Restricting foreign investment:
Estimating the costs of Canada’s foreign property rule*

Kevin Milligan†
and
Michael Smart‡

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Abstract
International portfolio investment of Canadian pension funds is restricted to 20 per cent of the value of assets. This policy increases non-diversifiable risk in portfolio returns, which entails costs for pension plan beneficiaries. We provide new estimates of the cost of this policy by comparing performance of mutual funds that are constrained by the limit and those that are not. We further examine the effects of an increase in the limit, from 10 to 20 per cent, during the early 1990s. The evidence suggests the policy has had a significant negative effect on risk-adjusted returns to retirement savings in Canada. Our estimates imply that a movement in the limit from 20 per cent to 30 percent would increase average annual returns by 0.31 per cent, which corresponds to an increase of 9.9 per cent in terminal portfolio wealth over a 20-year period.

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†Department of Economics, University of Toronto. Corresponding author: 140 St. George St., Toronto ON M5S 3G6 Canada, e-mail: kmilliga@chass.utoronto.ca
‡Institute for Policy Analysis, University of Toronto
1 Introduction

The book value of assets in tax-exempt pension plans in Canada stood at C$822 billion in 1997 — 31 per cent of household net worth. At present, no more than 20 per cent of these assets may be held in foreign investments. These restrictions on international portfolio investment constrain diversification opportunities, which imposes costs on pension plan beneficiaries. We present new estimates of the cost of the Canadian limit using data on mutual fund performance. Our estimates imply that a movement in the limit from 20 per cent to 30 percent would increase average annual returns by 0.31 per cent, which corresponds to an increase of 9.9 per cent in terminal portfolio wealth over a 20-year period.

Many previous researchers have measured the cost of international capital controls simply by comparing returns to two portfolios—one purely domestic and the other internationally diversified. The cost of the restriction is estimated by some measure of distance between the two portfolio frontiers. For example, De Santis and Gerard (1997) use reference portfolios to estimate the expected gains from international diversification for a US investor. In Canada, Ambachtsheer (1995) calculates the cost of the foreign property rule to be 0.2 percentage points per year by comparing the returns of a constructed Canadian portfolio to those from a portfolio investing in internationally diversified indexes.

We adopt a different approach to estimating the cost of this policy. A subset of publicly traded mutual funds in Canada are subject to the 20 per cent foreign property rule, whereas others are not. We exploit this feature of the regulations by comparing the returns to the two classes of funds. Furthermore, because the allowable foreign content limit increased through our sample period, we observe the returns of funds subject to the regulation at different values of the limit. Thus, we can use the unconstrained funds as a control group for the constrained funds. The exogenous policy change and the presence of a control group fortify the credibility of our inferences by aiding the identification of the policy’s effect.

Our approach offers several advantages over existing practice, which generally uses historical returns to national market indexes to estimate the cost of the policy. In so doing, the researcher implicitly assumes that observed returns to domestic and foreign securities reflect equilibrium expected returns. In particular, previous estimates of the costs of the limit have assumed that the poor performance of Canadian stocks in recent years will persist. If the subjective expectations of returns held by domestic investors differ from recent historical means, then this portfolio approach will lead to bias. In contrast, our estimation technique incorporates explicit controls for the return to domestic stocks during the sample period, allowing for differences in systematic risk of portfolios in the sample. We believe estimates based on this approach are likely to be superior, since they do not confuse short-run differences in returns with long-run effects of the policy.

Moreover, the standard approach assumes that investors hold diversified portfolios of national market indexes. But evidence suggests that investors holding assets not subject to foreign property restrictions still have a strong preference for domestic assets. In the presence of this “home bias”, investors will not hold idealized reference portfolios, which makes it difficult to predict their portfolio choices under alternative foreign property regimes. As well, investors typically do not invest passively in national market indexes, but instead follow a more active investment strategy. Assuming that they do invest passively will bias estimates of the gains to diversification if the return to passive versus active investment strategies differs in international and domestic markets.

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1 Indeed, the portfolio allocations of domestic investors suggest a considerable divergence of beliefs from historical returns. See French and Poterba (1991) for a discussion.

2 See the discussion in French and Poterba (1991), Gordon and Bovenberg (1996), or Lewis (1999).
By using actual portfolios, our estimates reflect the behaviour of real fund managers, and so may better measure the true cost of the restrictions.

The plan of the paper is as follows. We next describe in Section 2 the retirement savings institutions in Canada, the mechanics of the foreign property limit, and the liberalizing reforms of the early 1990s. Section 3 presents a model of optimal portfolios in the presence of quantity restrictions on foreign asset holdings, which forms the basis for our estimation strategy. Section 5 describes the data and results of estimation, and Section 6 concludes the paper.

2 Pensions and foreign property restrictions in Canada

In Canada, pension savings receive special regulatory treatment when held in registered plans. These plans take two forms. Registered Pension Plans (RPPs) are pensions provided by an employer, while Registered Retirement Savings Plans (RRSPs) are discretionary, individual savings plans. Contributions to both RPPs and RRSPs are initially tax-deductible for plan beneficiaries (and employers in the case of RPPs), accrued income is tax-exempt, and withdrawals are taxable as ordinary income. Under current regulations, each taxpayer is entitled to contribute up to the lesser of C$13,500 or 18 per cent of earned income to RRSPs and RPPs combined. In 1997, the book value of assets in RPPs stood at C$587 billion, with C$235 billion in RRSPs Canada (1996).

