluctuations of the financial markets after 1997 (Calmès and Liu 2009). Data actually suggest the presence of a structural break around this date[4]. The post 1997 increased volatility of *snonin* series is also apparent if banks are considered individually[5]. As an illustration, Figure 3 provides a comparison of *snonin* for three Canadian banks differing by size: a relatively small-sized bank, the National Bank of Canada (NBC); a medium-sized bank, the Toronto-Dominion Bank (TD), and the largest Canadian bank, the Royal Bank of Canada (RBC)[6]. Contrary to the RBC noninterest income share, the NBC and especially the TD shares have become very volatile since 1997. The *snonin* of NBC has remained on a volatile upward trend, before collapsing on the fourth quarter of 2007, while the TD share has decreased

[3] Due to the high correlation of *ROE* and *ROA* we only study the behaviour of *ROA* in this paper.
[4] We run a Chow test confirming this structural break. See also Calmès and Théoret (2009). Additional tests follow.
[5] Note however that there is some evidence of the benefit of relying on banks aggregate data for macroprudential analysis. For example, return on equity is among the best indicators of the raise in systemic banking instability (Cihàk and Schaeck 2007).
[6] For the second quarter of 2010, total assets of NBC, TD and RBC amounted to 134 billion $, 567 billion $, and 659 billion $ respectively. Their relative shares in the assets of the pool of the eight domestic banks were 5.0%, 21.5%, and 25.0%. For the fourth quarter of 1996, these respective shares were: 5.5%, 13.5% and 23.9%.

substantially since 2000. The dispersion in banks *snonin* has also greatly increased since 1997, maybe suggesting a less herd like behavior, and perhaps a sign of improvement in the diversification of the banking industry.

Insert Figures 2 and 3 about here

Insert Table 1 about here

Since the volatility of *snonin* contributes to the volatility of operating income, we should consequently expect an increase in Canadian banks net operating income volatility after 1997. Of course, the financial turmoil in the Asian markets and the high-tech bubble can be partly accountable for such fluctuations. But the increasing share of noninterest income is surely another important factor to understand the change in banks net operating income. In this respect, the adoption, in 1997, of the VaR as the standard banks risk measure has likely contributed to the increased income growth volatility because of the tendency of this risk measure to underestimate the negative impact of fat tails. Table 1 provides the decomposition of the variance of net operating income growth over the whole sample period and over the two subperiods 1988-1996 and 1997-2010. On the first subperiod net interest income contributes the most to the variance of net operating income. However after 1997, the rise in the variance of bank net operating income is due, for the most part, to the increased volatility of noninterest income. For instance, from the subperiod 1988-1996 to the subperiod 1997-2010, the variance of net operating income growth increased from 11 to 66.3, and the absolute contribution of noninterest income increased from 3.0 to 63.4. Relatedly, Figure 4 illustrates the behaviour of the variance moving average of net operating income growth and its two components, net interest income and noninterest income, over the period 1983 to 2010. While the volatility of net operating income growth is relatively stable before 1997, it is no longer the case after, as the fluctuations of the variance of the net operating income growth sharply increase.

To further describe where the change is coming from, Table 2 provides the descriptive statistics of the components of Canadian banks noninterest income over the period 1997-2010[7]. Observe that the components which have the highest standard deviations are those related to market-oriented activities, mainly the capital markets and the trading revenues. The average share of these two components is almost 50% over the period 1997-2010. Figure 5 confirms that these two components indeed drive the fluctuations of the variance moving average of noninterest income growth over the period 1997-2010. Relatedly, Table 3 also provides the decomposition of the variance of noninterest income growth over this period. On a total variance of 3016.3, the absolute contribution of the trading income component is as high as 2929.9, which represents a relative contribution of 97% to the total variance, although the relative share of trading income to noninterest income only amounts to 11%. The remaining variance is mainly explained by the capital market income component. In other words, the fluctuations of the noninterest income growth are, to a large extent, explained by the two components related to banks market-oriented activities. Note that the high relative contribution of these two inter-related components to the covariance of noninterest income suggests that the additional diversification benefits that these components could bring might be low. In terms of diversification benefits, *ceteris paribus*, it is securitization and insurance revenues which actually seem to offer the better perspective, with relative contributions to covariance equal respectively to -24.7 and -6.8, and a total contribution to variance equal to -23.4 and -1.7.

[7] Statistics are not available before this date.

Insert Figures 4 and 5 about here

Insert Tables 2 and 3 about here

To conclude on this change in the banking activities mix, note that after 1997 the volatility of *snonin* increases in conjunction with the Canadian stock market index, i.e., S&P/TSX (Figure 6), and with the fluctuations of banks stock trading portfolio (Figure 7). A closer look at Figure 6 actually suggests that there might be a cointegration relationship between *TSX* and *snonin* over the sample period. We run an augmented Dickey-Fuller test which seems to suggest that both *snonin* and *TSX* are two I(1) variables (Table 6), so that they can potentially be cointegrated. Table 4 reports the results of a Johansen's cointegration test between these two variables. When the variables are both expressed in levels, the test indicates a conintegration relationship between the two variables over the period 1988-2010 at the 10% threshold. More importantly the test identifies a tighter cointegration relationship after 1997, while the test fails to reject the hypothesis of no conintegration relationship over the first subperiod (1988-1996). We also perform the test by taking the logarithm of *TSX* and obtain the same kind of result. The growing importance of capital markets and trading income might thus well be related to this tighter cointegration of *snonin* and TSX, which might have contributed to the growth in operating income volatility at that time. Moreover, this tighter cointegration can also partly explained the increased procyclicality observed in the banking sector over the last decade (Calmès and Théoret 2010, Nijskens and Wagner 2011).

Insert Figure 6 and 7 about here

Insert Tables 4 and 5 about here

Naturally, the greater volatility of banks operating income observed after 1997 should be associated with a higher expected *ROA,* as finance theory would predict. However, in practice, the evidence is rather mixed. For example, Stiroh and Rumble (2006) and Calmès and Liu (2009) do not find clear diversification benefits associated with OBS activities, whereas Nijskens and Wagner (2011) finds a positive diversification effect, but associated with an increased systemic risk. Table 5 suggests that the Spearman rank-order correlation between *ROA* and *snonin* seems to be moderately negative for the aggregate of the eight major Canadian banks and our three banks between 1988 and 1996. On the other hand, we find that it becomes strongly positive after 1997 (cf. Figure 8). Relatedly, as banks increase their involvement in OBS activities, their loan loss provisions *(LLP)* decrease, both in level and volatility (Figure 9). This trend might be explained by a new type of banking strategy aiming at transferring bank risk off-balance-sheet (Brunnermeier 2009). Nevertheless, since the *ROE* and *ROA* volatility also increase with noninterest income volatility, the change in banks returns cannot be attributed to *LLP,* at least in the second part of our sample. Overall, these preliminary results may constitute a first set of evidence that banks have likely changed their business model. The next sections are intended to thoroughly investigate the extent to which this is the case.

Insert Figure 8 and 9 about here

3. Empirical Framework

3.1. The banks returns model

The general formulation often used to describe banks performance and noninterest income can be expressed as:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 snonin_t + \mathbf{X}_t \alpha + \varepsilon_t \quad (1)$$

where y_t stands for an accounting measure of bank performance – e.g., the return on equity (*ROE)* or the return on assets (*ROA)* –, $\mathbf{X_t}$ is a vector of control variables, and ε_t is the innovation, or error term. $\mathbf{X_t}$ controls for factors that impact banks performance (e.g. riskiness of loans or spread between the yield on loans and the cost of funds).

In its canonical form, however simple, this model presents a little complication since *snonin* is usually considered endogenous. Based on first principles and accounting identities, the endogeneity of *snonin* seems fairly non-controversial. The decision to diversify in OBS activities *is* endogenous (Campa and Kedia 2002, Baele et al. 2007, Laeven and Levine 2007, De Jonghe 2009). Banks returns on assets (*ROA*) may well be a function of the share of noninterest income (*snonin*), but *snonin* may itself be a function of *ROA* (Goddard et al. 2008). OBS activities could give raise to diversification benefits, which tends to increase *ROA*, and in this case the relation between *ROA* and *snonin* should be positive, but at the same time, a decrease in *ROA* might also induce banks to take more risk by increasing their involvement in OBS activities, and then the relation between *ROA* and *snonin* would be negative. *ROA* and *snonin* are thus two interactive banks decision variables, so that the associated endogeneity can possibly bias the estimation of the sensitivity of *ROA* to *snonin.* To illustrate this issue more precisely, consider the two following simultaneous equations:

$$ROA_t = \alpha_1 snonin_t + \beta_1 z_{1t} + \mu_{1t} \qquad (2)$$

$$snonin_t = \alpha_2 ROA_t + \beta_2 z_{2t} + \mu_{2t} \quad (3)$$

where z_{1t} and z_{2t} are two exogenous variables, and μ_{1t} and μ_{2t} are the innovations.

Equation (2) is a simplified version of the model we use in this article. If OBS activities lead to diversification benefits, then $\alpha_1 > 0$. However, we must account for the counter effect described by equation (3). We can suspect that $\alpha_2 < 0$ since banks would increase *snonin* in reaction to the decrease in *ROA* we usually associate with the decline of traditional banking (Boyd and Gertler 1994)[8]. If we estimate equation (2) by OLS, we are thus confronted with a simultaneity or endogenous bias. Obtaining the direction of the bias for the α_1 coefficient is generally complicated. The asymptotic bias of α_1 is equal to:

$$p\,lim\,\hat{\alpha}_{1,OLS} - \alpha_1 = \frac{cov(snonin, \mu_1)}{var(snonin)} \qquad (4)$$

where $\hat{\alpha}_{1,OLS}$ is the estimation of α_1 obtained by applying OLS to equation (2). According to equation (4), the sign of the bias depends on the covariance between $snonin$ and μ_1. To compute this covariance, we can simplify equation (2) by dropping z_{1t}, making this equation exactly identified. Assume that μ_{1t} and μ_{2t} are uncorrelated, then the covariance between $snonin_t$ and μ_{1t} is:

$$cov(snonin, \mu_1) = \frac{\alpha_2}{1-\alpha_1\alpha_2}\sigma^2_{\mu_1} \qquad (5)$$

In this case the asymptotic bias (or inconsistency) in the OLS estimation of α_1 has the same sign as $\frac{\alpha_2}{1-\alpha_1\alpha_2}$. Consequently, if $\alpha_2 < 0$, and if $\alpha_2\alpha_1 < 1$, the asymptotic bias is negative and the estimation of α_1 is biased downward. This downward bias means that a

[8] Actually, this could have been the main motive for banks to invest in OBS activities (Calmès and Liu 2009).

conventional OLS estimation could underestimate the impact of *snonin* on *ROA*, or, more specifically, the diversification benefits due OBS activities.

