Off-Balance-Sheet Activities and the Shadow Banking System: *An Application of the Hausman Test with Higher Moments Instruments*

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Abstract

The noninterest income banks generate from their off-balance-sheet activities contributes greatly to the volatility of their operating revenues. Using Canadian data, we apply a modified Hausman procedure based on higher moments instruments and revisit this phenomenon to establish that the share of noninterest income (*snonin*) is actually endogenous to banks returns. In 1997, after the adoption of the Value at Risk (VaR) as a measure of banks risk, the *snonin* sign turns positive in the returns equations, indicating the emergence of diversification gains from banks non-traditional activities. ARCH-M estimations corroborate the idea that banks have gradually adapted to their new business lines, with an adjustment process begun even before 1997. However, the banks risk premium associated to OBS activities has continuously increased since that date.

JEL Classification: G20; G21; C32.

Keywords: Bank Risk Measures; Diversification; Noninterest income; Hausman test; Endogeneity; ARCH-M.

Résumé

Les revenus autres que d'intérêt que les banques dérivent de leurs activités hors-bilan contribuent grandement à la volatilité de leurs revenus d'opération. A partir d'une base de données canadiennes, nous recourons à une procédure modifiée d'Hausman basée sur les moments supérieurs pour établir que la part des revenus autres que d'intérêt (*snonin*) est endogène dans notre modèle de rendements bancaires. En 1997, après l'adoption de la VaR par les banques comme mesure de risque, le signe de *snonin* est devenu positif dans les équations de rendement, ce qui indique l'apparition de bénéfices de diversification dans les activités non traditionnelles des banques. Nos estimations ARCH-M corroborent l'idée que les banques se sont adaptées graduellement à leurs nouvelles activités, le processus d'ajustement s'étant même enclenché avant 1997. Toutefois, la prime de risque de marché associée aux activités hors-bilan s'est accrue continuellement depuis cette date.

Classification JEL : G20 ; G21 ; C32.

Mots-clefs **:** Mesures du risque bancaire ; Diversification ; Revenu autre que d'intérêt ; Test d'Hausman ; Endogénéité ; ARCH-M.

1. Introduction

The recent credit crisis is mainly attributable to banks off-balance-sheet (OBS) activities, and in particular securitization, which has fuelled the last lending boom by enabling banks to increase their operational funding. This eventually led to a standard maturity mismatch and a liquidity crisis (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) call Shadow Banking[2], presents a considerable challenge to regulators. In the context of this new banking era, it becomes crucial 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 strong influence on banks risk-return trade-off (DeYoung and Roland 2001, Estrella 2001, Acharya 2002, Clark and Siems 2002, Stiroh 2004a, Stiroh and Rumble 2006, Baele et al. 2007, De Jonghe 2009). It triggered a substantial increase in the volatility of banks net operating revenue growth, without a compensated increase in their returns (Stiroh 2004a, Baele et al. 2007, De Jonghe 2009).

In this paper we go a step further, analysing whether banks decision to diversify with OBS activities could actually be endogenous. Most studies of the banking literature assume this decision to be exogenous[3]. However, since it is the by-product of a classical dynamic profit optimization problem, the decision to diversify should be considered endogenous instead (Campa and Kedia 2002, Baele et al. 2007, DeJonghe 2009). Indeed, changes in the risk-return trade-off ought to lead banks to adjust their share of noninterest income. As a matter of fact, there is some empirical evidence that banks rely more on OBS activities when they need extra funding[4]. Banks could also use the diversification gains they make from their market-oriented activities to decrease the marginal cost of their risk and engage in riskier activities (Demsetz and Stahan 1997). For example, in the USA, larger banks are more diversified but many use their noninterest income as leverage to pursue riskier activities.

Needless to say, an investigation of the cyclical properties of *snonin* would certainly be worthwhile, however, our primary concern is not to build a financial soundness indicator (FSI) encompassing OBS activities[5], but simply to first check whether the noninterest income is indeed endogenous, and second, if such is the case, to propose alternative measures of banks risk more suited for modern macroprudential analysis. Our starting point is the idea that the interaction between banks risk measures and the noninterest income share might generate a serious simultaneity bias. On this matter Stiroh (2004b, p. 148) notes that, "This (*interaction*) clouds the interpretation between diversification and risk and could lead the results to be biased upwards". One possible consequence of this bias might be that banks risk-return trade-off could spuriously indicate an overstated deterioration following the increase in noninterest income[6]. In the context of banks diversification, this endogeneity issue is thus a very important aspect. Neglecting it might lead to spurious correlations between this variable and unobservables not accounted for in banks returns equations (Stiroh 2004a). In particular, the resulting non-orthogonality of the non-

[1] See Shin (2009) for a detailed explanation. For Canadian evidence, see also Calmès (2004). For the U.S. evidence, see, for example, Boyd and Gertler (1994).

[2] Securitized banking is also an expression used in the literature to refer to the banks non-traditional activities. Gorton and Metrick (2009) call securitized banks the institutions whom finance their portfolios of securitized bonds via repos (obviously a very risky collateral for repos), as opposed to traditional depository institutions that are regulated.

[3] Few noticeable exceptiona include Baele et al. (2007) (which uses a lagged regressor) and Busch and Kick (2009) study based on a TSLS approach.

[4] For example, this could be the case for mismanaged financial institutions (or the whole banking system, in periods of expected downturns). On this matter, see, for example, Farhi and Tirole (2009).

[5] Note also that banks may succomb to the temptation of creative accounting in difficult times (Chihak and Schaeck 2007).

[6] This is obviously the case when using estimation techniques such as OLS, which do not account for endogeneity

interest income share with the innovation in the returns equations can cause serious biases in the parameters estimates.

To tackle this issue, Stiroh and Rumble (2006) and Baele et al. (2007) introduce fixed effects or lagged explanatory variables in their panel regressions. However, this approach does not completely alleviate the problem. One of the main contributions of this paper is to propose an econometric framework to study this endogeneity question directly and more systematically. First, to analyse the endogeneity of noninterest income, we complete Stiroh (2004a) and Stiroh and Rumble (2006) by introducing a risk premium in the banks returns equations. Generally, banks risk is accounted for by scaling returns measures by their respective volatilities[7]. Instead, we propose the direct incorporation of various banking risk measures in the returns equations, as generally done in the asset pricing literature. In this context, we introduce two alternative measures of banking risk, (i) one based on the market risk premium, the usual systemic risk variable appearing in banks returns equations; (ii) the other directly related to the *volatility* of banks returns. According to Stiroh (2004a), returns volatility is more relevant than the traditional market risk premium. It is all the more relevant considering that VaR, the standard risk measure adopted by most banks since 1997, is itself a measure based on returns volatility[8].

Second, we do not test for the endogeneity of *snonin* with the standard Hausman (1978) test. Since the *h* test tends to be less reliable when the covariance matrix is not positive definite, we instead propose a transformed version of this test. This modified *h* test is based on an artificial regression, equivalent to a TSLS procedure incorporating a direct measure of the biases in the endogenous variables coefficients. The instruments used to run this test are based on robust instruments built with the higher moments of the explanatory variables (Fuller 1987, Lewbel 1997, Racicot and Théoret 2008).