Assets held in RPPs and RRSPs are subject to foreign property restrictions. Prior to 1990, no more than 10 per cent of wealth in these plans could be held in assets deemed to be foreign property. The limit was subsequently raised in two per cent annual increments until reaching 20 per cent in 1994, where it has remained since. RRSP assets are often held in publicly traded mutual funds, which receive special regulatory treatment. If a mutual fund holds less than 20 per cent of its assets in domestic property, then the fund is deemed “eligible”. If a fund exceeds the 20 per cent limit, then it is deemed “ineligible”. Eligible funds count as wholly domestic property for an individual’s foreign content limit, while ineligible funds count as 100 per cent foreign property. So, eligible mutual funds face the same restrictions as individual tax-exempt investors and pension funds, whereas ineligible funds do not.

3 A model of investment under a foreign property limit

How does regulation of outward foreign investment affect the behaviour of investors and the equilibrium returns to constrained portfolios? Here we briefly present an extension of the Capital Asset Pricing Model that incorporates quantitative restrictions on portfolio shares such as the foreign property rule. The model provides a basis for empirical specifications of the effect of the limit estimated in Section 5.

Let \( \tilde{R} \) be a \( K \)-vector of risky security returns with \( \tilde{R} \sim N(R,V) \). There exists a safe asset which pays a return \( r \). Investors are divided into two groups. If \( i \in C \), then the investor is

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3 Similar restrictions exist in other countries. In Germany, for example, foreign property may not exceed 30 per cent of pension funds’ assets. In Japan, the limit is six per cent. See Laboul (1998) and Davis (1995) for more comprehensive lists of these regulations, and Mathieson and Rojas-Suarez (1994) for a survey of their causes and consequences.

4 It should be noted that it is not possible for investors to offset the behaviour of eligible funds with “homemade foreign content,” i.e. by holding additional foreign assets directly. Since constrained funds count as 100 per cent domestic property under the regulations as long as they respect the 20 per cent limit, an investor can in fact obtain 36 per cent foreign content by holding foreign assets to the limit on personal account and then investing the balance of the portfolio in mutual funds with maximal foreign content.

5 Our model is related to the treatment of international transactions taxes in Black (1974) and Stulz (1981). See also Gouriéroux and Jouneau (1999) for an analysis of econometric tests of the CAPM in the presence of affine constraints on portfolios.
constrained to hold a maximum of \( L \) per cent of assets in foreign property. If \( i \in \bar{C} \) then the portfolio is unconstrained. Similarly, let \( F \subset \{1, \ldots, K\} \) be the index set of foreign securities, and its complement be the set of risky domestic securities. The safe asset is considered domestic property. Let \( w \) be the \( K \)-vector of portfolio weights on the \( K \) risky assets; that is, \( w_k \) is the fraction of each dollar in wealth that is invested in security \( k \). A portfolio then satisfies the foreign content restriction if and only if\(^6\)

\[
\sum_{k \in F} w_k \leq L. \tag{1}
\]

Each investor \( i \) has constant absolute risk aversion preferences for aggregate returns to their portfolios, with a utility function \( u_i(x) = -\exp(-\rho_i x) \) and initial wealth \( A_i \). Investor \( i \) chooses a vector of portfolio weights \( w_i \), and end-of-period wealth is a random variable \( \tilde{R}_i = A_i \sum_k w_{ik} \tilde{R}_k \). Given the investor’s preferences, an optimal portfolio maximizes the certainty equivalent rate of return

\[
\max_{w^i} \mu_i(w^i) - \frac{\rho_i}{2} \sigma_i^2(w^i) \tag{2}
\]

subject to (1), where

\[
\mu_i(w^i) = A_i \sum_k w_k (R_k - r) \tag{3}
\]

\[
\sigma_i^2(w^i) = A_i^2 \sum_k \sum_l V_{kl} w_k^i w_l^i \tag{4}
\]

are the expected value and variance of terminal wealth respectively.

**Proposition 1** Let \( \mu_m \) equal the expected rate of return to the world market portfolio. The equilibrium expected rate of return \( \mu_i \) to the optimal portfolio of investor \( i \) satisfies

\[
\mu_i - r = \begin{cases} 
\beta_i (\mu_m - r - \alpha) + \gamma L & \text{if } i \in C \\
\beta_i (\mu_m - r) & \text{otherwise}
\end{cases} \tag{5}
\]

where \( \alpha \), \( \gamma \), and \( (\beta_i) \) are scalar parameters, independent of the policy variable \( L \).

*Proof.* See appendix.

When \( \gamma > 0 \), a constrained investor \( i \in C \) chooses a portfolio on the interior of the mean–variance frontier, and the standard two-fund separation property of the Capital Asset Pricing Model is violated. The proposition states that the departure from two-fund separation is a simple one, however, as the expected return to the constrained portfolio increases linearly in the foreign property limit \( L \), as long as the constraint is binding. Thus the slope parameter \( \gamma \) can be interpreted as the marginal expected return for investors associated with increasing the limit. Estimating this parameter is the goal of our empirical research reported below. Moreover, when \( \alpha > 0 \), the proposition states that the cost of the constraint is larger for more aggressive (high beta) investors. This is as expected, since investors with high risk tolerance hold less of their wealth in the safe

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\(^6\)In this formulation, short positions in foreign assets count as negative foreign content. Under Canada’s foreign property rules, in contrast, short positions receive a weight of zero and thus are simply excluded from the foreign content calculation. The analysis could be extended to deal with this regime, albeit with considerable complications. See Stulz (1981) for a related analysis of asymmetric taxes on international portfolio transactions.
asset, which is domestic property. These investors must therefore accept greater distortions in their portfolios, relative to those predicted by the CAPM, in order to satisfy the foreign property limit.