The motivation of this study comes from the idea that the endogeneity of *snonin* can lead to a severe underestimation of α_1 in equation (2), i.e. the sensitivity of *ROA* to *snonin*. In the next subsections, we propose a rigorous treatment of this endogeneity issue and detail how to construct the higher moments instruments we use to endogeneize the *snonin* variable.

3.2. Robust higher moments instruments

Fuller (1987) shows how the higher moments of the explanatory variables may be used as instruments. To explain his developments in a simple setting, consider a two variables model such that: $y_t = \alpha + \beta x_t + \varepsilon_t,\ t = 1,2,...,n$, where $\varepsilon \sim N(0,\sigma^2)$, and assume that $E\{x_t\varepsilon_t\} \neq 0$, i.e. x_t, not being orthogonal to ε_t, can be considered endogenous. Assume also that there exists a variable z_t which satisfies the two following conditions, $E\{z_t x_t\} \neq 0$, and $E\{z_t\varepsilon_t\} = 0$. Then z_t may be used as an instrumental variable for x_t. Suppose that the distribution of x_t is not normal but asymmetric and leptokurtic. Since the distribution of x_t is asymmetric, we have $E\left\{(x_t - \mu_x)^3\right\} \neq 0$, with μ_x, the expected value of x. Let us set $z_t = (x_t - \bar{x})^2$, a potential instrumental variable, where $\bar{x}$ stands for the mean value of x. Then $E\{(x_t - \mu_x)(z_t - \mu_z)\} = (1 - n^{-1})E\left\{(x_t - \mu_x)^3\right\} \neq 0$, and in accordance with the properties of the normal distribution: $E\{z_t\varepsilon_t\} = 0$. Thus, the second-order moment $(x_t - \bar{x})^2$ qualifies as an instrumental variable for x_t. By the same token, if the distribution of x_t is leptokurtic, the third-order moment $(x_t - \bar{x})^3$ also qualifies as an instrumental variable. According to

Fuller (1987), the co-moment $(y_t - \overline{y})(x_t - \overline{x})$ and the second-order moment of the dependent variable $(y_t - \overline{y})^2$ may also be used as instruments.

Two key advantages of using these higher-moments instruments is that (i) they are robust in the sense that their correlation with the endogenous variable is high while they are orthogonal to the equation residuals, and (ii) they are based on the variables of the model itself, thus requiring no extraneous information. In the context of our model, resorting to higher moments instruments of this nature delivers a consistent estimator of β_2, the *snonin* coefficient of our model (equation (1)). For the treatment of *snonin* endogeneity, we thus use the following set of instruments:

$$\mathbf{Z} = \left\{ x_{t-1,} \left(x_t - \overline{x}\right)^2, \left(x_t - \overline{x}\right)^3, \left(y_t - \overline{y}\right)^2, \left(y_t - \overline{y}\right)^3 \right\} \qquad (6)$$

where x_t represents any of the explanatory variables of the banks returns model.

3.3. A modified TSLS regression incorporating an Hausman endogeneity test

To test for the endogeneity of *snonin* we do not rely on the standard Hausman (1978) test but rather a transformed version of this test based on an artificial (auxiliary) regression. The standard Hausman test, i.e. the *h* test, is based on the following *h* statistic: $h = \left(\hat{\beta}_{IV} - \hat{\beta}_{OLS}\right)^T \left[Var\left(\hat{\beta}_{IV}\right) - Var\left(\hat{\beta}_{OLS}\right)\right]^{-1} \left(\hat{\beta}_{IV} - \hat{\beta}_{OLS}\right) \sim \chi^2(g)$, where $\hat{\beta}_{OLS}$ is the OLS estimator of the parameters vector; $\hat{\beta}_{IV}$, the corresponding instrumental variable (IV) estimator; $Var\left(\hat{\beta}_{OLS}\right)$ and $Var\left(\hat{\beta}_{IV}\right)$ the respective variances of the estimated parameters, and *g* the number of explanatory variables. The standard Hausman test measures the significance of the distance vector $\left(\hat{\beta}_{IV} - \hat{\beta}_{OLS}\right)$. If the *p*-value of the test is less than 5%, the hypothesis H0 of no-endogeneity is rejected for a confidence level of 95%. However, as noted by McKinnon (1992), when the weighting matrix of the test $\left[Var\left(\hat{\beta}_{IV}\right) - Var\left(\hat{\beta}_{OLS}\right)\right]$ is not positive definite, the *h* test is problematic. Moreover, the standard *h* test does not directly provide coefficients adjusted for endogeneity. To address these issues, we resort to an alternative Hausman test. The modified version of the *h* test we introduce is directly related to the work of Hausman (1978), Spencer and Berk (1981), McKinnon (1992) and Pindyck and Rubinfeld (1998)[9]. To implement this version of the Hausman test, we first rewrite the banks returns model (equation (1)) as:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 snonin_t + \mathbf{X}_t \alpha + \varepsilon_t \quad (7)$$

Since $E\left(snonin_t, \varepsilon_t\right) \neq 0$, *snonin* is an endogenous variable. A consistent estimator can be found if we can identify an instrument data matrix $\mathbf{Z} = \left\{z_1, z_2, ..., z_k\right\}$ – k being the number of instruments – to treat the *snonin* endogeneity. As discussed earlier, in our case, this instrument set is the vector of higher moments **Z** (equation (6)). The higher moments Hausman test is then implemented in two steps. First, using the instrument set **Z,** we compute the fitted value of $snonin_t$, noted $sn\hat{o}nin_t$. Thus we regress $snonin_t$ on the instruments vector $\mathbf{Z}_t$ to obtain $sn\hat{o}nin_t$,

$$snonin_t = \hat{c}_0 + \mathbf{Z}_t \hat{\rho} + \hat{w}_{snonin_t} = sn\hat{o}nin_t + \hat{w}_{snonin_t} \quad (8)$$

where $\hat{w}_{snonin_t}$ is the innovation resulting from the regression of *snonin* on the instruments set **Z**. Then, we substitute $sn\hat{o}nin_t$ to $snonin_t$ in the banks returns model (equation (7)). This way we can obtain consistent estimates of the coefficients of the returns equations.

[9] For an application to hedge funds see also Racicot and Théoret (2008).

In a second step, provided that there is no endogeneity concern, we can substitute equation (8) in equation (7) to obtain the following artificial (or auxiliary) regression

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 sn\hat{o}nin_t + \mathbf{X}_t\alpha + \beta_2 \hat{w}_{snonin_t} + \varepsilon_t \quad (9)$$

Finally, using equation (9), we can build our endogeneity Hausman test with higher moments. Despite the evidence gathered so far, let assume for a moment that we do not know a priori whether *snonin* is endogenous or not, so that the coefficients of $sn\hat{o}nin_t$ and $\hat{w}_{snonin_t}$ are not necessarily the same. In this case, we have to replace the coefficient β_2 attached to $\hat{w}_{snonin_t}$ by θ, a mute coefficient, in equation (9), and thus we have:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 sn\hat{o}nin_t + \mathbf{X}_t\alpha + \theta \hat{w}_{snonin_t} + \varepsilon_t \quad (10)$$

With $snonin_t = sn\hat{o}nin_t + \hat{w}_{snonin_t}$, we can reformulate equation (10) as follows:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 snonin_t + \mathbf{X}_t\alpha + \varphi \hat{w}_{snonin_t} + \varepsilon_t \quad (11)$$

where $\varphi = \theta - \beta_2$.

The endogeneity test can then be described as follows. If there is no endogeneity problem, then $\varphi = 0$, or equivalently $\theta = \beta_2$. On the other hand, if *snonin* happens to be endogenous, then φ is significantly different from zero, that is to say $\theta \neq \beta_2$ in equation (10).

Compared to the standard *h* test, one crucial advantage of this procedure is that, besides providing an endogeneity test, it can also be used to gauge the *severity* of the endogeneity problem. Define $\hat{\varphi} = f\left(\hat{\beta}_2 - \hat{\beta}_2^*\right)$, with $f' > 0$, $\hat{\beta}_2$ the coefficient estimated by OLS, and $\hat{\beta}_2^*$ the coefficient estimated with the two-step Hausman procedure just described. According to equation (11), if $\hat{\varphi}$ is significantly positive it indicates that the coefficient of *snonin* is overstated in the OLS regression, i.e. $\hat{\beta}_2 > \hat{\beta}_2^*$. As implied by the definition, the severity of the endogeneity problem increases with $\hat{\varphi}$. The opposite argument holds true if $\hat{\varphi}$ is significantly negative. Finally, if $\hat{\varphi}$ is not significantly different from zero, then $\hat{\beta}_2 = \hat{\beta}_2^*$ and there is no clear evidence of an endogeneity problem in this case.

As a final remark note that, as implicitly suggested by Spencer and Berk (1981) and Pindyck and Rubinfeld (1998), the coefficients estimated with the auxiliary regression (11) are the same as those obtained from a standard TSLS procedure based on the instruments used for the $\hat{w}_{snonin_t}$ computation. If $\hat{\varphi}$ is not significantly different from zero (i.e. the case of no endogeneity), the OLS estimator obtains and equation (11) becomes:

$$\left(y_t\right)_{OLS} = \hat{\beta}_0 + \hat{\beta}_1 y_{t-1} + \hat{\beta}_2 snonin_t + \mathbf{X}_t \hat{\alpha} + \varepsilon_t \quad (12)$$

However, if $\hat{\varphi}$ is significantly different from zero, the TSLS estimator obtains and equation (11) reads:

$$\left(y_t\right)_{TSLS} = \hat{\beta}_0^* + \hat{\beta}_1^* y_{t-1} + \hat{\beta}_2^* snonin_t + \mathbf{X}_t \alpha^* + \hat{\varphi}\hat{w}_{snonin_t} + \varepsilon_t \quad (13)$$

where the coefficients are starred to indicate that they are equivalent to those obtained from a TSLS procedure. Consequently, our endogeneity indicator may also be rewritten as $\hat{\varphi} = f\left(\hat{\beta}_{2,OLS} - \hat{\beta}_{2,TSLS}\right)$, where $\hat{\varphi}$ becomes an indicator of the distance between the OLS and the TSLS *snonin* coefficients.

