Estimating the Effect of the “Zero Lower Bound” on the Term Structure of Interest Rates in Canada

by

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Abstract

Employing a novel econometric methodology developed by Swanson and Williams (2014), I estimate how the Bank of Canada’s ability to effectively conduct unconventional monetary policy was affected when the zero lower bound (ZLB) was binding in Canada. To undertake this task, I proxy the potential effectiveness of unconventional monetary policy by the sensitivity of the term structure of interest rates to macroeconomic news, over the 16-year period 2001 to 2016. I find that, for the majority of the time Canada spent at the ZLB, the Bank of Canada could have conducted unconventional monetary policy through the medium- to long-end of the yield curve without constraint.

*Keywords:* zero lower bound, unconventional monetary policy, Bank of Canada, macroeconomic news

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1 Introduction

The zero lower bound on nominal interest rates arises from traditional economic theory that precludes the possibility of a central bank setting its key policy rate below zero. The rationale behind this non-negativity constraint stems from cash – coins and banknotes – offering a zero nominal rate of return (I refer the reader to section 2 for a primer on the theory of the effect of the zero lower bound on nominal interest rates). Though the rigidity of the zero lower bound (ZLB) has been put into question in recent years, estimating the effect of the ZLB on the efficacy of unconventional monetary policy, essentially the only policy available to a central bank while operating at the ZLB, has become a growing topic of interest in macroeconomics. I contribute to the literature by estimating the effect the ZLB had on the sensitivity of the term structure of interest rates to macroeconomic news in Canada, using a novel empirical framework developed by Swanson and Williams (2014a). Following their framework, I use the sensitivity of yields to macroeconomic news as a proxy of the Bank of Canada's ability to effectively conduct unconventional monetary policy, while operating at the ZLB from April 2009 to May 2010.

Until Japan began successively cutting its key policy rate throughout the 1990s, the concept of the ZLB was only theoretical and of little practical significance (see, for example, Eggertsson & Woodford, 2003). When Japan’s overnight rate was cut to zero in 1999, it became clear that stabilizing the economy in the face of the ZLB – a question first explored by John Maynard Keynes in the context of the liquidity trap – was a very real predicament, with there being minimal research, theoretical and empirical, on the phenomenon. Today, this predicament is more prevalent than ever before. For instance, Canada operates within 50
basis points\(^1\) (bps) of the ZLB and the United States has just recently exited a seven-year period in which the federal funds target rate was between zero and 25bps. The Bank of England cut its official bank rate to 25bps in August 2016, and the European Central Bank’s refinancing rate now rests at zero. Markets all across the world have been humming the mantra lower-for-longer.

Despite the zero lower bound’s growing significance, its notion as a true lower bound on nominal interest rates has been put into question as certain central banks have set their key policy rates below zero. At current, Japan is targeting a range of zero to \(-10\)bps. Sweden now operates at \(-50\)bps, and has been in the negative territory since February 2015. Switzerland has maintained negative rates since December 2014, and now operates at \(-75\)bps – the lowest in the world. Lastly, Denmark, which has been experimenting with negative rates for the longest (since 2012), operates at \(-65\)bps. The average twentieth century economist would have been skeptical – to say the least\(^2\) – of such a state of the world occurring, as traditional theory ruled out the possibility of negative interest rates: Why earn a negative nominal rate of return when you can just hold cash? However, central banks have been forced to consider it as a tool in the fight against stagnant growth and deflationary pressures that have persisted in the aftermath of the Great Recession (2008/09). It has simply become another unconventional monetary policy tool, which to much disbelief, has not necessitated large institutional changes as traditionally thought would be the case (e.g., see Rognlie, 2015).

\(^1\) A basis point refers to one one-hundredth of a percentage point, i.e. 100 basis points is equivalent to one percent.

\(^2\) See Ilgmann and Menners’ (2011, p.385) history of negative nominal interest rates. They note, “during the post-war era, orthodox authors have paid little attention to the possibility of negative interest rates, and scholars of the history of economic thought have called [Silvio] Gesell [credited as the first proponent of taxing money to overcome the zero lower bound] a ‘typical monetary crank’.”
Breaching the zero lower “bound” has altered the way in which economists, among others, interpret the ZLB theory (e.g., see Rognlie 2015), but it has not eliminated the need to understand monetary policy and the macroeconomy in the presence of it. Negative interest rates, though prevalent, might still be regarded as an experiment and much debate remains as to whether they are an effective monetary policy tool, and whether the risks and uncertainties outweigh the benefits (e.g., see Hannoun, 2015). With the recent influx of data from economies operating at or near the zero lower bound, a unique opportunity has presented itself to improve our understanding of the limitations the ZLB might impose on monetary policy. Applying empirical analysis to estimate the experience of economies at the ZLB is quite a void in the literature, with only few such studies existing. A seminal study, fundamental for my work, was conducted by Swanson and Williams (2014a).

Swanson and Williams, who focus on the United States, devise a novel empirical framework for indirectly quantifying the effectiveness of unconventional monetary policy in the presence of the ZLB. They do so by estimating the responsiveness of U.S. Treasury yields of various maturities to U.S. macroeconomic announcements (news) over time, motivating their empirical work within a standard New Keynesian framework. The framework suggests that the ZLB will constrain short-term yields from responding to macroeconomic news if the natural rate of interest is sufficiently below zero and the best the central bank can do is respond by lowering the key policy rate to zero (see Swanson and William, 2014a). The length of duration that the natural rate is below zero, and consequently the policy rate at zero, influences how much of the yield curve is constrained. By the expectations hypothesis, this constraint will be greatest at the short-end of the yield curve and largely unnoticeable at the long-end (so long as the duration is sufficiently short).

Swanson and Williams argue that in so far as the central bank has the capacity to manipulate
the medium- to long-end of the yield curve through unconventional tools, monetary policy can still be effective so long as the targeted portion of the yield curve remains unconstrained, that is, sensitive to news.

The authors find that while the federal funds target rate was resting on the ZLB in late 2008, the 1-year and 2-year yields remained largely unconstrained until 2010 and 2011 respectively. The 5-year and 10-year yields did not begin to exhibit constrained behaviour until the last few weeks of 2012 (the end of their sample period). From these observations, they conclude that the Federal Reserve’s ability to conduct monetary policy out a year or more on the yield curve (e.g., through quantitative easing) retained the effectiveness of such tools being implemented throughout “normal” periods; i.e., periods when the ZLB is not binding. The results have important implications for understanding the efficacy of monetary policy at the ZLB and for determining which portion of the yield curve unconventional tools should target.

Following Swanson and Williams econometric methodology, I estimate how the sensitivity of Government of Canada treasury bill and bond yields (i.e., the term structure of interest rates) has changed to macroeconomic news over the 16-year period 2001 to 2016. Though Canada has not had quite as prolonged of an experience at the ZLB as the U.S., the Bank of Canada did maintain the overnight rate at the effective lower bound (25bps) for a little over 13 months following the onset of the Great Recession (see Witmer & Yang, 2016). In this paper, I exploit that 13-month period (April 2009 to May 2010) to estimate the degree to which the ZLB constrained Canadian yields. Looking at numerous points along the yield curve from the 3-month treasury bill to the 30-year Government of Canada bond, I explore how far out the yield curve the Bank of Canada would have had to reach in order to
influence the economy through unconventional monetary policy. To the best of my
knowledge, only one other paper attempts to quantify the effect of the ZLB in Canada.

Richhild Moessner (2014), using a more restricted econometric methodology than
Swanson and Williams (2014), measures the sensitivity of Government of Canada bond
yields to both U.S. and Canadian macroeconomic news over the period 1998 to 2013. He
finds that the ZLB dampened the sensitivity of the Canadian 1-year yield – the shortest
maturity considered in his study – to domestic macroeconomic news by 77%, and by 49% to
U.S. news. The yields on 2-, 5-, and 10-year bonds, however, remained unconstrained. Thus,
Moessner concludes that monetary policy, in so far as it could affect yields out two years or
more on the yield curve, would have been as effective as normal while the ZLB was binding
in Canada.

Similar to Moessner, I account for both U.S. and Canadian macroeconomic news as
drivers behind daily changes in Canadian yields. Canada is a small open economy heavily
influenced by foreign shocks, predominantly by the United States, making it essential to
widen the scope to include U.S. news in my analysis. In fact, evidence suggests that Canadian
yields have historically been equally or more responsive to U.S. news than to domestic news
(e.g., see Gravelle & Moessner, 2002). Complementing Moessner’s work, I apply a less
restrictive econometric methodology using more recent data, the results of which suggest
that Moessner’s model was perhaps too restrictive. Employing Swanson and Williams’
framework, I estimate the stability of the sensitivity of yields to macroeconomic news by
estimating rolling regressions that allow the sensitivity parameter to take on a different value
every 10 days over the entire sample window (398 estimates). This contrasts with Moessner,
who limits the sensitivity parameter to only two values (i.e., one when the ZLB was binding
and one for the rest of the sample). As explored at length in section 4, the data suggests that
these time-varying sensitivity parameters are likely highly unstable, and that even allowing them to take on different values on an annual basis would be too restrictive.

To preview my results, I find that Canada’s 13-month experience at the zero lower bond is likely characterized by two starkly different intervals. For the first four months, from April to July (2009) all yields, from the 3-month to the 30-year, are found to be unresponsive to macroeconomic news. These results form somewhat of an anomaly, unexplained by Swanson and Williams’ illustrative New Keynesian framework and inconsistent with the alternative explanation I provide to motivate the behaviour of yields exhibited throughout all other periods. The cause of the behaviour that I observe over these four months remains to be understood. For this reason, the effectiveness of unconventional monetary policy over this window cannot be inferred from my results.

For the following 10-month period, from August 2009 to May 2010, the empirical analysis I conduct suggests that the Bank of Canada retained considerable power over medium- and longer-term interest rates. In so far as the Bank has the ability to manipulate these interest rates through unconventional monetary policy tools (e.g., quantitative easing), my findings suggests that the Bank would have been able to do so without constraint. In this regard, my findings support Moessner’s claim that unconventional monetary policy targeted at bonds with greater than one year to maturity would have retained normal effectiveness. That said, my analysis is not only limited to the period in which the zero lower bound was binding. Exploring the sensitivity of yields to macroeconomic news post-May 2010, my outcomes suggests that perhaps the true cause of the constraint may not actually be the zero lower bound.

In terms of the short-end of the yield curve, my analysis suggests that the yield curve remains largely insensitive to macroeconomic news, at least over the seven-year period 2010
to 2016. Given that the Bank of Canada lifted the key policy rate off the lower bound in May 2010, the ZLB explanation of this insensitive behaviour does not seem to be consistent with my results. This conclusion is conceivably further corroborated by the lack of response short-end yields exhibited post-December 2015, when the Bank of Canada announced that Canada’s effective lower bound is –50bps, no longer 25bps (see Witmer & Yang, 2016). If the ZLB was the reason for the constrained behaviour of yields to macroeconomics news, the additional 75bps buffer for conventional monetary policy to take place should have provided some relief. For these reasons, I propose an alternative explanation that perhaps better describes the underlying phenomenon. Namely, I propose that the insensitivity of short-end yields, implied by the tests conducted in this paper, are the result of those yields exhibiting substantially lower volatility from April 2009 onward. I elaborate on this purposed explanation in the thesis.