4 Data
To study the effect of the foreign property limit on returns we employ a sample of 979 equity mutual funds taken from the Financial Post Mutual Fund data base. The sample includes only those funds in operation in July 1998. The Financial Post data base reports the net asset value per share, inclusive of reinvested dividends, for each fund for every month from the fund's inception (or January 1961) to July 1998. The data base also reports several fund characteristics such as the fund's RRSP eligibility status, management expense ratio, and percentage of the portfolio invested in foreign assets. Most of these characteristics are reported only for July 1998 - no longitudinal information is available. The return to the market portfolio is proxied by the Toronto Stock Exchange (TSE) 300 total return index. To calculate excess returns, we use the Government of Canada 30-day treasury bill rate as the risk-free rate.

Several funds were eliminated before arriving at the data set used for estimation. Funds with fewer than 36 monthly observations, and funds missing any monthly observations of net asset value per share were eliminated to maintain the panel structure of the data set. In order to focus on diversified equity funds, we eliminated 32 sector-specific funds. We also removed 51 funds that follow an avoidance strategy using derivatives of foreign assets. Because this strategy allows fund managers to increase effective foreign content above the limit while maintaining RRSP eligibility, these funds do not fit within our model. These selection criteria left 58119 monthly observations from 510 mutual funds spanning a period from 1961 to 1998. Of these, 206 were eligible for treatment as domestic content, with the remaining 304 being deemed ineligible.

The unbalanced structure of this panel may pose problems for estimation. Some funds first appear only after the foreign content limit was raised to 20 per cent, while others existed both before and after the reform. If the “new” funds differ in unobservable ways from the “established” funds then it may be difficult to separate the effect of the limit from the effect of the unobservable differences. For this reason, we extracted a subsample of funds for which reported returns are available for a period spanning January 1986 to July 1998. This balanced panel contains 151 monthly observations per fund for 125 funds, 76 of which were eligible for treatment as domestic content. Because all funds in the balanced panel exist throughout the reform period, we avoid any problems associated with the inclusion of new funds.

Figure 1 presents a histogram of the 510 funds in our full sample, classed by their reported percentage of portfolio in foreign assets in July 1998. Perhaps most striking is that 122 of the 206 constrained funds held less than 10 per cent of their assets abroad. Because international diversification is maximized only if the constraint is binding, the behaviour of these mutual funds remains a puzzle. We further explore the behaviour of these funds in presenting our empirical results below.

5 Estimates of the policy’s impact
To begin our analysis of the impact of the foreign property rule on mutual fund performance, we examine some summary statistics of our data set. Table 1 presents measures of return and risk for the two classes of funds. The first three rows compare the mean monthly excess return to eligible

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7Our results is potentially susceptible to survivorship bias. However, this becomes a concern only to the extent that a fund’s probability of survival was influenced by reforms to the foreign content limit.

8Burgess and Fried (1999) describe in detail the financial techniques used by these funds.
funds, ineligible funds, and the TSE 300 index before and after the reform, as well as for the full sample period. Before the reform, eligible funds had an average return 9.3 basis points (b.p.) lower than the average return of ineligible funds. In the period following the reform, the gap between the average returns fell to 3.5 b.p. The average market beta of eligible and ineligible funds is reported in the fourth row of Table 1. The mean beta of eligible funds is much higher than that of ineligible funds, meaning that eligible funds on average hold portfolios exposed to more systematic risk than ineligible funds. So, there is some evidence that the gap in the performance of eligible and ineligible funds closed following the reform to the foreign property limit, as predicted by the theory.

Are these differences caused by the foreign property limit? If eligible and ineligible funds differ in ways unrelated to the policy but which affect returns, inferences about the effect of the limit require a more sophisticated empirical strategy. We therefore extend these comparison-based estimates to a regression framework, progressively adding controls for other factors. We first examine returns using a simple indicator variable for eligible funds. Next, we add the excess return to the market portfolio to the regression to control for systematic risk. This leads to an examination of the influence of observable fund characteristics on fund returns. Finally, at the end of this section, we estimate our structural model, based on (5).

The first row of Table 2 shows the results of a regression of excess fund return on an indicator for RRSP eligibility status. The coefficient can be interpreted as the average difference in monthly return over the entire sample period. For the balanced panel, this difference is 17.8 b.p., while for the full sample the average difference is 12.1 b.p.\(^9\) Both of these differences are significant at the 5\(^9\)

\(^9\)Pink (1989) also measures the cost of the policy by comparing returns to eligible and ineligible mutual funds, for
Table 1: Risk and return for constrained and unconstrained funds

<table>
<thead>
<tr>
<th></th>
<th>Eligible Funds</th>
<th>Ineligible Funds</th>
<th>Market Index</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean monthly excess return (basis points):</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Pre-reform period (1961-89)</td>
<td>25.5</td>
<td>34.8</td>
<td>32.6</td>
</tr>
<tr>
<td>Post-reform period (1994-98)</td>
<td>43.7</td>
<td>47.2</td>
<td>71.7</td>
</tr>
<tr>
<td>All years</td>
<td>26.0</td>
<td>37.0</td>
<td>32.2</td>
</tr>
<tr>
<td>Mean market beta</td>
<td>0.824</td>
<td>0.568</td>
<td>1</td>
</tr>
</tbody>
</table>