In summary, the Hausman procedure we propose can be seen as a modified TSLS directly incorporating an endogeneity test. This correspondence between the Hausman artificial regression and the TSLS is often overlooked in the econometric literature. Maybe researchers do not realize that, by using this kind of modified procedure they can directly obtain an indication of the acuity of the endogeneity problem. Obviously, for the estimation of equation (1), the standard TSLS procedure and this Hausman procedure are interchangeable. The estimated coefficients of the explanatory variables are the same in both cases. However, the motivation to favour the latter is that it provides a crucial information on endogeneity, namely, it helps assess the severity of the biases.

4. Empirical results

Insert Table 6 about here

In this section we discuss the empirical results of the various experiments we just described, beginning with those of the estimation method most commonly used in the literature, i.e. the OLS. Note however that we first need to examine the stationarity of the time series used in our model in order to avoid spurious results. Table 6 provides the Augmented Dickey-Fuller unit root test for the time series used in this study. Over the sample period, the test indicates that only the *snonin* variable seems to have a unit root, the *p*-value of the test being equal to 0.276 for this variable, so the hypothesis of the presence of a unit root cannot be rejected at the usual thresholds. To make the *snonin* variable I(0) we thus express it in first-differences in our experiments.

4.1. OLS results

Table 7 reports the results of the OLS estimation of equation (1) where banks returns are proxied by *ROA*, a standard approach in the literature. We call this version Model 1, where the ratio of loan loss provisions to total assets is the only significant control variable, so that Model 1 reads:

$$ROA_t = \gamma_1 + \gamma_2 d(snonin)_t + \gamma_3 LLP_t + \gamma_4 ROA_{t-1} + \zeta_t \quad (14)$$

where *ROA* is the return on assets, *d(snonin)* is the first-difference of *snonin*, *LLP* are loan losses provisions and ζ is the innovation[10].

The fit of the model seems quite good over the whole sample period, the adjusted R^2 being 0.62. Consistent with the idea that loan loss provisions ought to lower profits, the coefficient of the ratio of loan loss provisions to total assets, at -0.50, is found significantly negative. Since the ratio of loan loss provisions increases during recessions, it magnifies the procyclicality of *ROA*.

Insert Table 7 about here

Table 7 shows that the risk-return trade-off improves throughout the sample period. The coefficient of *d(snonin),* i.e. *snonin* expressed in first-differences, significant at the 95% confidence level, is 1.28. To illustrate the evolution of the *snonin-ROA* relationship, it is much instructive to run a recursive regression over the whole period. Figure 10 reveals a regime shift in the sensitivity of *ROA* to *d(snonin)* around 1997, which corroborates our previous findings about the presence of a structural break. According to the results derived from this recursive regression, the sensitivity of *ROA* to *d(snonin)* appears larger after 1997. We find this relationship both positive and much more significant. In this respect Figure 10 suggests a narrowing of the confidence interval of the *d(snonin)* coefficient after 1997. The N-step forecast of *ROA* also confirms the presence of a structural break[11].

Insert Figure 10 about here

[10] To check the robustness of the model we also consider a second version of the specification, Model 2, where we introduce risk premia.

[11] A rolling regression of fifteen quarters on Model 1, which provides a more precise estimation than a recursive regression, also confirms that the sensitivity of *ROA* to *snonin* turns from negative to positive around 1997. This supports the emergence of the diversification gains associated with market-oriented activities.

Because of the growth in the banks new business lines, we should expect a deterioration of the model performance in the second subperiod. Indeed, it is during this second subperiod that banks begin to integrate their new banking business to their traditional bank lending activities. Our experiments suggest that the risk prevailing in the second subperiod, as implied by the volatility of the banks income growth, is actually more pronounced, and feeds into the innovation term of the equation[12]. The data track this change in the banking environment quite well. In the *ROA* equation, the adjusted R^2 is equal to 0.87 over the first period, and then falls to 0.38 in the second subperiod, corroborating the deterioration of the model fit (Table 7).

More importantly, while banks non-traditional activities were developing, we also observe a change in the sign of the *d(snonin)* coefficient. Since banks optimize their profits, the shift from lending activities to OBS ones has to be motivated by expectations of higher returns, and eventually translates into a positive impact of *snonin* on banks performance. As expected, we indeed find that *d(snonin)* is negative (-0.17), although insignificant, during the subperiod 1988-1996, but becomes significantly positive (1.71) after 1997.

4.2. Hausman artificial regression results

Insert Table 8 about here

We report the results of the Hausman estimation of the banks returns model (equation (11)) in Table 8. As previously mentioned, the Hausman procedure is very similar to a regular TSLS estimation[13]. However, the Hausman regressions offer the

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[12] In figure 10, note that the volatility of the residuals of the recursive regression of equation (14) is much higher after 1997.

[13] Since the results obtained for the TSLS and the Hausman procedure are essentially the same we only report the Hausman procedure findings.

advantage of directly embedding an endogeneity test based on the significancy of w_{snonin}, as measured by its *t*-statistic[14]. Furthermore, this particular method also provides an indication of the severity of the endogeneity issue with the level of the w_{snonin} coefficient. What Table 8 first confirms is that the endogeneity seriously biases the estimated coefficient of *d(snonin)*. Over the whole sample, the coefficient of *d(snonin)* is equal to 1.28 when estimated by OLS, but to 2.50 when estimated with the Hausman procedure. As a matter of fact, the coefficient of *d(snonin)* appears to be globally underestimated when the endogeneity bias is ignored. The coefficient of $w_{d(snonin)}$ is equal to -2.59 for the whole estimation period, and significant at the 99% confidence level. Being negative and high in absolute value, the coefficient of $w_{d(snonin)}$ strongly suggests that the coefficient of *d(snonin)* is significantly understated in the OLS run. During the first subperiod, the coefficient of *d(snonin)* estimated by OLS is equal to -0.17, but it becomes -0.90 if we account for the endogeneity of *snonin*. The coefficient of $w_{d(snonin)}$, although insignificant, is equal to 1.20, which suggests that OLS overstates the impact of *d(snonin)* over the first subperiod, in line with the results of Stiroh and Rumble (2006).

More importantly, during the second subperiod, the coefficient of $w_{d(snonin)}$, at -3.93, is much higher in absolute value than over the whole sample period, and becomes significant at the 95% confidence level. This result strongly suggests that the endogeneity is more pronounced during the second subperiod. Compared to the whole sample period, the underestimation of the positive effect of *d(snonin)* on *ROA* is particularly severe in this period. This result indicates that the sensitivity of *ROA* to *d(snonin)* has increased after 1997, a fact consistent with the idea of a better integration of OBS activities to traditional business lines – i.e., the rise of shadow banking. To confirm this finding, it is

[14] i.e. the *t* test constitutes the Hausman test.

much instructive to run a recursive regression. In Figure 11 note that the confidence interval of the coefficient of $w_{d(snonin)}$ shrinks greatly through time. This indicates that *snonin* endogeneity issue becomes more important *pari passu* with the increased involvement of banks in market-oriented business lines. In this respect, the spike of the $w_{d(snonin)}$ coefficient during the subprime crisis might also suggest that the endogeneity issue is actually more acute during turbulent times.

Insert Figure 11 about here

Overall, our results suggest an important understatement of the coefficients of *d(snonin)* in the OLS regressions over the whole sample period, and especially for the second subperiod. Taking endogeneity carefully into account reveals that, with a better integration of traditional and OBS activities, the negative sensitivity of *ROA* to *d(snonin)* detected during the first subperiod progressively decreases to actually become positive during the 1997-2010 subperiod. After 1997, controlling for endogeneity the way we do translates in substantial gains in terms of estimation, and unveils a clear, positive influence of *d(snonin)* on returns, the coefficient more than doubling, from 1.71 to 3.79. What is crucial here is not merely the fact that the positive influence of OBS on returns obtains when controlling for endogeneity, but the fact that, unless endogeneity is treated seriously, this positive influence can be significantly underestimated. In this respect, the findings we obtain are quite natural if we consider that the endogeneity of *snonin* must evolve along with the involvement of banks in market-oriented banking. In other words, it should not be too much surprising to find that *snonin* endogeneity becomes more severe with the rise of shadow banking.

4.3. Results robustness

4.3.1. Estimation with Sharpe ratios

It is interesting to first check if our results are robust to a change in the way we account for risk. We thus express *ROA* with a risk-adjusted measure based on the Sharpe ratio. This ratio is defined as: $Sharpe_t = \frac{ROA_t - R\bar{O}A}{\sigma_{t,ROA}}$, where $R\bar{O}A$ is the mean value of *ROA* over the whole sample period, and $\sigma_{t,ROA}$, the standard deviation of *ROA*, is represented by a moving average computed on a rolling window of four quarters. The numerator of the ratio is thus the return of *ROA* at time *t* expressed in deviation from the mean value, and it is scaled by a moving average of the standard deviation to arrive at a risk-adjusted measure. We estimate Model 1 using this Sharpe ratio as the dependent variable. The estimated equation becomes:

$$Sharpe_t = \gamma_1 + \gamma_2 d(snonin)_t + \gamma_3 LLP_t + \gamma_4 Sharpe_{t-1} + \zeta_t \quad (15)$$

Insert Table 9 and 10 about here

The results of the OLS estimation of equation (15) are provided in Table 9, and the corresponding Hausman regression results are reported in Table 10. These tables indicate that the results are almost identical with an alternative measure of returns adjusted for risk. The OLS regression underestimates substantially the coefficient of *d(snonin)* over the whole sample period, and particularly so after 1997, the coefficient being actually insignificant over the subperiod 1988-1996. Over the whole sample (1988-2010), the coefficient of *d(snonin)*, significant at the 1% threshold, is equal to 11.32 when estimated by OLS, but to 19.10, also significant at the 1% threshold, when estimated with our Hausman procedure. The coefficient of $w_{d(snonin)}$, significant at the 10% threshold, is also high, at -15.08, which confirms the large underestimation of the coefficient of *d(snonin)*. The same results obtain over the period 1997-2010. Figure 12 displays the behaviour of the coefficient of *d(snonin)* when we run a recursive regression on the Sharpe ratio version of Model 1 over the period 1988-2010. The figure gives a clear picture of the evolution of the impact of OBS activities on banks performance. The coefficient touches a low of -27.4 in the third quarter of 1993. It increases progressively thereafter, and turned positive in the second quarter of 1995. From 1997 onwards, it stabilizes around a level of 12. As for the *ROA* Model 1, the confidence interval of the *d(snonin)* coefficient is much narrower after 1997, which suggests that the diversification benefits of banks OBS activities improve substantially after the structural break.