There are broadly three bodies of literature relevant to this paper: (1) the body of research measuring the effect of the ZLB on the term structure of interest rates in Canada; (2) the studies measuring the effect of the ZLB on the term structure of interest rates outside of Canada; and (3) the empirical work on estimating the effect of macroeconomic news on bonds and other financial assets. As noted (and summarized) above, Moessner’s work on Canada, to the best of my knowledge, is the total extent of the research measuring the effect of the ZLB in Canada. Thus, in the remainder of this section, I touch on the other two relevant research areas to provide spatial context for where this paper resides among its surrounding literature.

As alluded to in the beginning of this section, the issue of the zero lower bound has been in no way contained to only Canada and the United States. It is a pervasive predicament. This has created the opportunity for researchers to estimate the effect of the
ZLB on interest rates in numerous economies. For example, Moessner, Haan and Jansen (2016) measure the effect of the ZLB on interest rate swaps in Sweden during 2009 to 2010 and again from 2014 to 2015 – two separate periods in which the ZLB was binding. Their methodology closely parallels Moessner (2014). The authors conclude that only short-term maturities were constrained in 2009/10. During Sweden’s second bout of the ZLB in 2014/15, all maturities remained unconstrained, possibly the result of markets understanding that the policy rate had room to go negative (which it eventually did).

In another paper by Swanson and Williams (2014b), the authors apply the methodology of Swanson and Williams (2014a) – employed in this paper – to both yields and exchange rates for the U.K. and Germany. They find that both exchange rates paired against the U.S. dollar were unaffected by the zero lower bound, consistent with the theory that they use to motivate their empirical work. UK gilts and German bunds, on the other hand, did experience periods of being constrained, though the behaviour exhibited by these yields were not as clearly explainable as the results obtained from their analysis of the United States.

Several more recent papers have added even greater depth to the U.S. literature. Carvalho, Hsu and Nechio (2016) modify Swanson and Williams’ (2014a) methodology to take a more direct approach, i.e. by looking directly at Federal Reserve communications, to measure the effect of the ZLB in the United States. In a similar paper, Wu (2016) estimates the time-varying sensitivity of monetary policy shocks on the broader financial markets in the U.S., including corporate bond markets and stock markets. Analyzing the responsiveness of real interest rates, as opposed to nominal interest rates, is the focus of Zhang (2016), who looks at how Treasury Inflation-Protected Securities responded to macroeconomic and monetary policy news while at the ZLB. Arguably, the United States has made for the most
interesting case study given the length of time spent at the ZLB and the accessibility of data on the economy (an issue in analyzing Japan’s experience, for instance).

Outside the context of the ZLB, there is a wide body of literature using econometric models to estimate the effect of macroeconomic news on bond yields and other financial returns and assets. Even within the scope of Canada, this body of work has relative depth. Gravelle and Moessner (2002) use sample data from 1995 to 2000 to measure the sensitivity of interest rates to both U.S. and domestic macroeconomic news to understand how the Bank of Canada’s (the Bank) steps towards improving transparency over the 1990s influenced the market (i.e., to see if market participants better understood the Bank’s monetary policy reaction function). Their model closely resembles the first step of the empirical work that I undertake in this thesis.

Nicolas Parent (2002), adopting a similar approach to Gravelle and Moessner (2002), uses two more years of data to explore the effect of the Bank’s introduction of fixed announcement dates (often seen as another step towards greater transparency). A generalized autoregressive conditional heteroskedastic model is examined by Hayo and Neuenkirch (2010), who look more broadly at the effect of Canadian and U.S. central bank communications, along with macroeconomic news, on Canada’s financial markets and their volatility over the period 1998 to 2006. Their study extends Parent (2002) by widening the scope of the analysis to examine stocks, bonds and foreign exchange markets.

Several other researchers explore the effect of macroeconomic news and central bank communications with a cross-country perspective. For example, Kearns and Manners (2006) use intraday data to analyze the effect of domestic monetary policy surprises on exchange rates in Canada, Australia, New Zealand and the United Kingdom. They find that, on average, a 100bps increase in the key policy rate of a country leads to a one and a half
percent appreciation in their exchange rate (paired against the U.S. dollar). Zettelmeyer (2004) follows a similar line of research (excluding the U.K.), concluding that a 100bps increase in the domestic 3-month interest rate, caused by a domestic monetary policy shock, results in the exchange rate appreciating by two to three percent. The impact of U.S. macroeconomic announcements on stock market returns in Canada, Britain, Germany, Hong Kong and Singapore is analyzed by Lucey et al. (2008). As likely expected, they find that U.S. macroeconomic news plays a role in the stock market movements of these countries.

The body of literature encompassing similar works untied to Canada is unsurprisingly vast. Bond markets, stock markets, and foreign exchange markets each have their own domain in the literature, which one could further dichotomize into research on those markets’ sensitivities to central bank communications and to macroeconomic news. I do not attempt to summarize this body of work as even a brief overview could fill several papers. I merely raise this research to indicate the breadth and depth of the studies interested in related topics and highlight that there is a long history of econometric methods used in the literature, some of which resembles the baseline empirical work I undertake in my thesis. For a history of the analytical tools used to measure the effect of macroeconomic news on stocks and bonds see, for example, Fleming and Remolona (1997). Additionally, for an influential paper on how foreign exchange markets respond to macroeconomic news, with implied links to economic fundamentals, I refer the reader to Andersen, Bollerslev, Diebold and Vega (2003).

The remainder of this paper is structured as follows. Section 2 provides a simple theoretical motivation for the empirical work I conduct, which is centered around the expectations hypothesis of the term structure of interest rates. The empirical strategy I
employ to estimate the time-varying sensitivity of yields to U.S. and domestic macroeconomic news is outlined in section 3, along with preliminary results that preface the focal point of this paper. The main results of the fundamental empirical research that I undertake are presented in section 4. Section 5 provides brief details of several sensitivity explorations, and section 6 concludes. Finally, an Appendix provides greater detail on the data used in the empirical analysis.

2 Theoretical Motivation

The theory of the zero lower bound on nominal interest rates is a basic tenet of monetary economics, which states that the central bank cannot lower its key policy rate below zero. The rationale behind this non-negativity constraint stems from cash – coins and banknotes – offering a zero nominal rate of return. In the retail banking sector, the theory suggests that an interest rate below zero (e.g., on saving accounts) would cause households to withdraw deposits and hold cash to avoid earning a negative nominal rate of return (e.g., see Rognlie, 2015). The incentive for households to do so is clear: a nominal return of zero is more attractive than a negative one. In the more complex sectors of the banking industry, the story is roughly the same. Hoarding cash, along with disruptive incentives and structural institutional changes, are all key issues that arise from violating the zero lower bound (e.g., see Witmer & Yang, 2016). As explained in the introduction though, negative rates are a reality in several parts of the world.

Unsurprisingly, storing wealth in banknotes and coins is risky. Proper safeguards to preserve the integrity of cash over time, including storage facilities and insurance, are expensive. Moreover, in modern economies, paying bills in cash is a cumbersome alternative to using available electronic technologies. Combined, these points provide rationale for why
nominal interest rates can in fact fall below zero without causing demand for cash to skyrocket. As such, central banks sometimes refer to their estimated “effective lower bound” (ELB), which in essence attempts to account for such costs that firms and households are willing to incur. The Bank of Canada, for example, currently estimates that the effective lower bound for the Canadian economy is –50bps (Witmer & Yang, 2016). The important point is that despite the prevalence of negative rates in the global economy, there remains a lower bound. At some point, the cost of holding negative nominal rate of return assets surpasses the cost of safeguarding cash.

Over the last two years, a number of central banks have taken advantage of this additional buffer and lowered their key policy rate below zero. For those that have, conventional monetary policy has continued to provide monetary stimulus. However, this buffer below zero is limited, and once the key rate bumps up against the effective lower bound, conventional policy is likely to become impotent. Therefore, the key concepts related to the issue of the zero lower bound on nominal interest rates is still a concern, despite the lower bound of some countries in reality being effectively below zero.

Without room to lower the key interest rate, the central bank cannot follow its monetary policy rule to offset shocks to the economy. A monetary policy rule can simply be thought of as a mathematical equation that the central bank follows to adjust its key interest rate to economic conditions in pursuit of some goal. The Bank of Canada’s goal, for instance, is to stabilize inflation at two percent. The rule’s simplicity allows markets to form expectations about how the central bank will respond to the state of the economy, which is perpetually in flux as a result of current shocks hitting the economy and past shocks propagating through the economy (e.g., refer to Gali, 2008). As an example of a rule, Swanson and Williams (2014a) (hereafter referred to as SW) use a Taylor-type rule (Taylor,
1993) – see equation (1) below – to illustrate how a central bank in pursuit of an inflation target would become powerless at the ZLB. It takes the form:

\[
i_t = \max \{0, \pi_t + r_t^* + 0.5(\pi_t - \bar{\pi}) + 0.5\bar{y}_t\}
\]

where \(i_t\) is the one-period nominal interest rate set by the central bank (e.g., the overnight rate in Canada), subject to a non-negativity constraint. The parameter \(\pi_t\) is the rate of inflation at time \(t\), \(\bar{\pi}\) is the targeted rate of inflation, \(r_t^*\) is the natural interest rate at time \(t\) and \(\bar{y}_t\) is the output gap at time \(t\).³ One way to think about modelling shocks to the economy would be to consider output shocks (as opposed to inflation shocks) that are embodied by shocks to the natural rate of interest, \(r_t^*\). To further illustrate, SW define \(r_t^*\) to follow an AR(1) process of the form:

\[
r_t^* = \rho r_{t-1}^* + e_t; \quad e_t \sim iid(0, \sigma^2)
\]

\[0 < \rho < 1\]

The central bank is assumed to adjust \(i_t\) using equation (1) in response to shocks that perturb the natural rate of interest and propagate throughout the economy, as defined by equation (2). When the ZLB is not an issue, it can exactly offset such shocks. For example, if \(r_t^*\) is shocked and falls from 10% to 9%, then the central bank would initially lower \(i_t\) by 1%, and then gradually increase it as the shock dissipates.⁴ As equation (1) makes clear though, there is an issue if a shock pushes the natural rate of interest sufficiently below

³ The output gap is the difference between an economy’s actual output and its estimated potential output. The natural rate of interest is not as easily defined. For most purposes, the rate can be interpreted as the real rate of interest that closes the output gap, stabilizes inflation and results in full employment. A large body of literature on the natural rate of interest has accumulated since its introduction in 1898 by Knut Wicksell (Holston, Laubach and Williams, 2016). For a modern take on estimating the natural rate of interest in countries including Canada see, for instance, Holston, Laubach and Williams (2016).

