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The Rise of Market-oriented Banking and the Hidden Benefits of Diversification

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Abstract

The diversification benefits associated with banks' off-balance-sheet activities (OBS) is a question much debated in the literature. The emergence of market-oriented banking greatly contributes to the volatility of bank operating revenues, but the link between market-based activities and accounting returns is less clear (Stiroh and Rumble, 2006). In this paper, we apply a new empirical framework to a Canadian dataset and confirm that OBS activities have indeed fuelled the surge in bank income volatility and the increase in the risk-adjusted return on assets (*ROA*). However we argue that this result is not only explained by the presence of a cointegrating relationship between financial market returns and the share of non-interest income (*snonin*) OBS generate, but also by the endogenous link between *ROA* and *snonin*. Our main contribution is to show that the positive impact of *snonin* on *ROA* is reinforced when the endogeneity issue is properly accounted for. In particular, we find that the interdependence between *snonin* and *ROA* has clearly increased with the progressive diversification of banks in market-oriented business lines.

Keywords:

Market-oriented banking, Non-interest income, Hausman test, Structural break, Endogeneity.

JEL classification:

C32; G20; G21.

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El auge de la banca 'Market-Oriented' **y los beneficios ocultos de la diversificación**

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Resumen

Los beneficios de la diversificación de las actividades de los bancos fuera de balance (OBS) es una cuestión ampliamente debatida en la literatura. La aparición de las actividades bancarias "market-oriented" contribuye en gran medida a la volatilidad de los ingresos operacionales de los bancos, pero el vínculo entre las actividades "market-oriented" y los retornos contables es menos claro (Stiroh and Rumble, 2006). En este artículo se aplica un nuevo marco empírico a una base de datos canadiense y se confirma que las actividades fuera de balance, de hecho, han avivado el aumento de la volatilidad de los ingresos bancarios y el incremento de la rentabilidad ajustada al riesgo sobre los activos financieros (*ROA*). Sin embargo, sostenemos que este resultado no solo se explica por la presencia de una relación de cointegración entre los rendimientos de los mercados financieros y la proporción de los ingresos no financieros (*snonin*) que generan las actividades fuera de balance, sino también por el vínculo endógeno entre *ROA* y *snonin*. Nuestra principal contribución es mostrar que el impacto positivo del *ROA* en el *snonin* se ve reforzado cuando se tiene en cuenta, de manera apropiada, la relación endógena anteriormente mencionada. En particular, encontramos que la interdependencia entre *snonin* y *ROA* ha aumentado claramente con la progresiva diversificación de los bancos en las líneas de negocio orientadas al mercado.

Palabras clave:

Banca "market-oriented", ingresos no financieros, test de Hausman, cambio estructural, endogeneidad.

1. Introduction

Banks' off-balance-sheet (OBS) activities (e.g., securitization and loan commitments) have fuelled the last lending boom, enabling banks to increase their operational funding. This eventually led to a typical 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. This new type of banking presents a considerable challenge to central banks and regulators. In the context of the new banking era, it becomes crucial for practitioners to fully understand the behaviour of OBS activities. What we know so far is that the increase in bank non-traditional activities has had a significant influence on bank risk-return trade-off (DeYoung and Roland, 2001; Estrella, 2001; Acharya 2002; Clark and Siems, 2002; Stiroh, 2004; Vander Vennet *et al.*, 2004; Stiroh and Rumble, 2006; Baele *et al.*, 2007). International evidence suggests that it triggered a substantial increase in the volatility of bank net operating revenue growth, but the associated increase in returns is less clear (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 of the revenues generated from market-oriented activities, i.e. non-interest income, rapidly became a key, procyclical determinant of bank profits after 1997 (Calmès and Théoret, 2010).

However, the diversification benefits provided by these activities still remain a debatable question. For example, Slaikouras and Wood (2003), Smith *et al.* (2003), De Young and Rice (2004), Chiorazzo *et al.* (2008) and Lepetit *et al.* (2008) find a negative correlation between non-interest income and net interest income, while other authors report a positive link between these two income streams (e.g., Vander Vennet *et al.*, 2004; Stiroh, 2004; Stiroh and Rumble, 2006; Mercieca *et al.*, 2007; Calmès and Liu, 2009). If non-interest income is imperfectly correlated with traditional intermediation revenue, it could potentially deliver some diversification benefits. Nevertheless, several considerations might hide the true benefits of diversification. First, banks have expanded in OBS activities partly to react to the declining performance of their traditional activities (Boyd and Gertler, 1994; Smith *et al.*, 2003; Laeven and Levine, 2007; Busch and Kick, 2009; Cardone-Riportella *et al.*, 2010). Second, the diversification benefits associated with the rise of *snonin*[1] may also induce banks to take additional risks since market-oriented banking and financial engineering enhance their risk-sharing and risk-shifting capacities (Demsetz and Strahan, 1995; Vander Vennet *et al.*, 2004; Wagner and Marsh, 2004; Stiroh, 2006; Shin, 2009)[2]. In other words, there is a "reverse causality" problem at

[1] *To be precise, snonin is a measure of the degree of functional diversification (Vander Vennet et al., 2004).*

[2] *Note also that OBS activities tend to be more leveraged than traditional ones, which per se increases bank risk (DeYoung and Roland, 2001). For instance, fee-based activities require less regulatory capital and also increase banks' fixed costs via labour expenses (rising operational leverage).*

play (Stiroh and Rumble 2006). By intensifying the interaction between bank profits and financial markets, OBS activities increase the fluctuations of bank revenues and do not necessarily translate into stability gains for their bundled income (Wagner, 2006, 2010; DeJonghe, 2009). To shed new light on the diversification debate, it is thus crucial to carefully examine this simultaneity bias and the endogeneity of the *snonin-ROA* relationship.

To our knowledge, the literature does not provide any rigorous evidence about the evolution of the link between the share of non-interest income (*snonin*) and bank returns during the transition period the banking business 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 market-oriented banking, is associated with a change in the degree of OBS endogeneity. More precisely, if bank non-traditional activities are better integrated and provide diversification gains to the banking business, we should expect a closer link between *snonin* and *ROA*. To check this conjecture, we introduce a new approach based on a modified version of the Hausman test specifically designed to gauge the *changes* in the endogeneity of bank decision to expand their market-based activities.

Another important motivation for adopting this new approach comes from the fact that treating endogeneity too casually can leave spurious correlations between *snonin* and unobservables not accounted for in bank returns equations[3]. In particular, the remaining non-orthogonality of *snonin* with the innovation in the returns equations can cause serious biases in the parameters estimates and may even yield misleading results (Campa and Kedia, 2002; Stiroh, 2004). 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 rely on various control variables and other techniques to deal with this issue (Graham *et al.*, 2002; Villalonga, 2004; Laeven and Levine, 2007; DeJonghe, 2009)[4]. However, this kind of approach does not completely alleviate the problem. In particular, it is not suited to investigate the *changes* in the relative contribution of non-interest income to bank profits. To our knowledge, the studies of DeYoung and Rice (2004), Goddard *et al.* (2008) and Busch and Kick (2009) are the only ones using conventional instrumental variables to treat the endogeneity of *snonin*. We depart from their method by introducing an *h* test based on an artificial regression equivalent to a two-stage least squares (TSLS) procedure. The key advantage of this procedure is that it provides a direct measure of the changing biases in the endogenous variable coefficient associated with non-interest income, while also delivering robust instruments built

[3] *Unobservables like managerial performance, see Campa and Kedia (2002).*

[4] *See also Fluck and Lynch, 1999; Chevalier, 2000; Lamont and Polk, 2001; Maksimovic and Phillips, 2002; and Schmid and Walter, 2009.*

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with the higher moments of the explanatory variables (Fuller, 1987; Lewbel, 1997; Racicot and Théoret, 2008, 2012; Meng *et al.*, 2011).

We apply this framework to a Canadian dataset running from the first fiscal quarter of 1988 to the second fiscal quarter of 2010[5]. 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 risk-adjusted returns on assets and equity after 1997. This structural break coincides with a sharp increase in the volatility of bank net operating revenues growth and in the ratio of non-interest income. The year 1997 is a plausible break since it is around this date that banks begun to modify significantly the mix of their OBS activities in Canada, giving a much greater weight to their market-based operations like trading and capital markets operations. More importantly however, compared to the previous studies, our main result suggests that endogeneity, which was negligible before 1997, *increases* substantially thereafter, a fact so far overlooked in the literature. More precisely, the link between *ROA* and *snonin* becomes stronger and much more significant after the structural break, a result reinforced when endogeneity is properly accounted for. The ordinary least-squares (OLS) estimations tend to understate the sensitivity of *ROA* to *snonin* during the second subperiod, 1997-2010.

