Working Paper 2000-8 / Document de travail 2000-8

Testing the Pricing-to-Market Hypothesis: Case of the Transportation Equipment Industry

by

Lynda Khalaf and Maral Kichian

Bank of Canada Banque du Canada

ISSN 1192-5434
ISBN 0-662-28959-5

Printed in Canada on recycled paper

Bank of Canada Working Paper 2000-8

May 2000

Testing the Pricing-to-Market Hypothesis: Case of the Transportation Equipment Industry

by

Lynda Khalaf* and Maral Kichian+

*Département d'économique, Université Laval, Ste. Foy, Canada

+Research Department, Bank of Canada
Ottawa, Canada K1A 0G9
mkichian@bank-banque-canada.ca

Contents

Acknowledgements

The authors would like to thank Larry Schembri, Christos Constantatos, Phillippe Barla, and seminar participants at the Bank of Canada and Université Laval for useful comments and suggestions. All remaining errors are our own.

Abstract

Pricing-to-market (PTM) theory suggests that monopolistic firms which export adjust their destination-specific markups in reaction to exchange rate shocks. These adjustments limit changes in the price of their exports. Thus, important movements in the bilateral nominal exchange rate between two countries that trade are not necessarily fully reflected in the price of imports.

Evidence in favour of PTM has been mostly obtained through hypothesis testing on the OLS, instrumental variable (IV), or single-equation error-correction estimates of partial-equilibrium models. However, we know from the recent econometric literature that Wald tests applied to some of these estimates may give erroneous results in the presence of endogeneity and weak instruments. In this paper, we examine the reliability of the evidence supporting the hypothesis of pricing-to-market using LIML-based LR Monte Carlo tests. These tests, developed by Dufour and Khalaf (1998), have good power and, unlike the Wald test, also have the correct test size.

We find that the size-correct Monte Carlo LR-based test reverses half of the results obtained from the popular Wald test, indicating that PTM may not be as widespread as previously believed. In addition, our results support the view that PTM behaviour is likely to be present in the same industry across different countries and that pass-through is possibly higher with a larger market share of exports.

The above findings are illustrated using the model developed by Marston (1990) and our analysis is conducted for export pricing firms in the transportation equipment industry for three country pairs: Canada exporting to the United States, the United States exporting to Canada, and Japan exporting to (mainly) the United States.

JEL classification: C12, C15, L11, L16
Bank classification: Econometric and statistical methods; Market structure and pricing

Résumé

Selon la théorie voulant que les prix soient adaptés au marché (*pricing-to-market theory)*, les entreprises monopolistiques ajusteraient les marges de profits applicables à leurs différents marchés d'exportation en réaction aux chocs de taux de change. Ces ajustements auraient ainsi pour effet de limiter les variations du prix des exportations. Par conséquent, les mouvements importants du taux de change nominal entre deux pays qui commercent ensemble ne se répercuteraient pas forcément dans leur intégralité sur le prix des importations.

La plupart des résultats obtenus à l'appui de cette théorie proviennent de l'estimation de modèles d'équilibre partiel au moyen des moindres carrés ordinaires, des doubles moindres carrés ou d'une méthode faisant appel à une équation de correction des erreurs. Or, des travaux économétriques récents montrent que l'application de tests de Wald à certaines de ces estimations peut donner des résultats erronés en présence d'endogénéité ou d'instruments de piètre qualité. Dans l'étude dont il est question ici, les auteures ont recours à des tests de Monte-Carlo fondés sur la statistique du rapport des vraisemblances calculée selon une méthode à information limitée pour examiner la validité des résultats qui appuient cette théorie. Ces tests, qui ont été élaborés par Dufour et Khalaf (1998), sont relativement puissants, et, contrairement au test de Wald, leur niveau est également adéquat.

Les auteures constatent que, dans la moitié des cas où le test de Wald n'avait pas entraîné le rejet de l'hypothèse nulle, le test de Monte-Carlo fondé sur la statistique du rapport des vraisemblances aboutit, lui, au rejet. L'adaptation des prix au marché ne serait donc pas aussi répandue qu'on le croyait jusqu'ici. En outre, les résultats obtenus étayent l'opinion voulant qu'un tel comportement soit observé au sein d'une même branche d'activité dans divers pays et que le pays exportateur détenant une part plus importante du marché puisse répercuter ses hausses de coûts plus facilement que les autres.