⁴ In the complex New Keynesian framework, the central bank’s response is not exactly one-for-one with the shock. The central bank must additionally account for changes in inflation expectations. I refer the reader to Gali (2008, p.54) for a robust discussion of a central bank’s reaction to a technology shock. For our purposes though, it is sufficient to think of the responses as a one-for-one relationship.
zero. Though this is the case of interest, it is illustrative to initially work through the case in which the ZLB is not an issue.

Imagine that the Canadian economy is hit with a very large negative output shock that causes the natural rate of interest ($r^*_t$) to fall from 10% to 5%. Assuming that the key interest rate is well above zero (say 14%), the central bank has considerable room to lower the rate (in this case to 9%) to offset the shock, giving it the power to keep inflation on target at two percent and the economy close to potential GDP. The market, understanding the Bank of Canada’s monetary policy rule, can reasonably estimate that the Bank will lower the overnight rate in an effort to maintain target inflation. Subsequently, asset prices respond to the central bank’s monetary policy reaction (e.g., see Gravelle & Moessner, 2002). This response is perhaps easiest to see in nominal yields on Government of Canada bonds and treasury bills, which are closely tied to the Bank’s overnight rate (e.g., see Lange, 2005).

The expectations hypothesis of the term structure of interest rates suggests that yields on long-term bonds are the average of current and expected future short rates (the Bank’s overnight rate in this example), which are themselves functions of the real interest rate and expected inflation. An approximation of the formal term structure of interest rates relationship typically takes the form of equation (3) (Romer, 2011, p.519):

$$i^m_t = \frac{i_t^1 + E_t i_{t+1}^1 + E_t i_{t+2}^1 + \cdots + E_t i_{t+m-1}^1}{m} + \theta_{m,t}$$

where $i^m_t$ is the nominal interest rate at time $t$ on an m-period maturity bond. The formula states that the nominal interest rate on an m-period bond ($i^m_t$) is equal to the average of the current and expected future one-period interest rates (short-rates) plus a term premium that depends on the maturity of the bond. In the above formula, the expectations of future short
rates are conditional on information at time $t$ and denoted as $E_t$ (e.g., $E_t i^1_{t+1}$ is the expectation conditional on information at time $t$ of the short-rate one period ahead).

A shock, which lowers the natural rate, prompts the central bank to lower the overnight rate in an effort to lower the real interest rate in the economy – down with the natural rate – causing certain nominal interest rates to fall. Which nominal rates would be affected? The answer depends on how persistent the shock is. We can think of the short-rates that appear in equation (3) as the overnight rate or one-day interest rate. For concreteness, I assume that the market believes that the negative shock will have a relatively transitory effect, in other words its persistence is rather weak, and say disappears in one-year following the AR(1) process described by equation (2). In this case, the market would revise downward its expectations on short-rates over the coming year (i.e., $E_t i^1_{t+1}, E_t i^1_{t+2}, ..., E_t i^1_{t+364}$) consistent with the markets expectations on how the central bank will adjust the overnight rate to offset the gradually dissipating shock. Expectations on short-rates past one-year (i.e., $E_t i^1_{t+365}, E_t i^1_{t+366}, ..., E_t i^1_{t+m-1}$), however, would remain unchanged. Thus, the largest fall in interest rates will occur on the short-end of the yield curve. To illustrate this, I formulate equation (3) for a 1-year, 2-year and 30-year bond, interpreting the short-rate as the overnight rate, assuming 365-day years, and ignoring the term premium for simplicity. This leads to:

\begin{align*}
\hat{i}_{t}^{1\text{ year}} &= \frac{i^1_t + E_t i^1_{t+1} + E_t i^1_{t+2} + \cdots + E_t i^1_{t+364}}{365} \\
\hat{i}_{t}^{2\text{ year}} &= \frac{i^1_t + E_t i^1_{t+1} + E_t i^1_{t+2} + \cdots + E_t i^1_{t+729}}{729} \\
\hat{i}_{t}^{30\text{ year}} &= \frac{i^1_t + E_t i^1_{t+1} + E_t i^1_{t+2} + \cdots + E_t i^1_{t+10,949}}{10,949}
\end{align*}
Equation (4) makes it clear that the 1-year nominal interest rate is the average of current and expected future short-rates over the coming year. Given that the market revised downward its expectations on each individual short-rate, the entire expression is affected by the shock. The 2-year rate, represented by equation (5), is also affected by the shock given that it is in part determined by the same expectations on short-rates over the coming year. However, the 2-year rate is also determined in part by the expectations on short-rates realized greater than a year from now, which are unaffected by the shock. Therefore, the 2-year interest rate is affected, but not to the same degree as the 1-year rate. This feature can be seen even more obviously in the thirty-year bond rate – equation (6) – in which the 364 lower expectations on short-rates over the coming year are swamped by the 10,585 unaffected expectations for the following 29-years. The pattern is evident – the longer is the maturity of the bond, the less affected is the yield to the shock. Equivalently, the greater the persistence of the shock, the further out the yield curve that shock would be felt. A similar story could be told with respect to a positive shock. The case of interest here though, is what happens to yields when the central bank is operating at the zero, or effective, lower bound.

Initially, imagine that the economy is in equilibrium with the central bank’s optimal policy being to keep the overnight rate exactly at one percent; i.e., $i_t^1 = 1\%$. For simplicity, let us assume that the economy’s effective lower bound is the zero lower bound. Now suppose that the economy is hit with a large negative shock, causing the natural rate of interest to fall to $-5\%$. This creates a problem. As noted above, the job of the central bank is to adjust the real interest rate in the economy – targeting the natural rate – by means of adjusting the overnight rate. If the shock is sufficiently large enough to cause the entire expression $\pi_t + r_t^* + 0.5(\pi_t - \bar{\pi}) + 0.5\hat{y}_t$ to become negative, then in absence of the zero lower bound, the central bank would want to set the overnight rate below zero. Equation (2)
explicitly restricts this option though, because of the zero lower bound. Therefore, the bank’s key policy tool becomes ineffective at stabilizing the economy.

Though, perhaps, the illustrative story sounds a little bit unrealistic, it roughly parallels an experience of the United States’ during the 2008/09 Financial Crisis. At the peak of the Crisis, standard monetary policy rules suggested that, in the absence of the zero lower bound, the Federal Reserve – if it were to exactly follow such a rule – would have reduced its federal funds target rate to $-5\%$ (Rudebusch, 2009). With the zero lower bound in its way, of course, it was unable to implement this adjustment. The best option available to the Federal Reserve, in terms of using conventional monetary policy to stabilize the economy, was to keep the rate at zero. In my example, $i_t^l$ would be immediately lowered to zero; however, this would not be enough to completely offset the shock. Thus, expectations on future short-rates would be revised down to zero for as long as the market believes the shock will keep the expression $\pi_t + r_t^* + 0.5(\pi_t - \bar{\pi}) + 0.5\bar{y}_t$ negative. If the shock is expected to persist for several years, and over this period the expected optimal policy remains as $i_t^l = 0\%$, the shock would drag down a substantial portion of the yield curve. Thus, similar to the case in which the ZLB is not an issue, the shock causes markets to revise expectations on future short-rates downward (to the extreme of zero in this case), which reverberates along the yield curve in exactly the same manner. That is, the longer the maturity the less affected by the shock, or equivalently, the greater the persistence, the further out the yield curve the shock will be felt.

One could further imagine that another negative shock\(^5\) hits the economy, causing the natural rate of interest to fall even further below zero, say $-7\%$. How would yields

\(^5\)This also holds true for positive shocks that are not sufficiently large enough to induce the optimal overnight rate to become positive; i.e., $\pi_t + r_t^* + 0.5(\pi_t - \bar{\pi}) + 0.5\bar{y}_t$ remains below zero.
respond now that the central bank is already operating at the zero lower bound? The answer depends on the yield, and once again on the persistence of the current (and previous) shock. If the current and previous shocks are expected to persist, for instance, for around three years, then short-end yields, such as the 3-month, 6-month, 1-year and 2-year, are likely to be completely unresponsive to the current shock. This is because the market is aware of the central bank’s limitations and knows that the central bank cannot lower the overnight rate any further. Therefore, expectations on future short-rates for the next three years will not change because of the shock, as these rates have reached a minimum. In a more extreme scenario, if the market believes the persistence of the previous and current shocks are long enough, then then the 5-year, 10-year and even the 30-year yield could be substantially constrained from going down in response to one or more shocks. The long-horizons on such bonds, however, give rise to larger term premiums and possibly larger expected inflation components that result in an effective lower bound on these yields that is likely above zero.

Swanson and Williams (2014a) formalize this illustrative story in a standard New Keynesian framework. I refer the reader to their paper for a detailed outline of the theoretical implications. As the empirical work in this paper only requires a heuristic understanding of why interest rates along the yield curve would become unresponsive to economic shocks in the presence of the ZLB, I have only provided a sketch of the possible mechanisms at play.

The aim of my research is to estimate the degree to which the ZLB constrained yields in Canada, and to highlight how much of the yield curve was constrained. Despite conventional monetary policy being ineffective at the ZLB, unconventional tools, such as forward guidance and quantitative easing, can still be effective so long as such policies target,
and condition expectations, on the parts of the yield curve that remain unconstrained. As such, quantifying the effect of the zero lower bound, undertaken in the following sections, has considerable implications for monetary policy.

3 Empirical Framework & Preliminary Results

I employ Swanson and Williams’ (2014) empirical framework to estimate the time-varying sensitivity of Canadian yields to U.S. and domestic macroeconomic news. It consists of three components. In the first component, I establish a baseline estimate of the effect of news on the term structure of interest rates during a period in which the zero lower bound is not an issue (i.e., a “normal” period). In the second component, I estimate how the effects vary over time with a particular interest on the behaviour of yields throughout April 2009 to May 2010 – when the ZLB was binding in Canada. Preliminary results for the first and second steps are described alongside their methodology to preface the core empirical results undertaken in the last step, and reported on in the main results section. In this last component, I explore the stability of the time-varying sensitivity parameter at a more granular level, as compared to the second component, by giving the parameter the freedom to vary every 10 days over the entire sample window. I refer the reader to section 5 for a brief discussion on several sensitivity explorations considered for this empirical analysis.

i. Data

Throughout my study, two sets of daily data are employed: yield data and macroeconomic news data. The yield series is made up of 3- and 6-month Government of Canada treasury bill yields along with 1-year, 2-year, 5-year, 10-year and 30-year Government of Canada bond yields. All series reflect end of day observations and have been sourced
from Bloomberg (the data appendix provides explicit information regarding ticker IDs for both sets of data). The macroeconomic news series is itself comprised of two datasets. Consistent with the literature, “news” is defined as deviations from the market’s expectations. The rationale for such a definition stems from financial markets being forward looking and forming their own expectations about the path of the macroeconomy. Asset prices, which incorporate these expectations, therefore only respond to deviations from them (e.g., see Swanson and Williams, 2014a).

For each macroeconomic variable considered (e.g., real GDP growth) the news component is created by subtracting the market’s expectation from the realized value for each given release. For example, if quarter-over-quarter real GDP growth for the first quarter of 2016 was expected to be one percent and Statistics Canada announces to the market on June 3, 2016 that it was two percent, the “news” would be a positive one percent surprise in GDP. This approach regards this component as news because financial markets have already essentially responded to their expectations. Measuring expectations is, of course, difficult and subject to error and interpretation. For this study, I use the median forecast from Bloomberg’s survey of financial institutions and professional forecasters6 to proxy market expectations.