Insert Figure 12 about here

4.3.2. Adding risk premia

To check the robustness of the results obtained with our primary model, we can also define an augmented version of the model, Model 2, adding risk premia to the explanatory variables. In this case, equation (14) becomes:

$$ROA_t = \gamma_1 + \gamma_2 d(snonin)_t + \gamma_3 LLP_t + \gamma_4 r_{TSX,t-1} + \gamma_5 Spread_t + \gamma_6 ROA_{t-1} + \zeta_t \qquad (16)$$

where r_{tsx} is the return on the TSX index and *Spread* is the difference between the yield on loans and their funding cost. We expect a positive sign for γ_4, as an increase in the stock return should lead to an increase in *ROA*, especially in the new banking context, given the relative contribution of market-oriented activities to the banking business. The sign of γ_5 is less clear however. If supply-side effects dominate the loans market, the sign should be positive, an increase in *Spread* leading to a corresponding increase in

ROA. But demand-side effects might mitigate this relationship. An increase in *Spread* may thus induce a decrease in the demand for loans, and possibly a decrease in *ROA*. This counter-effect may also be more pronounced during the second part of our sample, since financial deepening accelerates and traditional activities lose steam. Moreover, the increase in *Spread* might also be symptomatic of an increase of credit risk, especially in the form of loans defaults, leading to a decrease in *ROA*.

Insert Tables 11 and 12 about here

Tables 11 and 12 report, respectively, the OLS and Hausman regressions results for Model 2. Not surprisingly the R^2 of the regressions are generally higher following the addition of risk premia. For instance, over the period 1988-2010, for the OLS regressions the R^2 without the risk premia is equal to 0.62, but it increases to 0.71 with the added variable. The increase of the R^2 is higher over the first period, the influence of the spread being much higher during this subperiod. The results for *d(snonin)* are essentially the same after the addition of risk premia, although the impact of this variable decreases somewhat. For example, in Model 2, the coefficient of *d(snonin)* is equal to 0.90 in the OLS estimation run over the whole sample period, while it is equal to 1.28 in Model 1. It also decreases from 1.71 to 1.37 over the subperiod 1997-2010. More importantly, the Hausman regressions indicate that the sensitivity of *ROA* to *d(snonin)* is understated over the whole period, and especially so during the second subperiod (1997-2010, Table 12). Over the whole sample, the coefficient of $w_{d(snonin)}$ is equal to -2.68 and significant at the 10% threshold. However, after 1997, the coefficient of $w_{d(snonin)}$, equal to -4.05 and significant at the 5% threshold, suggests a much larger underestimation of the coefficient of *d(snonin)*. Accordingly, when shifting from the OLS to the Hausman regressions, the coefficients of *d(snonin)* are revised from 0.90 to 2.29 over the period 1988-2010, and from 1.37 to 3.47 over the subperiod 1997-2010.

The coefficients of the r_{TSX} and *Spread* variables are both positive and significant at the 5% threshold in the OLS and Hausman regressions run over the whole sample, although the impact of r_{TSX} is much lower than the spread one. As expected, the influence of the spread is higher in the first subperiod, its coefficient being equal to 0.43 in the OLS regression and significant at the 1% threshold, and actually becomes insignificant in the second subperiod. If, during the first subperiod, an increase in the spread leads to an increase in *ROA* (supply-side effect), it exerts an opposite impact over the second period (demand-side effect). To confirm this idea, Figure 13 plots the behaviour of the spread coefficient when applying a recursive regression to Model 2. Note that the impact of the spread coefficient increases from 1988 to 1996, but decreases continuously thereafter, a switch which also accords with the structural break identified earlier. This result suggests that banks traditional activities recedes from 1997 onwards, as the financial deepening progressively unfolds. The finding can also be explained by increased competition, as demand-side effects could have begun to dominate supply-side ones in the loans markets.

Insert Figure 13 about here

Finally note that r_{TSX} has no impact on *ROA* in the first subperiod, whereas its impact becomes positive and significant over the second subperiod. This corroborates the view of a sharp regime shift in the banking industry. Before 1997 the fact that banks activities were more focused on traditional business lines explains the greater impact of the spread and the insignificant impact of stock returns. The situation clearly reverses thereafter[15].

4.3.3. A look at disaggregate data

Insert Tables 13 and 14 about here

Since idiosyncratic risk is diluted by diversification, it should have a lower influence at the aggregate level. Despite the data limitation, it is thus legitimate to wonder whether the results also hold at the disaggregate level. To check this, we consider the three banks selected for building Figure 3. In terms of assets, these three banks account for more than 50% of the Canadian banking system. The results for Model 1 are reported in Tables 13 and 14, for the OLS and Hausman regressions respectively. The analysis of the coefficients of $w_{snonin,it}$ shows that, not taking into account *snonin* endogeneity may also seriously bias the estimated coefficients at the disaggregate level. Although the results obviously differ from one bank to another, the OLS generally underestimate the impact of *d(snonin)*. The $w_{snonin,it}$ coefficients are high in absolute value and very significant for the Toronto-Dominion bank over the whole period and over the two subperiods, 1988-1996 and 1997-2010. These coefficients are respectively -9.11, -8.90 and -8.72 over these periods, which suggests a serious understatement of the *d(snonin)* coefficients over the three periods. Without accounting for endogeneity, the Toronto-Dominion *d(snonin)* coefficients are respectively -2.82, 4.63 and 4.13 over the periods 1988-1996, 1997-2010, and 1988-2010, and significant and the 5% threshold (Table 13), but when we account for endogeneity, the coefficients increase respectively to 5.12, 9.14 and 9.11 over the same

[15] As with Model 1, we also run Model 2 (equation (16)) with a Sharpe ratio. However the results remain fairly unchanged (cf. the appendix, Tables 17 and 18).

periods, again significant at the 5% threshold. Moreover, in the Hausman regressions, the sign of *d(snonin)* is already positive and significant over the first subperiod, which is not the case for the two other banks. The National Bank results remain aligned with those of the Toronto-Dominion, but they differ greatly for the Royal Bank. First, this bank displays less persistence in its *ROA*, as measured by the coefficient of ROA_{t-1} (Table 14). Second, as it is the case at the aggregate level, the financial performance of this bank has initially suffered from its increasing involvement in OBS activities, but in its particular case, there does not seem to be tangible benefits over the second subperiod either. Indeed, in the Hausman regression, the estimated coefficient of *d(snonin)* is equal to -0.64 and insignificant over the subperiod 1997-2010 (Table 14). Third, over the last subperiod, there is a positive comovement between the Royal Bank *LLP* and its *ROA*, which is not the case for the two other banks for which *LLP* coefficients are negative and significant. Indeed, for the Royal Bank, the coefficient of *LLP* is equal to 0.82 and significant at the 99% confidence level. This provides an example of some banks managing their provisions during the second subperiod, perhaps, progressively increasing their *LLP* to better reflect their rising exposure to OBS activities and less favourable risk-return trade-off (i.e. earnings management by income smoothing, Bikker 2005, Quagliariello 2008, Eickmeier and Hofmann, 2009, Nijskens and Wagner 2011).

Insert Tables 15 and 16 about here

Table 15 and 16 provides the estimation of Model 2 (i.e., Model 1, risk premia augmented) for the three banks with the OLS and Hausman regressions respectively. The model works quite well for the National Bank and the Toronto Dominion bank over the three periods, but its performance is rather poor for the Royal Bank over the whole subperiod, with a R^2 equal to 0.07. As shown in Table 16, the addition of risk premia in the *ROA* equation does not change the results regarding the behaviour of the *d(snonin)* coefficient, and in particular its undestatement over the whole sample period and the second subperiod. Consistent with what happens at the aggregate level, stocks returns have a negligible impact over the first subperiod for the National Bank and the Royal Bank, and a significant positive impact over the second period, although this impact is close to 0 and insignificant for the Toronto-Dominion Bank. Concerning the spread variable, the results also supports the findings obtained at the aggregate level, namely the declining influence of the spread through the sample period. However, the coefficient of the spread remains positive and significant for the National Bank during the second subperiod, which reflects its greater involvement in retail activities.

To summarize, although the results derived from a casual look at the disaggregate data have to be considered with caution, given the restricted sample size, the evidence we gather seems to confirm the progressive improvement of banks accounting returns associated with their expansion in market-oriented activities.

5. Conclusion

The change in the endogeneity of banks decision to invest in OBS activities may well be related to the fact that, due to the decreasing return on their traditional activities, banks had to resort to market-oriented activities as a way to increase their profitability (Boyd and Gertler 1994). Initially, a structural downward pressure on *ROA* could have led to a rise in *snonin* whose endogeneity thus mechanically increased through time. At first, when banks engaged in non-traditional activities, they were not necessarily aware of the increased risk they were taking however[16]. Our data reveal that, after 1987, with the successive waves of banking deregulation, and the financial deepening associated with the increased firms reliance on direct financing, it took almost ten years for banks to eventually record some diversification gains from OBS activities. After this maturation phase however, the change in the banking system, namely the emergence of shadow banking is clearly characterized by the growing share of market-oriented business lines in OBS activities, and a concomitant increase in operating revenue volatility, but also by the eventual pricing of the risk associated with the new business lines which gradually made the bulk of the banking business (Calmès and Théoret 2010, Nijskens and Wagner 2011).

Accordingly, in this paper we argue that the interdependence of *snonin* and *ROA*, and the resulting endogeneity of *snonin* have increased with the progressive diversification of banks in market-oriented business lines. Consistent with this view, the new Hausman procedure we introduce reveals that the endogeneity due to the dependence of *snonin* on *ROA* becomes much more significant during the last subperiod. The endogenity of OBS activities may not be much of a concern before 1997, but it increases substantially during the last decade. In this respect, neglecting endogeneity leads to a serious underestimation of the impact of noninterest income on *ROA*. Actually, the increase in *ROA* might be attributed to a risk premium required to price the increased risk associated with banks new activities, as evidenced by the jump in the volatility of banks net operating income growth after 1997.

The policy implications we can derive from our analysis are quite straightforward. First, given their high endogeneity degree, there is a need to better monitor OBS

[16] Comments can be found in the work of DeYoung and Roland (2001) about U.S. bankers initial thoughts on OBS activities.

activities. Banks should have the obligation to be more transparent about the involvement in these activities. Second, and more importantly, although the focus of the Bank of International Settlements (Basle II), the International Monetary Fund, and central banks in general has been mainly on credit risk analysis – i.e. the supervision of on-balance-sheet items and risk management – there is an obvious need to include more comprehensive measures of bank systemic risk, encompassing both the traditional measures of VaR and various regulatory measures of leverage, but also measures accounting for the risk inherent to OBS activities. In this respect, it is not clear whether the use of the standard measures of leverage, as those endorsed by Basle III, could account for the new cyclical aspects of banks systemic risk. In this sense, the research agenda could for example aim at building more general leverage measures and indicators of bank risk such as the ones proposed by DeYoung and Roland (2001), and Breuer (2002).