Regardless of the estimation period considered, this version of TSLS reveals that the coefficient of the share of noninterest income of banks net operating revenues (*snonin*) is indeed overstated by the usual OLS estimations. According to Canadian data, this problem seems to be more pronounced in the first subperiod we study (1988-1996), when *snonin* has a strong significant negative impact on banks returns. By taking the non-interest income endogeneity into account, our findings reinforce the results the literature document on a negative relationship between banks returns and *snonin* for the 1988-1996 subperiod, and in particular the OLS findings of Stiroh and Rumble (2006). These authors find a negative, not significant, impact of *snonin* on banks returns, but in light of the new evidence we gather, we are inclined to think that the lack of significance of *snonin* they find might simply be due to their general treatment of the endogeneity problem.

Studying the more recent period (1997-2007), we also find that *snonin* continues to have a strong, significant, impact on banks returns. However, this impact becomes *positive* in this period. This change in sign relates to the adjustment process the banking industry underwent, with the progressive integration of the new OBS activities into banks traditional business lines, what Adrian and Shin (2009) call the shadow banking system[9].

In other respect, taking into account the fact that *snonin* is endogenous also leads to the conclusion that banks risk-return trade-off did not deteriorate as much as previously thought. Indeed, using an ARCH-M procedure, we find that *snonin* endogeneity has a major impact on the risk premium required for the pricing of the risk associated to OBS activities. Whatever the banks performance measures we examine, neglecting the endogeneity of noninterest income results in a

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[7] Some authors use a measure of market risk as this is what investors care about; however, we favor accounting data, since, from the banks supervisors' point of view, they provide a more informative picture about the ex-post outcomes. For a similar view see Stiroh (2004b).

[8] Given the highly nonlinear pattern of the returns volatility, we estimate this measure of banking risk with an instrumental IV ARCH-M procedure (Engle et al. 1987), to our knowledge, a new approach in this literature.

[9] Loosely speaking, this new banking is the result of a maturation process understood both as a progressive change in the banks activities mixture, and a learning-by-doing or learning-by-observing adaptation to new business lines. For more on this, see Calmès (2003), and Delong and DeYoung (2007).

non significant risk premium, whereas it is actually found significant when *snonin* is properly specified.

This paper is organized as follows. In the next section, we present some stylized facts about the behaviour of noninterest income, and introduce a factor model to study *snonin* properties. Then, in section 3, we expose the banks returns model and the modified Hausman method we use to test the endogeneity of noninterest income. We also detail how to construct the higher moments instruments used in the IV estimations. The fourth and fifth sections detail our results, while the last section concludes with final considerations.

2. Some Stylized Facts about the Behaviour of Noninterest Income

2.1 The Change in the Noninterest Income Series

Figure 1 Share of noninterest income for the eight domestic banks, 1988-2007

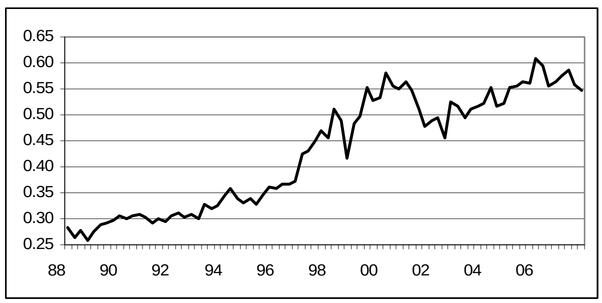


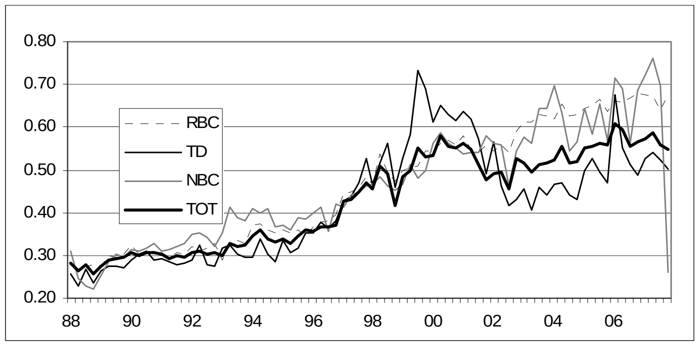
Figure 1 illustrates the growth in the noninterest income share of Canadian banks net operating revenue. By 2000, noninterest income accounted for 57% of net operating revenue, up from only 25% in 1988. However, this ratio seems to have stabilized thereafter, as the new banking businesses matured. The ratio recovered somewhat after the high tech bubble burst, culminating at 60% in the first quarter of 2006, but decreased again, particularly during the recent credit crisis. More importantly, note that the fluctuations of the share of noninterest income are much larger after 1997. In particular, *snonin* became increasingly sensitive to the fluctuations of the financial markets after 1997 (Calmès and Liu 2009)[10]. This tends to suggest the presence of a structural break around this date[11]. This change relates to the better integration of banks OBS activities, and the gradual adjustment in the way banks deal with these new non-traditional business lines. The gradualism with which this integration operated can be explained by standard arguments, such as learning by doing, time-to-build or adjustment costs (Caballero et al 2003) among others. It can also explain why, in the past, few evidence of diversification benefits of OBS activities have been found in the data (e.g. Stiroh and Rumble 2006). As a matter of fact, in the United-States, financial deregulation sometimes lagged its Canadian and European counterparts. For instance, the Gramm-Leach-Bliley Act (GLBA), which allowed American banks to engage in brokerage activities (one of the most sensitive OBS activity), was not adopted until 1999, while

[10] As a matter of fact, we find that there might be a cointegration relationship between *snonin* and the stock market index.

[11] The Chow tests we perform on the banks returns equations confirm the presence of a structural break around 1997. Liu et al. (2006) suggest another reason for the presence of a structural break around this date. Indeed, since 1996, the probability of banks failures has decreased in Canada because the OFSI has been given more powers to rescue a financial institution in difficulties, like the power to take control of an institution's asset without having to prove that it is insolvent Another regulatory change was the introduction of the Prompt Corrective Action in the mid-1990s. With the help of the launching of the VaR, these measures might have contributed to increase moral hazard in the Canadian banking system since 1997 and induced banks to take more risk, as shown in the tests appearing in our article.

the corresponding amendment to the Canadian Bank Act was promulgated almost ten years before, in 1987.