The macroeconomic variables included in my main analysis are chosen to be consistent with both Moessner (2014) and Swanson and Williams (2014a). With the exception of three variables7 used by Moessner that could not be obtained due to data limitations, all other variables are present, including Swanson and Williams’ U.S. regressors.

---

6 Bloomberg Financial Services and Money Market Services are the two most commonly used sources for market expectations in the literature, though Bloomberg has been cited as the superior source (see, for example, Ouadghi, Mignon and Boitout, 2015 for a comparison). Swanson and Williams (2014a, p. 3162) note that “these data pass standard tests of forecast rationality and provide a reasonable measure of ex-ante expectations of the data releases...”.

7 The Bloomberg surveys for Canadian retail sales, U.S. hourly earnings, and U.S. trade data seemed not to be available, despite an extensive search.
In total, twenty-five news variables are employed to conduct my main empirical analysis. I also obtained other variables not included in Moessner’s and Swanson and Williams’ regressions, as possible explanatory factors. The sensitivity of my results to these additional explanatory variables is explored in section 5.

Both datasets begin on January 2, 2001 and conclude on November 9, 2016. Given that bond markets are closed on weekends and holidays, these days have been excluded from the sample. Following Swanson and Williams (2014a), I also exclude days in which no macroeconomic announcements were made.\footnote{Similar to Swanson and Williams (2014a), I find that my results are not qualitatively different when including days without macroeconomics announcements. These additional results are available upon request.}

ii. The Effect of Macroeconomic News During “Normal” Times

Throughout the seven-year period 2001 to 2007, the Bank of Canada’s overnight rate fluctuated quite considerably. At a high of 5.75% in 2001, it fell to 2% in 2002 before jumping back up to 4.5% by 2007 (see Figure 1 below). Even at its low point of 2%, however, the policy rate retained considerable headroom above the zero lower bound. It was not until 2008 that the overnight rate began cascading down toward its eventual resting place at 0.25% in April 2009, where it stayed until May 2010. For this reason, along with data limitations imposed by Bloomberg surveys (many surveys only go back as far as 2000), I define 2001 to 2007 as the “normal” period in which the zero lower bound is not considered a concern.

To begin my analysis, over this normal period, I consider the following model to explain the effect of macroeconomic news on yields:

\[
\Delta y_t^m = \alpha^m + \sum_{j=1}^n (\beta_j^m x_{j,t}) + \varepsilon_t^m; \quad \varepsilon_t^m \sim (0, \sigma_{\varepsilon_t^m}^2)
\]
where $\Delta y_t^m$ refers to the one-day change in the yield on day $t$ of bond maturity $m$, $n$ refers to the number of macroeconomic news variables included in the model, and $x_{j,t}$ denotes the (standardized) macroeconomic news produced by variable $j$ on day $t$. Due to their differing units, news variables are standardized by dividing each observation by the respective series’ historical standard deviation. This allows for a common interpretation of $\beta_j^m$ as the basis for their estimated point per standard deviation effect of macroeconomic news variable $j$ on the one-day change in bond yield $m$. The regressors are assumed to be uncorrelated with the disturbance term $\varepsilon_t^m$ so that model (7) is estimated using ordinary least squares. Table 1 provides the estimated coefficients for the 3-month, 2-year and 10-year yields, along with heteroskedasticity-consistent t-statistics in parentheses.
The disturbance terms, $\varepsilon_t^m$, are assumed to be mean-zero and independent, but are not assumed to be identically distributed due to unconditional heteroskedasticity. Though it may be somewhat surprising that we need to allow for unconditional heteroskedasticity in the context of time series data, the rationale for doing so is rather clear. The volatility of an interest rate moves in tandem with its level (see, for example, Chan et al., 1992). SW (2014b, p.13) remark that low interest rates simply have “less room to run”, and hence, periods of low interest rates are typically associated with low volatility, and periods of high interest rates are usually coupled with high volatility. I come back to this feature of the data in the main results section. In regard to the independence of $\varepsilon_t^m$, I settled on this assumption only after experimenting with heteroskedasticity and autocorrelation consistent (HAC) standard errors (SES) in addition with the more restricted heteroskedasticity-consistent (HC) SES. Autocorrelation, though mildly present at shorter maturities, is not material enough to affect the SES.\footnote{The regression results using HAC standard errors are available upon request.} For this reason, I use HC SES throughout the study, which is the norm in the literature (see, for example Carvalho, Hsu, Nechio, 2016; Wu, 2016; Hayo & Neuenkirch, 2012; and, Fleming & Remolona, 1997).

The results are surprisingly similar to those found by Moessner (2014). For example, I estimate the effect of the Canadian CPI inflation rate on the 2-year yield as 2.37bps, with the t-ratio suggesting that the coefficient is significant at the one percent level. Moessner’s corresponding result from his nonlinear least squares regression is 2.58bps (similarly, significant at the one percent level). This is in spite of Moessner considering a different sample window (1998 to 2013) and including three additional regressors.

Moreover, the signs of my estimates are largely consistent with those reported by Moessner, and more importantly, consistent with sign predictions from economic theory.
Table 1 – Coefficient Estimates of $\beta_j^m$ from OLS Regression $\Delta y_t^m = \alpha^m + \sum_{j=1}^{n} (\beta_j^m \ x_{j,t}) + \epsilon_t^m$ over the Sample Window 2001 to 2007

<table>
<thead>
<tr>
<th></th>
<th>3-Month</th>
<th>2-Year</th>
<th>10-Year</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Canada</strong></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Capacity Utilization</td>
<td>0.43</td>
<td>1.24</td>
<td>(1.13)</td>
</tr>
<tr>
<td>CPI</td>
<td>-0.94</td>
<td>0.82</td>
<td>(1.65)</td>
</tr>
<tr>
<td>GDP</td>
<td>0.77</td>
<td>1.47</td>
<td>(0.84)</td>
</tr>
<tr>
<td>Core PPI</td>
<td>-0.03</td>
<td>0.35</td>
<td>(1.00)</td>
</tr>
<tr>
<td>Nonfarm Payrolls</td>
<td>1.53</td>
<td>4.52</td>
<td>(6.07)</td>
</tr>
<tr>
<td>Unemployment Rate</td>
<td>-1.65</td>
<td>-1.64</td>
<td>(-2.02)</td>
</tr>
<tr>
<td>ISM Manufacturing</td>
<td>1.81</td>
<td>2.30</td>
<td>(3.15)</td>
</tr>
<tr>
<td>Consumer Confidence</td>
<td>1.06</td>
<td>1.38</td>
<td>(1.96)</td>
</tr>
<tr>
<td>Retail Sales ex. autos</td>
<td>1.45</td>
<td>1.85</td>
<td>(2.48)</td>
</tr>
<tr>
<td>Initial Claims</td>
<td>-1.07</td>
<td>-1.50</td>
<td>(-5.12)</td>
</tr>
<tr>
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<td>(-0.18)</td>
</tr>
<tr>
<td>House Sales</td>
<td>0.18</td>
<td>0.25</td>
<td>(0.61)</td>
</tr>
<tr>
<td>Industrial Production</td>
<td>0.33</td>
<td>0.78</td>
<td>(0.92)</td>
</tr>
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<td><strong>United States</strong></td>
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<tr>
<td>Capacity Utilization</td>
<td>0.43</td>
<td>1.24</td>
<td>(1.13)</td>
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<tr>
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<td>0.33</td>
<td>0.78</td>
<td>(0.92)</td>
</tr>
<tr>
<td><strong>Observations</strong></td>
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<td>1,134</td>
</tr>
<tr>
<td><strong>H0: $\beta = 0$, p-value</strong></td>
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<td>0.00</td>
<td>0.00</td>
</tr>
</tbody>
</table>

Notes: The coefficient estimates are interpreted as basis point per standard deviation responses to macroeconomic news. Heteroskedastic consistent t-statistics are in parentheses. The hypothesis test is in reference to an F-test for the joint relevance of $\beta^m$.

For instance, considering the effects of U.S. and Canadian unemployment rates, we see that both variables have a negative effect on yields, irrespective of maturity – i.e., a surprise increase in the unemployment rate has the effect of lowering yields. The news causes financial markets to adjust their expectations on inflation downward – perhaps seen most clearly through the lens of the short-run Phillips curve – in light of the worse than expected
economic conditions. Interrelatedly, the effect can be interpreted as the market revising downward expectations on future short-rates. In contrast, the effect of news in the Canadian CPI inflation rate is consistently positive. This outcome suggests that a surprise increase in inflation causes the market to revise inflation (and short-rate) expectations upward, pushing nominal yields higher – illustrated by the Fisher equation.

Somewhat more remarkable is that Swanson and Williams’ (SW) results for the United States even bear resemblance to the Canadian estimates that I obtain. Most notably, the variables that have the greatest impact on U.S. yields tend to have the greatest impact on Canadian yields. As an example, SW estimate that the average 2-year yield basis point per standard deviation response to U.S. nonfarm payroll and ISM manufacturing is 4.56bps and 3.44bps, respectively. They are the two largest estimated effects on the U.S. 2-year yield reported by SW. My corresponding estimates for the impact of these U.S. explanatory variables on the Canadian 2-year yield are 4.52bps and 2.30bps, respectively. Out of all the U.S. regressors that I consider, these two explanatory factors produce the largest estimated effects. These similarities of my results to those of SW are found in spite of SW considering: (1) a different sample window (1990-2000), (2) U.S. yields and (3) only U.S. regressors. This perhaps supports the notion that it is reasonable to assume that the $\beta$ vector is stable over time (and perhaps stably related to its U.S. counterpart). My outcomes also reinforce the findings of Gravelle and Moessner (2002) that U.S. news tends to be equally or more important than domestic news for Canadian yields. As an example of this feature, the single largest estimated effect on the Canadian 2-year, 5-year and 10-year yield is U.S. nonfarm payroll.

It is reasonable to question whether these estimated impacts are practically significant. After all, a four or five basis-point movement certainly does not appear to be all
that interesting. There are several reasons, however, that we would expect these micro-effects, and that they are practically relevant. First, the underlying macroeconomic conditions of Canada – or for any country – are hardly characterized by any single statistic. The unemployment rate, for example, though important, is but one statistic among hundreds that financial markets use in an attempt to decipher the true state of the economy. Moreover, the majority of the economic news variables considered in this study, are so-called “headline” statistics. In the reports that accompany each headline statistic are a myriad of other economic statistics that give it context. A one standard deviation surprise in the unemployment rate (the headline labour statistic), for example, may not be all that concerning if it was caused by an increase in the participation rate (found in the Statistics Canada Labour Force Survey). So it is important to keep in mind that the news variables considered in this study only give part of the story. Due to the lack of market expectations data for many non-headline variables, this limitation is common throughout the literature.