In order to exploit the intertemporal variation in the foreign property limit, we constructed a variable representing the actual limit for each fund in each year. Thus, \( L = 1 \) for ineligible funds in all periods, and \( L \in [0.1, 0.2] \) for eligible funds. Row 2 of the table presents the estimated coefficient for the limit variable \( L \). This specification therefore captures the effect of the reform on returns to eligible funds, controlling for common temporal effects in returns using observed returns to ineligible funds. The estimate is a significant 26.2 b.p. in the balanced panel and 17.6 b.p. in the full sample. As an additional control for temporal effects, we include a full set of year indicator variables. The coefficient on \( L \) falls to 20.8 b.p. for the balanced panel and an insignificant 5.2 b.p. for the full sample. Because the full sample has more observations for later periods than for earlier periods (as a consequence of the introduction of new funds), less weight is put on earlier periods relative to later periods when temporal controls are excluded. In contrast, the balanced panel has an equal number of funds in each period by construction. When temporal controls are included, the effect of the heavier weighting on later periods no longer affects the estimated effect of the limit. This may explain why the estimate shows greater sensitivity to the inclusion of temporal controls in the full sample.

We observed earlier that the average market beta of eligible funds exceeds that of ineligible funds. To control for the possibility that our estimate of the effect of the policy is affected by differences in systematic risk, we next estimated

\[
\tilde{R}_{it} - r_t = \alpha + \beta_i (\tilde{R}_{mt} - r_t) - \gamma (1 - L_{it}) + \epsilon_{it}
\]

for all funds. By including the excess return to the market portfolio as a regressor, this specification controls for differences in systematic risk and the impact of abnormal market returns on the two classes of funds. For example, if the domestic market underperformed during the reform period, excluding this variable would lead to bias. This specification conforms to our theoretical model (presented in Section 3 and summarized by (5)). However, we depart from the theory by including ineligible funds in the sample. These observations control for unobserved changes in fund returns contemporaneous with the reform that are not reflected in the return to the market index, but that are common to the two classes of funds. Row 4 of Table 2 displays estimates of \( \gamma \) using (6). The results are very similar to the specification which includes only \( L \) — 26.4 b.p. in the balanced panel

\[\text{1974-85 period. Because he employs a different methodology, our results are not directly comparable, but appear to be similar in magnitude.}\]

\[10\text{In all cases, we report standard errors adjusted for arbitrary forms of heteroskedasticity, using the Huber–White variance estimator. OLS standard errors are similar.}\]

\[11\text{The use of year indicator variables matches the periodicity of the foreign property restriction. An alternative would be to insert controls for each period. This changes the estimates only slightly.}\]
### Table 2: Estimated effect of foreign property limit

<table>
<thead>
<tr>
<th>Sample:</th>
<th>Balanced Panel</th>
<th>Full Sample</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. Eligibility Indicator</td>
<td></td>
<td></td>
</tr>
<tr>
<td>ELIGIBLE</td>
<td>17.8* (6.6)</td>
<td>12.1* (3.6)</td>
</tr>
<tr>
<td>2. Limit alone</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$L$</td>
<td>26.2* (7.8)</td>
<td>17.6* (4.3)</td>
</tr>
<tr>
<td>3. Limit with year indicators</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$L$</td>
<td>20.8* (7.7)</td>
<td>5.2 (4.3)</td>
</tr>
<tr>
<td>4. Limit with fund betas</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$L$</td>
<td>26.4* (5.6)</td>
<td>16.4* (3.1)</td>
</tr>
</tbody>
</table>

**Notes:** Estimated coefficient on $L$ in regressions, measured in basis points per month. Estimated robust standard errors in parentheses.  
*: Significant at 5 per cent level.

compared to 26.2 b.p. above — although the inclusion of controls for systematic risk improves the precision of the estimates. This suggests that differences in systematic risk did not underlie the estimated effect of the policy using $L$ alone.

For a number of reasons, the use of ineligible funds as a control group may be inappropriate to the data. First, even if returns for the funds in the two classes do not differ systematically in ways unrelated to the policy, the marginal effects of policy reforms may differ at different levels of the foreign property limit. That is, the linear specification in (6) implies that a ten percentage point increase in the limit, from 20 to 30 per cent, would have an effect on fund expected returns equal to the effect of an increase from 90 to 100 per cent. Since, in the presence of home bias, few funds would likely become fully diversified following such a reform, our linear specification may underestimate the impact of small increases in the current limit. Second, fund characteristics affecting returns may vary systematically across the two classes of funds. In this case, the efficacy of the ineligible funds as a control group diminishes, since any observed differences between eligible and ineligible funds might be in reality attributable to the characteristics rather than the foreign property limit. The results presented in Table 3 explore the sensitivity of our estimates to three different approaches to this problem.

In Row 1 of Table 3, we present results for a specification that includes quadratic terms in the limit variable (labelled $L$ and $LSQ$). This relaxes the assumption of a linear effect of the policy, allowing different marginal effects of the policy at different levels of the limit. For the balanced panel, the coefficient on the limit variable is now 32.5 b.p. and the coefficient on the quadratic term is -7.1 b.p. This implies that the marginal effect of a change in the foreign content limit at $L = 0.2$ is 31.1 b.p., and 25.4 b.p. at $L = 1$. However, an $F$-test comparing the linear and the quadratic specifications strongly rejects the inclusion of the quadratic term (the $F$-statistic is 0.02 for the balanced panel and 0.01 for the full sample). So, allowing the limit to have a quadratic effect on returns has only a small and insignificant impact on the estimate of the cost of the policy relative to the linear specification.