Les auteures font appel au modèle de Marston (1990) aux fins de leur démonstration et analysent le comportement des entreprises qui exportent du matériel de transport. Elles étudient plus particulièrement les exportations touchant trois paires de pays : celles du Canada à destination des États-Unis, celles des États-Unis vers le Canada et celles du Japon à destination (principalement) des États-Unis.

Classification JEL : C12, C15, L11, L16
Classification de la Banque : Méthodes économétriques et statistiques; Structure de marché et fixation des prix

1. Introduction

Empirical econometricians are often confronted with technical difficulties when estimating and conducting inference on partial-equilibrium models. In the absence of a fully specified model, estimation is carried out using limited-information methods, an example of which is the two-stage–least-squares or instrumental-variable technique. In this context, important issues have recently been raised with respect to inference when the instruments used are weak. A major concern is that poor instruments can lead to serious identification problems. In particular, inference using a Wald test can lead to spurious rejections of the null hypothesis when applying instrumental-variable estimation in presence of such instruments. This over-rejection of the null can have serious implications for the conclusions drawn about the economic relation being examined. Therefore, testing strategies must be employed that are not dependent on the quality of instruments used. This is especially important when one tests hypotheses in single-equation economic models.

This paper is concerned with the application of a size-correct test, which is immune to poor instruments, to a partial-equilibrium model of pricing-to-market (PTM). Pricing-to-market theory was advanced by Krugman (1987) as a possible explanation for the empirical observation that important movements in the bilateral nominal exchange rate between two countries that trade are not necessarily fully reflected in the price of imports. In particular, the theory suggests that monopolistic firms which export adjust their destination-specific markups in reaction to exchange rate shocks. These adjustments limit changes in the price of their exports. In addition, if export-good invoicing is done in the currency of the destination country, this behaviour brings about short-term nominal price stickiness, or local currency price stability (hereafter LCPS) in the importing country for these goods.

Evidence in favour of PTM exists largely through the application of Wald tests to the parameters of various single-equation models, some of which were estimated with instrumental methods to account for endogeneity. The extent of pricing-to-market behaviour in a country's imports and exports has implications that are important not only for explaining deviations from the law of one price and of purchasing power parity in the short run, but also for understanding the international monetary transmission mechanism. Using general-equilibrium models with sticky prices and LCPS, Betts and Devereux (1998, 2000) show that, in the absence of PTM, a positive monetary innovation in the home country will lead to an increase in home output and a decrease in foreign output. On the other hand, when exporting firms in both countries practise PTM and invoice in the buyer's currency, both outputs increase. In addition, these authors show that exchange rate volatility increases markedly when there is no PTM. In another general-equilibrium study, Obstfeld and Rogoff (1999) argue that nominal wage rigidity in conjunction with PTM behaviour under uncertainty creates price stickiness in the price of tradable and non-tradable goods, but that this stickiness is in domestic currency terms. In this case, the international monetary transmission mechanism follows the well-known Mundell-Fleming explanation whereby nominal exchange rate changes will cause important changes in the terms of trade and in the real exchange rate.

The above studies, along with others like Bergin and Feenstra (1999), Betts and Devereux (1996), Faruqee (1995), and Cheung and Chinn (1999) highlight the dynamic interactions between the various microeconomic and macroeconomic variables of the economy and indicate that PTM, ideally, should be estimated in a general-equilibrium model. At present, however, existing estimates of exchange rate pass-through and statistical evidence in favour of PTM come from partial-equilibrium models. These models typically focus on a monopolist who sells a differentiated good locally and in the export market and who maximizes profits by choosing the prices it will charge in the different countries. The first-order condition of this problem yields that the export price to each destination equals the product of the common marginal cost and a destination-specific markup so that the econometric equivalent features regressors that usually consist of the domestic price of the same good that is exported, the exchange rate of the two countries trading, cost factors, and demand shifters.

The evidence coming from such models largely supports the PTM+LCPS hypothesis. For example, in the case of Japanese manufacturing firms exporting to the United States, Feenstra (1989) and Marston (1990) find the pass-through to be 50 to 80 per cent, depending on the industry examined. In the case of other industrialized countries, Knetter (1993) and Feenstra, Gagnon, and Knetter (1996) find that German, French, U.K., U.S., and Swedish exports also exhibit substantial PTM in a variety of countries. These authors find, however, that pass-through is higher with German, U.K., and U.S. firms, especially in the automobile industry. Additional evidence for PTM comes from Giovannini (1988), Hooper and Mann(1989), Kasa(1992), Gagnon and Knetter (1995), and Feenstra and Kendall (1997), to mention a few.