Second, when interpreting the seemingly rather small coefficient estimates reported in Table 1, we need to consider that a one-standard deviation surprise in any of the macroeconomic news variables does not necessarily suggest markets should adjust their expectations by all that much. Markets are prone to making errors in their forecasts and are (we hope) aware of their limitations. As a result, they should be expecting forecast error, and consequently, should not really be all that surprised if their forecasts are inaccurate. In other words, a one standard deviation news surprise might not be all that surprising!

Last to consider is that it takes a lot to move interest rates. Gurkaynak, Sack and Swanson (2005) show that in normal times the Federal Reserve would have to unexpectedly cut the federal funds target rate by 100bp in order to manufacture a 20bps decline in medium- and longer-term government yields. Given that the news variables considered in
this study clearly have a less direct effect on yields than a central bank’s key policy rate, it is not unexpected that the surprise effect of any given variable is only a few basis points. I refer the reader to Carvalho, Hsu and Nechio (2016) for a discussion and estimation of the effect of central bank communication on U.S. yields (using SW’s methodology).

Thus the effects produced by macroeconomic news, though small, are to be expected, and their practical significance should not be disregarded. To bond traders and money market fund managers, a few basis point movement in the wrong direction can wipe out profits. To financial economists, the market response, in and of itself, supports the theory that asset prices are linked to fundamentals (e.g., see Anderson et al., 2003 for a discussion in the context of exchange rates). Perhaps, most importantly, to the average individual that holds their savings in bonds, a few basis points compounded over time can materially affect their wealth.\(^{10}\) This leads me to conclude that the practical significance of the coefficient estimates in Table 1 are not negligible, particularly as they are one-day effects.

A natural question that follows from a discussion on practical significance, is how can we interpret the statistical significance of the coefficients in Table 1? The answer is, with caution. Because the regressors are likely correlated, the standard errors (SES) may be larger than desired. Specifically, high correlation among the regressors (multicollinearity) makes it difficult for any estimator to separate out the individual effects of the explanatory variables so that SES are high, which results in small t-statistics. As such, regressors that appear to be statistically insignificant may in fact be economically relevant. This likely problem of multicollinearity suggests that it is perhaps prudent to focus on the statistical significance of

\(^{10}\) Consider a simple example in which an investor holds $100,000 in bonds earning a 5% annual return, ignoring reinvestment of coupons and principal over time. At the end of 25 years the investment grows to $338,635. A small, 10bps decline in the return, i.e., reconsidering the example at a 4.9% rate, would yield $330,664. Though not drastic, the example illustrates that small movements can have material effects.
all the variables as a group, rather than individually. On the other hand, the high frequency of the daily data creates a rather conducive environment to generate statistical significance. Looking at the regressors together in a joint test of relevance, one thing is clear: the regressions are highly significant. Reported at the bottom of Table 1, the p-values for corresponding F-statistics are exceptionally small.

iii. Estimating the Time-Varying Sensitivity of Yields to Macroeconomic News

Although of interest, the focal point of this research is not the estimated effects of news on yields per se, but rather the sensitivity of those yields to news over time. Specifically, the interest is in how the yields respond to macroeconomic news around the time that the ZLB was binding in Canada (i.e., April 2009 to May 2010). Following SW’s methodology, I estimate the time-varying sensitivity of yields to macroeconomic news by estimating a scaling parameter that captures this degree of sensitivity. The model of interest for yield \( m \), of which all coefficients and the error term are maturity-specific, is:

\[
\Delta y^m_t = \sum_{i=2001}^{2016} \gamma_i^m D_i + (7 - \sum_{i=2002}^{2007} \delta_i^m) D_{2001} \beta^m X_t + \sum_{i=2002}^{2016} \delta_i^m D_i \beta^m X_t + \epsilon^m_t; \\
\epsilon^m_t \sim (0, \sigma_{\epsilon_t^m}^2)
\]

where \( \delta_i^m \) is a scalar that amplifies or dampens \( \beta^m \), the coefficient vector for yield \( m \). As explained in the subsequent paragraph, the relative magnitude of the elements in \( \beta^m \) are assumed to be constant over time. Only the impact of the macroeconomic variables as a group is allowed to vary. Given \( \beta^m \), the implication is that we only need to estimate the single scalar parameter, \( \delta_i^m \) for each year, \( i = 2001, 2002, \ldots, 2016 \), which eliminates the small-sample problems related to alternative specifications (i.e., allowing every element of \( \beta^m \) to vary in each year). The vector \( X_t \) is a column vector of macroeconomic news.
variables, described in the data section above, which includes exactly the same set of variables used in model (7). Both $y^m_i$, the intercept, and $\delta^m_i$, the time-varying sensitivity parameters that measure the impact of macroeconomic news on the change in yield for maturity $m$, are given the freedom to vary in each year, reflected by the dummy variable $D_i$ – one of the key differences between my specification and Moessner’s (2014). The dummy variable takes on one for any given year $i$, and simultaneously zero for all other years $j \neq i$.

There is one exception, however. In order to identify both $\beta^m_i$ and $\delta^m_i$, given the nonlinear nature of the model, a normalization is required. Following SW, I elect to normalize on one $\delta^m_i$, specifically for 2001 (the first year of the sample). As such, I identify $\beta^m_i$ in 2001\textsuperscript{11} and then hold the relative magnitude of the coefficients constant over time. The normalization chosen sets the average of the sensitivity parameters over the normal period equal to one, which can be seen in equation (2) as $(7 - \sum_{i=2002}^{2007} \delta^m_i)$ – effectively defining $\delta^m_{2001}$.

Therefore, a value of $\delta^m_i$ equal to one suggests the yield remains as responsive to news as “normal.” SW refer to such a yield as being completely “unconstrained.” Intuitively, a $\delta^m_i$ of zero suggests that the yield is completely unresponsive to news during the year, and hence is referred to as being completely constrained. For all $\delta^m_i$ between zero and one, the yield is interpreted as being partially constrained. Lastly, a $\delta^m_i$ greater than one suggests the yield was more sensitive to news than normal, which I refer to as hyper-sensitivity.

The rationale behind holding $\beta^m_i$ constant throughout time is supported by SW’s work, both through their empirical work and theoretical motivation, which I reinforce from my own analysis. They argue that the ZLB should dampen the effect of all news in an

\textsuperscript{11} Despite identifying $\beta^m_i$ in 2001, the nonlinear least squares regression employs the entire sample throughout the iterative estimation procedure. Matlab estimates nonlinear equations using the Levenberg-Marquardt algorithm, which minimizes the sum of the squared residuals by pushing estimates in the direction of the gradient.
indiscriminate manner, as their illustrative model used to motive their empirical work views news homogenously. The only assumption they make, is that all news – viewed as exogenous output and inflation shocks in their New Keynesian framework – propagates through the economy with the same persistence. Under this assumption, shocks from all sources can be scaled through a single parameter that augments $\beta^m$, i.e., $\delta^m_t$, which proportionally dampens (or amplifies) the effect of all news. In their words (2014, p.3164), “…if a Treasury security’s sensitivity to news is reduced because its yield is starting to bump up against the zero bound, then we expect that security’s responsiveness to all macroeconomic data releases to be damped by a roughly proportionate amount.” As mentioned above, the similarity of my OLS coefficient estimates to those estimated by SW and Moessner over different sample windows perhaps provides further support for this being a reasonable assumption. The same logic applies to the coefficient estimates from the nonlinear least squares regressions for the 3-month, 2-year and 10-year yields that appear in Table 2 below. As evident in the number of observations reported at the bottom of Table 2, the regressions use the entire sample from 2001 to 2016 to estimate $\beta^m$; once again producing similar results for a different window. More formally, I use the method of moments $J$-test of overidentifying restrictions,\footnote{I refer the reader to Swanson and Williams (2014, p.3167; footnote 14) for a detailed description of the test. Following their approach, I employ a continuous-updating method of moments estimator proposed by Hansen, Heaton and Yaron (1996).} proposed by SW in this context, to test the null that $\beta^m$ is constant over time, against the alternative hypothesis that the elements of $\beta^m$ vary individually. The p-values associated with the J-statistics – reported at the bottom of Table 2 – are all greater than 0.50 giving more formal support to the assumption that the relative magnitude of the elements in $\beta^m$ are constant over time.
whether the time constant and equal to one in all years. The coefficient estimates are interpreted as basis point per standard deviation responses to macroeconomic news. Heteroskedastic consistent t-statistics are in parenthesis. The hypothesis tests are as follows: $H_0: \beta$ constant tests the null that the coefficients in $\beta$ are constant over time; $H_0: \delta$ symmetric tests whether $\delta_{i}^{m}$ is the same for both positive and negative news; $H_0: \delta$ constant tests whether $\delta_{i}^{m}$ is constant and equal to one in all years. See the text for details.

To further test the validity of the restrictions inherent to equation (8), I also test whether the time-varying sensitivity parameters should be allowed to take on different values.

<table>
<thead>
<tr>
<th>Country</th>
<th>Government of Canada Treasury Bill/Bonds Yields</th>
<th>3-Month</th>
<th>2-Year</th>
<th>10-Year</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Canada</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Capacity Utilization</td>
<td>-1.42 (2.34)</td>
<td>1.17 (1.89)</td>
<td>0.75 (1.35)</td>
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<td>CPI</td>
<td>1.63 (3.45)</td>
<td>2.43 (5.05)</td>
<td>0.72 (1.98)</td>
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<td>GDP</td>
<td>1.36 (2.30)</td>
<td>2.37 (4.47)</td>
<td>0.97 (2.07)</td>
<td></td>
</tr>
<tr>
<td>PPI</td>
<td>-1.09 (-1.36)</td>
<td>0.20 (0.37)</td>
<td>-0.18 (-0.53)</td>
<td></td>
</tr>
<tr>
<td>Net Change in Employment</td>
<td>0.91 (1.41)</td>
<td>2.97 (3.50)</td>
<td>1.56 (2.96)</td>
<td></td>
</tr>
<tr>
<td>Unemployment Rate</td>
<td>-1.18 (-1.51)</td>
<td>-2.24 (-2.63)</td>
<td>-0.65 (-1.35)</td>
<td></td>
</tr>
<tr>
<td>Ivey Index</td>
<td>-0.52 (-0.54)</td>
<td>0.09 (0.15)</td>
<td>-0.13 (-0.32)</td>
<td></td>
</tr>
<tr>
<td>Survey of Manufacturers</td>
<td>-0.32 (-1.13)</td>
<td>0.33 (0.98)</td>
<td>-0.01 (-0.05)</td>
<td></td>
</tr>
<tr>
<td>Raw Material Prices</td>
<td>-0.97 (-0.97)</td>
<td>-0.93 (-1.30)</td>
<td>-0.68 (-1.32)</td>
<td></td>
</tr>
<tr>
<td>Trade Balance</td>
<td>1.11 (1.09)</td>
<td>0.59 (1.35)</td>
<td>0.37 (1.28)</td>
<td></td>
</tr>
<tr>
<td>Housing Starts</td>
<td>-0.41 (-0.75)</td>
<td>-0.75 (-1.47)</td>
<td>-0.38 (-1.19)</td>
<td></td>
</tr>
<tr>
<td>Current Account Balance</td>
<td>-0.41 (-0.72)</td>
<td>0.71 (1.06)</td>
<td>0.63 (1.18)</td>
<td></td>
</tr>
</tbody>
</table>