Figure 2 Share of noninterest income in net operating revenue, three Canadian domestic banks, 1988 – 2007

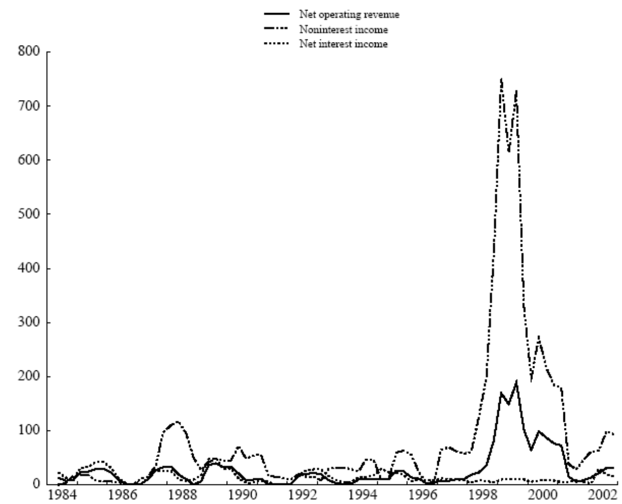


Note: RBC: Royal Bank of Canada; TD: Toronto Dominion Bank; NBC: National Bank of Canada; TOT: eight Canadian domestic banks.

The post 1997 increased volatility of *snonin* series is more apparent if individual banks are considered instead of the aggregate sample of Canadian banks[12]. As an illustration, Figure 2 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)[13]. Contrary to the RBC 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 substantially since 2000. The dispersion in banks *snonin* has also greatly increased since 1997, perhaps an indication of improved diversification of the Canadian banking industry as a whole.

[12] Due to data limitation, our study is only restricted to the analysis of aggregate data and we only display this figure as an illustration. In this respect, note also that there is some evidence for the benefit of relying on banks data on the aggregate level for macroprudential analysis. For example, return on equity is among the best indicators of the raise in systemic banking instability (Cihàk and Schaeck, 2007).
[13] For the fourth quarter of 2007, total assets of NBC, TD and RBC amounted to 113 billion $, 422 billion $, and 600 billion $ respectively. Their relative share in the assets of the pool of the eight domestic banks were 4.5%, 18.5%, and 26.2%.

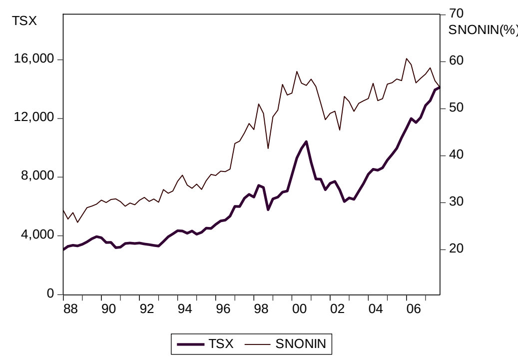
Figure 3 Volatility of banks net operating revenue growth, 1984-2002



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. Going back to aggregate data, Figure 3 obviously supports this idea. Of course, the 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 current instability 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 banks income growth volatility because of its tendency to underestimate the negative impact of fat tails. Indeed, to compute their VaR, banks have been heavily inclined to rely on the middle range of their returns distributions, where most of the losses are located. But this induced banks to blindly engage in riskier activities, particularly OBS activities, and consequently increased their total leverage. It certainly explains a great deal of the increased banks income volatility, as banks overlooked the surging risk of their operations, in the immediate years following 1997, with VaR understating their true risk. VaR underestimates risk because the central limit theorem no longer holds in the presence of non-normal returns (Gabaix 2009). Hence, idiosyncratic risk is not fully diversifiable in this case. There are actually incompressible grains which reduce the diversifiability of idiosyncratic risk. If the returns are not Gaussian, VaR should thus be adjusted by a factor which Ebert and Lütkebohmert (2009) call a granular adjustment. This adjustment may likely boost the VaR.

2.2 The *Snonin* Factor Model

Figure 4 TSX and Canadian banks share of noninterest income, 1988-2007



Studying the share of noninterest income (*snonin*), first note that this variable is strongly influenced by financial and macroeconomic factors such as financial markets conditions, the stock market index, the interest rate and the industrial production index. Given these characteristics, we adopt a portfolio approach and specify *snonin* with a factor model much in the spirit of Stiroh and Rumble (2006). Among the factors influencing Canadian banks noninterest income, Calmès and Liu (2009) find TSX – the Toronto Stock Exchange index – to be one of the most significant explanatory variables. Plotting *snonin* against TSX, Figure 4 actually displays an apparent cointegration between the two variables[14]. To confirm this, we run a cointegration test, first computing the residuals (*epsi*) associated to an OLS regression of *snonin* on *TSX* over the whole sample period (1988 to 2007), and then testing with the following regression

$$d(epsi) = c + \gamma\, epsi_{t-1} + \xi_t \quad (1)$$

where *d(epsi)* are the residuals expressed in first difference. The estimated coefficient γ of the lagged residuals is found significantly negative, suggesting that the residuals follow a stochastic mean-reverting process, and that *snonin* and *TSX* could be cointegrated. Consequently, we estimate the *snonin* factor model using the *levels* of these variables in equation (2). In addition to the *TSX*, the explanatory variables also include macroeconomic and financial factors such as *GDP* or the subprime crisis. More precisely, the *snonin* equation reads:

$$snonin_t = \theta_0 + \theta_1 \ln(TSX)_t + \theta_2 robg_{soc.t} + \theta_3 d\ln gdptrim_t + \theta_4 snonin_{t-1} + \theta_5 subpri + \varepsilon_t \quad (2)$$

where $robg_{soc}$ represents the corporate bond interest rate; *dlngdptrim*, the quarterly Canadian GDP growth rate, and *subpri* is a dummy variable accounting for the subprime crisis. Note that when the *TSX* is expressed in logarithm, the R^2 of the regression improves. This confirms the nonlinear relationship between *snonin* and *TSX*.

We estimate equation (2) with a sample composed of the eight major Canadian domestic banks, running from the first quarter of 1988 to the fourth quarter of 2007. At first, 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. However, our methodological choice, based on aggregate time series comprising more than 80 observations and a very parsimonious model, is more than enough to derive reliable results. Besides, the

[14] Not surprisingly, Figure 1 suggests that the cointegration between these two variables deteriorated in 2007, when the subprime crisis unfolded.

Canadian banking system has the reputation of being one of the most robust system in the world (Ratnovski and Huang 2009) so that any trace of bank instability there is likely indication of a widespread phenomenon. In other words, if we successfully document the emergence of a risk-premium associated to OBS activities, it should not be too surprising, applying the same methodology, to find the same pattern occurring elsewhere[15].

Data come from the Canadian Bankers Association and the Office of the Superintendant of Financial institutions (Canada). Since we identify a structural break in 1997, we also estimate equation (2) over the subsamples 1988-1996 and 1997-2007. Table 1 provides our results for the estimation of equation (2) using a GARCH(1,1) procedure. We test for other econometric specifications of conditional volatility, like GARCH(p,q), TARCH, EGARCH and PARCH, also using different distributions for the error term (normal, Student, and generalized error). However, based on the traditional measures of fit (e.g. Akaike, Schwarz and Hannan-Quinn criteria), we find that the standard GARCH(1,1) specification offers the best fit of conditional volatility. The estimated residuals confirm the presence of conditional heteroskedasticity. In line with the results obtained when testing for the cointegration between *snonin* and the *TSX,* the R^2 of the regression is very high for the whole sample (0.97). Note that, when regressing with first-order integrated explanatory variables, spurious results may unfold since the R^2 could be high, while, at the same time, the *DW* statistic would be low. However, in the case of our *snonin* regression, the *DW* is equal to 2.15, suggesting the absence of autocorrelation in the residuals. As expected, the stock market index is the most significant variable in equation (2). Over the whole sample, *snonin* appears very sensitive to the stock market conditions. The estimations thus suggest that banks revenues could be increasingly related to the stock market index.