By contrast, the new evidence we gather based on TSLS indicates a marked increase in the endogeneity of non-interest income, strongly supporting the view that the bank regime shift might be more persistent than previously thought. On the one hand, our findings indicate that banks have benefited from this business trend, registering greater risk-adjusted returns. On the other hand, given the volatility of *snonin*, the fact that bank returns are increasingly tied to *snonin*, and the related evolution of *snonin* endogeneity provide a direct empirical evidence of the changes in bank systemic risk. Indeed, compared to the existing literature (e.g., Stiroh and Rumble, 2006; Lepetit *et al.*, 2008; Busch and Kick, 2009), our detailed decomposition of *snonin* reveals that in times of financial crisis the benefits of diversification seem to vanish (although securitization and insurance still act as buffers). Consequently, we are inclined to think that regulatory agencies should scrutinize OBS activities more carefully to systematically monitor their dynamic interaction with traditional banking activities.

This paper is organized as follows. In the next section we present the data and some basic stylized facts about the evolution of non-interest income. In section 3 we expose the bank returns model and the modified Hausman method we introduce to monitor

[5] *Note that the involvement of Canadian banks in OBS activities was quite restricted before 1987, banks not being allowed to get involved in investment banking until this date. For example, before 1987, Canadian banks reported very low commissions. For more details see Théoret (2000) and Calmès (2004).*

the change in the endogenous link between non-interest income and *ROA*. The fourth section details our results and the fifth section provides a discussion before concluding with some straightforward policy implications.

2. The Change in Non-interest 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 reasonable number 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 domestic banks, which, taken together, account for 90 percent of the banking business. All of them are chartered banks, i.e. commercial banks regulated by the Bank Act, running a broad range of activities, from loan business to investment banking, fiduciary services, financial advice, 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[6]. 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 market-oriented 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 U.S. 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 and very parsimonious models, is more than enough to derive robust results.

[6] *For this reason, relying on the "state of the art" z-score to assess risk-adjusted returns could prove misleading. Since the conventional measures of z-scores are close to Sharpe ratios and standard measures of risk-adjusted returns, we rely on these measures instead.*

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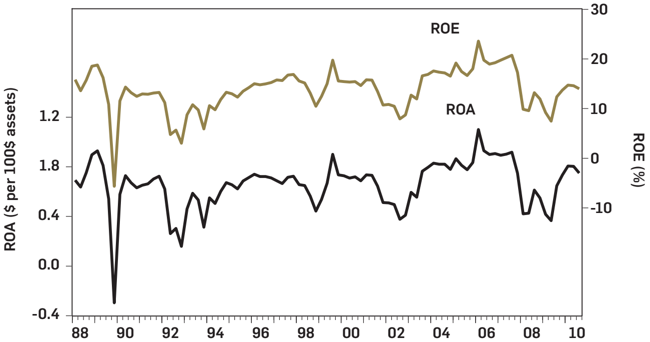
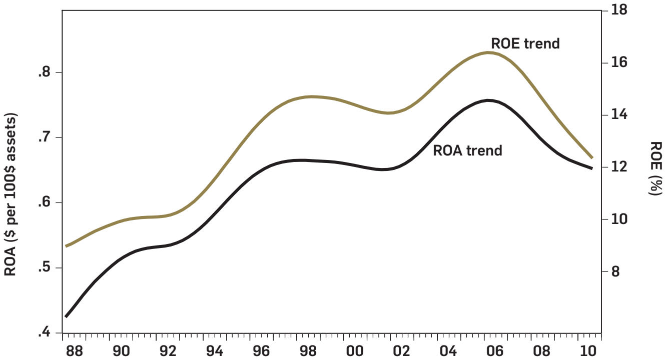
2.2. The Evolution of the Non-interest Income Series

■ Figure 1. ***ROA* and *ROE* Levels and Trends.**

Levels



Hodrick Prescott Trends



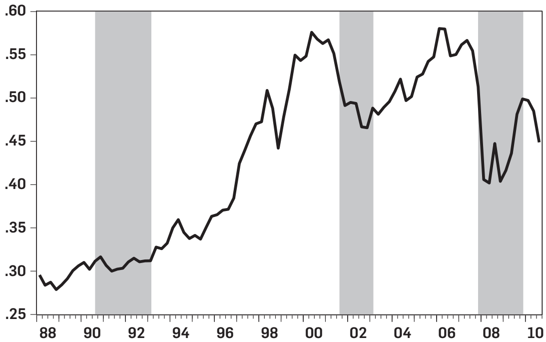
SOURCE: CANADIAN BANKERS ASSOCIATION.

Figure 1 presents the evolution of the performance of the banking system during the whole sample period. First note that bank returns, as measured by the return on assets (*ROA*) and the return on equity (*ROE*) share a very close relationship[7]. 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 also presumably relates to the better integration of bank traditional and OBS activities. Consistent with this conjecture, Figure 2 illustrates the growth in the bank non-interest income share (in net operating revenues). By 2000, non-interest income accounts for 57% of net operating revenue, up from only 25% in 1988. This ratio seems to stabilize thereafter, as the new banking business lines matured. More importantly, note that the fluctuations of *snonin* are

[7] *Given the high correlation between ROE and ROA we only report the ROA results.*

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[8].

■ Figure 2. **Share of Non-interest Income, 1988-2010**



Note: Shaded areas correspond to periods of contractions or marked economic slowdown.
SOURCE: CANADIAN BANKERS ASSOCIATION.

● 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 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**
Non-interest 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 $w^T\Omega w$ *, where* w *is the vector of the respective shares of net interest income and non-interest income in bank net operating revenue, and* Ω *is the variance-covariance matrix of net interest income growth and non-interest income growth.*
DATA SOURCE: CANADIAN BANKERS ASSOCIATION AND BANK OF CANADA.

Since the volatility of *snonin* contributes to the volatility of operating income, we should expect an increase in bank 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. The adoption in 1997 of the Value-at-Risk (VaR) as the standard bank risk measure might also contribute to the increased income growth volatility because of the tendency of this risk measure to underestimate the negative impact of fat tails[9]. But the increasing share of non-interest income is surely another important factor to understand the change in bank net operating

[8] *We run a Chow test confirming this structural break. See also Calmès and Théoret (2009). A discussion of additional tests follows.*

[9] *Fat tails risks, which are related to the kurtosis of the returns distributions, are generally much higher than the risks associated with the variance (DeJonghe, 2009).*

income (Stiroh, 2004; Lepetit *et al.*, 2008). As a matter of fact, during the first subperiod, net interest income contributes the most to the variance of net operating income, but after 1997 the rise in the variance of bank net operating income is due for the most part to the increased volatility of non-interest income (Table 1). For instance, from the subperiod 1988-1996 to the subperiod 1997-2010 the variance of net operating income growth increases from 11 to 66.3, and the absolute contribution of non-interest income increases from 3.0 to 63.4. Relatedly, Figure 3 illustrates the behaviour of the moving average variance of net operating income growth and its two components (net interest income and non-interest income) and reveals that 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, we follow Stiroh (2004) and provide the descriptive statistics of the components of bank non-interest income over the period 1997-2010[10] (Table 2). First, 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 4 confirms that these two components indeed drive the fluctuations of the moving average variance of non-interest income over the period 1997-2010[11]. Note also that while the variances of the two components are relatively moderate in general, they increase sharply during contraction episodes. This is the reason why it is often assumed that more diversified activities do not necessarily make banks safer (Vander Vennet *et al.*, 2004; Vives, 2010; Wagner, 2006, 2010; De Jonghe, 2009).

When looking at the decomposition of the variance of non-interest income growth, Table 3 shows that 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 non-interest income only amounts to 11%. The remaining variance is mainly explained by the capital market income component. In other words, the fluctuations of non-interest income growth seem to be mainly explained by the two components most related to bank market-oriented activities. Importantly, note that the high relative contribution of these two components suggests that the diversification benefits that they could still bring might be low[12]. In terms of diversification benefits, *ceteris paribus*, securitization and insurance revenues should offer the best perspective.[13]

[10] *Statistics on the components of non-interest income are not available before this date.*

[11] *Due to the presence of negative numbers and the sharp fluctuations of the trading income series, we compute the moving average variances on the levels of the series rather than on their growth rates.*

[12] *This conjecture is analyzed in more details in the discussion section.*

[13] *Insurance is often mentioned as an activity which provides substantial diversification benefits to banks (Boyd and Graham, 1988; Kwan and Laderman, 1999; Estrella, 2001; Vander Vennet et al., 2004; DeJonghe, 2009; Schmid and Walter, 2009).*