| **United States** |                                               |              |              |              |
| Capacity Utilization | 0.29 (0.38) | 2.38 (1.91) | 0.96 (1.60)  |              |
| Core CPI           | -0.55 (-0.84) | 0.64 (1.38) | 0.29 (0.82)  |              |
| GDP               | 0.28 (0.26) | 0.96 (1.14) | 0.64 (1.20)  |              |
| Core PPI           | 0.33 (0.64) | 0.82 (1.77) | 0.86 (2.77)  |              |
| Nonfarm Payrolls  | 1.50 (2.05) | 4.63 (8.27) | 3.42 (7.37)  |              |
| Unemployment Rate | -0.86 (-2.29) | -1.37 (-3.04) | -0.82 (-2.80) |              |
| ISM Manufacturing  | 1.91 (3.57) | 2.27 (4.32) | 1.94 (5.10)  |              |
| Consumer Confidence | 0.50 (1.01) | 0.92 (1.52) | 0.51 (1.16)  |              |
| Retail Sales ex. autos | 1.43 (3.75) | 2.43 (4.91) | 2.06 (5.28)  |              |
| Initial Claims    | -0.58 (-2.13) | -1.21 (-5.01) | -0.84 (-4.64) |              |
| Leading Indicators | 1.40 (1.49) | -0.04 (-0.09) | 0.10 (0.32)  |              |
| House Sales       | -0.13 (-0.28) | 0.07 (0.16) | 0.21 (0.65)  |              |
| Industrial Production | 0.98 (1.42) | -0.29 (-0.37) | -0.15 (-0.30) |              |

| Observations      | 2,594 | 2,594 | 2,594 |
| $H_0: \beta$ constant, p-value | 0.68 | 0.51 | 0.80 |
| $H_0: \delta$ symmetric, p-value | 0.08 | 0.16 | 0.72 |
| $H_0: \delta$ constant, p-value | 0.00 | 0.00 | 0.00 |

**Notes:** The coefficient estimates are interpreted as basis point per standard deviation responses to macroeconomic news. See the text for details.

<table>
<thead>
<tr>
<th>Country</th>
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<tbody>
<tr>
<td><strong>Table 2</strong></td>
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<td></td>
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</tr>
</tbody>
</table>
| $\Delta y_t^m = \Sigma_{i=2001}^{2007} y_t^m \Delta_i + (7 - \Sigma_{i=2002}^{2007} \delta_{i}^{m}) \Delta_{2001} y_t^m \delta_{i}^m X_t + \Sigma_{i=2002}^{2016} \delta_{i}^{m} \Delta_i \beta_t^m X_t + \epsilon_t^m$ over the entire sample window 2001 to 2016.
for positive and negative surprises, that is whether $\delta^m_i$ is symmetric. To do this, the data is separated into negative and positive surprises, and equation (8) is estimated allowing for a different $\delta^m_i$ for both positive and negative surprises in each given year. The null of the hypothesis is that the two parameters in each year are equal; the alternative is that they are not. The p-values reported at the bottom of Table 2 suggest slightly mixed results, with increasing statistical evidence at longer maturities.

The last hypothesis test conducted on equation (8) is that the time-varying sensitivity parameters are equal across time, being one given the normalization. If they were equal, then this would suggest that yields have remained consistently sensitive (at a normal level) to macroeconomic news over the period 2001 to 2016. As expected, the data strongly rejects the null of $\delta^m_i$ being constant, with estimates showing that in fact these sensitivity effects vary quite considerable from year to year (I explore this further below). These findings suggest that Moessner’s model, which restricts $\delta^m_i$ to take on only two values over his entire 16-year sample, is likely too restrictive.

Overall, the three hypotheses tests yield similar results to those reported by Swanson and Williams (2014) for U.S. yields. I now follow their approach further by ascertaining more granular estimates of the time-varying sensitivity parameter $\delta^m_i$ in pursuit of exploring its stability properties in a less restricted manner.

iv. Rolling Regressions to Explore the Stability of the Sensitivity Parameter $\delta^m$

To further explore the stability of $\delta^m$ over time, I complete the last step in Swanson and Williams (2014a) empirical framework, which is to estimate a rolling regression that generates more granular estimates of $\delta^m$. To distinguish between $\delta^m$ in model (8) and the same parameter in this model, model (9), I denote it as $\delta^m_r$. The rationale for distinguishing
between them, is that in model (8) the time-varying sensitivity parameter is restricted to take on only one value in each given year \( i = 2001, 2002, \ldots, 2016 \). In model (9), the regression rolls over 262-day windows – approximately the number of business days in any given year – at 10 day increments resulting in 398 estimates of the time-varying sensitivity parameter, i.e. \( \tau = 1, 2, \ldots, 398 \). The model takes the form:

\[
\Delta y_t^m = \gamma_t^m + \delta_t^m \sum_{j=1}^{n} (\hat{\beta}_j^m x_{j,t}) + \varepsilon_t^m ; \quad \varepsilon_t^m \sim (0, \sigma_{\varepsilon_t^m}^2)
\]

where \( \hat{\beta}_j^m x_{j,t} \) is a generated regressor (discussed below) created from the nonlinear least squares estimates of \( \beta^m \) in equation (8) multiplied by the same regressors \( x_{j,t} \) that appear in both models (7) and (8). The parameter \( \gamma_t^m \), the intercept term, along with \( \delta_t^m \), the time-varying sensitivity parameter – both for yield \( m \) – are given the freedom to vary every 10 days. As noted previously, the data suggests that forcing them to be constant, even if only for a year, is likely too restrictive. I come back to this question in the next section. The assumptions on \( \varepsilon_t^m \) are not changed: it is mean-zero and independent, but not identically distributed due to unconditional heteroskedasticity.

The use of generated regressors creates an additional level of uncertainty in model (9) that necessitates a manual correction to the standard errors (see, for example, Oxley and McAleer, 1993, for a formal overview of the issues caused by generated regressors). Before touching on the correction, it may be useful to explain the issue at hand. In theory, there should be one regressor in model (9): an index that encapsulates macroeconomic news. This index is unobservable, however, and therefore to estimate the model, a proxy must take its place. Following SW, the chosen proxy is an index built using the estimated effects of news from model (8). It simply sums up all of the predictions of macroeconomic news on any day.
\( t \), weighted by their estimated impact. The index, being the sum of coefficient estimates multiplied by exogenous regressors, is consequently subject to measurement error, given it is a proxy, and sampling error due to the use of the sample estimates from model (8). The implication is that the index – a generated regressor – is stochastic, which is not accounted for in the standard errors of the ordinary least squares regression of model (9). Not accounting for this additional randomness leaves the standard errors understated, hence the required correction. Several correction methods have been proposed (see, for example, Gawande, 1997), however, I opt to follow SW in their approach, which scales up the variance estimates by comparing the standard errors from model (8) with those from model (9). See SW (2014, p.3165; footnote 12) for detail.

Intuitively, the aim of the model is to ascertain how \( \Delta y^m_t \) responds to the proxy index of weighted macroeconomic news. If it responds as exactly predicted over a ten-day period, \( \delta^m_t \) would take on the value of one and the yield is referred to as completely unconstrained. If it does not respond at all, \( \delta^m_t \) takes on the value of zero and the yield is referred to as being completely constrained. For \( \delta^m_t \) between zero and one, it is referred to as being partially constrained. Lastly, for \( \delta^m_t \) greater than one, I refer to it as being hyper-sensitive. In the interpretation of \( \delta^m_t \), model (9) exactly parallels model (8). I now turn to the estimates of \( \delta^m_t \) which are the focal points of this research.

4  Main Results: The Stability of the Sensitivity Parameters \( \delta^m \)

Figure 2 plots the time-varying sensitivity parameter, \( \delta^m_t \), from equation (9), over the period 2008 to 2016 for each yield in descending order from the 3-month to 30-year. The
dotted lines surrounding the point estimates are plus- and minus-two standard-error bands generated from estimating equation (9) and adjusted for two-stage sampling uncertainty, using the modification adopted by SW. Most importantly, periods in which $\delta_{T}^{m}$ is not statistically different from zero are shaded red, and periods in which it is significantly less than one are shaded yellow. These shaded regions, respectively, highlight when the yields are identified as being completely and partially constrained.

The most notable feature of both the 3- and 6-month yield plots of $\delta_{T}^{m}$ is the amount of red in the graphs. Over almost the entire nine-year period both yields remained completely unresponsive to macroeconomic news. The only period not shaded red or yellow, is between the start of 2008 and early 2009, which is when the ZLB was not yet binding in Canada. In April 2009, however, just around the time that the Bank cut the overnight rate to 25bps, the sensitivity of the yields plummet – seen most clearly in the 3-month plot. From that point on, the yields remain constrained. Even when the Bank pushed the overnight rate back up to 50bps in May 2010, the yields continued to remain unresponsive to macroeconomics news. Before discussing the possible reasons for this behaviour, which is only partially consistent with the theory, it is enlightening to examine the plots for the 10- and 30-year yields.

The 10- and 30-year plots tell a different story than with the shorter yields. With the exception of a very brief period of being constrained in 2009, both the 10- and 30-year yields remain as sensitive to news as normal throughout the nine-year period. This stark contrast to the 3- and 6-month plots is consistent with the theory. Long-term yields are the average of current and expected future short-rates, and hence with the expectation of the zero lower bound only binding for a short period, longer-term yields would be expected to remain as sensitive to news as normal. The shorter the maturity, the more affected the yield. Filling in
the gap between the short-end and long-end of the yield curve, the 1-year, 2-year and 5-year plots of the estimated $\delta_i^m$ coefficients illustrate this dampening effect.

Starting with the 1-year plot, the estimated $\delta_i^m$'s suggest evidence of the yield being partially and completely constrained; however, the red-shaded periods take up substantially less of the graph than compared to the 3- and 6-month plots. The 2-year yield shows even less evidence of being constrained, with only minimal amounts of red, and periods following the ZLB in which the yield responds to news as normal. Turning to the 5-year yield, the pattern seems to become evident with a gradual weakening of the constraint as one moves from the short- to long-end of the yield curve. In this regard, this is exactly what the expectations hypothesis would predict. But, a question remains: if the zero lower bound was only binding between April 2009 and May 2010, why are the yields of any maturity exhibiting constrained behaviour post-May 2010? Moreover, with the Bank of Canada announcing in December 2015 that the country’s effective lower bound (ELB) rests somewhere around –50bps, why is the short-end of the yield curve – 100bps above the ELB – still constrained?

It is important to caution that the zero lower bound is not the only possible reason to explain the constrained behaviour of yields evident in Figure 2. Another cause, with a slightly simpler explanation, is that the “constrained” behaviour is really just the result of lower volatility in the yields. For some time, research has advocated that the volatility of interest rates falls with the level of the interest rate (for example, see, Chan et al., 1992 or Kim & Singleton, 2012). Therefore, feasibly the dampened sensitivity to macroeconomic news is capturing the fact that very low interest rates have little room to fluctuate.