Table 1 Estimation of the *snonin* equation by a GARCH (1,1) procedure

	1988-1996	**1997-2007**	**1988-2007**
c	-40.08**	-68.30**	-64.38***
ln(TSX)	6.95**	11.68***	10.28***
$robg_{soc}$	-0.31	-0.20	-0.43*
dlngdptrim	-0.20	0.41	0.04
$snonin_{t-1}$	0.55***	0.33**	0.52***
DUM2Q	-0.73	1.41	-0.13
DUM3Q	0.60	-3.83*	-1.63
DUM4Q	0.94	-3.57*	-1.58
SUBPRI		-3.44*	-3.75***
Adjusted R^2	0.88	0.78	0.97

Note: The explanatory variables are the following: *ln(TSX):* logarithm of the Toronto Stock Exchange index; $robg_{soc}$: Canadian corporate bond interest rate; *dlnGDPtrim*: quarterly growth rate of the Canadian GDP; $snonin_{t-1}$: share of noninterest income lagged one period; *DUMiQ*: seasonal dummies taking the value of 1 on the i[th] quarter and 0 otherwise (the first quarter being omitted); subpri: dummy variable accounting for the subprime crisis. This equation is estimated with a GARCH(1,1) procedure due to the presence of conditional heteroskedasticity.

The results also indicate that *snonin* is negatively related to the level of the corporate bonds interest rate. Indeed, the coefficient of $robg_{soc}$ is equal to -0.43 in the *snonin* equation, and significant at the 90% confidence level. Increases in the interest rate tend to depress bond issues with a corresponding decrease in banks revenues from service fees. They also slow down

[15] In this respect, Ratnovski and Huang (2009) mention that ‘’the Basel Accord requires international active banks to hold tier-one capital of at least 4 percent and total capital of at least 8 percent of risk weighted assets. Canada imposes capital requirements targets that are higher than the Basel minima: tier-one capital of 7 percent and total capital of 10 percent.’’

securitization, an important component of OBS activities. This negative sensitivity of *snonin* to the corporate bond rate might explain the observed negative correlation between the growth rates of net interest and noninterest revenues observed over the subperiods 1993-1997 and 2003-2007 (Calmès and Théoret 2009). It suggests that traditional banking activities may act as a buffer during periods of adverse returns on OBS activities. Relatedly, the sensitivity of *snonin* to the interest rate variable remains negative in the most recent period, although not significant over any particular subperiod.

In other respect, over the whole sample period, *GDP* growth does not seem to significantly influence *snonin*, even if its coefficient is positive. This suggests a procyclical behaviour for *snonin*. Incidentally, the variable associated to economic activity, *dlngdptrim*, is not significant in the *snonin* equation, regardless of the estimation period considered. This might be explained by the fact that the influence of the real sector variables is already incorporated in the stock market index – a leading indicator of economic activity. Not surprisingly, the 2007 subprime crisis has a very negative impact on *snonin*, especially for banks whose exposure to the asset backed securities market was high.

Turning to the comparison of the results obtained from the two selected subsamples, on the basis of the adjusted R^2, we first note that the performance of the *snonin* equation is lower in the second subperiod. This can be attributed to the increase in the banking idiosyncratic risk component observed since 1997, as *snonin* became more sensitive to the stock market during the second subperiod. Indeed, the estimated coefficients of the stock market variable are respectively 6.95 and 11.68 for the subperiods 1988-1996 and 1997-2007, and they are both significant at the 95% confidence level. The fact that *snonin* is increasingly sensitive to the stock market might suggest that Canadian banks are more exposed to systemic risk than they used to be[16]. This is symptomatic of the "new banking" era which emerged after the successive waves of financial deregulation the Canadian banking system underwent.

3. The Banks Returns Model and *Snonin* Endogeneity

3.1 The Banks Returns Model

Following Stiroh (2004a), the general formulation of the relationship between banks performance and noninterest income we study can be expressed as:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 snonin_t + \mathbf{X}_t \alpha + \varepsilon_t \quad (3)$$

where y_t is 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. bank size, riskiness of loans or asset growth). As a robustness check, and for the sake of consistency, equation (3) is first estimated on a risk-adjusted basis as was originally done in Stiroh (2004a) and Calmès and Liu (2009). For these regressions, y_t is divided by a fourth-quarter moving average of its standard deviation. But since one of the main arguments we develop in this study is the idea that the increased volatility of bank operating revenues might have given rise to the emergence of a risk premium in banks

[16] This fact has already been widely documented for other countries. cf. for example: Houston and Stiroh (2006), Wagner(2008, 2009), and Coval (2009a, 2009b). In particular, Wagner (2009), p.33, mentions that diversification in OBS activities ''makes banks more similar to each other by exposing them to the same risk.''He shows that this increases banking systemic risk, as when one bank fails it makes it more likely that others will follow.

returns, we also consider a modified version of equation (3) with a specification of the risk-return trade-off according with finance theory, such that

$$r_t = \psi_1 + \psi_2 risk_t + \mu_t \quad (4)$$

where r_t stands for return, $risk_t$ for a factor of risk to which r_t is exposed, and ψ_2, its degree of exposure.

For the latter set of experiments, two measures of banking risk are considered. First, in equation (3), we introduce the market risk premium, that is, the spread between the stock market return and the risk-free rate, so that equation (3), combined with equation (4), becomes

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 snonin_t + \mathbf{X}_t \alpha + \beta_3 (r_{TSX} - r_{TBILL}) + \varepsilon_t \quad (5)$$

where r_{TSX} stands for the return on the stock market index, and r_{TBILL} represents the Treasury bill interest rate. On the other hand, banks measure risk with VaR models. Consequently, we also consider an alternative specification of equation (5), where risk is measured by the conditional volatility of returns. To capture nonlinearities[17], this measure of banking risk is then analyzed with an IV ARCH-M procedure. More precisely, we introduce returns conditional volatility in equation (3) such that:

$$y_t = \phi_0 + \phi_1 y_{t-1} + \phi_2 snonin_t + \mathbf{X}_t \upsilon + \phi_3 \sigma_{c,t} + \varepsilon_t \quad (6)$$

where $\sigma_{c,t}$, the conditional volatility, is computed using the following equation

$$\sigma_{c,t}^2 = \pi_0 + \pi_1 \sigma_{c,t-1}^2 + \pi_2 \varepsilon_{t-1}^2 \quad (7)$$

3.2 The Treatment of *Snonin* Endogeneity

In this subsection, we detail how we construct the higher moments instruments endogenizing the *snonin* variable. Since the main focus of this paper is to study whether *snonin* endogeneity plays any role in explaining banks returns, we first want to introduce higher moments instruments for *snonin*, and then modify the returns model in several steps as discussed below.