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Appendix

Estimation of Model 2 with a Sharpe ratio

Table 17
OLS estimation

Variables	1988-1996	1997-2010	1988-2010
c	-7.60	0.36	-1.31
	-1.60	*0.11*	*-0.81*
d(snonin)	-1.20	7.19	7.12
	-0.07	*2.10*	*2.02*
LLP	-2.29	-5.87	-2.33
	-4.71	*-3.31*	*-4.15*
$r_{TSX(-1)}$	0.04	0.04	0.05
	1.07	*1.76*	*2.39*
Spread	2.32	0.48	0.72
	1.87	*0.47*	*1.55*
$Sharpe_{t-1}$	0.55	0.35	0.47
	6.00	*3.60*	*5.49*
Adjusted R^2	0.66	0.33	0.38
***DW* stat.**	1.88	1.99	2.09

Notes: The dependent variable is the excess return of *ROA*, defined as the difference between *ROA* and its expected value, scaled by a rolling *ROA* standard deviation of four quarters. *ROA*, return on assets ; *d(snonin)*, first-difference of the share of noninterest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets; r_{TSX}, stock return index; *Spread*, difference between the yield on loans and the cost of funds. The *t* statistics are reported in italics.

Table 18

Hausman regression

Variables	1988-1996	1997-2010	1988-2010
c	-7.44	0.84	-1.20
	-1.72	*0.23*	*-0.74*
d(snonin)	25.23	15.14	15.15
	1.11	*2.82*	*3.50*
LLP	-2.27	-5.86	-2.31
	-5.64	*-3.16*	*-4.42*
$r_{TSX(-1)}$	0.04	0.05	0.06
	0.89	*1.85*	*2.47*
Spread	2.26	0.32	0.67
	2.03	*0.29*	*1.47*
$Sharpe_{t-1}$	0.59	0.35	0.48
	6.73	*3.33*	*5.27*
$w_{d(snonin)}$	-26.90	-16.41	-16.01
	-1.52	*-1.63*	*-2.10*
Adjusted R^2	0.68	0.33	0.39
***DW* stat.**	1.84	2.01	2.11

Notes: The dependent variable is the excess return of *ROA*, defined as the difference between *ROA* and its expected value, scaled by a rolling *ROA* standard deviation of four quarters. *ROA*, return on assets ; *d(snonin)*, first-difference of the share of noninterest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets; r_{TSX}, stock return index; *Spread*, difference between the yield on loans and the cost of funds. The *w* variable is the residuals obtained with a regression of *d(snonin)* on instruments, which constitutes the Hausman test. The *t* statistics are reported in italics.

Tables

Table 1
Decomposition of the variance of net operating income growth

	1988-1996			1997-2010			1988-2010		
	Average share	**Variance**	**Contribution to variance**	**Average share**	**Variance**	**Contribution to variance**	**Average share**	**Variance**	**Contribution to variance**
Net operating revenue		11.0			66.3			33.3	
Net interest income	0.67	13.6	**6.1**	0.49	17	**4.1**	0.57	15.2	**4.9**
Noninterest income	0.33	27.7	**3.0**	0.51	243.7	**63.4**	0.43	153.4	**28.4**
Covariance		4.3	**1.9**		-2.2	**-1.1**		0.05	**0.0**
Correlation		0.22			-0.03			0.01	

Note: The variance decomposition is obtained by using the simple portfolio variance formula, which is $Variance = \mathbf{w}^T \mathbf{\Omega} \mathbf{w}$, where **w** is the vector of the respective shares of net interest income and noninterest income in banks net operating revenue, and $\mathbf{\Omega}$ is the variance-covariance matrix of net interest income growth and noninterest income growth.
Source of data: Canadian Bankers Association and Bank of Canada.

Table 2
Descriptive statistics of the components of noninterest income, 1997-2010

	non-interest income	**capital markets**	**income wealth mgt**	**retail**	**insurance**	**trading**	**securitization**	**other**
Level (end-of-period, million $)	9244859	2513622	2039389	1522159	2017460.6	130621	723671	297936
Mean (million $)	7952304	2692371	1345518	1039965	1070721	939987	403025	556378
Median (million $)	7789256	2647529	1297605	1021984	889839	1293237	334615	554668
Std. Dev. (million $)	1679514	556968	430527	275342	545111	1191256	220719	204013
Share (start-of-period)		0.37	0.14	0.13	0.05	0.14	0.01	0.16
Share (end-of-period)		0.27	0.22	0.16	0.22	0.01	0.08	0.03
Average share		0.35	0.17	0.13	0.11	0.11	0.05	0.08
Skewness	-0.09	0.76	0.10	0.15	0.80	-1.84	2.18	-0.72
Kurtosis	2.51	4.52	1.75	2.08	2.91	6.24	9.11	3.15

Notes: Capital markets comprises the global wholesale banking business providing corporate , public sector and institutional clients with a wide range of products and services. *Income wealth management* designates a full range of investment, trust and other wealth management, and asset management products and services provided to high net worth clients. *Retail* income includes personal and business retail banking operations like mutual funds, services fees and credit cards management. *Insurance* comprises life and health, home, auto and travel insurance products. *Trading* comprises trading and distribution operations largely related to fixed income, foreign exchange, equities and derivative products. *Securitization* refers to the securitization process of credit card receivables and residential mortgages primarily used to diversify banks funding sources and enhance liquidity positions.
Source of data: Bank of Canada.

Table 3
Decomposition of the variance of noninterest income growth, Canadian Banks, 1997-2010

	Average share	Variance	Contribution to variance	Covariance	Contribution to covariance	Total contribution
Noninterest income		**3016.3**				**3016.3**
Components						
capital market income	0.35	342.7	**42.0**	147.9	**51.8**	**93.7**
income wealth-mgt income	0.17	44.3	**1.3**	24.5	**4.2**	**5.4**
retail income	0.13	98.9	**1.7**	65.2	**8.5**	**10.1**
insurance income	0.11	399.4	**4.8**	-59.6	**-6.6**	**-1.7**
trading income	0.11	238112.0	**2881.2**	443.2	**48.7**	**2929.9**
securitization income	0.05	522.9	**1.3**	-494.4	**-24.7**	**-23.4**
other income	0.08	18.9	**0.1**	26.3	**2.1**	**2.2**
Total			**2932.3**		**84.0**	**3016.3**

Notes: The variance decomposition is obtained by using the simple portfolio variance formula, which is $Variance = \mathbf{w}^T \mathbf{\Omega w}$, where **w** is the vector of the respective shares of the components of noninterest income, and $\mathbf{\Omega}$ is the variance-covariance matrix of the components expressed in growth rates. The contribution of component *i* to the total variance and covariance is computed with the following derivative: $\frac{\partial var iance}{d\mathbf{w}} = 2\mathbf{\Omega w}$, where the relative contribution of component *i* is equal to $2\mathbf{\Omega}_i \mathbf{w}$ with $\mathbf{\Omega}_i$ the i^{th} line of the $\mathbf{\Omega}$ matrix.

Table 4
Johansen's cointegration test for *snonin* versus *TSX*

	periods	test *p*-value	number of cointegrating equations	Normalized cointegrating coefficients	
TSX **(level)**	1988-1996	0.6067	0	***snonin***	***TSX***
	1988-2010	0.0601	1	1.000	-3.00E-05***
	1997-2010	0.0096	1	1.000	-6.49E-06***
TSX **(log)**	1988-1996	0.8662	0		
	1988-2010	0.0965	1	1.000	-0.197***
	1997-2010	0.0377	1	1.000	-0.054***

Notes: A *p*-value equal to 0.05 indicates the presence of a cointegrating relationship between *snonin* and *TSX* at the 95% confidence level and a *p*-value equal to 0.10 signals the presence of a conintegrating relationship at the 90% confidence level. The table reports the cointegrating vector when a cointegrating relationship is detected. *, ** and *** denote statistical significance at the 10%, 5% and 1% levels, respectively.

Table 5
Spearman rank-order correlations of *ROA* and *snonin*

	NBC	RBC	TD	8 domestic banks
1988-1996	-0.32**	-0.19	-0.41***	-0.25
1997-2010	0.66***	0.39***	0.20	0.65***

Notes: NBC: National Bank of Canada; RBC: Royal Bank of Canada Financial Group; TD: Toronto Dominion Bank Financial Group. *, ** and *** denote statistical significance at the 10%, 5% and 1% levels, respectively.

Table 6
Augmented Dickey Fuller test for the model explanatory variables

	test *p*-value					
	ROA	***snonin***	***LLP***	***Spread***	***TSX***	r_{TSX}
level	0.000	0.276	0.002	0.024	0.333	0.000
first-differences		0.000			0.000	

Note : A *p*-value of 0.05 leads to the rejection of the H0 hypothesis (presence of a unit root) at the 95% confidence level. In the table, the variables *snonin* and *TSX* present a unit root. They become I(0) when expressed in first-differences.

Table 7
Model 1: OLS estimation of *ROA*

Variables	1988-1996	1997-2010	1988-2010
c	0.93	0.53	0.77
	23.90	*3.87*	*15.06*
d(snonin)	-0.17	1.71	1.28
	-0.14	*4.34*	*2.57*
LLP	-0.57	-0.46	-0.50
	-20.81	*-2.85*	*-11.67*
ROA_{t-1}	0.02	0.37	0.10
	0.27	*2.51*	*0.13*
Adjusted R^2	0.87	0.38	0.62
***DW* stat.**	0.64	2.08	1.36

Notes: ROA, return on assets ; *d(snonin)*, first-difference of the share of non-interest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets. The *t* statistics are reported in italics.

Table 8
Model 1: Hausman regression of *ROA*

Variables	1988-1996	1997-2010	1988-2010
c	0.93	0.44	0.86
	26.97	*4.57*	*27.91*
d(snonin)	-0.90	3.79	2.50
	-0.39	*3.48*	*4.96*
LLP	-0.58	-0.43	-0.61
	-23.15	*-3.41*	*-16.05*
ROA_{t-1}	0.01	0.50	0.51
	0.13	*5.26*	*5.32*
$w_{d(snonin)}$	1.20	-3.93	-2.59
	0.48	*-2.11*	*-3.61*
Adjusted R^2	0.89	0.52	0.75
***DW* stat.**	0.80	2.41	2.10

Notes: ROA, return on assets; *d(snonin),* first-difference of the share of non-interest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets. The *w* variable is the residuals obtained with a regression of *d(snonin)* on robust instruments. The *t* statistics are reported in italics.