Given the partial-equilibrium nature of these models, however, the above regressions may be subject to endogeneity of regressors. This could thus lead to possible mismeasurement of the extent of pass-through and to erroneous interpretations of the outcomes. Indeed, the estimation methods applied above consist mainly of ordinary least squares and single-equation error-correction methods, both of which lead to inconsistent estimators when endogeneity is present and not accounted for. In the few cases where instrumental variable estimation was used to address this issue,[1] Wald tests were applied to test the PTM hypothesis. Yet recent econometric studies show decisively that hypothesis tests in IV-regressions, as well as the usual Hausman specification tests applied to regressors in view of establishing exogeneity, are both subject to severe size distortions in the presence of poor instruments.[2]

In this paper, we focus on the endogeneity problem,[3] and while we do not construct a general-equilibrium model, we ask whether such a modelling strategy is necessary. We answer this question by comparing evidence for short-run PTM obtained from the usual testing methods to more reliable evidence obtained from applying tests that are not dependent on the "quality" of instruments used. We illustrate our findings using the well-known PTM+LCPS model developed by Marston (1990). Our analysis is conducted for export pricing firms in the transportation equipment industry for three country pairs:[4] Japan exporting to (mainly) the United States, Canada exporting to the United States, and the United States exporting to Canada. We thus show that the accumulated evidence above is not entirely reliable and that more adequate testing of the hypothesis of pricing-to-market needs to be undertaken.

In the next section, we describe Marston's model. Section 3 explains the data and presents estimation results. In Section 4, we describe the Dufour (1987) test for exogeneity and present the results of its application to our equations. Sections 5 and 6 document IV estimation and Wald and LR-based test results on the PTM hypothesis. The final section concludes.

2. The model

To illustrate our arguments, the Marston (1990) model is used where it is assumed that there are two countries consisting of monopolistic firms, each of which sells a differentiated product. These firms can charge different prices for their product in these two locations. Thus, at time t, a domestic firm in industry i can charge a price of P_{it} for its product when it is sold locally and Q_{it}, denominated in foreign currency, when it exports it. Taking the exchange rate and the demand functions in both countries as given, profit maximization for this firm implies the following first-order conditions:

---

1. Feenstra (1989) is one example where instruments are used for wages and income in the model. Klitgaard (1999), on the other hand, uses dynamic ordinary least squares prior to using an error-correction methodology to take this problem into account.
2. See, for instance, Staiger and Stock (1997), Dufour (1997), Wang and Zivot (1998).
3. We do not deal with some of the other concerns that have been raised in the literature such as measurement error biases and the fact that LCPS may not be the correct assumption in the case of all countries.
4. There is extensive partial-equilibrium evidence for the presence of PTM in this industry.

$$P_{it} = C_1\left(\frac{W_{it}}{P_t}, \frac{P_t^m}{P_t}, TD_t\right)M\left(\frac{P_{it}}{P_t}, Y_t\right) \tag{1}$$

$$S_t Q_{it} = C_1\left(\frac{W_{it}}{P_t}, \frac{P_t^m}{P_t}, TD_t\right)N\left(\frac{Q_{it}}{Q_t}, Z_t\right) \tag{2}$$

where S_t is the nominal exchange rate (domestic/foreign), P_t and Q_t are the domestic and foreign general price levels respectively, Y_t and Z_t are domestic and foreign real incomes in these countries. $C_1(\ )$ represents marginal costs of the firm and is a function of nominal wages W_{it}, the price of raw materials P_t^m, and total demand for the firm's output, TD_t. Finally, $M(\ )$ and $N(\ )$ are the markups of domestic and foreign prices over marginal costs.

Totally differentiating equations (1) and (2), defining $X_{it} = (S_t Q_{it})/P_{it}$ and the real exchange rate as $R_t = (S_t Q_t)/P_t$, and combining terms, an expression is obtained for the percentage change in the relative price of the good in the two locations as

$$\left(\frac{dX_{it}}{X_{it}}\right) = \alpha_1\left(\frac{dR_t}{R_t}\right) + \alpha_2\left\{d\left(\frac{W_{it}}{P_t}\right)\Big/\frac{W_{it}}{P_t}\right\} + \alpha_3\left(\frac{dY_t}{Y_t}\right) + \alpha_4\left\{d\left(\frac{P_t^m}{P_t}\right)\Big/\frac{P_t^m}{P_t}\right\} + \alpha_5\left(\frac{dZ_t}{Z_t}\right), \tag{3}$$

where W_{it}/P_t represents real wages in this industry and P_t^m/P_t is the real price of raw materials.