If ineligible funds are inappropriate as a control group, an alternative approach is simply to discard observations on ineligible funds and identify the impact of the policy by studying eligible funds through the 1990-94 reform period. This follows from a strict interpretation of the model as applying only to funds constrained by the limit. This is appropriate if the two classes of funds are entirely incomparable, because of unobservable yet relevant fund characteristics, or differences in
Table 3: Sensitivity of estimates to assumptions about the control group

<table>
<thead>
<tr>
<th>Sample:</th>
<th>Balanced Panel</th>
<th>Full Sample</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. Limit quadratic with fund betas</td>
<td>L 32.5 (42.0)</td>
<td>19.6 (28.8)</td>
</tr>
<tr>
<td></td>
<td>LSQ -7.1 (47.8)</td>
<td>-3.7 (33.0)</td>
</tr>
<tr>
<td>2. Eligible funds only, with fund betas</td>
<td>L 12.2 (39.9)</td>
<td>7.6 (27.6)</td>
</tr>
<tr>
<td>3. Eligible funds only, with fund betas and fixed effects</td>
<td>L 12.2 (39.9)</td>
<td>-31.5 (30.9)</td>
</tr>
</tbody>
</table>

Notes: Estimated coefficient on $L$ in regressions, measured in basis points per month. Estimated robust standard errors in parentheses.

*: Significant at 5 per cent level.

the impact of the policy on the two classes. Here we estimate $\gamma$ with OLS on eligible funds only, using policy variation alone to identify $\gamma$. With this specification, the market return $\tilde{R}_{mt}$ serves as the primary control for unobserved temporal effects in fund returns, rather than the returns to the control group of ineligible funds. Estimating the policy’s effect without the ineligible funds relieves the need to make assumptions about their appropriateness as a control group. However, discarding the control group introduces a different set of complications. Any trends that affected mutual fund returns relative to the market index may be wrongly attributed to the change in the foreign property limit over time. For example, if fund companies were changing managers over the reform period, any increase in fund returns attributable to better portfolio management could result in an upward bias in the estimate of $\gamma$. Results for this specification are reported in row 2 of Table 3. The estimated effect of the policy is smaller than the estimates including the ineligible funds, for both the balanced panel and the full sample. With the ineligible funds as a control group, the estimated coefficient on $L$ was a significant 26.4 b.p. in Table 2 for the balanced panel. Without the control group of funds, the estimate falls to an insignificant 12.2 b.p. For the full sample, the point estimate drops to -31.5 b.p., although this is not significantly different from zero.

Similarly, one might allow for unobservable permanent differences in returns to funds which are unrelated to the foreign property restrictions per se. Several studies of mutual funds performance have found persistent differences in returns to individual funds (so-called “hot hands”; see Carhart 1997 for a review of the literature). To examine the robustness of our results to these considerations, we simply added fund fixed effects to the preceding regression equation. If (6) is the correct specification, then it should not be possible to reject the hypothesis that the fixed effects are jointly equal to zero. Estimates of $\gamma$ for the fixed effects regressions are reported in Row 3 of Table 3. The fixed effects for both data sets are jointly significant. For the balanced panel, the estimated coefficient on $L$ is necessarily the same as the previous specification. Each fund appears in each period, so there will be no correlation between the fixed effects and $L$, which

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12Grinblatt and Titman (1989) observe that estimated persistent differences in performance can reflect market-timing strategies of managers, among other factors.

13Our results in this case would be unchanged if ineligible funds were included in the estimation. Since $L = 1$ over the entire sample period for ineligible funds, the effect of the limit on the returns of ineligible funds is indistinguishable from the fixed effect.

14The F-statistic for fixed effects using the balanced panel is 2.25. Since there are 75 restrictions, the five per cent critical value is 1.28. For the full sample with 205 restrictions, the F-statistic is 2.71 and the five per cent critical value is 1.17.
leaves the estimate of $\gamma$ unchanged by the fixed effects. In contrast, the full sample includes funds that first appear only after the limit reaches 20 per cent. For these funds $L$ is collinear with the fixed effect. The point estimate for this specification is much lower than without the fixed effects, but is insignificantly different from zero. With both data sets, the 95 per cent confidence interval of the estimates contains the corresponding point estimate from the OLS specification in Row 4 of Table 2, but the fixed effect point estimate is much lower. Thus we conclude that these estimates are not inconsistent with the cross-sectional estimates.

Our data include several fund characteristics that may be related to fund performance, which allows us to develop the fixed effect approach further. Previous research (e.g., Carhart, 1997) has suggested that funds with high management expense ratios and large asset pools under management tend to underperform the average fund. The inclusion of fund characteristics permits a richer examination of potentially influential differences among funds than possible with a simple fixed effect specification. Moreover, the results for regressions including characteristics suggest an explanation for the lower point estimates from regressions on eligible funds alone. We extend (6) to include the recorded management expense ratio (EXPRAT) in our data set, usually for July 1998, and the lagged value of the fund’s total net assets (TNA). Additionally, we investigate the impact of a fund’s foreign content on returns. Our data set records the actual percentage of a fund’s assets invested in foreign property for the most recent month available. Our formal model assumes that the foreign property constraint is binding on all RRSP-eligible portfolios at all dates. In fact, Figure 1 makes clear that many funds in the sample report foreign property well below the current 20 per cent limit. To investigate the influence of these funds on aggregate results, we created an indicator variable (DOM) equal to one for funds with reported foreign content less than one per cent, and we include DOM and its interaction with the limit variable in the regressions. (There are 28 such funds in the balanced panel and 89 in the full sample.) We report coefficient estimates for $L$ and the characteristics variables in Table 4. We report results for the balanced panel and full sample, both with and without the ineligible funds included in estimation.