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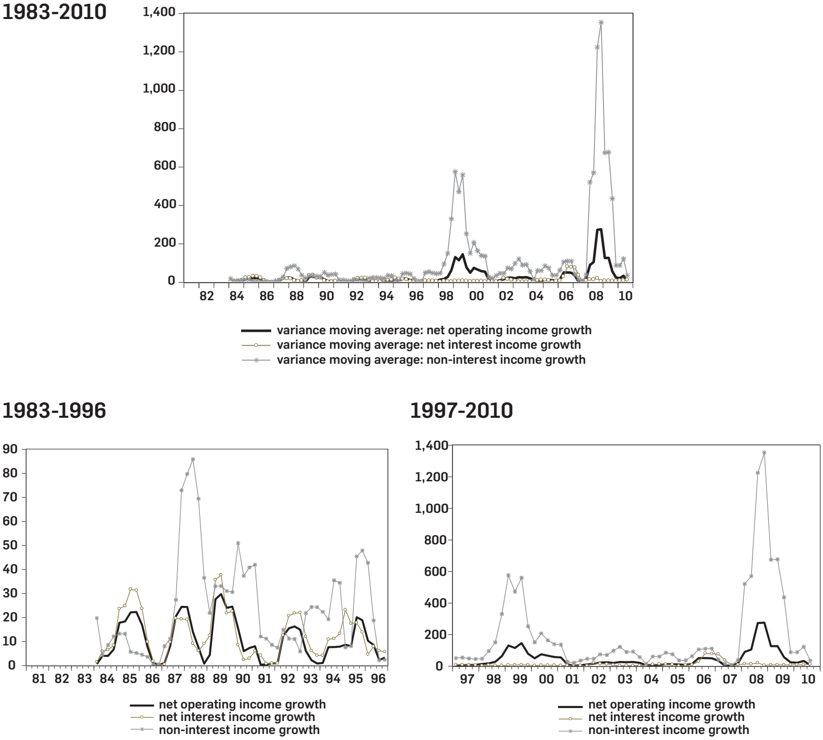
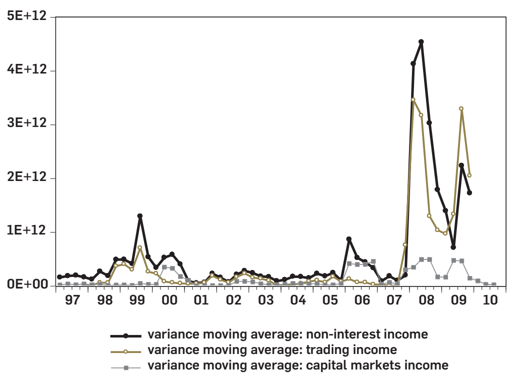
Figure 3. Variance of Net Operating Income Growth and its Components



Note: The variance is a rolling variance computed over four quarters.
SOURCE: BANK OF CANADA.

Figure 4. Variance of Non-interest Income and of its Two most Molatile Components, Trading Income and Capital Markets Income, 1997-2010



Note: The variance is a rolling variance computed over four quarters.
SOURCE: BANK OF CANADA.

Table 2. Components of Non-interest Income, 1997-2010

	Non-interest income	Capital markets	Income wealth mgt	Retail	Insurance	Trading	Securitization	Other
Level (end-of-period, thousand $)	9244859	2513622	2039389	1522159	2017461	130621	723671	297936
Mean (thousand $)	7952304	2692371	1345518	1039965	1070721	939987	403025	556378
Median (thousand $)	7789256	2647529	1297605	1021984	889839	1293237	334615	554668
Std. Dev. (thousand $)	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 bank funding sources and enhance liquidity positions.

DATA SOURCE: BANK OF CANADA.

Table 3. Decomposition of the Variance of Non-interest Income Growth, 1997-2010

	Average share	Variance	Contribution to variance	Covariance	Contribution to covariance	Total contribution
Non-interest 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 $Variance = w^T \Omega w$*, where* w *is the vector of the respective shares of the components of non-interest income, and* Ω *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:* $\frac{\partial\, variance}{dw} = 2\Omega w$*, where the relative contribution of component* i *is equal to* $2\Omega_i w$ *with* Ω_i *the* i^{th} *line of the* Ω *matrix.*

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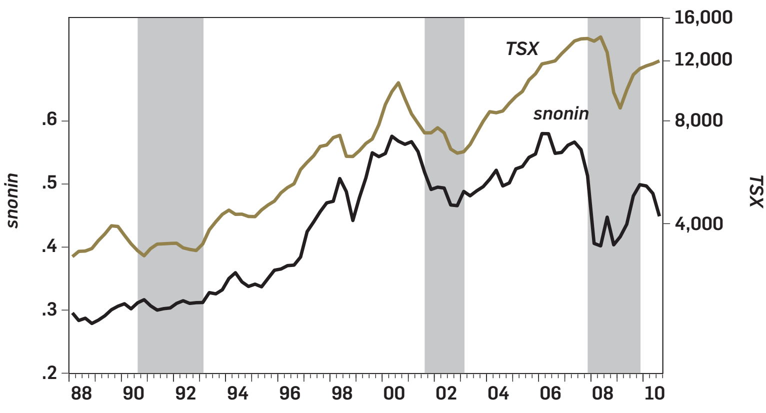
To conclude on this change in the banking activities mix, note that after 1997 the volatility of *snonin* increases in conjunction with the stock market index (S&P/TSX), and with the fluctuations of bank stock trading portfolio. A close look at Figure 5 actually suggests that there might be a cointegration 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 I(1) variables, so that they can potentially be cointegrated. The results of a Johansen's cointegration test confirm this conclusion. When the variables are both expressed in levels the test indicates a cointegration relationship between the two variables over the period 1988-2010 at the 10% threshold. The test identifies a tighter cointegration relationship after 1997, while the test fails to reject the hypothesis of no cointegration over the first subperiod (1988-1996)[14]. The growing importance of the income generated by capital markets and trading might thus partly be related to the tighter cointegration of *snonin* and *TSX*, which in turn would directly contribute to the growth in operating income volatility. Actually, this tighter cointegration might explain the increased procyclicality observed in the banking sector over the last decade (Calmès and Théoret, 2010; Nijskens and Wagner, 2011).

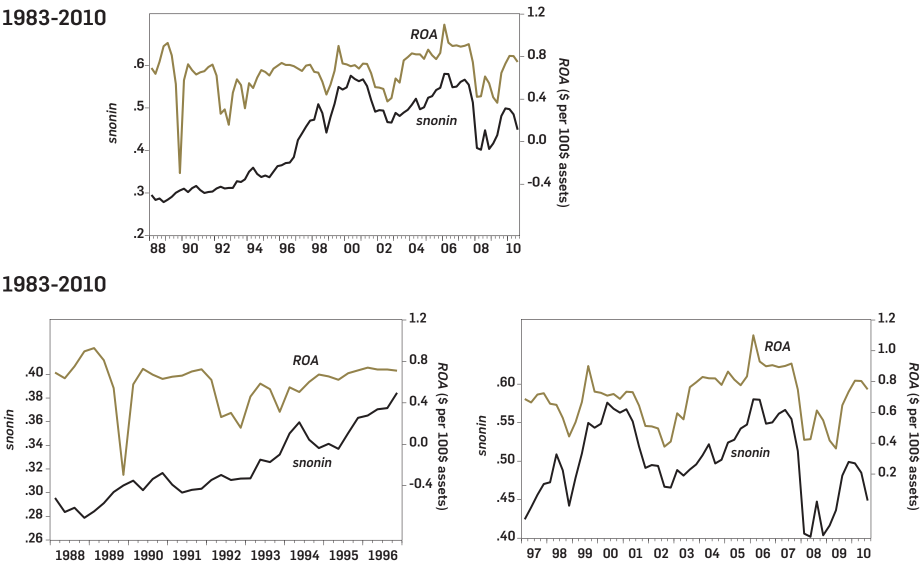
Figure 5. *TSX* and *snonin*, 1988-2010



Shaded areas correspond to periods of contractions or marked economic slowdown in Canada.
SOURCE: STATISTICS CANADA (CANSIM) AND CANADIAN BANKERS ASSOCIATION.

[14] *We also perform the test by taking the logarithm of TSX and obtain the same kind of results. Our findings are available upon request.*

Figure 6. Return on Assets and Share of Non-interest Income



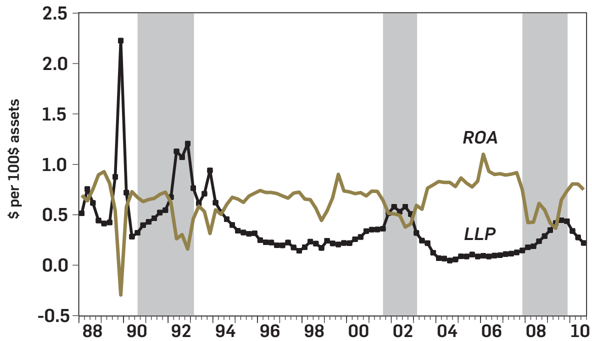
SOURCE: CANADIAN BANKERS ASSOCIATION.