Equation (3) shows that the firm will systematically change its export price relative to its domestic price in reaction to two effects. One change is in response to real exchange rate shocks where markups will be adjusted to limit the impact of exchange rate changes on the competitiveness of the firm. This response is given by the parameter α_1, the so-called PTM elasticity. The other change is in reaction to changes in costs and incomes where, for example, changes in costs of production will lead to unequal adjustments of the firm's profit margins if income elasticities of demand are not similar in the two countries.

Marston shows that the PTM elasticity can be expressed as $\alpha_1 = 1 + \beta_1 - \beta_2$, where β_1 is the elasticity of the foreign currency export price with respect to the exchange rate (commonly known as the pass-through coefficient, this elasticity is negative). β_2 is the elasticity of domestic price with respect to the exchange rate, with a sign that depends on the derivative of marginal cost with respect to output. For example, assuming constant marginal costs, that is $\beta_2 = 0$, full pass-through ($\beta_1 = -1$) implies no pricing-to-market so that α_1 equals zero. At the other extreme, with constant marginal costs, zero pass-through yields $\alpha_1 = 1$.

Marston also shows that α_1 can be written as a function of the elasticities of domestic and foreign markups with respect to prices. Since these reflect the curvature of the respective demand curves, they can be positive, negative, or zero, so that the PTM elasticity can take a number of values. In fact, there is pricing-to-market, that is $\alpha_1 \neq 0$, as long as these markup elasticities are non-constant. In particular, when the demand curves are less convex than a constant elasticity curve, it can be shown that $0 < \alpha_1 < 1$, whether marginal costs are increasing or constant.

Note that, even in the absence of PTM, if firms preset their export prices in the currency of the destination country, unanticipated changes in the nominal exchange rate lead to proportional changes in the relative price ratio.[5] Marston, therefore, writes his estimation equation to also take into account price presetting and exchange rate surprises.[6] This is given in equation (4) below where variables are first expressed in logs and then their differences are taken.

$$\Delta x_{it} = \alpha_0(\Delta s_t - \Delta s_{t-1}) + \alpha_1 \Delta r_t + \alpha_2 \Delta\left(\frac{w_{it}}{p_t}\right) + \alpha_3 \Delta y_t + \alpha_4 \Delta\left(\frac{p_t^m}{p_t}\right) + \alpha_5 \Delta z_t + u_t \qquad (4)$$

From equation (4) then, planned changes in the relative price ratio occur directly through the PTM coefficient and indirectly through the coefficients on real cost and income changes.[7] The remaining fluctuation in the dependent variable is due to surprises to the nominal exchange rate and is captured by the α_0 coefficient. This means that, if foreign currency export prices are preset, unanticipated changes in exchange rates (that is, innovations in the change of the nominal exchange rate) will also move the dependent variable. This effect is thus distinct from the planned pricing strategy of PTM.

Finally, lags of regressors are also allowed in the estimation equation to account for the fact that prices for these may have been preset at earlier times.

5. This point was initially made by Giovannini (1988).
6. Notice that the econometric specification that is used to account for these issues might not strictly distinguish between price-presetting, exchange rate surprises, and PTM. This is because the theoretical model underlying the estimating equation does not derive an optimal price-setting rule in a dynamic, general-equilibrium context where costs of adjusting prices are explicitly considered.
7. See Marston for more details on these indirect effect.

3. Data and estimation

We estimate the above empirical model for our three country pairs using OLS and tabulate the results in Table 1 below. Japanese export estimations are over the 1980:1 to 1987:12 period, that for Canadian exports is 1983:1 to 1998:9, while U.S. export estimations are carried out over the 1981:1 to 1998:9 span. See Appendix 1 for a detailed description of the data.

In the U.S. and Canadian cases, we estimate equations where, along with the contemporaneous values, we include two lags of the exchange rate regressors and one lag of each of the remaining regressors on the right-hand side. In the Japanese case, because of less data, we chose to include only two lags for the real exchange rate, one lag for wages, and no lags for the other regressors.