The first section of Table 4 displays the results for regressions including both eligible and ineligible funds. As expected, management expense ratios and total assets are negatively related to fund excess returns. In the balanced panel, a one percentage point increase in a fund’s annual management expense ratio is estimated to reduce returns by 12.0 b.p. per month. To put this figure in perspective, observe it implies that a 100 b.p. increase in the MER decreases gross annual returns by approximately 44 b.p. The estimated effect of fund asset size is negative but relatively small in magnitude: an increase in a fund’s total net assets of $1 billion is estimated to decrease monthly returned by an insignificant 3.1 b.p. in the balanced panel. Coefficients for DOM and its interaction with the limit are not significantly different from zero. Thus we infer that inclusion of these specialized domestic funds in the sample has negligible impact on our results.

The estimated coefficient on $L$ is a significant 30.0 b.p. in the balanced panel, which is slightly higher than the 26.4 b.p. for the regression without fund characteristics in Row 4 of Table 2. When the characteristics are excluded, the estimate of $\gamma$ will be influenced by any correlations between the excluded characteristics and $L$. In the balanced panel, eligible funds have a smaller average TNA than ineligible funds (C$181 million compared to C$223 million). As well, eligible funds

\footnote{The one period lag of TNA should obviate any endogeneity of TNA with returns. Substitution of the one-period lag with lags of one quarter, one year, or 2 years changed the results little.}

\footnote{TNA was constructed from the reported total net assets in each period. TNA is not reported for 6,035 observations (10.4 per cent) in the full sample, and 596 (3.1 per cent) in the balanced panel. For these periods, the TNA variable was constructed using the reported shares outstanding and net asset value per share. Regressions (not reported here) using samples that exclude the incomplete observations gave similar results.}
Table 4: Estimated effect of foreign property limit and fund characteristics

<table>
<thead>
<tr>
<th>Sample:</th>
<th>Balanced Panel</th>
<th>Full Sample</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. Eligible and Ineligible Funds</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$L$</td>
<td>30.0* (5.6)</td>
<td>17.6* (3.3)</td>
</tr>
<tr>
<td>EXPRAT</td>
<td>-12.0* (3.2)</td>
<td>-16.5* (1.9)</td>
</tr>
<tr>
<td>TNA</td>
<td>-3.1 (3.3)</td>
<td>-3.2 (2.5)</td>
</tr>
<tr>
<td>DOM</td>
<td>-18.9 (59.3)</td>
<td>-16.8 (20.5)</td>
</tr>
<tr>
<td>DOM*L</td>
<td>-20.4 (69.8)</td>
<td>-8.7 (24.1)</td>
</tr>
<tr>
<td>2. Eligible Only</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$L$</td>
<td>41.4 (50.0)</td>
<td>59.1 (37.0)</td>
</tr>
<tr>
<td>EXPRAT</td>
<td>-5.2* (2.4)</td>
<td>-5.6* (1.8)</td>
</tr>
<tr>
<td>TNA</td>
<td>-11.7* (3.9)</td>
<td>-17.8* (2.7)</td>
</tr>
<tr>
<td>DOM</td>
<td>-21.7 (72.3)</td>
<td>-24.5 (48.7)</td>
</tr>
<tr>
<td>DOM*L</td>
<td>-24.5 (84.9)</td>
<td>-16.2 (56.7)</td>
</tr>
<tr>
<td>3. Fixed effect with TNA</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$L$</td>
<td>61.3 (42.4)</td>
<td>44.4 (33.1)</td>
</tr>
<tr>
<td>TNA</td>
<td>-28.3* (6.6)</td>
<td>-32.8* (3.8)</td>
</tr>
</tbody>
</table>

Notes: Estimated coefficients on $L$ and fund characteristics in regressions, measured in basis points per month. EXPRAT is the fund management expense ratio in July 1998, measured in percentage points. TNA is the fund's total net assets, measured in billions of dollars. DOM is an indicator variable equal to one for funds with actual foreign content in July 1998 less than one per cent, and equal to zero otherwise. Estimated robust standard errors in parentheses.

*: Significant at 5 per cent level.

have smaller management expenses, 1.92 per cent on average compared to 2.21 for ineligible funds. Because these characteristics vary across the two classes of funds, excluding them will in general lead to bias.

The second section of Table 4 displays results from regressions on the eligible funds only. Compared to the corresponding estimates in row 2 of Table 3, the estimates of $\gamma$ here remain insignificant, but are much larger. Without the characteristics, any trend in risk-adjusted returns through the reform period will be ascribed to the increase in the foreign content limit. In particular, this period saw a large increase in the average TNA. (In the balanced panel, the average TNA of an eligible fund was C$116 million in 1990 and C$212 million in 1994.) Because TNA decreases returns, excluding TNA from the regressions may confuse the return-enhancing effect of the foreign content limit liberalization with the return-diminishing effect of the trend in TNA. In contrast, when TNA is included, the effect of the trend in TNA is removed from the estimate of the policy’s impact. The last section of Table 4 shows a similar effect on the estimated coefficient on $L$ when TNA is included in the fixed effect estimations. This emphasizes the value of using the ineligible funds as a control group. When the control group is included in the regression, trends that are common to mutual funds of both classes (such as the growth in average TNA through the reform period) do not affect estimates of $\gamma$.