Theoretically, the greater volatility of bank operating income observed after 1997 should be associated with a higher expected *ROA*: an additional risk premium must be added to the expected *ROA* to optimally account for the greater risk (DeYoung and Roland, 2001; Stiroh, 2006; Laeven and Levine, 2007). 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. Busch and Kick (2009) are more positive but they ignore the issue of strategic complementarities and the leverage effect of OBS. DeYoung and Rice (2004), Chiorazzo *et al.* (2008), Lepetit *et al.* (2008), and Nijskens and Wagner (2011) find a positive diversification effect, but associated with an increased systemic risk. In this study, we find that the Spearman rank-order correlation[15] between *ROA* and *snonin* is moderately negative before 1997, but actually becomes positive after 1997 (Figure 6)[16]. As banks increased their involvement in OBS activities, their loan loss provisions ratio (*LLP*) decreased, both in level and volatility (Figure 7). 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* volatilities also increase with non-interest income volatility, the change we observe in bank risk-adjusted returns cannot be attributed to *LLP*, at least in the second part of our sample. Instead, based on this preliminary evidence, we can conjecture that banks might have changed their business model.

[15] Compared to the standard (Pearson) correlation coefficient, the Spearman correlation coefficient mitigates the impact time series outliers.

[16] For more details on these correlations see also Calmès and Théoret (2011).

Figure 7. Return on Assets and Loan Loss Provisions Ratio



Shaded areas correspond to periods of contractions or marked economic slowdown in Canada.
SOURCE: STATISTICS CANADA (CANSIM) AND CANADIAN BANKERS ASSOCIATION.

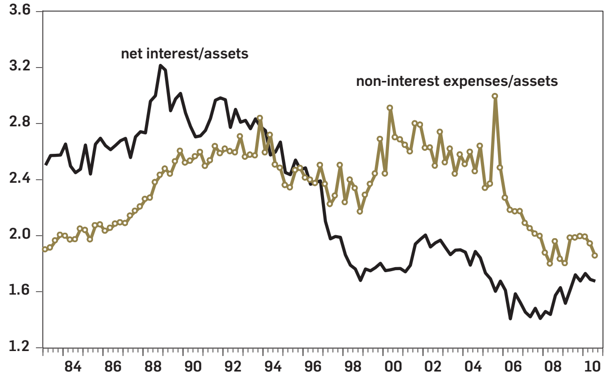
3. Empirical Framework

3.1. Diversification and Endogeneity

Based on first principles and accounting identities the endogeneity of *snonin* is 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). As argued by Campa and Kedia (2002), a firm's choice to diversify is likely to be a reaction to exogenous forces which also impact firm's value. In that respect, bank returns on assets (ROA) may well be a function of the share of non-interest income (*snonin*), but *snonin* is also a function of *ROA* (Demsetz and Strahan, 1995; Goddard *et al.*, 2008; Busch and Kick, 2009). OBS activities could generate 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 however, a decrease in performance might also induce banks to take more risk by increasing their involvement in OBS activities (Boyd and Gertler, 1994; Estrella, 2001; Vander Vennet *et al.*, 2004; Busch and Kick, 2009), and then the relation between *ROA* and *snonin* would be negative[17].

[17] *Incidentally, Canadian banks' net interest margin moves on a downward trend since 1987 and does no longer cover the non-interest expenses ratio after 1993 (Figure 8).*

Figure 8. Canadian Bank Net Interest and Non-interest Expenses as % Assets, 1983-2010



In brief, *ROA* and *snonin* are two interactive bank decision variables, so that the associated endogeneity can possibly bias the estimation of the sensitivity of *ROA* to *snonin* (Stiroh, 2004; Stiroh and Rumble, 2006). To illustrate this issue more precisely consider the two following simultaneous equations:

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

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

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

Equation (1) is a simplified version of the canonical model used in most studies. If OBS activities lead to diversification benefits then $\alpha_1>0$. However, we must account for the "reverse causality" described in Equation (2). In particular, we can suspect that $\alpha_2<0$, because banks might have increased *snonin* in reaction to the decrease in *ROA* associated with the decline of traditional banking (Boyd and Gertler, 1994; Busch and Kick, 2009)[18]. If we estimate Equation (1) 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 complicate. The asymptotic bias of α_1 is equal to:

$$plim\hat{\alpha}_{1,OLS} - \alpha_1 = \frac{cov\,(snonin, \mu_1)}{var\,(snonin)} \quad (3)$$

where $\hat{\alpha}_{1,OLS}$ is the estimation of α_1 obtained by applying OLS to Equation (1). According to Equation (3), the sign of the bias depends on the covariance between *snonin* and μ_1. To compute this covariance we can simplify Equation (1) 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:

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

$$cov\ (snonin, \mu_1) = \frac{\alpha_2}{1-\alpha_1\alpha_2}\ \sigma^2_{\mu_1} \tag{4}$$

In this case the asymptotic bias (or inconsistency) in the OLS estimation of α_1 has the same sign as $\frac{\alpha_2}{1-\alpha_1\alpha_2}$. In the case where $\alpha_2<0$ and $\alpha_2\alpha_1<1$, i.e. the likeliest scenario we face, the asymptotic bias is negative and we should expect the estimation of α_1 to be biased downward. This downward bias means that a conventional OLS estimation underestimates the impact of *snonin* on *ROA*, or, more specifically, hides the diversification benefits associated with OBS activities.

The motivation of our study comes precisely from the idea that the endogeneity of *snonin* can lead to an underestimation of α_1, i.e. the sensitivity of *ROA* to *snonin*. In the next subsection we thus introduce a rigorous treatment of this endogeneity, and detail in appendix how to construct the higher moment instruments we use to endogeneize the *snonin* variable.

3.2. The Modified *TSLS* Regression Incorporating an Hausman Endogeneity Test

To treat 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 *H*0 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 drawbacks we resort to an alternative Hausman test. The modified version of the *h* test we introduce is directly related to Hausman (1978), Spencer and Berk (1981), McKinnon (1992) and Pindyck and Rubinfeld (1998)[19]. To implement this version of the Hausman test we first write our bank returns model as:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 snonin_t + \mathbf{X}_t\alpha + \varepsilon_t \tag{5}$$

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 bank performance (e.g. riskiness of loans or spread between the yield and funding cost of

[19] *For an application to hedge funds see also Racicot and Théoret (2008, 2012).*

loans). 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 $\mathbf{Z} = \{z_1, z_2, \ldots, z_K\}$ — k being the number of instruments — to treat the *snonin* endogeneity. As discussed in the appendix, in our case this instrument set is the vector of higher moments $\mathbf{Z}$. The higher moment Hausman test is then implemented in two steps. First, regressing $snonin_t$ on the instrument set $\mathbf{Z}_t$, we compute the fitted value of $snonin_t$, noted $sn\hat{o}nin_t$:

$$snonin_t = \hat{c}_0 + \mathbf{Z}_t \hat{\rho} + \hat{w}_{snonin_t} = sn\hat{o}nin_t + \hat{w}_{snonin_t} \tag{6}$$

where $\hat{w}_{snonin_t}$ is the innovation resulting from the regression of *snonin* on the instrument set $\mathbf{Z}_t$. Then, in a second step we substitute $sn\hat{o}nin_t$ to $snonin_t$ in the bank return model (Equation 5). This way we can obtain consistent estimates of the coefficients of the returns equations. Provided that there is no endogeneity concern we can substitute Equation (6) in Equation (5) to obtain the following artificial (or auxiliary) regression:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 sn\hat{o}nin_t + \mathbf{X}_t \alpha + \beta_2 \hat{w}_{snonin_t} + \varepsilon_t \tag{7}$$

Finally, using Equation (7) 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, and thus Equation (7) reads:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 sn\hat{o}nin_t + \mathbf{X}_t \alpha + \theta \hat{w}_{snonin_t} + \varepsilon_t \tag{8}$$

With $snonin_t = sn\hat{o}nin_t + \hat{w}_{snonin_t}$, we can reformulate Equation (8) as follows:

$$y_t = \beta_0 + \beta_1 y_{t-1} + \beta_2 snonin_t + \mathbf{X}_t \alpha + \varphi \hat{w}_{snonin_t} + \varepsilon_t \tag{9}$$

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

Our 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 (8).

Compared to the standard h test, one crucial advantage of our 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 (9) 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 (9) 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 (9) becomes:

$$(y_t)_{OLS} = \hat{\beta}_0 + \hat{\beta}_1 y_{t-1} + \hat{\beta}_2 \, snonin_t + \mathbf{X}_t \hat{\alpha} + \varepsilon_t \tag{10}$$

However if $\hat{\varphi}$ is significantly different from zero the *TSLS* estimator obtains and Equation (9) reads:

$$(y_t)_{TSLS} = \hat{\beta}_0^* + \hat{\beta}_1^* y_{t-1} + \hat{\beta}_2^* snonin_t + \mathbf{X}_t \alpha^* + \hat{\varphi}\, \hat{w}_{snonin_t} + \varepsilon_t \tag{11}$$

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 interpreted 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 (5) the standard *TSLS* procedure and this Hausman procedure are interchangeable. However, our motivation to favour the latter is that it provides a crucial information on endogeneity, namely, it helps assess the biases changes.