The OLS estimation results in Table 1 reveal that, for the three country pairs, the PTM elasticity for all industries is statistically greater than zero,[8] with the estimates ranging from 0.39 to 1.39. Thus, it would seem that these firms react systematically to limit changes in the foreign-currency price of exports in the face of a movement in the exchange rate. In other words, there is substantial pricing-to-market in the transportation equipment industries of Japan, Canada, and the United States.[9] There is also evidence of indirect pass-through in one case, the category consisting of U.S. heavy truck exports to Canada where the sum of coefficients on non-exchange rate variables is -1.13 and significant. Interestingly, eight of the nine α_0 coefficients are also significant, indicating that, because of export price pre-setting in foreign currency, nominal exchange rate surprises also influence the relative price variable in the short run. These results are, therefore, largely consistent with the evidence obtained from similar models mentioned above. The question is: Are they reliable?

---

8. Notice that these estimates are very similar to those obtained originally by Marston in the case of Japanese exports to the United States.
9. We should mention that pricing decisions for the Canada-U.S. transportation equipment probably also involve some transfer pricing qualifications because of the existence of the Auto Pact Agreement between the two countries. However we abstract from such issues in the present study.

Table 1: OLS estimation results

	Industry group	OLS estimation results								
		constant		α_0 [a]		α_1 [b]		β [c]		
		Estimate	Test stat. (p-value)	Estimate	Test stat. (p-value)	Estimate	Test stat. (p-value)	Estimate	Test stat. (p-value)	R^2
JPUS[d]	Passenger cars	0.0004	0.2587 (0.7965)	-0.6374	-2.7387 (0.0075)	0.8756	87.0862 (0.0000)	0.0415	0.0126 (0.9110)	0.841
	Small passenger cars	0.0016	1.2657 (0.2092)	-0.3716	-2.2392 (0.0278)	0.46425	48.1194 (0.0000)	-0.2712	1.0582 (0.3066)	0.764
	Trucks	0.0007	0.7638 (0.4471)	-0.0581	-0.4854 (0.6286)	0.3852	68.3629 (0.0000)	-0.1325	0.4625 (0.4983)	0.700
CAUS	Automobiles	0.0012	1.2904 (0.1986)	1.0194	62.5591 (0.0000)	0.9153	36.0129 (0.0000)	-0.1889	0.8860 (0.3479)	0.518
	Trucks	0.0004	0.4603 (0.6459)	1.0889	70.7642 (0.0000)	1.0465	46.0789 (0.0000)	-0.0308	0.0237 (0.8777)	0.526
	Heavy trucks	-0.0003	-0.3903 (0.6967)	0.8336	66.2332 (0.0000)	0.8602	49.7225 (0.0000)	-0.0892	0.3183 (0.5733)	0.583
USCA	Automobiles	-0.0009	-0.7483 (0.4552)	0.9562	25.6019 (0.0000)	1.0249	32.9774 (0.0000)	-0.2231	0.2415 (0.6237)	0.385
	Trucks	0.0000	0.0125 (0.9899)	0.7945	12.4950 (0.0005)	1.0227	23.2101 (0.0000)	-0.5979	1.2255 (0.2697)	0.258
	Heavy trucks	0.0023	1.6153 (0.1079)	0.4414	3.6342 (0.0581)	1.3948	40.6921 (0.0000)	-1.1319	4.1389 (0.0432)	0.276

a. The sum of coefficients on the difference of changes in the nominal exchange rate
b. The sum of coefficients on the changes in the real exchange rate
c. The sum of coefficients on the changes in other variables (the real wages, the real oil prices, the real domestic income, and the real foreign income)
d. JPUS is Japanese exports to mainly the United States, CAUS is Canadian exports to the United States, and USCA is U.S. exports to Canada.

4. Testing for exogeneity

The traditional test for exogeneity of regressors is the Chi-squared Hausman test. However, the recent econometric literature has indicated that the outcome of this test is unfortunately dependent on the choice of instruments. Thus, one may inadvertently draw an erroneous conclusion from the testing exercise based on whether these instruments are really relevant or not to the regression framework. Naturally, the solution is to pick the instruments judiciously. The difficulty is that, in some situations, there is little theory to guide this choice. Furthermore, even if one has some idea of the best variables to choose as exogeneous regressors, data availability often limits the researcher's options.

To avoid falling into the "weak instruments" trap, we apply the exogeneity test statistic proposed by Dufour (1987). Although this test is highly related to that of Hausman, its null distribution has a cut-off point that is valid irrespective of the "quality'' of the available instruments. The test procedure is described in the last section of Appendix 2.