While these cross-sectional and intertemporal comparisons are a natural way to investigate the effect of the policy, estimation of (6) does not precisely implement the relationship between the policy and expected excess returns to constrained portfolios implied by our theoretical model in
Table 5: Non-linear least squares estimates

<table>
<thead>
<tr>
<th>Coefficient:</th>
<th>Balanced Panel</th>
<th>Full Sample</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. No characteristics</td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \gamma )</td>
<td>-18.9 (19.5)</td>
<td>11.4 (11.0)</td>
</tr>
<tr>
<td>( \alpha )</td>
<td>0.3 (3.1)</td>
<td>0.2 (1.5)</td>
</tr>
<tr>
<td>2. With characteristics</td>
<td></td>
<td></td>
</tr>
<tr>
<td>( \gamma )</td>
<td>40.6 (52.1)</td>
<td>59.1 (37.3)</td>
</tr>
<tr>
<td>( \alpha )</td>
<td>-21.2 (13.2)</td>
<td>4.6 (9.7)</td>
</tr>
</tbody>
</table>

Notes: Estimates of equation (7), using sample of constrained funds. Asymptotic standard errors in parentheses. For extreme parameter starting values, the estimation algorithm showed some signs of local convergence.

*: Significant at 5 per cent level.

Section 3. To provide estimates of \( \gamma \) based on the structural model, we next add expectational errors to (5) and estimate

\[
\tilde{R}_{it} - r_t = \beta_i (\tilde{R}_{mt} - r_t - \alpha) + \gamma L_{it} + \epsilon_{it} \tag{7}
\]

for eligible funds,\(^{17}\) where \( \gamma \), \((\beta_i)_{i \in C}\) and \( \alpha \) are parameters to be estimated by non-linear least squares. Observe that (7) is in fact a special case of the fixed effects specification previously reported, in which fund fixed effects \((-\beta_i \alpha)\) are constrained to be collinear with fund betas.

Estimates of \( \gamma \) and \( \alpha \) are presented in Table 5, measured in b.p. per month. The first section contains estimates from the balanced panel and full sample without fund characteristics. The point estimates of both \( \gamma \) and \( \alpha \) are small and not significantly different from zero at the five per cent level. Because these estimates are from regressions on eligible funds only, the same potential bias from trends in other factors affecting mutual fund performance over the reform period could influence results. The second section of the table reports results from regressions including measures of observable fund characteristics. With the characteristics included, the estimates of \( \gamma \) become 40.6 b.p. for the balanced panel and 59.1 b.p. in the full sample. With the other fund characteristics included, the estimates for \( \alpha \) change considerably. This is to be expected, since the inclusion of fund characteristics changes the interpretation of the fund fixed effects.

In summary, our estimates of the effect of the foreign property restriction suggest a significant cost of the policy to pension beneficiaries. When controlling for systematic risk and using ineligible funds as controls, we estimate the cost of the policy to be 26.4 basis points per month. We explore assumptions about the comparability of the two classes of funds, and find that the results are consistent with our main estimates, although they are sensitive to the exclusion of variables that may affect returns. Structural estimation of the exact non-linear model produces estimates of similar magnitude.

5.1 Simulations

Our estimates of \( \gamma \) can be used to calculate the impact of incremental policy reforms on returns. In Table 6, we examine an increase in the foreign property limit from 20 per cent to 30 per cent, using

\(^{17}\)Since (5) makes no prediction about expected returns to unconstrained funds, we excluded ineligible funds from estimation here.
Table 6: Long-term effects of limit increase from 20 per cent to 30 per cent

<table>
<thead>
<tr>
<th>γ</th>
<th>annual</th>
<th>20 year</th>
</tr>
</thead>
<tbody>
<tr>
<td>12.1</td>
<td>0.14%</td>
<td>4.4%</td>
</tr>
<tr>
<td>16.4</td>
<td>0.20%</td>
<td>6.0%</td>
</tr>
<tr>
<td>26.4</td>
<td>0.31%</td>
<td>9.9%</td>
</tr>
<tr>
<td>41.4</td>
<td>0.49%</td>
<td>15.9%</td>
</tr>
</tbody>
</table>

Notes: The values of γ are taken from Table 2 and Table 4. “Annual” refers to the implied increase in annual return, measured in percentage points. “20 year” is the increase in terminal portfolio wealth after 20 years, measured in percentage points.

selected point estimates of γ. The implied increase in annual return, as well as the return over a 20-year horizon are reported. The γ estimate of 26.4 b.p. implies an increase in annual returns of 0.31% for a limit increase to 30 per cent. To place the magnitude of this estimate in a broader context, we observe that this 0.31% increase means a C$310 million increase in annual returns for pension beneficiaries for every C$100 billion held in equities through eligible pension funds. Since pension funds collect contributions to fund obligations decades in the future, we also present portfolio returns over a longer horizon. The implied increase in 20-year return corresponding to the same γ estimate is 9.9%. If 20-year returns on a portfolio increase by 9.9%, then the same initial investment could fund an annuity 9.9% larger. These simulations suggest that even small values of γ could have a large impact on the retirement income provided by registered plans.

6 Conclusion

This paper presents a model of mean-variance efficient portfolios when outward foreign investment is constrained, which serves to motivate our estimation of the cost of Canada’s 20 per cent foreign content restriction on pension assets. We study mutual fund performance through a period of regulatory reform, examining a number of different samples and specifications. A significant advantage of our approach is that it draws inferences from observed returns to real-world portfolios following the increase in Canada’s foreign property limit in the early 1990s. We believe our approach is likely to be superior to the standard methodology, which measures the cost of the policy using hypothetical portfolios. In contrast, our estimates reflect actual behaviour of fund managers, such as home bias in portfolio choices and active trading strategies, and so should generate better estimates.