4. Empirical Results

In this section we discuss the empirical results of the various experiments we just described, beginning with those of the estimation method often used in the literature, i.e. the *OLS*, which serves as a benchmark to our Hausman procedure. Note that we first need to examine the stationarity of the time series used in our model in order to avoid spurious results. We thus run an Augmented Dickey-Fuller unit root test for the time series used in this study. Over the sample period the test indicates that the *snonin* variable displays a unit root[20]. To make the *snonin* variable $I(0)$ we thus express it in first-differences throughout the experiments.

4.1. OLS Results

To estimate Equation (5) we include the ratio of loan loss provisions to total assets as it is the only significant control variable, and our benchmark model reads:

$$ROA_t = \gamma_1 + \gamma_2\, d(snonin)_t + \gamma_3\, LLP_t + \gamma_4 ROA_{t-1} + \zeta_t \qquad (12)$$

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

The fit of the model seems quite reasonable over the whole sample period, the adjusted R^2 being 0.62 (Table 4). 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 this ratio increases during recessions, it also magnifies the procyclicality of *ROA*.

Table 4. 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

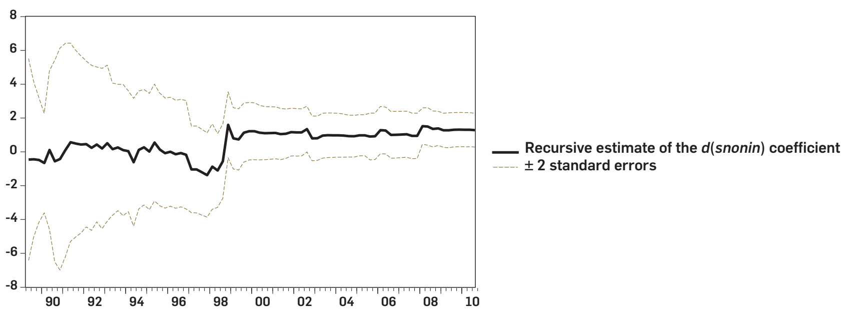
Notes: ROA, return on assets; d(snonin), first-difference of the share of non-interest income in net operating revenue; LLP, ratio of loan loss provisions over total assets. The t statistics are reported in italics.

As the literature suggests, Table 4 confirms that the risk-return trade-off improves throughout the sample period. The coefficient of *d*(*snonin*), significant at the 95% confidence level, is 1.28. However, since we are primarily interested by the changes in

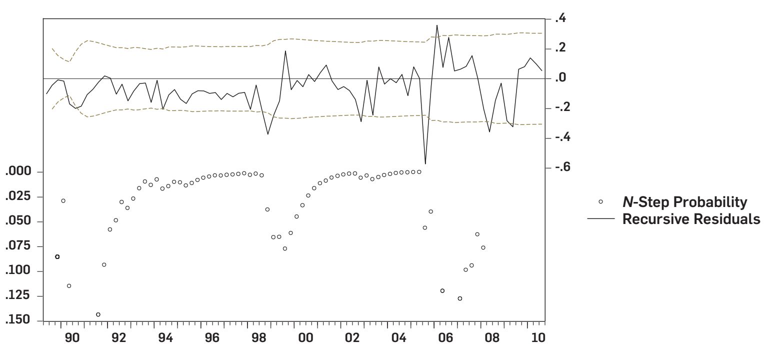
[20] *The Dickey-Fuller tests are available upon request.*

the *snonin-ROA* relationship, it is also instructive to run a recursive regression over the whole period. Figure 9 actually 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 higher after 1997. We find this relationship both positive and much more significant. In this respect Figure 9 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. A rolling regression of fifteen quarters, which provides a more precise estimation, corroborates that the sensitivity of *ROA* to *d*(*snonin*) turns from negative to positive around 1997. This indicates *a priori* the emergence of some diversification gains associated with market-oriented activities. 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 bank performance. As conjectured, 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[21].

Figure 9. **Recursive Estimate of the *d*(*snonin*) Coefficient in the *ROA* Model**



***N*-step *ROA* Forecast**



[21] *This result is also consistent with the time-to-build story evoked in Busch and Kick (2009) and Nguyen (2012).*

Because of the growth in the bank new business lines we should also expect a deterio-ration of the model performance in the second subperiod. It is during this period that banks begin to integrate their new banking business to their traditional lending activities more systematically. Our experiments suggest that the risk prevailing in the second subperiod, as implied by the volatility of the bank income growth, is indeed more pronounced, and feeds into the innovation term of the equation. More precisely, the volatility of the residuals of the recursive regression (Equation 12) is much higher after 1997 (Figure 9). Hence, the data seem to track the change in the banking environment quite well. In the *ROA* equation, the adjusted R^2 is equal to 0.87 over the first subperiod, and then falls to 0.38 in the second one. This corroborates the deterioration of the model fit when the endogeneity issue is ignored (Table 4).

Summarizing, it took ten years before banks could properly integrate their new business lines to their traditional activities, but the bank risk-return trade-off seems to improve after 1997. However, we argue that to get a more definite assessment of this maturation process we must properly account for the evolution of the endogenous link between *snonin* and *ROA*, as exposed below.

4.2. Hausman Artificial Regression Results

Table 5. **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 the robust instruments defined in the Appendix. The t statistics are reported in italics.

We report the results of the Hausman estimation (Equation 9) in Table 5. As previously mentioned, the Hausman procedure is very similar to a standard *TSLS* estimation[22]. However, the Hausman regressions offer the key advantage of directly embedding an endogeneity test based on the significancy of $w_{d(snonin)}$, as measured by its t-statistic[23]. More importantly, this particular method provides an indication of the severity of the simultaneity bias with the level of the $w_{d(snonin)}$ coefficient. As expected, the endogeneity significantly biases the estimated coefficient of $d(snonin)$. Over the whole sample, the coefficient of $d(snonin)$ is equal to 1.28 when estimated with the usual *OLS* method, but to 2.50 with the Hausman procedure (Table 5). In other words, 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 $w_{d(snonin)}$ coefficient indicates that the impact of $d(snonin)$ is significantly understated in the *OLS* run.

During the first subperiod, the coefficient of $d(snonin)$ estimated with *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* actually overstates the impact of $d(snonin)$ over the first subperiod, in line with the results of Stiroh (2004) and Stiroh and Rumble (2006). Moreover, consistent with our conjecture, the coefficient of $w_{d(snonin)}$, at –3.93, is much higher in absolute value during the second subperiod. After 1997, controlling for endogeneity the way we do translates in substantial gains in terms of estimation, and clearly suggests a significant, positive influence of $d(snonin)$ on returns, the coefficient more than doubling, from 1.71 to 3.79. The adjusted R^2 also improves when endogeneity is accounted for, increasing from 0.62 to 0.75 for the whole sample and from 0.38 to 0.52 for the latter period.

To summarize, the underestimation of the positive effect of $d(snonin)$ on *ROA* is particularly important in the last period. This result is important because it clearly suggests 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. symptomatic of the presence of increasing diversification gains. To confirm this finding, it is much instructive to run a recursive regression. In Figure 10 note that the confidence interval of the coefficient of $w_{d(snonin)}$ shrinks greatly through time. This indicates that the *snonin* endogeneity issue indeed becomes more important *pari passu* with the increased involvement of banks in market-oriented business lines. 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 (DeJonghe, 2009)[24].

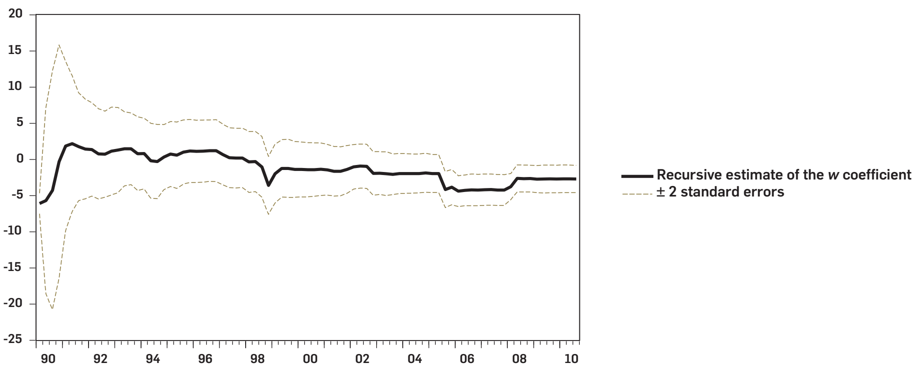
[22] *Since the results obtained with the TSLS and the Hausman procedure are essentially the same we only report the Hausman procedure findings.*

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

[24] *For more details on this, see the discussion section.*

What is crucial here is not merely the fact that the positive influence of OBS on returns obtains when controlling for endogeneity. This result has been often reported in the literature (Campa and Kedia, 2002; Busch and Kick, 2009; Schmid and Walter, 2009) regardless of the way endogeneity is accounted for. The novelty is rather the fact that, when accounting for endogeneity, the Hausman procedure shows that this positive influence is actually increasing through time[25]. The natural intuition behind this finding is that the *snonin-ROA* link ought to evolve along with the involvement of banks in market-oriented banking. In other words, the *snonin* endogeneity becomes more severe with the maturation of market-oriented business lines.