3.2.1 Higher Moments Instruments

Fuller (1987) has shown how the higher moments of the explanatory variables may be used as instruments. To explain his developments in a simple setting, consider a one variable model such that: $y_t = \alpha + \beta x_t + \varepsilon_t,\ t = 1,2,...,n$, where $\varepsilon \sim N(0,1)$, and assume that $E\{x_t \varepsilon_t\} \neq 0$, i.e. x_t not being orthogonal to ε_t, and can be considered as endogenous. Assume 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. Assume also that the distribution of x_t is not normal but asymmetric and leptokurtic. Since the distribution of x_t is asymmetric, we have $E\left\{\left(x_t - \mu_x\right)^3\right\} \neq 0$, with μ_x, the expected value of x. This information may be used to build an explanatory variable for x. If we set $z_t = \left(x_t - \bar{x}\right)^2$, $\bar{x}$ being the mean value of x, then $E\left\{\left(x_t - \mu_x\right)\left(z_t - \mu_z\right)\right\} = \left(1 - n^{-1}\right) E\left\{\left(x_t - \mu_x\right)^3\right\} \neq 0$, and in accordance with the properties of the normal distribution: $E\{z_t \varepsilon_t\} = 0$. Thus, the second-order moment $\left(x_t - \bar{x}\right)^2$ qualifies as an

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[17] Regarding nonlinearities, note that there might exist an optimal level of the share of noninterest income. Indeed, the relationship between this variable and the banks idiosyncratic risk feeding the error term is nonlinear (cf. Baele et al. 2007, and Wagner 2008, 2009).

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

One key advantage of using these higher-moments instruments is that 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 (in equations (3) and (5)). For the treatment of *snonin* endogeneity, we thus use the following set of instruments **Z** = $\left\{ x_{t-1,} (x_t - \bar{x})^2, (x_t - \bar{x})^3, (y_t - \bar{y})^2, (y_t - \bar{y})^3 \right\}$, where x_t represents any of the explanatory variables of the banks returns model.

3.2.2 A Modified TSLS with the Hausman Endogeneity Test

To test for the endogeneity of *snonin*, we do not use the standard Hausman (1978) test but 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 IV estimator; and $Var\left(\hat{\beta}_{OLS}\right)$ and $Var\left(\hat{\beta}_{IV}\right)$, the respective variances of the estimated parameters, and *g* being the number of explanatory variables. The Hausman test measures the significancy 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. To avoid this issue, we resort to an alternative Hausman test. The modified version of the *h* test we propose is directly related to the works of Hausman (1978), Spencer and Berk (1981), McKinnon (1992) and Pindyck and Rubinfeld (1998). To implement this version of the Hausman test, we first rewrite the banks returns model without a risk premium:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 snonin_t + \mathbf{X}_t \alpha + \varepsilon_t \quad (8)$$

Since $E(snonin_t, \varepsilon_t) \neq 0$, *snonin* is an endogenous variable. A consistent estimator can be found if we can identify an instrument data matrix $\mathbf{Z} = \{z_1, z_2, ..., z_k\}$ – *k* being the number of instruments – to treat the *snonin* endogeneity. As discussed previously, this instrument set is the vector of higher moments **Z**.

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 $sn\hat{o}nin_t$,

$$snonin_t = \hat{c}_0 + \mathbf{Z}_t \hat{\rho} + \hat{w}_{snonin_t} = sn\hat{o}nin_t + \hat{w}_{snonin_t} \quad (9)$$

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 (equations (8)). This way we obtain consistent estimates of the coefficients of the returns equations.

In a second step, provided that there is no endogeneity concern, we can substitute equation (9) in equation (8) to obtain the following artificial (or auxiliary) regression

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

Finally, using equation (10), we can build our endogeneity Hausman test with higher moments. Despite the evidence gathered in section 2, 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 (10), and thus have

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

With $snonin_t = sn\hat{o}nin_t + \hat{w}_{snonin_t}$, we can reformulate equation (11) as follows:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 snonin_t + \mathbf{X}_t\alpha + \varphi \hat{w}_{snonin_t} + \varepsilon_t \quad (12)$$

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

The endogeneity test is then expressed as follows. If there is no endogeneity problem, then $\varphi = 0$, or equivalently $\theta = \beta_2$. On the other hand, if *snonin* is endogenous, then φ is significantly different from zero, that is to say $\theta \neq \beta_2$ in equation (11). One advantage of this procedure is that, beside 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 one estimated with the two-step Hausman procedure described above. According to equation (12), if $\hat{\varphi}$ is significantly positive, this is an indication 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 (12) are the same as those obtained from a TSLS procedure based on the same 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) in equation (12), the OLS estimator obtains, and equation (12) thus becomes

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

Instead, if $\hat{\varphi}$ is significantly different from zero (the case of endogenous issue), the TSLS estimator obtains and equation (12) reads:

$$\left(y_t\right)_{TSLS} = \hat{\beta}_0^* + \hat{\beta}_1^* y_{t-1} + \hat{\beta}_2^* snonin_t + \mathbf{X}_t\alpha^* + \hat{\varphi}\hat{w}_{snonin_t} + \varepsilon_t \quad (14)$$

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

$\hat{\varphi} = f\left(\hat{\beta}_{2,OLS} - \hat{\beta}_{2,TSLS}\right)$, which sates that $\hat{\varphi}$ is an indicator of the distance between the OLS and 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 (3), the standard TSLS procedure and this Hausman procedure are interchangeable. The estimated coefficients of the explanatory variables are the same in both cases. However, we favor the latter because it provides additional information about the endogeneity problem.

4. Empirical Results

In this section, we discuss the empirical results of the various experiments just described above, beginning with those of the estimation method most commonly used in the literature, namely the OLS, before presenting the results obtained from the IV method.

Table 2 OLS estimation of the banks returns models: 1988-2007

Canadian domestic banks

	***ROE*(1)**	***R0E*(2)**	***ROE*/$\sigma_{t,uc}$**	***ROA*(1)**	***ROA*(2)**	***ROA*/$\sigma_{t,uc}$**
c	0.24***	0.24***	125.16***	1.03***	1.04***	13.90***
y(t-1)	0.14**	0.14**	0.07***	0.11*	0.1	0.66***
snonin	0.11**	-0.11**	-1.63**	-0.39**	-0.43**	-0.16
LLP	-0.14***	-0.14***	-0.79***	-0.55***	-0.57***	-0.08***
$r_{TSX} - r_{TBILL}$		0.02			0.09	
Adjusted R^2	0.68	0.69	0.64	0.70	0.71	0.50

Note: Explanatory variables: y_{t-1}, lagged dependent variable; *snonin*, share of noninterest income in net operating revenue; *LLP*, ratio of loan loss provisions over total assets; $\sigma_{t,uc}$, unconditional volatility of the dependent variable computed using a rolling window of four quarters; r_{TSX}, return of the TSX index; r_{TBILL}, Canadian Treasury Bills interest rate. *ROE*(1) and *ROA*(1) are models without the market risk premium; *ROE*(2) and *ROA*(2) are models incorporating the market risk premium (r_{TSX} - r_{TBILL}). *Seasonal dummies were also included as regressors.* Finally, asterisks indicate the significance levels: * stands for 10%, ** stands for 5%, and *** stands for 1%.