Conceivably, this explanation is consistent with the data rather than the ZLB justification. As the Bank reached the zero lower bound in April 2009, this period marked a point in time in which conventional monetary policy became relatively ineffective, as well as a point in time in which interest rates reached their lowest historical levels. When the zero lower bound was
no longer binding in May 2010, interest rates continued to remain at historically low levels with minimal volatility, and despite creeping up ever so slightly over the past six years, they remain historically low (see Figure 1).

The gradual weakening of the constrained behaviour of yields to macroeconomic news as one moves from the short- to long-end of the yield curve is consistent with this volatility premise, as the level of yields increase with maturity (see Figure 1), allowing more room for yields to fluctuate. The very low yields on the 3- and 6-month treasuries over the past nine years have correspondingly lower volatility, with consequently dampened sensitivity to macroeconomic news over the nine-year period. Given how low the yields have been, complete insensitivity to news, which is what the plots exhibit, is not surprising. Breaking the data into pre- and post-zero lower bound periods, and measuring the volatility of yields using their historical standard deviation, highlights this decline in volatility due to the yields lower levels. For example, the 3-month yield, over the period October 2001 to April 2009, has a standard deviation of 93bps. For the equal-sized period following this (i.e., May 2010 to November 2016), the standard deviation is only 28bps – down by a factor of three. Applying the same analysis to the 6-month, 1-year and 2-year yields reveals a similar disparity. However, for yields longer than 2-years, the standard deviation across these two subsamples is strikingly similar. The standard deviation on the 5-year yield actually increases by 2bps post-April 2009 (66bps and 68bps in each respective period).

Swanson and Williams (2014a) acknowledge the volatility phenomenon as a possible explanation for their results, which are qualitatively similar to those that I present here. The sample window they use in their analysis, however, does not extend beyond the zero lower bound period, which, for the United States did not end until December 2016. As a result, SW cannot explore how the sensitivity of yields to macroeconomic news changes in the aftermath of the ZLB. Consequently, I suggest that their results provide little convincing
Figure 2: Estimated $\delta_t^m$'s from Rolling Regression $\Delta y_t^m = y_t^m + \delta_t^m \sum_{j=1}^m (\hat{\beta}_j x_{j,t}) + \epsilon_t^m$ over 2008 to 2016

A. 3-Month Yield – Time Varying Sensitivity to News

B. 6-Month Yield – Time Varying Sensitivity to News
C. 1-Year Yield – Time Varying Sensitivity to News

D. 2-Year Yield – Time Varying Sensitivity to News
E. 5-Year Yield – Time Varying Sensitivity to News

F. 10-Year Yield – Time Varying Sensitivity to News
evidence in favor of the reason behind the lack of sensitivity to macroeconomic news; i.e.,
whether it be the zero lower bound and/or lower volatility that led to the change in the
sensitivity of yields to macroeconomic news. In my analysis, however, given the data I have
available, I hypothesize that the empirical evidence in support of a particular constraint is
perhaps slightly clearer. My conjecture is that the zero lower bound camouflages the
constraint on yields caused by their exceptionally low levels and lower volatility; only, when
the ZLB is regarded as no longer binding, with yields remaining low, do the results suggest
that it is changes in volatility that is the likely cause of the constrained behaviour.

In addition, given that the Bank of Canada announced in December 2015 that the
effective lower bound in Canada is –50bps – no longer 25bps – and yields continue to
exhibit the same behaviour in terms of response to macroeconomic news is conceivably
further evidence that the zero, or effective, lower bound is probably not the cause of the constrained responses. If the key were the zero lower bound, then knowledge of an additional 75bps buffer above the ELB would likely have improved the sensitivity of yields to news. However, the estimates reported in the plots provided in Figure 2 indicate that the Bank’s announcement did not fundamentally alter the sensitivity of yields to news.

Despite the lower volatility explanation perhaps providing a better description of the observed results, two points must be noted. First, there is no definitive way of determining the true cause of the constraint. Moreover, there could be more than one cause. The empirical work and hypotheses tests that I conduct in this research aims to detect any period in which yields are insensitive to news (partially or otherwise). The methodology does not enable me to differentiate between the ZLB, lower volatility or any other possible source, or some combination of more than one basis for the lack of sensitivity to news. Second, the lower volatility explanation fails to explain the constrained behaviour evident in the 10-year and 30-year yields (the last two plots reported in Figure 2). Both yields were hovering around four percent at the time of the Financial Crisis (see Figure 1), which is when my analysis determines them to be completely constrained. Given that four percent is only marginally below that of the yields’ pre-crisis levels, a sudden decline in volatility would not be consistent with the explanation. As noted above, there could be multiple reasons for a yield to lack sensitivity to macroeconomics news and perhaps this period in the data is evidence of just that possibility.

It could even be that the Bank of Canada itself was partly responsible for the constrained behaviour. By issuing forward guidance as an extraordinary measure (Carney, 2012) at the same time as setting the overnight rate on the lower bound in 2009, perhaps markets were unsure of what to make of the unprecedented monetary stimulus. Moreover,
markets may have been unsure of what other tools the Bank was willing to employ. After all, the United States was already in its first round of quantitative easing at the time (Charbonneau & Rennison, 2015). This additional uncertainty, on top of the great deal of unrest caused by the Financial Crisis, may have led the market to take a “wait-and-see” approach for a brief period, unsure of the central bank’s possible reactions to macroeconomic indicators. Though a conceivable rationale for the observed outcomes, these two particular cases of constrained behaviour (i.e., on the 10- and 30-year yields) remain to be empirically explained.

With the stability of the time-varying sensitivity parameters now presented, I return to the main question of interest: with the zero lower bound immobilizing the Bank of Canada’s key policy tool from April 2009 to May 2010, how effective could have been the Bank’s unconventional monetary policy? Following SW (2014a) and Moessner (2014), I use the estimated time-varying sensitivity coefficients as a proxy of the Bank’s ability to effectively conduct unconventional monetary policy. With this paradigm, the degree to which a yield is constrained reflects the degree to which the Bank can influence that yield (through its arsenal of unconventional tools). A completely constrained yield is, hence, considered to be an ineffectual target for monetary policy, as it does not respond to manipulation; in the same fashion, a completely unconstrained yield is judged to produce, if targeted, the same effect as monetary policy would have had in a “normal” period (recall, I define the normal period as 2001 to 2007). Intuitively, partially constrained yields fall somewhere between the two extremes of being ineffectual targets and producing “normal” effects. On this basis, the estimates reported in Figure 2 provide insight into the efficacy of monetary policy in the face of the zero lower bound, or perhaps more generally, in the face of very low interest rates. However, I acknowledge that there are several limitations of such
interpretations if the true cause of the constraint(s) is neither the ZLB nor lower volatility. Without knowing the underlying reason(s) for sensitivity or insensitivity, the implications for monetary policy are difficult to characterize. To continue with the analysis in a tractable manner, however, I limit the deemed grounds of any constrained reaction to news to be either lower volatility or the ZLB (with the exception of during the brief period that seems unexplained by either reason). My rationale is that both of these possible causes carry similar implications for monetary policy, given that they originate from the same source – exceptionally low interest rates.

My empirical results suggest that the episode in which the zero lower bound was binding in Canada can be roughly separated into two distinct intervals: (1) the first four months (April to July, 2009) and (2) the last 10 months (August 2009 to May 2010). In the first four months of the impact of the zero lower bound, no yield of any maturity seems to have retained its normal effectiveness. This can be seen in Figure 2 as the initial red- and yellow-shaded periods that appear in every plot from the 3-month yield to the 30-year yield. Without a definitive explanation as to the reason behind the constrained response of yields to news over this four-month interval, it would not be reasonable to infer the implications for the effectiveness of unconventional monetary policy. Simply surmising that the Bank’s ability to influence the economy was severely limited at that time, from the observed feature that all yields seem completely or partially unresponsive to news, does not account for the possibility that the constrained sensitivities to news may have been able to be altered through some Bank action. Unlike the ZLB or lower volatility justifications (both of which cannot be resolved without the central bank methodically choosing to increase its overnight rate), a constraint of some other cause, perhaps endogenously arising from the central bank’s policy, may have had a more straightforward resolution. In any case, the implications for monetary
policy at the time are unclear. Over the last 10 months in which the ZLB was binding, however, my analysis seems more fruitful.

From August 2009 to May 2010, the sensitivity of all yields to macroeconomic news on bonds with maturities over one year is estimated to be normal. By implication, in so far as the Bank of Canada can influence yields on bonds with maturities greater than one year through unconventional monetary policy tools, the efficacy of such policies would be expected to be normal. In other words, my findings suggest that Canadian policymakers continued to have considerable power over medium- and longer-term interest rates, despite the zero lower bound being binding. This finding is analogous to that reported by SW for the United States. The short-end of the yield curve may have been constrained by the ZLB, but monetary policy – in its broadest sense – was not impotent. Moessner (2014) details similar results as well with his restrictive model. In his outcomes for Canada, the only yield constrained over the entire ZLB period was on the 1-year Government of Canada bond – the shortest maturity considered in his study. Despite the key interest rate stuck on the ZLB, unconventional tools such as quantitative easing and credit easing that target the mid- to long-end of the yield curve could have perhaps been effective tools, at least in so far as they theoretically would be during normal times. I say “theoretically” because such tools have not been adopted by the Bank of Canada during normal times, as there is no need when the overnight rate is free to take on its role as the key policy tool.

The analysis I have undertaken has important implications for policymakers. For instance, using such an empirical framework in real-time could perhaps help guide the central bank when conducting unconventional monetary policy while operating at the zero lower bound, or perhaps more generally, while functioning in a low interest rate environment. By estimating which portion of the yield curve seems constrained by the zero
lower bound or lower volatility, such policies could be specifically targeted toward those yields less constrained, with the greatest potential to transmit the intended stimulus. In another light, it is helpful estimating how much room the Bank had to conduct monetary policy over the last decade. In this regard, my research contributes to a better understanding of the limits of monetary policy in Canada.

5 Robustness Analysis

In this section, I explore how robust my estimates of the time-varying sensitivity parameters \( \delta^m_t \) are to alternative specifications of model (8) and (9). My focus is on varying the set of explanatory variables to give a cursory account of the model uncertainty inherent in my initial specification used for the results reported in the previous sections. As noted in the data part of my thesis, I chose the macroeconomic news variables (the regressors in the models) to be consistent with much of the previous research. These works, however, does not provide robust rationale for why the chosen explanatory variables are employed. It is possible that data limitations regarding the length of time series data on Bloomberg surveys, and the variables surveyed, contributed to the choice of variables used by Swanson and Williams (SW) (2014a) and Moessner (2014). Even if it did not, data limitations did partially shape my set of regressors, as I was unable to obtain three variables considered by SW and Moessner. Given that the true data generating process is unobservable, and that the theory does not exactly clarify the appropriate macroeconomic news variables, it is highly likely – almost surely – that the models I used are misspecified. Whether any such misspecification materially alter my quantitative or qualitative results is an
empirical issue, despite, of course, affecting the theoretical properties of the considered estimators and tests.