Table 9
Model 1: OLS estimation of the *ROA* Sharpe ratio

Variables	1988-1996	1997-2010	1988-2010
c	1.30	2.00	1.03
	4.33	*2.94*	*3.57*
d(snonin)	8.63	10.06	11.32
	0.59	*2.73*	*3.32*
LLP	-1.87	-6.14	-1.87
	-4.92	*-3.17*	*-4.19*
$Sharpe_{t-1}$	0.56	0.35	0.49
	6.49	*3.35*	*4.94*
Adjusted R^2	0.58	0.33	0.35
DW stat.	2.05	2.02	2.09

Notes: The dependent variable is the excess return of *ROA*, defined as the difference between *ROA* and its expected value, scaled by a rolling *ROA* standard deviation of four quarters. The explanatory variables are: *d(snonin)*, first-difference of the share of noninterest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets, and the Sharpe ratio lagged one period. The *t* statistics are reported in italics.

Table 10

Model 1: Hausman regression of the *ROA* Sharpe ratio

Variables	1988-1996	1997-2010	1988-2010
c	1.19	2.01	1.01
	4.24	*2.80*	*3.48*
d(snonin)	39.25	17.67	19.10
	1.52	*3.20*	*3.53*
LLP	-1.88	-6.30	-1.88
	-5.50	*-3.18*	*-4.23*
$Sharpe_{t-1}$	0.60	0.37	0.51
	5.12	*3.17*	*4.83*
$w_{d(snonin)}$	-32.83	-15.40	-15.08
	-1.57	*-1.64*	*-1.68*
Adjusted R^2	0.61	0.33	0.35
***DW* stat.**	1.96	2.05	2.12

*Note*s: The dependent variable is the excess return of *ROA*, defined as the difference between *ROA* and its expected value, scaled by a rolling *ROA* standard deviation of four quarters. The explanatory variables are: *d(snonin),* first-difference of the share of noninterest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets, and the Sharpe ratio lagged one period. The *w* variable is the residuals obtained with a regression of *d(snonin)* on the robust instruments. The *t* statistics are reported in italics.

Table 11
Model 2: OLS estimation of *ROA*

Variables	1988-1996	1997-2010	1988-2010
c	-0.75	0.40	0.23
	-7.25	*1.66*	*1.49*
d(snonin)	0.80	1.37	0.90
	1.10	*3.44*	*2.35*
LLP	-0.63	-0.42	-0.61
	-42.32	*-3.20*	*-23.97*
$r_{TSX(-1)}$	0.00	0.01	0.01
	0.30	*2.53*	*2.71*
Spread	0.43	0.04	0.16
	16.89	*0.79*	*3.85*
ROA_{t-1}	0.02	0.35	0.12
	0.59	*2.61*	*1.46*
Adjusted R^2	0.97	0.41	0.71
***DW* stat.**	1.07	2.07	1.44

Notes: ROA, return on assets; *d(snonin)*, first-difference of the share of non-interest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets; r_{TSX}, stock return index; *Spread*, difference between the yield on loans and the cost of funds. The *t* statistics are reported in italics.

Table 12

Model 2: Hausman regression of *ROA*

Variables	1988-1996	1997-2010	1988-2010
c	-0.70	0.38	0.25
	-6.58	*1.45*	*1.54*
d(snonin)	0.40	3.47	2.29
	0.44	*3.32*	*2.44*
LLP	-0.63	-0.37	-0.61
	-32.81	*-3.44*	*-21.10*
$r_{TSX(-1)}$	0.00	0.01	0.01
	-0.25	*3.37*	*3.49*
Spread	0.43	0.02	0.14
	16.02	*0.23*	*3.21*
ROA_{t-1}	0.01	0.47	0.15
	0.46	*5.13*	*1.42*
$w_{d(snonin)}$	1.16	-4.05	-2.68
	1.07	*-2.24*	*-1.75*
Adjusted R^2	0.98	0.57	0.74
***DW* stat.**	1.24	2.45	1.41

Notes: ROA, return on assets; *d(snonin)*, first-difference of the share of non-interest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets; r_{TSX}, stock return index; *Spread*, difference between the yield on loans and the cost of funds. The *w* variable is the residuals obtained with regression of *d(snonin)* on the robust instruments. The *t* statistics are reported in italics.

Table 13
Model 1: OLS estimation of *ROA*: Three Canadian individual banks

	National Bank of Canada			Toronto-Dominion Bank			Royal Bank of Canada		
Variables	**1988-1996**	**1997-2010**	**1988-2010**	**1988-1996**	**1997-2010**	**1988-2010**	**1988-1996**	**1997-2010**	**1988-2010**
c	0.98	0.47	0.80	0.35	0.51	0.48	0.96	0.57	0.80
	17.09	*4.47*	*13.38*	*2.98*	*7.25*	*7.55*	*14.77*	*10.81*	*9.46*
d(snonin)	-1.45	1.40	1.12	-2.82	4.63	4.13	-2.55	-0.74	-0.59
	-1.09	*4.70*	*3.84*	*-2.50*	*3.13*	*2.93*	*-1.88*	*-1.33*	*-0.61*
LLP	-0.71	-0.23	-0.58	-0.26	-0.51	-0.35	-0.46	0.82	-0.22
	-14.66	*-1.57*	*-12.43*	*-2.32*	*-5.03*	*-3.43*	*-3.80*	*3.84*	*-3.45*
ROA_{t-1}	0.08	0.38	0.07	0.71	0.43	0.48	-0.01	-0.02	-0.06
	1.36	*2.72*	*1.13*	*6.65*	*7.58*	*8.55*	*-0.25*	*-0.36*	*-0.61*
Adjusted R^2	0.86	0.29	0.65	0.67	0.51	0.46	0.73	0.51	0.09
DW stat.	1.49	2.22	1.49	2.50	1.96	2.05	1.63	1.25	2.06

Notes: *ROA*, return on assets; *d(snonin)*, first-difference of the share of noninterest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets. The *t* statistics are reported in italics.

Table 14
Model 1: Hausman regression of *ROA*: Three Canadian individual banks

	National Bank of Canada			Toronto-Dominion Bank			Royal Bank of Canada		
Variables	**1988-1996**	**1997-2010**	**1988-2010**	**1988-1996**	**1997-2010**	**1988-2010**	**1988-1996**	**1997-2010**	**1988-2010**
c	0.95	0.39	0.77	0.40	0.38	0.36	0.96	0.57	0.79
	15.62	*4.10*	*8.48*	*5.03*	*4.66*	*6.07*	*14.37*	*10.62*	*9.23*
d(snonin)	0.75	2.24	1.83	5.12	9.14	9.11	-3.17	-0.64	-1.79
	0.38	*6.54*	*3.72*	*3.57*	*10.44*	*7.86*	*-2.39*	*-0.57*	*-0.75*
LLP	-0.70	-0.23	-0.59	-0.26	-0.44	-0.29	-0.46	0.82	-0.22
	-14.30	*-1.83*	*-6.79*	*-3.45*	*-3.50*	*-3.06*	*-3.73*	*3.82*	*-3.39*
ROA_{t-1}	0.07	0.50	0.10	0.63	0.57	0.60	-0.01	-0.02	-0.05
	1.14	*3.95*	*1.11*	*8.50*	*7.02*	*10.16*	*-0.22*	*-0.36*	*-0.46*
$w_{d(snonin)}$	*-1.13*	*-2.18*	*-1.83*	*-8.90*	*-8.72*	*-9.11*	*0.78*	*-0.14*	*1.79*
	-0.85	*-3.83*	*-3.04*	*-6.40*	*-7.08*	*-6.27*	*0.64*	*-0.09*	*0.61*
Adjusted R^2	0.87	0.44	0.71	0.85	0.75	0.73	0.72	0.50	0.08
DW stat.	1.43	2.12	1.48	2.16	1.98	1.99	1.66	1.25	2.04

Notes: *ROA*, return on assets; *d(snonin)*, first-difference of the share of noninterest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets. The *w* variable is the residuals obtained with a regression of *d(snonin)* on the robust instruments. The *t* statistics are reported in italics.

Table 15

Model 2: OLS estimation of *ROA*: Three Canadian individual banks

	National Bank of Canada			Toronto-Dominion Bank			Royal Bank of Canada		
Variables	**1988-1996**	**1997-2010**	**1988-2010**	**1988-1996**	**1997-2010**	**1988-2010**	**1988-1996**	**1997-2010**	**1988-2010**
c	-1.21	-0.34	-0.19	-0.87	1.37	0.11	-0.19	1.29	0.88
	-2.63	*-1.07*	*-0.95*	*-3.17*	*3.84*	*0.49*	*-0.54*	*2.76*	*2.86*
d(snonin)	-1.52	1.29	1.15	-1.85	4.76	4.21	-3.17	-0.38	-0.57
	-1.00	*2.64*	*2.47*	*-1.20*	*3.06*	*2.78*	*-1.84*	*-0.54*	*-0.41*
LLP	-0.76	-0.33	-0.71	-0.39	-0.54	-0.46	-0.49	0.84	-0.21
	-9.16	*-2.05*	*-7.12*	*-3.97*	*-7.06*	*-4.29*	*-3.91*	*4.68*	*-3.01*
$r_{TSX(-1)}$	*-0.01*	*0.01*	*0.01*	*0.01*	*-0.01*	*-0.01*	*0.01*	*0.01*	*0.01*
	-1.54	*2.22*	*1.75*	*0.31*	*-0.86*	*-0.74*	*1.07*	*2.68*	*0.44*
Spread	*0.57*	*0.30*	*0.29*	*0.35*	*-0.25*	*0.11*	*0.29*	*-0.23*	*-0.02*
	4.85	*2.95*	*4.00*	*4.07*	*-2.37*	*1.77*	*2.88*	*-1.55*	*-0.29*
ROA_{t-1}	0.08	0.21	0.10	0.57	0.37	0.48	-0.02	-0.05	-0.06
	1.45	*1.96*	*2.15*	*5.43*	*4.41*	*9.89*	*-0.49*	*-0.83*	*-0.63*
Adjusted R^2	0.91	0.43	0.75	0.72	0.51	0.47	0.75	0.57	0.07
DW stat.	2.08	1.98	1.79	2.63	1.92	2.02	1.99	1.44	2.06

Notes: *ROA*, return on assets; *d(snonin)*, first-difference of the share of noninterest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets; r_{TSX}, stock return index; *Spread*, difference between the yield on loans and the cost of funding them. The *t* statistics are reported in italics.