4.1 OLS Results

Table 2 reports the results obtained from the OLS estimation of the banks returns model were returns here are proxied by either *ROE* or *ROA*, a standard approach in the literature. Since the results are very comparable with or without the introduction of a risk premium, we only discuss the case of the risk premium (the equation (5) specification of the returns model). We only keep the ratio of loan loss provisions to total assets as a control variable because the other ones are found not significant. The fit of the model seems quite good over the whole sample period, the adjusted R^2 being generally in the neighbourhood of 0.70. Consistent with the idea that loan loss provisions ought to lower profits, the coefficient of the ratio of loan loss provisions to total assets is found significantly negative in both returns equations, and equal to -0.14 for *ROE* and to -0.57 for *ROA*, this last coefficient being about four times higher than the former, as expected. This could be expected since the ratio of *ROA* to *ROE* is mean-reverting to a level of 4 over the period we analyse[18], and also because of the low volatility of this ratio. Since the ratio of loan loss

[18] For instance, for the fourth quarter of 2007, the *ROE* and *ROA* of the eight major Canadian banks were respectively 0.18 and 0.79. Given the tight relationship between the estimated coefficients of the *ROE* and *ROA* equations, we often limit our attention to *ROA*.

provisions jumps during recessions, it magnifies the procyclicality of *ROA*, which, in other respects, become more procyclical after 1997, following the banks increased involvement in OBS activities.

Table 2 also shows that the risk-return trade-off worsened throughout the sample. The coefficients of *snonin*, significant at the 95% confidence level in both *ROE* and *ROA* equations, are respectively -0.11 and -0.43[19]. These results echoe those of Stiroh and Rumble (2006) regarding the sensitivity of banks returns to *snonin*, a priori casting doubts on the belief that noninterest income activities can lead to better bank performance through activities diversification (reduction in risk and/or higher returns).

In other respect, the coefficient of the market risk premium is positive but not significant for both *ROE* and *ROA*. At first, this result might appear surprising in light of the observed cointegration between *snonin* and the stock market index. In fact, this lack of significance of the market risk premium is directly related to the endogeneity of *snonin*, precisely the aspect hard to capture with the OLS benchmark estimation. In this form, the model accounts for systemic risk, but a preferable procedure consists in deflating the banks performance measures with their respective volatilities (cf. Stiroh and Rumble 2006). Doing so, Table 2 reveals that the sensitivity of banks returns to *snonin* is negative for both the *ROA* and *ROE* risk adjusted measures, although significant for *ROE* only. This confirms that directly integrating a risk measure in the banks returns equations, and especially a global risk measure, might be a more convenient way to account for banks risk.

Table 3 OLS estimation of the banks returns models: 1988-1996

Canadian domestic banks

	***ROE*(1)**	***R0E*(2)**	***ROE*/$\sigma_{t,uc}$**	***ROA*(1)**	***ROA*(2)**	***ROA*/$\sigma_{t,uc}$**
c	0.45***	0.47***	18.62	1.85***	1.94***	21.17
y(t-1)	0.03	0.02	0.76***	0.04	0.03	0.75***
snonin	-0.71***	-0.76***	-0.35	-2.79***	-3.02***	-0.40
LLP	-0.15***	-0.16***	-10.32***	-0.61***	-0.62***	-0.11***
$r_{TSX} - r_{TBILL}$		0.04***			0.14**	
Adjusted R^2	0.95	0.96	0.60	0.96	0.96	0.59

Note: see table 2

Table 4 OLS estimation of the banks returns models: 1997-2007

Canadian domestic banks

	***ROE*(1)**	***R0E*(2)**	***ROE*/$\sigma_{t,uc}$**	***ROA*(1)**	***ROA*(2)**	***ROA*/$\sigma_{t,uc}$**
c	-0.02	-0.02	20.19	-0.01	0	4.80
y(t-1)	0..21	0.21	0.80***	0.08	0.09	0.65***
snonin	0.39**	0.37**	1.32	1.59***	1.53***	0.02
LLP	-0.12***	-0.12**	-3.04	-0.52***	-0.45**	0.01
$r_{TSX} - r_{TBILL}$		0.01			0.05	
Adjusted R^2	0.47	0.46	0.67	0.45	0.45	0.44

Note: see table 2

[19] As expected, the estimated coefficient of ROA is again four times higher than the one of ROE.

Since the data suggest the presence of a structural break in 1997, we also run our regressions over two subsamples, 1988-1996 and 1997-2007. Tables 3 and 4 provide the results obtained for these two subperiods. To begin with, note that the growth in the banks new business lines should imply a deterioration of the model performance when shifting from the first subperiod to the second. Indeed, we observe that the idiosyncratic risk prevailing in the second subperiod, as suggested by the volatility of the banks income growth, is much more pronounced, and feeds into the innovation term of the equation. It is during this second subperiod that Canadian banks begun to better integrate their traditional bank lending activities with their new banking business. When banks begun to operate these non-traditional activities, their risk-return trade-off worsened. It took them almost ten years to finally adjust to the new banking environment. The data track this change in the banking environment quite well. For example, in the *ROA* equation, the adjusted R^2 is equal to 0.96 over the first period, and then falls to 0.45 in the second subperiod, corroborating the deterioration of the model fit.

On the other hand, while banks non-traditional activities were developing, we also observe a change in the sign of the *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. Consistent with this argument, we find that *snonin* is only significantly negative during the subperiod 1988-1996, and becomes significantly positive after 1997.

Finally, in line with what we find over the whole sample, note that the sensitivity of *ROA* to the market risk premium is significant, but quite low in the first subperiod, and no longer significant in the second one. Actually, adding the market risk premium in the *ROA* equation does not improve the fit of the regression in any subperiod. We argue that this low sensitivity of *ROA* to the market risk premium is related to the endogeneity of *snonin*[20].

Table 5 Comparison of OLS and GMM for estimating *ROA* equation with the market risk premium

	OLS		GMM	
	1988-1996	**1997-2007**	**1988-1996**	**1997-2007**
c	1.94***	0.00	2.08***	0.02
β_2	-3.02***	1.53***	-3.39***	1.45***
β_3	0.14**	0.05	0.08***	0.15***
Adj R^2	0.96	0.45	0.93	0.37

Note: the estimated coefficients in this table correspond to equation (5).