As a robustness check to my main results, I estimate several alternative specifications and (in the interest of space) report three\(^{13}\) that may be of economic interest. Namely, I report estimates of \(\delta^m_t\) using (1) a group of Canadian-only macroeconomic news variables, (2) a set of U.S.-only macroeconomic new variables and (3) a collection of seven additional regressors\(^{14}\) in combination with the complete set used in the initial specifications.

Where feasibly comparable, the estimates of the time-varying sensitivity parameters seem to be quite robust to the (approximately 20) alternative specifications of models (8) and (9) that I estimated. This is illustrated in Figure 3, which provides plots of the estimated time-varying sensitivity parameters from model (9), \(\delta^m_t\), for the 3-month yield over the period 2008 to 2016 for the alternative specifications just mentioned. Panel A plots the estimates using the set of Canadian-only regressors, which, in comparison to Panel A in Figure 2, yields qualitatively similar results. The outcomes continue to imply that this yield became completely unresponsive to macroeconomic news around the time that the zero lower bound was considered binding in Canada, with this lack of reaction to news remaining until recently. Panels B and C, provided in Figure 3, give plots, respectively, of the estimated \(\delta^m_t\)'s for the 3-month yield using only the U.S. regressors and the set of additional explanatory variables (in combination with the complete set originally included). These too qualitatively indicate the same constrained behaviour to news. I have only reported the results for the 3-month yield in interest of space, however, the estimates of the sensitivity parameters for the yields on longer maturities appear to be just as robust to the alternative

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\(^{13}\) The results for all of these other alternative specifications not reported here are available upon request.

\(^{14}\) The additional regressors include Canadian building permits and housing prices, and U.S. auto-sales, business inventories, housing starts, personal consumption, and the University of Michigan Consumer Sentiment Index.
specifications. My findings align with the small robustness explorations of both SW and Moessner to alternative specifications. Despite the time-varying sensitivity parameter being the focal point of my research, the robustness of the coefficient estimates are perhaps also of interest.

For the three alternative specifications, the coefficient estimates from model (9) are qualitatively comparable with those obtained using the initial specification, though there are some quantitative differences (these results have not been included in interest of space, but are available upon request). For both specifications that include only Canadian news variables and only U.S. factors, the explanatory variables with the largest (in magnitude) coefficient estimates in the initial specification of model (9) continue to have the largest coefficient estimates in these additional regressions. In certain cases, the estimates are quite similar in magnitude. For example, in the initial specification of model (9), U.S. ISM Manufacturing has a coefficient estimate of 1.91 bps per standard deviation response on the 3-month yield. In the U.S.-only specification, it has a coefficient estimate of 2.04. From a qualitative perspective, the signs on the coefficient estimates in the three additional specifications are identical (as comparable) to those from the initial specification. Though this may provide support for the robustness of my original estimates, this is by no means a comprehensive robustness check.

A more comprehensive approach to capturing the model uncertainty inherent to my initial specifications would include pursuing a model averaging approach. Model averaging, in its many forms, essentially would involve estimating models (8) and (9) using all feasible combinations of regressors and then averaging across the full set of obtained estimates using a criterion of some sort to weight the relevance of each model (e.g., the AIC criterion or a Bayesian approach). For any model, with $x$ number of regressors there is always $2^x$ possible
Figure 3: Estimated $\delta^m_t$’s for the 3-month yield from model (9), $\Delta y^m_t = y^m_t + \delta^m_t \sum_{j=1}^n (\hat{\beta}^m_j x_{j,t}) + \varepsilon^m_t$ that explores three alternative specifications of the set of macroeconomic news regressors

A. Canadian Macroeconomic News Variables Only

B. U.S. Macroeconomic News Variables Only
C. All Core Macroeconomic News Variables plus 7 Additional

Notes: The solid purple line depicts the time-varying sensitivity parameter from equation (9) with \( \pm 2\)-standard-error bands as dotted lines, which use the heteroskedasticity-consistent standard errors that account for two-stage sampling uncertainty (see the text for details). A value of 1 indicates a period in which the respective yield was as sensitive to macroeconomic news as normal. Periods shaded red are periods in which the time-varying sensitivity parameter is not statistically different from zero (completely constrained). Periods shaded yellow are periods in which the time-varying sensitivity parameter is significantly less than 1 (partially constrained). The set of Canadian and U.S. regressors employed in Panel A and B exactly match those used in model (9) (refer to Table 2 for a list of the exact variables). Panel C includes the entire set of macroeconomic news variables employed in model (9) along with seven additional regressors (refer to footnote 14).

Combinations. In my case, using the entire set of 32 regressors I was able to obtain (i.e., including the seven additional regressors), there are 4.2 billion feasible combinations of models to estimate. Even if I were to dichotomize my set of regressors into those that must appear in every model (i.e., the “focus” variables) and those that might be regarded as auxiliary, the number of possible models makes it computationally difficult to execute. Therefore, despite being a reasonable approach to accounting for model uncertainty, the computational difficulties put this methodology outside of the reach of this thesis. I refer the reader to Moral-Benito (2015) for a survey of the model averaging approaches commonly applied in empirical economic research.
There are several other robustness tests that may be of interest aside from varying the set of regressors in my models. SW’s two main robustness tests include (1) replacing the U.S. yields used in their models with Eurodollar futures and (2) employing a more structural specification of the time-varying sensitivity parameter, which controls for monetary policy uncertainty (see SW, 2014a, p. 3177). Their choice of the first robustness test is rooted in the feature that Eurodollar futures provide an alternative means of proxying market expectations on future short rates, giving them another channel to test for the constrained behaviour of yields. SW find that their results are robust in this direction. Their second robustness check considers the possibility that the time-varying sensitivity of yields to macroeconomic news is explained by changes in monetary policy uncertainty. Interestingly, in this case, they detect some lack of robustness in that outcomes are partially explained by changes in monetary policy uncertainty over time. This is likely evidence to support my proposition that there are several reasons (aside from the ZLB) that may alter the sensitivity of yields to macroeconomics news. If appropriate data had been available, I would have been interested to pursue similar robustness checks.

6 Conclusion

In this thesis, I use a novel empirical framework developed by Swanson and Williams (2014a) to estimate the potential effectiveness of unconventional monetary policy while the zero lower bound was considered binding in Canada. To do this, I estimate how the sensitivity of the term structure of interest rates to domestic and U.S. macroeconomic news has evolved over the period 2001 to 2016, with a focus on the period April 2009 to May 2010 – when the ZLB was binding in Canada. Following the literature, I use the estimated
sensitivities of yields to news as a proxy for the Bank of Canada’s ability to effectively conduct unconventional monetary policy.

My empirical results suggest that the period in which the ZLB was considered binding can be roughly separated into two distinct intervals: (1) the first four months (April 2009 to July 2009) and (2) the last 10 months (August 2009 to May 2010). Despite being able to provide only limited insight into the first four-month interval given its incongruence with the possible explanations I examine, the outcomes of my work suggests that the last 10 months could have been conducive to effective unconventional monetary policy, so long as it targeted yields with a year or more to maturity.

I conjecture, however, that the root cause of the lack of sensitivity of yields to macroeconomic news is likely not the zero lower bound, but rather reduced (compared with normal) volatility associated with exceptionally low interest rates. Whether the reason lies with the impact of the zero lower bound or the observed lower volatility, or perhaps both, the policy insights from my work remain the same. My results showed that the Bank of Canada retained control over the medium- to long-end of the yield curve, to the extent that their unconventional tools can influence such yields, when the key policy rate became an ineffective tool for conducting monetary policy. My recommendation from my research is that an empirical framework of similar sorts to that employed in this paper – used in real-time – could perhaps help guide the central bank in conducting unconventional monetary policy, while operating at the zero lower bound, and, more generally, in an environment of low interest rates that exhibit less than normal volatility.
References


http://www.bankofcanada.ca/2012/12/guidance/


Appendix

Information about the macroeconomic news variables and yields considered in this thesis, including their Bloomberg Ticker IDs (in square brackets), are given below.

Canadian Macroeconomic Variables

1. GDP, QoQ%, Quarterly [CGE9ANN]
2. CPI, MoM%, Monthly [CACPICHG]
3. Raw Material Price Index, MoM%, Monthly [CARAMOM]
4. Industrial Product Price Index (PPI), MoM%, Monthly [CAIPMOM]
5. Unemployment Rate, % of Labour Force, Monthly [CANELXEMR]
7. Total Housing Starts, ‘000s of houses, Monthly [CAHSTOTL]
8. Current Account, CAD$ Billions, Quarterly [CACURENT]
9. Capacity Utilization, % of Total Capacity, Quarterly [CACAPUTL]
10. Merchandise Trade Balance, CAD Billions, Monthly [CATBTOB]
11. Survey of Manufacturing Sales, MoM%, Monthly [CAMFCHNG]
12. Leading Indicators, MoM%, Monthly [CACOICHG]
13. Wholesale Trade, MoM%, Monthly [CAWTMOM]
14. New Housing Prices, Mom%, Monthly [CAHUMOM]
15. Ivey PMI Index, % Balance/Diffusion Index, Monthly [IVEY]
16. Building Permits, MoM%, Monthly [CAHOMOM]

US Macroeconomic Variables

1. Non-Farm Payroll, ‘000s of workers, Monthly [NFP TCH]
2. Core CPI, MoM%, Monthly [CPUPXCHG]
3. GDP, QoQ%, Quarterly [GDP CQOQ]
4. Core PPI, MoM%, Monthly [PXFECHNG]
5. Capacity Utilization, % of Total Capacity, Monthly [CPTICHNG]
6. Conference Board Consumer Confidence, Index, Monthly [CONCCONF]
7. Unemployment Rate, % of Total Labour Force, Monthly [USURTOT]
8. NAPM/ISM PMI, Index, Monthly [NAPMPMI]
9. Initial Jobless Claims, ‘000s of workers, Weekly [INJCJC]
10. Retail Sales Less Autos, MoM%, Monthly [RSTAXMOM]
11. New One Family Houses Sold, ‘000s of houses, Monthly [NHSLTOT]
12. New Privately Owned Housing Units Starts, ‘000s of houses, Monthly [NHSPSTOT]
13. Personal Consumption, QoQ%, Quarterly [GDPCTOT%]
14. Auto Sales, Millions Newly Registered, Monthly [SAARTOTL]
15. Leading Indicators Index, MoM%, Monthly [LEI CHNG]
16. Industrial Production, MoM%, Monthly [IP CHNG]
17. Business Inventories, MoM%, Monthly [MTIBCHNG]
18. University of Michigan Consumer Sentiment Index, Index, Monthly [CONSSENT]
19. Current Account, USD Billions, Quarterly [USCABAL]

Government of Canada Yields

1. 3-Month Treasury Bill [C1033M]
2. 6-Month Treasury Bill [C1036M]
3. 1-Year Government of Canada Bond [C1031Y]
4. 2-Year Government of Canada Bond [GCAN2YR]
5. 5-Year Government of Canada Bond [GCAN5YR]
6. 10-Year Government of Canada Bond [GCAN10YR]
7. 30-Year Government of Canada Bond [GCAN30YR]