Figure 10. **Recursive Estimate of the *w* Coefficient in the *ROA* Model**



4.3. Robustness Check and Additional Results

It is interesting to check if our results are robust to a change in the way we account for risk. We thus express *ROA* on a risk-adjusted basis. This ratio is defined as: $RA\_ROA_t = \frac{ROA_t}{\sigma_{t,ROA}}$, where $\sigma_{t,ROA}$, the standard deviation of *ROA*, is represented by a moving average computed on a rolling window of four quarters[26].

We estimate Equation (12) using *RA_ROA* as the dependent variable. The estimated equation becomes:

$$RA\_ROA_t = \gamma_1 + \gamma_2\, d(snonin)_t + \gamma_3\, LLP_t + \gamma_4 RA\_ROA_{t-1} + \zeta_t \qquad (13)$$

[25] *It is interesting to note that when accounting for the endogeneity of the decision to diversify, the diversification discount disap-pears or even turns to a premium in the study of Campa and Kedia (2002). Although we describe a more dynamic phenomenon, the general idea is in the same vein.*

[26] *We also considered z-score measures, like in Lepetit et al. (2008), but the correlation between the z-scores and risk-adjusted returns is near 1. Relying on Sharpe ratios also delivers very similar results.*

Table 6. *OLS* Estimation of the Risk-adjusted *ROA*

Variables	1988-1996	1997-2010	1988-2010
c	0.66	1.14	1.19
	6.17	*2.66*	*4.75*
d(*snonin*)	1.56	14.60	12.18
	0.44	*4.19*	*4.13*
LLP	–0.72	–1.28	–1.19
	–6.66	*–1.33*	*–4.28*
$RA\_ROA_{t\text{-}1}$	0.83	0.81	0.77
	27.32	*12.73*	*15.28*
Adjusted R^2	0.76	0.74	0.81
DW stat.	2.59	2.45	2.34

Notes: The dependent variable is ROA scaled by a rolling ROA standard deviation of four quarters. The explanatory variables are: d(snonin), first-difference of the share of non-interest income in net operating revenue; LLP, ratio of loan loss provisions over total assets, and RA_ROA_{t-1}*, risk-adjusted ROA lagged one period. The t statistics are reported in italics.*

Table 7. Hausman Regression of the Risk-adjusted *ROA*

Variables	1988-1996	1997-2010	1988-2010
c	1.25	1.20	0.79
	12.10	*4.41*	*8.84*
d(*snonin*)	10.54	14.93	13.33
	1.78	*5.78*	*23.29*
LLP	–1.15	–1.14	–0.94
	–16.28	*–1.93*	*–10.74*
$RA\_ROA_{t\text{-}1}$	0.68	0.77	0.83
	19.88	*14.13*	*37.84*
$w_{d(snonin)}$	–32.43	–8.39	–12.22
	–2.20	*–1.60*	*–7.70*
Adjusted R^2	0.81	0.58	0.79
DW stat.	1.92	2.34	2.38

Notes: The dependent variable is ROA scaled by a rolling ROA standard deviation of four quarters. The explanatory variables are: d(snonin), first-difference of the share of non-interest income in net operating revenue; LLP, ratio of loan loss provisions over total assets, and RA_ROA_{t-1}*, risk-adjusted ROA lagged one period. The w variable is the residuals obtained with a regression of d(snonin) on the robust instruments defined in the Appendix. The t statistics are reported in italics.*

The results of the *OLS* estimation of Equation (13) are provided in Table 6, and the corresponding Hausman regression results are reported in Table 7. These tables indicate that our findings are almost unaffected by the introduction of an alternative measure of returns adjusted for risk. The *OLS* regression underestimates the coefficient of *d*(*snonin*) over the whole sample period, and particularly so after 1997, the coefficient being actually insignificant at the 5% threshold over the subperiod 1988-1996.

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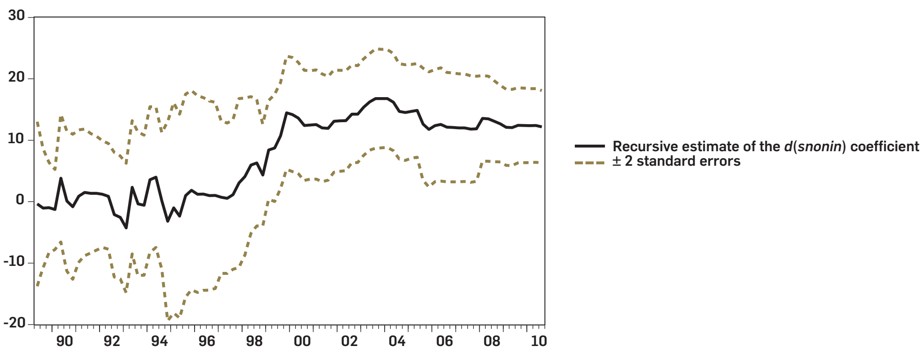
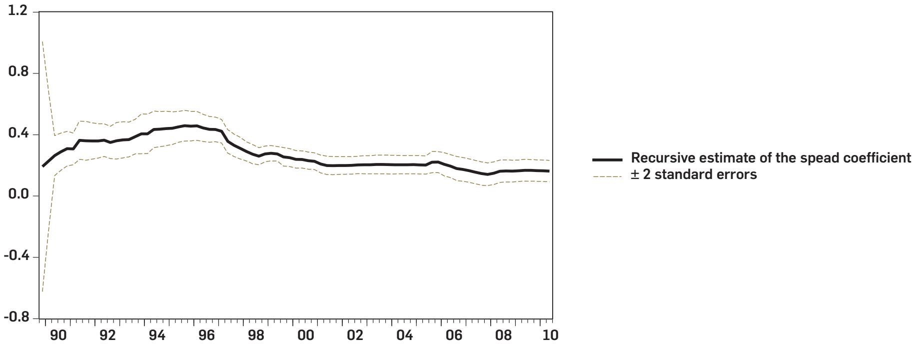
Over the whole sample (1988-2010), the coefficient of $d(snonin)$, significant at the 1% threshold, is equal to 12.18 when estimated with OLS, but to 13.33, also significant at the 1% threshold when estimated with our Hausman procedure. The coefficient of $w_{d(snonin)}$, significant at the 1% threshold, is also high, at –12.22, which confirms the underestimation of the coefficient of $d(snonin)$. When we run a recursive regression on *RA_ROA* over the period 1988-2010, this coefficient touches a low of –4.3 in the first quarter of 1993 and begins to increase progressively thereafter (Figure 11). As for the estimation of Equation (12), the confidence interval of the $d(snonin)$ coefficient is also narrowing after 1997, which confirms that the diversification benefits of banks' OBS activities are improving after the structural break. More precisely, even if banks had used up their internal diversification advantages, we still find that bank risk-adjusted returns improve through time (Demsetz and Strahan, 1995; DeYoung and Roland, 2001; Stiroh, 2006; Altunbas *et al.*, 2007).

Figure 11. **Recursive Estimate of the $d(snonin)$ Coefficient in the Risk-adjusted Return Model**



In other respects, Lepetit *et al.* (2008) and Busch and Kick (2009) argue that banks might use their diversified activities to charge lower interest margins. Consequently, we check the robustness of the results by considering two additional risk premia, namely the return on the *TSX* index and the spread between the yield and funding cost of loans. Following these modifications we can corroborate the cross-selling effect but our main results remain essentially the same. In particular, the sensitivity of *ROA* to the spread, when applying a recursive regression to our augmented *ROA* model including risk premia, increases from 1988 to 1996, but decreases continuously thereafter, a shift which accords with the structural break identified earlier (Figure 12).

Figure 12. Recursive Estimate of the Spread Coefficient



Note: The spread is the difference between the yield and funding cost of loans.

To our knowledge, the modified Hausman procedure we introduce is the first attempt to analyze bank diversification in a dynamic setting. Some authors (e.g., Stiroh and Rumble, 2006; Laeven and Levine, 2007) compare their results before and after the correction for the endogeneity but generally find their results robust to endogeneity. The studies of DeYoung and Rice (2004), Goddard *et al.* (2008) and Busch and Kick (2009) use standard instruments to tackle the simultaneity bias between *ROA* and *snonin*, however they do not measure the extent of the endogeneity biases and they account for endogeneity upfront. Yet, the seminal study of Campa and Kedia (2002) on the diversification discount in industrial conglomerates show the importance of performing Hausman tests. In their study, when accounting for endogeneity with this kind of procedure, the diversification discount disappears and even turns into a premium in some cases. The new set of results we derive from this kind of procedure is in the same spirit since endogeneity appears symp-tomatic of the increasing link between non-interest income and banks' risk-adjusted returns.