Table 16

Model 2: Hausman regression of *ROA*: Three Canadian individual banks

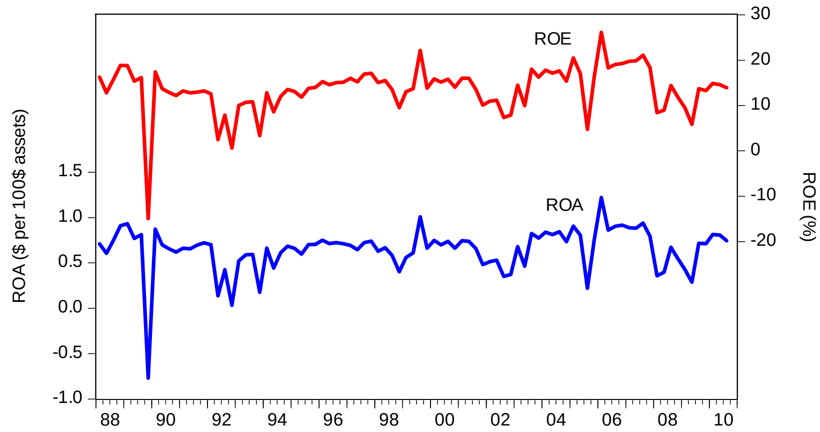
	National Bank of Canada			Toronto-Dominion Bank			Royal Bank of Canada		
Variables	**1988-1996**	**1997-2010**	**1988-2010**	**1988-1996**	**1997-2010**	**1988-2010**	**1988-1996**	**1997-2010**	**1988-2010**
c	-1.24	-0.32	-0.13	-0.02	0.93	0.05	-0.19	1.29	0.88
	-3.61	*-1.37*	*-0.71*	*-0.06*	*2.29*	*0.43*	*-0.53*	*2.74*	*2.74*
d(snonin)	0.93	2.12	1.90	4.74	9.03	9.06	-3.79	-0.49	-1.66
	0.54	*3.43*	*3.91*	*2.25*	*6.39*	*7.46*	*-2.96*	*-0.61*	*-0.68*
LLP	-0.75	-0.32	-0.71	-0.30	-0.44	-0.36	-0.49	0.84	-0.21
	-8.81	*-2.15*	*-8.16*	*-3.20*	*-3.49*	*-3.26*	*-3.82*	*4.57*	*-2.98*
$r_{TSX(-1)}$	*-0.01*	*0.01*	*0.01*	*0.01*	*-0.01*	*0.01*	*0.01*	*0.01*	*0.01*
	-1.05	*4.44*	*2.80*	*0.40*	*-0.03*	*0.17*	*0.84*	*2.60*	*0.38*
Spread	*0.56*	*0.27*	*0.27*	*0.12*	*-0.16*	*0.10*	*0.29*	*-0.23*	*-0.02*
	5.96	*2.89*	*4.78*	*1.18*	*-1.40*	*2.34*	*2.92*	*-1.55*	*-0.28*
ROA_{t-1}	0.07	0.34	0.12	0.60	0.52	0.60	-0.02	-0.05	-0.05
	1.56	*2.14*	*1.98*	*5.89*	*5.67*	*9.79*	*-0.44*	*-0.71*	*-0.50*
$w_{d(snonin)}$	*-1.84*	*-2.14*	*-1.97*	*-8.07*	*-8.57*	*-9.13*	*0.73*	*0.17*	*1.63*
	-1.18	*-3.17*	*-3.22*	*-3.47*	*-4.99*	*-6.16*	*0.44*	*0.15*	*0.55*
Adjusted R^2	0.91	0.59	0.79	0.85	0.74	0.74	0.73	0.57	0.07
DW stat.	1.97	1.86	1.67	2.27	1.95	2.00	1.97	1.44	2.04

Notes: *ROA*, return on assets; *d(snonin),* first-difference of the share of noninterest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets; r_{TSX}, stock return index; *Spread*, difference between the yield on loans and the cost of funds. The *w* variable is the residuals obtained with a regression of *d(snonin)* on the robust instruments. The *t* statistics are reported in italics.

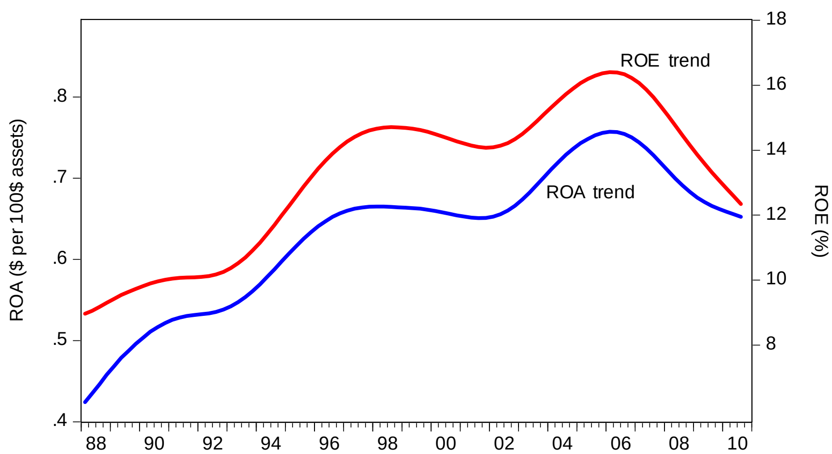
Figures

Figure 1 *ROA* and *ROE* for the eight major Canadian banks

Levels



Hodrick Prescott trends



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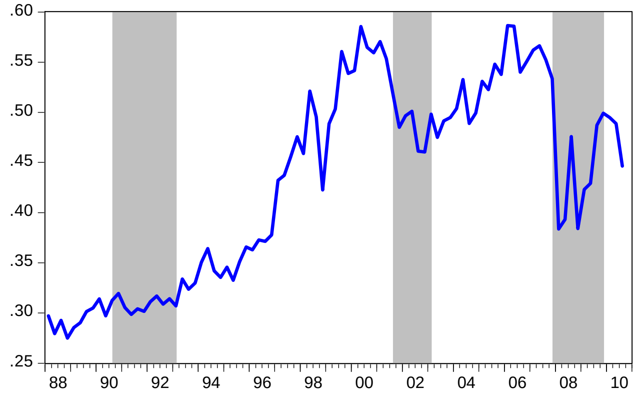
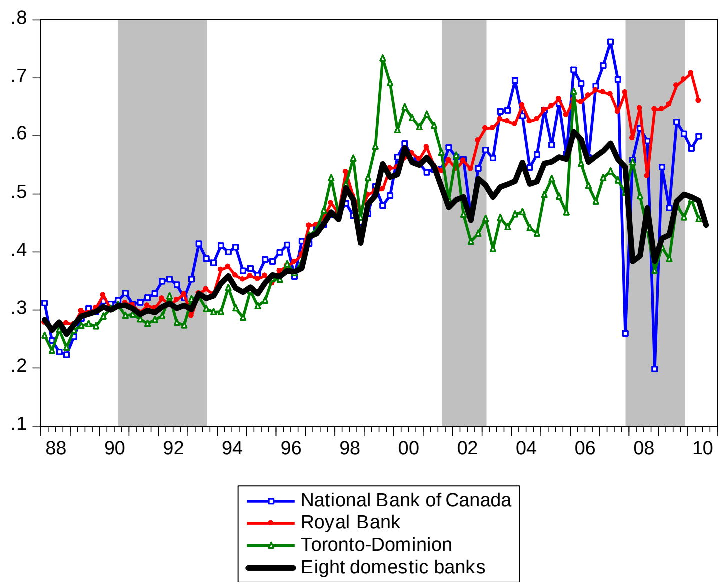
Source: Canadian Bankers Association.

Figure 2 Share of noninterest income (*snonin*) for the eight domestic banks, 1988-2010



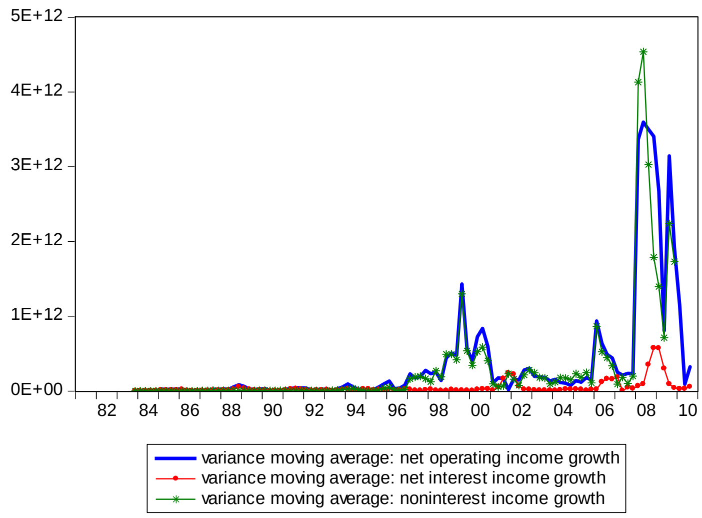
Note: Shaded areas correspond to periods of contractions or marked economic slowdown.
Source: Canadian Bankers Association.

Figure 3 Share of noninterest income in net operating revenue (*snonin*), three Canadian domestic banks, 1988-2010



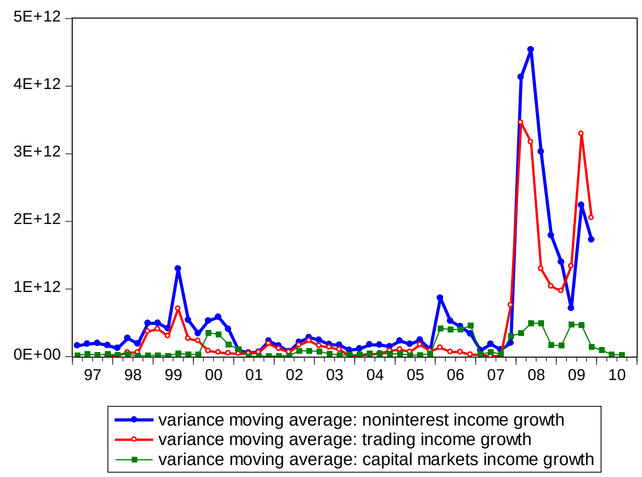
Note: Shaded areas correspond to periods of contractions or marked economic slowdown.
Source: Canadian Bankers Association.