4.2 Estimation of the *ROA* Equations with GMM

Table 5 compares the results of equation (5) obtained with OLS and GMM estimations. Focusing on the estimated coefficients of the share of noninterest income (*snonin*) and the market risk premium, the GMM estimation with higher moments instruments indicates that the sensitivity of ROA to *snonin* is equal to -3.39 during the first subperiod 1988-1996, a higher level compared to what obtains with the OLS benchmark. The GMM method thus suggests a greater deterioration of the banks risk-return trade-off in the first subperiod. Corroborating our previous findings, during the second subperiod, the coefficient of *snonin* becomes significantly positive and quite comparable to the one obtained with the OLS estimation. Overall, the two estimation methods confirm a deterioration of the banks risk-return trade-off, followed by an improvement, a phenomenon consistent with the idea that banks went through a period of adjustment when they slowly adapted to their changing environment and integrated their new lines of businesses.

[20] As exposed below, the alternative measure of risk we propose, i.e. the returns volatility, is much more adequate to account for banks risk exposure.

Note also that with OLS, the estimated coefficients of the market risk premium variable are significant only in the first subperiod, whereas the estimates obtained from GMM are significant for both subperiods. The sensitivity of ROA to the market risk premium has actually doubled in the second subperiod relative to the first one[21].

Our results compare well with those obtained on US data by Stiroh and Rumble (2006). However, in the regressions these authors run, the coefficients of the *snonin* variable are positive and insignificant when expressed in levels, but significantly negative when scaled by their volatilities. Instead, when we apply their equations on Canadian data over the period they use (1997-2002), we find positive and significant coefficients for the *snonin* regressor with returns expressed in levels, but *positive* and insignificant coefficients for returns scaled by their volatilities ─ although they are significant with the GMM procedure. At first, the results Stiroh and Rumble (2006) obtain from the regressions on scaled returns might seem somewhat puzzling. When moving from returns expressed in levels to scaled returns, since risk, although imperfectly, is taken into account in the endogenous variables, their negative sensitivity to the *snonin* ratio should decrease, not shift from insignificantly positive to significantly negative. The results of Stiroh and Rumble (2006) are likely attributable to the lack of stability and high short term volatility of scaled returns. Their findings on the insignificancy of the *snonin* coefficient for returns expressed in levels may also relate to their general treatment of the *snonin* endogeneity. As our results suggest, thoroughly accounting for endogeneity strengthens the impact of *snonin* on banks returns, leading to more significance and/or higher coefficients (in absolute value) for *snonin* in the returns equations.

4.3 *Snonin* Endogeneity and the TSLS Results

Regarding the endogeneity of the *snonin* variable, we report our results of the TSLS estimation for *ROA* (equation (12)) in Table 6. As previously explained, the Hausman procedure is very similar to a regular TSLS estimation[22]. However, the Hausman regressions are preferable since they directly embed an endogeneity test based on the significancy of w_{snonin}, as measured by its t-statistic, and provide an indication of the severity of the endogeneity issue with the level of the w_{snonin} coefficient.

Table 6 Modified Hausman endogeneity test for *ROA*

	1988-1996			1997-2007		
	OLS	HAUS	TSLS	OLS	HAUS	TSLS
c	1.94***	2.00***	2.00***	0.00	0.17	0.17
ROA_{t-1}	0.03	0.03	0.03	0.09	0.13	0.13
snonin	-3.02***	-3.19***	-3.19***	1.53***	1.16**	1.16**
LLP	-0.62***	-0.62***	-0.62***	-0.45**	-0.49***	-0.49***
$r_{TSX} - r_{TBILL}$	0.14**	0.15**	0.15**	0.05	0.06	0.06
w_{snonin}		4.93**			2.50*	
Adjusted R^2	0.96	0.97	0.96	0.45	0.48	0.44

[21] Our alternative risk measure based on the return volatility corroborates this increased *ROA* risk sensitivity from one subperiod to the other.

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

As reported in Table 6, the level of the w_{snonin} coefficient indicates that the estimated coefficient of *snonin* is significantly overstated by OLS, regardless of the period considered. This result is consistent with the view that the *snonin* variable is indeed endogenous. According to the result obtained from the Hausman artificial regression, this problem seems to be more important during the first subperiod, when the *snonin* variable impacts negatively banks returns. In the *ROA* regression, the coefficient of w_{snonin} is equal to 4.93 for this subperiod, compared to 2.50 for the second one. Furthermore, during the first subperiod, the coefficient of *snonin* is equal to -3.02 and becomes -3.19 if we account for the endogeneity of *snonin*. Taking endogeneity into account strengthens the results previously found in the literature, in particular the fact that *snonin* has a negative impact on banks returns in the 1988-1996 subperiod. More importantly, our results suggest that *snonin* has a positive impact in the more recent episodes. If endogeneity is considered, the impact of *snonin* is quite important in absolute value. Over the second subperiod (1997-2007), correcting for endogeneity slightly reduces the positive influence of *snonin* on returns, but its positive impact proves to be quite important and significant. This finding is consistent with the idea that the recent period has been characterized by a better integration of OBS activities with the traditional ones, and is also new evidence of the diversification benefits derived from banks non traditional operations.

5. Alternative Measures of Banks Risk

In the case of the banking industry, the idiosyncratic risks institutions usually face is not totally diversifiable (e.g. Stiroh 2004a) and thus must be priced accordingly. In this respect, a new strand of the financial literature argues that idiosyncratic risk is not totally diversifiable (Xu and Malkiel 2004, Hirth and Pandher 2005, Goyal and Santa-Clara 2007, Ebert and Lütkebohmert 2009, Gabaix 2009). Consequently, the returns volatility might be considered a better measure of risk than the market risk premium we used so far, because the latter only accounts for systemic risk, whereas the former provides a more comprehensive assessment of risk. According to Markowitz, returns volatility overstates risk in the case of an individual stock or of an undiversified portfolio. Yet, in the presence of granular idiosyncratic risk, returns volatility is still preferable to market risk premium to measure banking risk[23].

In addition to these motives for adopting returns volatility as an alternative measure of banks risk, another consideration which justifies its introduction comes from the fact that, in 1997, following the recommendations of the Basle Committee, VaR became the official banks risk measure. Since the most important element of the computation of this statistic is precisely the volatility of returns, to be consistent with banks view on risk, it is all the more relevant to consider an alternative measure of risk based on volatility. Incidentally, VaR computation can be extended to account for the "granularity" related to fat tails[24]. Insofar as VaR may be adjusted to become a more robust measure of banks risk, it is interesting to analyze the impact of returns volatility, viewed as an indicator of VaR, and analyze the impact of returns volatility on banks returns. We thus follow Calmès and Théoret (2009) and consider another measure of risk based on the conditional volatility of banks returns.

As we mentioned previously, the GMM estimations reveal that the impact of the market risk premium on ROA is quite small. This also tends to suggest that a risk measure accounting only for systemic risk is not ideal for explaining banks returns. For all these reasons,, we study the endogeneity of *snonin* with a modified ARCH-M procedure à la Engle et al. 1987, resorting to alternative measures of banking risk based on a more comprehensive view (i.e. the integration of the two dimensions of risk, the systemic and the idiosyncratic components). The ARCH-M approach allows the simultaneous estimation of the conditional return volatility and the coefficient

[23] However, note that volatility neglects fat-tails risk.