5. Discussion

5.1. The Relative Impact of *snonin* Components on Performance

To complete this study it is interesting to further examine the link between the non-interest income components (product-mix) and banks' performance. We first replicate Tables 4 to 7 for the components of non-interest income reported in Table 2 (Table 8). When we regress *ROA* on the shares of the components of non-interest income expressed in terms of operating income, and on the same control variables used to build Tables 4 to 7, we resort to two estimations methods: *OLS* and the Generalized Method of Moments (GMM), an IV estimation procedure[27].

[27] *To implement the GMM procedure we rely on the robust instruments presented in the Appendix.*

Table 8. Regression of *ROA* and Risk-adjusted *ROA* on the Components of *snonin*

Variables	ROA OLS	ROA GMM	RA_ROA OLS	RA_ROA GMM
c	0.45	0.38	1.38	1.07
	6.76	*4.85*	*9.02*	*2.09*
d(capital markets share)	2.98	1.67	5.16	23.04
	5.06	*4.61*	*3.18*	*2.27*
d(trading share)	2.05	2.15	9.87	17.32
	5.75	*6.56*	*16.07*	*2.81*
d(wealth manag.share)	2.65	8.57	39.78	15.64
	1.63	*5.74*	*13.14*	*0.57*
d(retail share)	8.34	14.59	48.22	82.34
	3.42	*8.12*	*9.00*	*2.09*
d(insurance share)	–0.23	-2.02	–5.13	–17.15
	–0.24	*-2.12*	*–4.71*	*–1.46*
d(securitization share)	–4.20	-3.16	–21.24	11.01
	–2.38	*-2.90*	*–6.26*	*0.47*
LLP	–0.37	-0.20	–1.79	–1.39
	–4.54	*-2.31*	*–15.95*	*–1.55*
y_{t-1}	0.47	0.55	0.76	0.85
	6.24	*6.40*	*17.77*	*8.05*
Adjusted R^2	0.47	0.52	0.44	0.71
DW stat.	2.25	2.23	2.24	2.43

Notes. RA_ROA is risk-adjusted ROA, i.e., ROA scaled by a rolling ROA standard deviation of four quarters. Consistent with snonin, the components of non-interest income are expressed as ratios of net operating income, the sum of net interest income and non-interest income. They appear in the regression as first-differences since snonin is expressed in first-differences in the previous regressions. y_{t-1} is the dependent variable lagged one period. The t statistics are reported in italics. Note that the estimation period runs from 1997 to 2010 since most of the components of non-interest income are not available before 1997.

Note that the components identified as providing the best diversification benefits in Table 3, i.e., the insurance and securitization shares, contribute negatively to *ROA* (Table 8). Their coefficients are negative in both the *OLS* and GMM estimations. This result is quite reasonable since these two components may be considered as an hedge from the point of view of banks' performance[28]. By contrast, the share of retail income contributes the most to *ROA* (Table 3). When estimated with *OLS*, its coefficient is equal to 8.34 significant at the 95% confidence level, and to 14.59 with GMM and significant at the 99% level. The three remaining components have quite comparable patterns, although GMM delivers a greater coefficient for the wealth management share than for the shares associated with bank market activities. Finally, in spite of their relative im-

[28] *For instance, securitization allows a better risk sharing across a larger and more diversified set of players.*

portance and volatility, the two components of *snonin* most related to bank market activities still display a moderate contribution to *ROA*.

Turning to the estimation of risk-adjusted *ROA*, the two components contributing negatively to the variance of the growth of non-interest income – securitization and insurance income – no longer impact bank performance in the GMM estimation. This result confirms the hedging capacity of these activities. As in the case of *ROA*, the retail share still contributes the most to the risk-adjusted *ROA*, its estimated coefficient being 48.22 in the *OLS* estimation and 82.34 for the GMM. In other words, consistent with the literature (e.g., Gallo *et al.*, 1996; Vander Vennet *et al.*, 2004; Busch and Kick, 2009), we find that retail is the activity which improves the most the risk-return trade-off. The contribution of the other business lines to the bank risk-return trade-off differs according to the estimation method, which signals an endogeneity issue. For example, as insurance and securitization, the coefficient of the wealth management share is no longer significant in the GMM estimation while the components of *snonin* most related to bank market-oriented activities still have an impact on risk-adjusted *ROA*, the coefficients estimated with GMM being 23.04 for the capital market share and 17.32 for the trading share, both significant at the 95% confidence level. These results are broadly consistent with the idea that a combination of commercial and investment banking may increase the performance of financial conglomerates (Krozner and Rajan, 1994; Vander Vennet *et al.*, 2004; De Jonghe, 2009; Schmid and Walter, 2009). Similar to our results, Stiroh and Rumble (2006) also find that fiduciary income seems to increase banks' risk adjusted returns.

However, the effect of trading activities on bank performance is less clear in the literature. Estrella (2001) and Stiroh and Rumble (2006) find that these activities decrease banks' risk-adjusted returns, while Lepetit *et al.* (2008) show that trading might reduce risk for small banks, and Krozner and Rajan (1994) and Schmid and Walter (2009) show that it could increase value. At a more consolidated level, Busch and Kick (2009) find that fee income activities rise German banks' risk-adjusted returns and Cardone-Riportella *et al.* (2010) show that Spanish banks rely on securitization to stop the decline of their risk-adjusted returns. To make sense of these mixed results, it is important to bear in mind that the diversification benefits associated with a specific component of non-interest income are likely time-varying (e.g., DeYoung and Rice, 2004; Stiroh and Rumble, 2006). Moreover, the empirical results may also be sensitive to the sample period considered, especially if the period corresponds to normal times versus crises (De Jonghe, 2009). For this reason, it is quite instructive to compute the conditional correlations between the banks' income streams.

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Table 9. Correlation between Net Interest Income and the Non-interest Income Components

	1997-2010
Capital markets	0.25
	0.07
Trading	0.32
	0.02
Wealth management	–0.51
	0.01
Retail	0.47
	0.00

	1997-2010
Insurance	–0.57
	0.00
Securitization	–0.16
	0.25
Other	0.76
	0.00

Notes. Since bank interest margin is defined in percent of assets, the time series composing non-interest income are also expressed on this basis. The p-values of the correlations are reported in italics.

5.2. The Conditional Correlations

Table 9 documents the comovements between the traditional and non-traditional business lines. Insurance and wealth management seem to provide the greatest benefits, with respective correlations of –0.57 and –0.51, followed by securitization, with a correlation of –0.16. Note that retail income is the most positively correlated with net interest income, a quite intuitive result given the interaction between deposit fees and net interest income (Lepetit *et al.*, 2008; Busch and Kick, 2009). Finally, the correlation between net interest income and the most market-oriented activities is moderate, the correlation to net interest income being 0.32 for trading and 0.25 for capital markets.

Table 10. GMM Regression of Net Interest Income on the Components of Non-interest Income

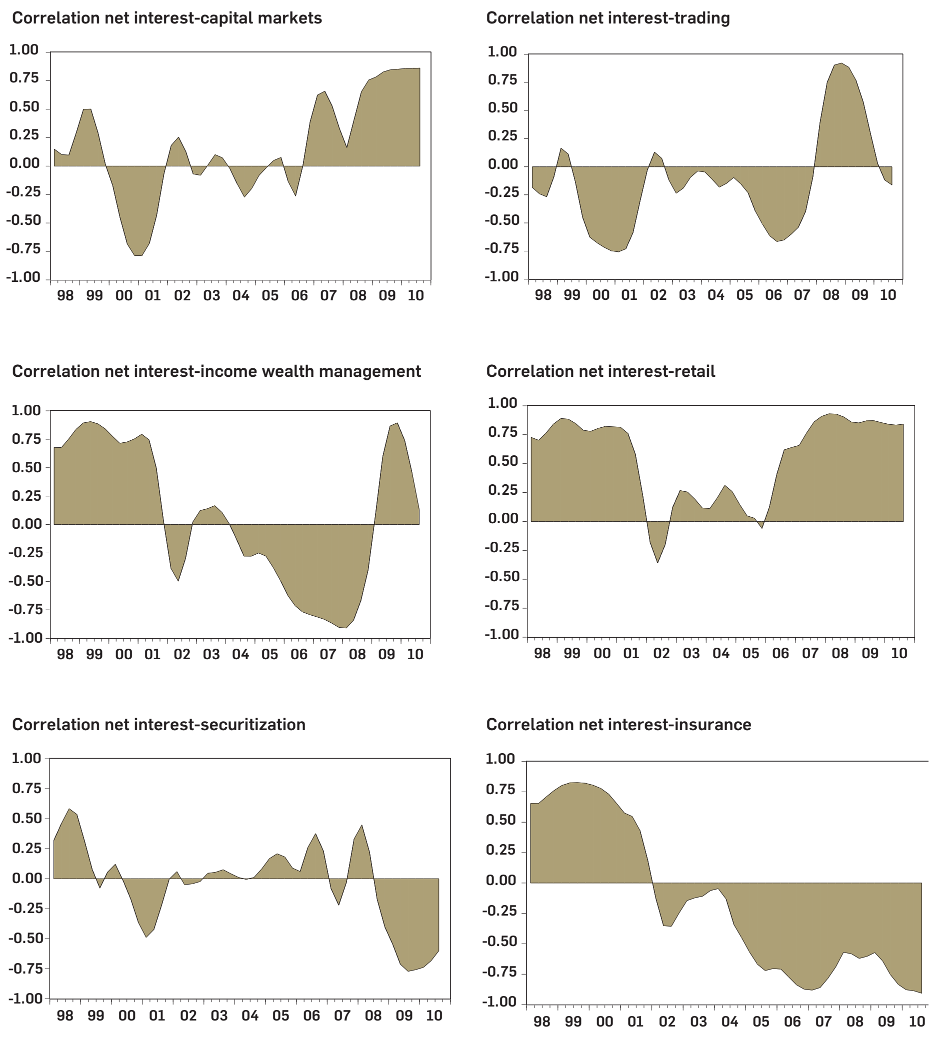
Variables	GMM
c	0.70
	5.68
Capital markets	0.02
	0.61
Trading	0.04
	1.82
Wealth manag.	–0.27
	–1.74
Retail	2.07
	8.07
Insurance	-0.93
	-7.57