Figure 4 Variance of net operating income growth and its components, 1983-2010



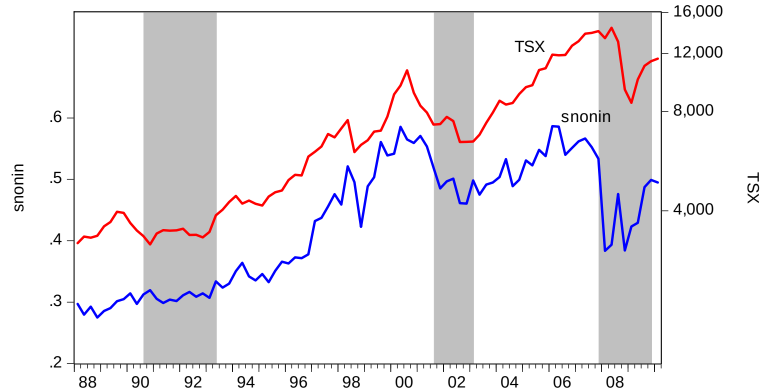
Note: The variance is a rolling variance computed on four quarters.
Source: Bank of Canada.

Figure 5 Variance of noninterest income growth and of its two most volatile components, trading income and capital markets income, 1997-2010



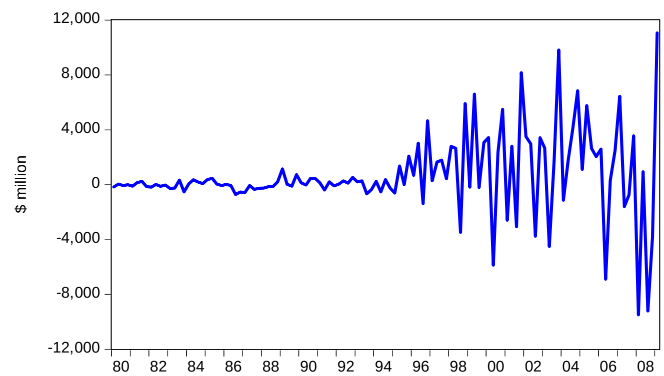
Note: The variance is a rolling variance computed on four quarters.
Source: Bank of Canada.

Figure 6 *TSX* and Canadian banks share of noninterest income (*snonin*), 1988-2010



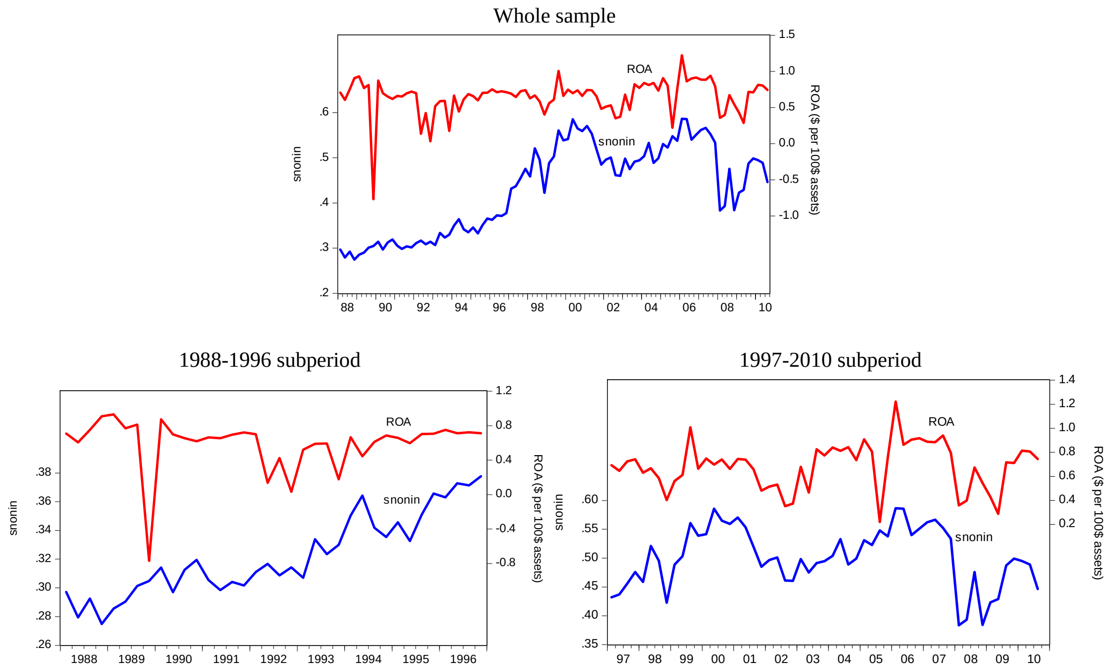
Note: Shaded areas correspond to periods of contractions or marked economic slowdown in Canada.
Source : Cansim, Statistics Canada and Canadian Bankers Association.

Figure 7 Quarterly changes in the Canadian bank stock portfolio, 1980-2009



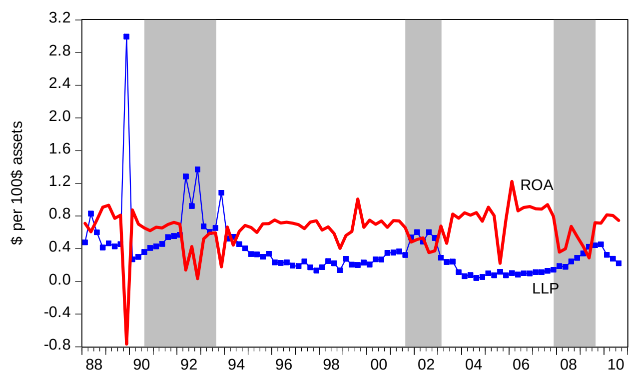
Source: Flow of funds accounts, Statistic Canada.

Figure 8 Return on assets (*ROA*) and share of noninterest income (*snonin*)



Source: Canadian Bankers Association.

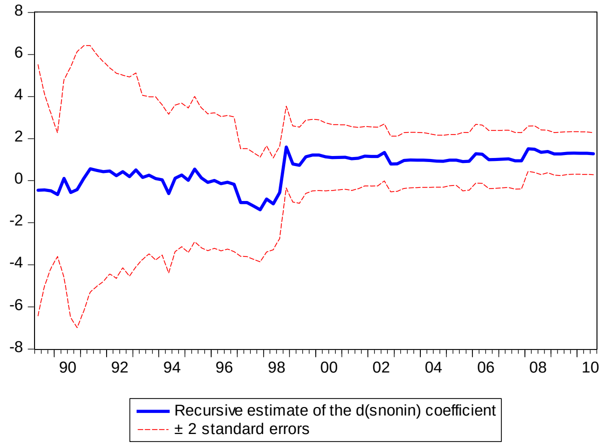
Figure 9 Return on assets (*ROA*) and loan loss provisions (*LLP*)



Note: Shaded areas correspond to periods of contractions or marked economic slowdown in Canada.

Source: Canadian Bankers Association.

Figure 10 Recursive estimate of the *d(snonin)* coefficient in *the ROA* Model 1, 1988-2010



N-step forecast of *ROA*

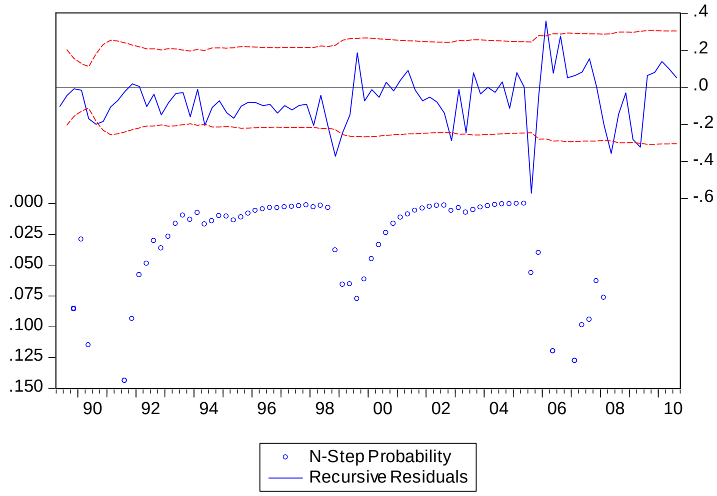


Figure 11 Recursive estimate of the *w* coefficient in the *ROA* Model 1, 1988-2010

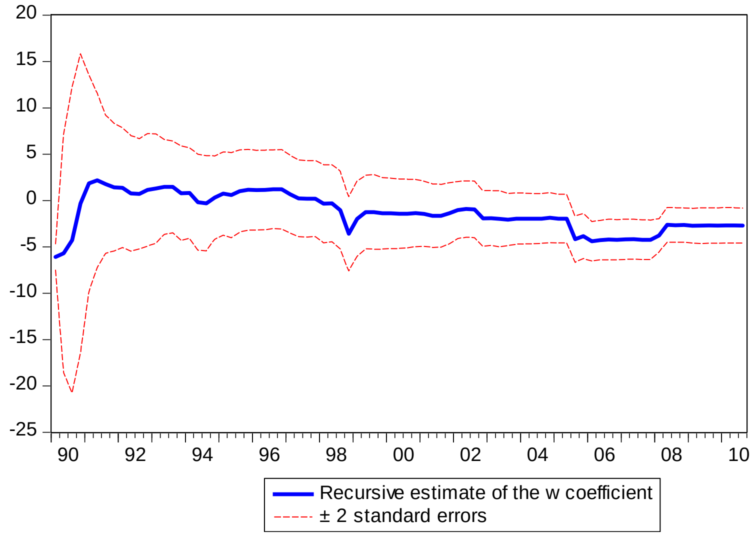


Figure 12 Recursive estimate of the *d(snonin*) coefficient in the Sharpe ratio version of Model 1

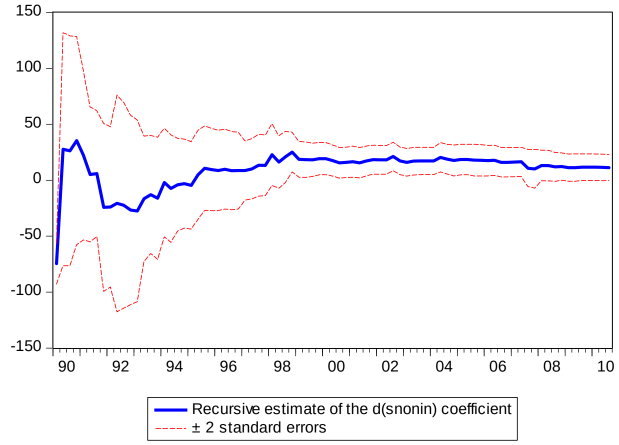


Figure 13 Recursive estimate of the *Spread* coefficient in the *ROA* Model 2