[24] Another statistic which adjust VaR for fat-tails is the expected tail loss (ETL). On this matter, see, for example, Fabozzi et al. (2007).

related to this volatility in the return equation. It captures the nonlinearities created by the increased idiosyncratic risks observed in the Canadian banking industry after 1997. This procedure is very appealing to estimate the risk premium in this context because it directly incorporates the conditional volatility, our measure of risk, in the return equation, instead of running a regression on returns defined on a risk-adjusted basis, i.e. a measure of return scaled down by an "ad hoc" measure of its volatility.

In the case of the banks measures of risk based on returns volatility, the new benchmark model features a standard ARCH-M estimation, which, in its basic form, does not treat endogeneity. To account for the endogeneity of *snonin*, we then modify the benchmark model in two steps. First, much in the spirit of what we do with the modified Hausman test, we regress the *snonin* variable on the higher moments instruments, **Z**, and then, we introduce the resulting fitted value of *snonin* obtained from this regression in the ARCH-M equation. Thus, the modified ARCH-M we propose is an IV ARCH-M procedure which accounts for the potential endogeneity of *snonin*.

Table 7 Comparison of ARCH-M estimations with *snonin* endogenous and exogenous

	ARRCH-M *snonin* exo		**ARCH-M *snonin* endo**	
	1988-1996	**1997-2007**	**1988-1996**	**1997-2007**
c	1.71**	-4.04***	1.90***	-1.96
ϕ_2	-2.69***	1.24***	-3.70***	1.97***
ϕ_3	1.26	5.62***	2.07***	3.87***
Adj R^2	0.95	0.43	0.98	0.52

Note: the estimated coefficients in this table correspond to equation (15).

Table 7 reports our estimation results with this IV ARCH-M technique. As the table shows, the results obtained are more significant than those reported in Table 5 for the GMM procedure. When shifting from the OLS estimation of the ROA equation to the IV ARCH-M, the adjusted R^2 increases from 0.45 to 0.52 over the subperiod 1997-2007. If estimated with an IV ARCH-M procedure instead of the basic OLS estimation, the *snonin* variable displays a greater impact on banks returns. For instance, over the 1988-1996 subperiod, the estimated coefficients for *ROA* are respectively, -3.70 and -2.79. Of course, a comparison of the results obtained with the IV ARCH-M procedure with those of the benchmark ARCH-M is more appropriate. Doing so, we can first note that, regardless the subperiod examined, the impact of *snonin* on *ROA* is found significantly greater with the IV ARCH-M. For instance, over the 1988-1996 subperiod, the coefficient of *snonin* is equal to -2.69 in the ROA regression with the benchmark estimation, but to -3.70 with the IV ARCH-M procedure. We observe the same phenomenon for the estimated coefficient of this variable over the 1997-2007 subperiod, and further confirm the findings obtained in the previous experiments, observing the same change in the *snonin* sign. The estimated coefficient is equal to 1.24 with the standard ARCH-M estimation, but 1.97 with the IV ARCH-M.

Regarding the estimation of the risk premium, the results reported in Table 7 also illustrate how good a job the ARCH-M can do for capturing the nonlinearities created by the volatility jumps observed in banks revenues between the 1988-1996 period and the more recent episode of 1997-2007. Indeed, these nonlinearities pile up in the error term of the returns equations, causing the residuals to be nonnormal and heteroskedastic. Consequently, the coefficient of the conditional volatility estimated with the IV ARCH-M procedure is much higher, going from 2.07

to 3.87 over the second subperiod, and significant at the 99% confidence level. This is indication that the Canadian chartered banks increased substantially the risk premium they associate to their OBS activities over the second subperiod[25]. Obviously, the results also relate to the jump in risk witnessed during the second subperiod (as measured by the increased volatility of noninterest income growth).

Finally note that accounting for endogeneity also increases the sensititivity of ROA to its conditional volatility in the *first* subperiod. Indeed, the coefficient of the conditional variance is not significant when no account is made for endogeneity, but becomes significant when endogenizing *snonin*. Therefore, ignoring the endogeneity of *snonin* creates the false impression that there was no significant risk premium associated to OBS activities over the first subperiod, despite the increase in the volatility of the income growth generated by these activities at the time. Allowing *snonin* to be endogenous leads to the opposite conclusion. Even if much smaller than during the recent period, a significant risk premium already prevailed in the 1988-1996 subperiod, already beginning to compensate for the higher volatility of noninterest revenues feeding banks income. In other words, if we account for the endogeneity of *snonin,* the deterioration observed in the banks risk-return trade-off appears less severe than previously thought.

6. Conclusion

In this paper we revisit the banks decision to diversify in OBS activities, considering as endogenous the share of noninterest income these new lines of business generate. The results obtained when applying an Hausman artificial regression based on higher moment instruments suggest that the endogeneity of *snonin* may change the nature of banks returns sensitivity to noninterest income. In particular, the influence of *snonin* on *ROE* and *ROA* is stronger than what was previously reported in the literature. We find that *snonin* has a more negative impact on returns over the first subperiod, and a more positive effect over the more recent period, suggesting that the banks risk-return trade-off was deteriorating over the first subperiod, before recovering in the second.

In other respects, we find that the risk premium, an alternative way of incorporating banks risk in the returns equations, has more impact on banks returns when considering *snonin* as endogenous. Not accounting for the endogeneity of this variable leads to weak and/or non-significant risk premia in the returns equations, while they become significant when considereing *snonin* endogeneity explicitly. The IV ARCH-M method we introduce to estimate this risk premium performs particularly well for capturing the nonlinearities created by the increased volatilities of *snonin* and banks income growth. Prolonging Stiroh (2004a) idea, focusing on a risk premium measured with the specific conditional volatility of banks returns is found more appropriate than using the market risk premium, so usually done to measure the sensitivity of banks returns to risk. In our case, the IV ARCH-M delivers higher and more significant risk premia if endogeneity is properly accounted for. Incidentally, when assuming *snonin* as exogenous, the risk premium is found not significant in the first subperiod (1988-1996), although the adjustment to the new banking environment was on its way. Instead, consistent with the idea that the shadow banking begun to emerge in the first subperiod, the IV ARCH-M regression suggests that, even if smaller, a risk premium was already present at the time, pricing the greater volatility of the OBS activities that were increasingly contributing to banking volatility.

As a follow-up to the question we address in this paper, it would be interesting to study the *snonin* endogeneity problem based on the components of noninterest income. We suspect that the income related to securities trading, which much depends on the stock market, is the principal

[25] Note that, although based on a different methodology, the results derived from these IV ARCH-M experiments can be considered as a corroboration of the results Stiroh (2006) obtains with U.S. data. In particular, he writes (p. 239) : ‘’The expansion of banking powers led equity market investors to become more cognizant with new and evolving activities and look off of the balance-sheet to identify them.’’

cause of the endogeneity of *snonin*. Another research avenue would focus on the nonlinearities created by the increased idiosyncratic risk generated by OBS activities. In this case, the Kalman filter could be used to analyze further the properties of the coefficients (their cyclical movement) of the returns regressions. These questions are left for future work.

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