Variables	GMM
Securitization	0.01
	0.04
r_{corp}	–0.03
	–4.50
r_{TSX}	–0.01
	–1.39
Net interest income$_{t-1}$	0.51
	11.82
Adjusted R^2	0.83
DW stat.	2.05

Notes: The times serie are expressed in percent of bank assets. rcorp is the 3-month prime corporate paper rate, an indicator of monetary conditions for the Bank of Canada, and r_{TSX} is the quarterly yield on the Canadian TSX stock market index. The t statistics are reported in italics. To implement the GMM regression, we resort to the robust instruments defined in the appendix.

If we regress net interest income on the *snonin* components of non-interest income and two control variables (the three-month prime corporate rate and the yield on the TSX stock market index), insurance and wealth management again appear the business lines contributing the most to diversification benefits, their respective coefficients being re-spectively –0.93 and –0.27 (Table 10). Consistent with the literature, we also find that retail business lines contribute the less with an estimated coefficient of 2.07.

■ Figure 13. **Conditional Correlations between Net Interest Income and the Components of Non-interest Income (as % Assets)**



Note: The conditional correlations are computed with a multivariate GARCH system based on a BEKK procedure (Engle and Kroner, 1995).

If we compute the conditional correlations between net interest income and the components of non-interest income with a multivariate GARCH system based on a BEKK procedure (Engle and Kroner, 1995), except for insurance and securitization, the conditional correlation jumps toward 1 during the last financial crisis (Figure 13). In other words, the benefits of diversification tend to disappear when they are needed the most for stabilizing income[29]. This phenomenon is related to the fat-tail risk associated with non-interest income, a risk not accounted for in the standard deviation but detected with the multivariate GARCH process.[30]

To further substantiate our previous results, the conditional correlation plots indicate that the benefits of diversification are upward trending for the wealth management component, and especially for the insurance component from 2000 to 2010. In normal times the diversification benefits associated with the capital markets component are also significant, and especially for the trading component, whose conditional correlation is indeed negative most of the time. Finally, we can confirm that the retail component displays very low diversification benefits, its conditional correlation with net interest income being often close to 1.

6. Conclusion

The intensification of the link between *ROA* and *snonin*, and the corresponding interest of banks to invest in OBS activities may well be related to the fact that financial institutions had to resort to market-oriented activities as a way to increase their profitability and compensate for the decreasing return on their traditional activities (Boyd and Gertler, 1994; Vander Vennet *et al.*, 2004; Busch and Kick, 2009). 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[31]. Our results suggest that, after 1987[32], with the successive waves of banking deregulation and the financial deepening associated with the increased firms' reliance on direct financing, it actually took ten years for banks to eventually record significant diversification gains from OBS activities. After this maturation phase, the change in

[29] *In line with our results Vander Vennet et al. (2004) find that European diversified banks did not perform better than specialized banks during the 2000-2003 downturn. However they do not qualify their findings with a non-interest income decomposition.*

[30] *Indeed, a GARCH process accounts for the kurtosis of a distribution, a higher moment which is related to rare events like crises. For an alternative approach and further details, see De Jonghe (2009).*

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

[32] *Canadian banks were allowed to engage in investment banking with the 1987 amendment to the Bank Act. Two years later the Second Bank Directive allowed European banks to combine commercial banking, investment banking, asset management, financial advisory activities and insurance underwriting. That may explain why our findings are closer to European studies than to the US ones, as in the latter case banks were forbidden to get involved in investment banking until the 1999 Gramm-Leach-Bliley Financial Institutions Modernization Act (DeYoung and Rice, 2004; Stiroh and Rumble, 2006; Busch and Kick, 2009).*

the banking system 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 (Stiroh and Rumble, 2006; Calmès and Théoret, 2010; Nijskens and Wagner, 2011).

In this paper, we argue that the interdependence of *snonin* and *ROA* has 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* (Equation 2) becomes more significant during the last subperiod. The endogeneity of OBS activities may not be much of a concern before 1997, but it increases substantially during the last period. In this respect, neglecting endogeneity leads to the underestimation of the positive impact of non-interest income on *ROA*, and thus of the recent improvements in bank diversification. In this respect, our analysis shows that insurance[33] and securitization present good diversification prospects for the banking industry. The market-oriented components of non-interest income also generate diversification benefits despite their relative volatility. However, these diversification benefits seem to be time-dependent and may evaporate during financial crises.

For decision makers, the policy implications we can derive from our analysis are quite straightforward. Despite the improvement in the risk-return trade-off, our study confirms that banking has become a riskier business. The volatility of bank revenues has greatly increased due to market-based activities and to the tighter link between *ROA* and *snonin*. In particular, in contractions, the variance of non-interest income has become a cause of concern since it is now much higher than the variance of net interest income. This underscores the importance of Basle-type capital adequacy rules to foster the stability of the banking system. In this respect however, it is not clear whether the use of the standard measures of leverage, as those suggested by Basle III, can really be effective for the assessment of bank systemic risk. The research agenda should aim at building more encompassing leverage measures such as the ones proposed by DeYoung and Roland (2001), and Breuer (2002).

More importantly, given their high endogeneity degree and the greater risk they pose, there is a need to better monitor OBS activities. In particular, banks should have the obligation to be more transparent about their involvement in these activities. Although the focus of the Bank for International Settlements, 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 – our study empha-

[33] *As noted previously, insurance has been identified as providing substantial diversification benefits for banks in many empirical studies. See for instance: Boyd and Graham, 1988; Kwan and Laderman, 1999; Estrella, 2001; Vander Vennet et al., 2004; De Jonghe, 2009; Schmid and Walter, 2009.*

sizes the need to also include more comprehensive measures of bank systemic risk, encompassing both the traditional measures of *VaR* and various regulatory measures of leverage, but also additional indicators of the risk inherent to OBS activities. For example, the new procedure we propose could be extended to analyze the evolution of the *ROA-snonin* relationship in other countries, or for financial intermediaries other than banks. All these issues are left for future work.

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Appendix. Robust Higher Moment 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- variable 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 $E\{(x_t-\mu_x)^3\}\neq 0$, with μ_x the expected value of x. Let us set $z_t=(x_t-\overline{x})^2$, a potential instrumental variable where $\overline{x}$ stands for the mean value of x. Then $E\{(x_t-\mu_x)(z_t-\mu_z)\}=(1-n^{-1})E\{(x_t-\mu_x)^3\}\neq 0$, and in accordance with the properties of the normal distribution, $E\{z_t\varepsilon_t\}=0$. Thus, the second-order moment $(x_t-\overline{x})^2$ qualifies as an instrumental variable for x_t. By the same token, if the distribution of x_t is leptokurtic the third-order moment $(x_t-\overline{x})^3$ 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-moment 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 moment instruments of this nature delivers a consistent estimator of β_2 , the *snonin* coefficient of our model (Equation 5). For the treatment of *snonin* endogeneity, we thus use the following set of instruments:

$$Z = \left\{x_{t-1,}(x_t-\overline{x})^2,(x_t-\overline{x})^3,\ (y_t-\overline{y})^2,(y_t-\overline{y})^3,\right\} \tag{14}$$

where x_t represents any of the explanatory variables of the bank returns model[34].

[34] *Note that when x_t is an exogenous variable, its value at time t also constitutes an instrumental variable.*