The estimates suggest the policy has a significant negative impact on risk-adjusted returns to retirement saving in Canada. Whatever objective it might be held to serve, the policy imposes costs on pension plan beneficiaries, lowering rates of return on saving and increasing non-diversifiable risk. Given the marked growth in registered pension assets in recent years, and the federal government’s nascent strategy of investing Canadian Pension Plan reserve funds in stock markets, the foreign property rule deserves greater scrutiny.

18 Our simulations assume that Canadian equity markets are globally integrated, meaning that policy changes do not change equilibrium returns on the capital market.

19 The simulations use the average monthly fund return of 94 b.p. as the base on which any increase implied by the estimated γ is added.
Define an indicator variable for foreign assets: \( f_k = 1 \) if \( k \in \bar{D} \) and \( f_k = 0 \) if \( k \in D \). The first-order conditions are, for each investor \( i \) and security \( k = 1, \ldots, K \),

\[
\sum_{j=1}^{K} w_{ij} V_{jk} = (\rho_i A_i)^{-1} (R_k - r - \theta_i f_k)
\]  

(8)

for some multipliers \( \theta_i \geq 0 \), with \( \theta_i = 0 \) if \( i \in \bar{C} \), or if \( i \in C \) and the constraint (1) is slack.

Multiplying (8) by initial wealth \( A_i \) and summing over \( i \in C \cup \bar{C} \) yields, for each \( k \),

\[
\sum_i \sum_j A_i w_{ij} V_{jk} = \left( \sum_i \rho_i^{-1} \right) (R_k - r) - \left( \sum_i \rho_i^{-1} \theta_i \right) f_k.
\]

(9)

Let \( A_m = \sum_i \sum_j A_i w_{ij} \) denote total world market capitalization and \( w_{mk} = \sum_i A_i w_{ik} / A_m \) denote the share of the world market portfolio of risky assets held in security \( k \). Define \( \bar{R}_m = \sum_j w_{mj} \bar{R}_j \) as the rate of return on the world market portfolio and

\[
\bar{\rho} = \left( \sum_i \rho_i^{-1} \right)^{-1}
\]

(10)

\[
\bar{\theta} = \rho \left( \sum_i \rho_i^{-1} \theta_i \right)
\]

(11)

\[
\bar{F} = \sum_j w_{mj} f_j.
\]

(12)

Dividing (9) by \( A_m \) and using (10)–(11) then yields

\[
\text{cov}(\bar{R}_k, \bar{R}_m) = \sum_j w_{mj} V_{jk} = \frac{R_k - r - \bar{\theta} f_k}{\bar{\rho} A_m}.
\]

(13)

When the expression is multiplied by \( w_{mk} \) and summed over \( k \), we have

\[
\text{var}(\bar{R}_m) \equiv \sum_j \sum_k w_{mj} w_{mk} V_{jk} = \frac{R_m - r - \bar{\theta} \bar{F}}{\bar{\rho} A_m}
\]

(14)

where \( R_m \) is the expected return to the world market portfolio. Equations (13) and (14) can then be combined to yield an expression for the world market beta of each security,

\[
\beta_k \equiv \frac{\text{cov}(\bar{R}_k, \bar{R}_m)}{\text{var}(\bar{R}_m)} = \frac{R_k - r - \bar{\theta} f_k}{R_m - r - \bar{\theta} \bar{F}}
\]

or,

\[
R_k - r = \beta_k (R_m - r - \bar{\theta} \bar{F}) + \bar{\theta} f_k.
\]

(15)

To characterize optimal portfolios of constrained investors, the first-order conditions (8) can be
stacked in vector form as

\[ Vw_i = (\rho_i A_i)^{-1}(R - r1 - \theta_i f) \]  

(16)

and then inverted to yield

\[ w_i = (\rho_i A_i)^{-1} \left[ V^{-1}(R - r1) - \theta_i V^{-1}f \right]. \]

(17)

Using the asset pricing expression (15),

\[ w_i = (\rho_i A_i)^{-1} \left[ \frac{R_m - r - \bar{\theta} \bar{F}}{\text{var}(R_m)} w_m + (\bar{\theta} - \theta_i) V^{-1}f \right]. \]

(18)

Define \( z = V^{-1}f \) and

\[ \beta_i = (\rho_i A_i)^{-1} \frac{R_m - r - \bar{\theta} \bar{F}}{\text{var}(R_m)}. \]

Then (18) is written more compactly as

\[ w_i = \beta_i w_m + (\rho_i A_i)^{-1}(\bar{\theta} - \theta_i) z. \]

(19)

The optimal portfolio for the constrained investor \( i \in C \) is therefore a linear combination of \( w_m \) and \( z \). (The latter portfolio is the minimum-variance portfolio that achieves a given level of foreign content \( F_z \equiv f^\top V^{-1}f \).) The investor holds \( w_m \) to the desired scale and then sells \( z \) short in order to achieve the required foreign content \( F \). To eliminate the unobserved multiplier \( \theta_i \) from the expression, recall \( w_i^\top f = F \) when \( \theta_i > 0 \); thus (19) implies

\[ \beta_i \bar{F} + (\rho_i A_i)^{-1}(\bar{\theta} - \theta_i)F_z = F \]

or

\[ w_i = \beta_i w_m + \frac{F - \beta_i \bar{F}}{F_z} z \]

(20)

Define \( \gamma = z^\top (R - r1) / F_z \), a measure of the risk-adjusted return to the adjustment portfolio \( z \), The expected excess rate of return to the constrained portfolio is then

\[ \mu_i - r = w_i^\top (R - r1) \]

\[ = \beta_i \left( R_m - r - \gamma \bar{F} \right) + \gamma F. \]

(21)

as required. \( \square \)
References