In this model, the real wage variable is clearly prone to endogeneity problems. Obviously, while the presence of a single endogenous right-hand-side variable will bias all of the parameters of the model, there is also a possibility that more than one regressor could be endogenous, in which case the bias would be exacerbated. That is, even if by design the right-hand-side variables in this model are assumed to be exogenous, empirical work should account for the fact that the data are more compatible with one where feedback effects exist. For instance, studies point out that the real exchange rate influences real income, especially in small open economies such as Canada's.[10] One of the channels through which this effect occurs is investment. We know, from Campa and Goldberg (1999), that the real exchange rate affects investment and that this depends on the competitive structure of the industry and its input markets. Another example is that, given that estimations are generally carried out on industry-level data, pricing strategies adopted by an important industry could conceivably directly influence the terms of trade and the real exchange rate. Faruqee (1995), for example, has shown that when nominal rigidities are present, PTM increases the persistence in the real exchange rate. Similarly, Cheung and Chinn (1999) demonstrate how, in an imperfectly competitive market environment, different PTM strategies by industries can be an important factor in causing deviations from PPP.

Given these possible interrelations for our data, we carry out our exogeneity tests simultaneously on the subset composed of the real exchange rate, real output, and real wages.[11] The

10. See, for instance, Ball (1998) and Campa and Goldberg (1999).
11. We ran exogeneity tests also on each variable alone and found that results were qualitatively similar.

instrument sets used for this purpose are specific to each exporting country and are the second and/or third lags of log differences of the following variables: for Canada—average hourly earnings and fixed-weight average hourly earnings in all sectors, average weekly earnings in the manufacturing sector, total employment, and real oil prices; for the United States—average hourly earnings in all sectors, average hourly earnings in the manufacturing sector, total employment, labour force participation, and real oil prices; for Japan—employment in the manufacturing sector, unit labour costs, labour force participation, and real oil prices.

The results of the test statistic and corresponding p-values are found in Table 2 along with the description of the instruments in detail. From here we can see that, except for Japanese exports of small passenger cars, exogeneity of the subset of tested variables is strongly rejected at the 5 per cent level.

5. IV estimations and Wald tests

Based on the above, it is therefore clear that, while we can be confident that there is pricing-to-market in Japanese exports of small passenger cars, the estimates for all the other cases may be subject to simultaneity bias. We therefore estimate our equations again, this time using instrumental variable methods, and where the same instrument sets as above are applied for this purpose. Testing the hypothesis of PTM is then equivalent to testing the significance of the sum of the coefficients on the real exchange rate changes in the regression equation (4).

Estimates of the PTM elasticity, the Wald test statistics, and their p-values are found in Table 3. For Japan, estimates are significantly different than zero and less than one, suggesting incomplete pass-through of exchange rate changes to Japanese export prices. As for Canada, estimates for the three categories of exports are also significant and above one.[12] For the United States, we notice that PTM elasticities of U.S. automobile and truck exports to Canada are no longer significant at the 10 per cent level, indicating that there is in fact no PTM behaviour in these

12. While remembering that the PTM elasticity is given by $\alpha_1 = 1 + \beta_1 - \beta_2$, and that this value could be above one for specific demand functions, we explicitly test in the next section whether PTM elasticities of Canadian exports are equal to one.

Table 2: Dufour exogeneity test results

	Industry group	Dufour exogeneity Hausman-type test results		
		Instrument A[a]	Instrument B[b]	Instrument C[c]
		F-stat (p-value)	F-stat (p-value).	F-stat (p-value)
JPUS	Passenger cars	3.6859 (0.0153)		
	Small passenger cars	0.2701 (0.8468)		
	Trucks	5.7893 (0.0012)		
CAUS	Automobiles		5.3805 (0.0015)	
	Trucks		2.7480 (0.0445)	
	Heavy trucks		7.6830 (0.0001)	
USCA	Automobiles			3.7354 (0.0122)
	Trucks			10.6324 (0.0000)
	Heavy trucks			3.7033 (0.0129)

a. Instrument set A includes second lags of each of these log difference Japanese variables: employment in manufacturing, seasonally-adjusted unit labour costs, labour force participation, and real oil prices.

b. Instrument set B includes second lags of each of these log difference Canadian variables: average hourly and fixed-weight average hourly earnings in all sectors, average weekly earnings in manufacturing, total employment, real oil prices. This instrument set also includes the third lag of the log difference in the real oil price.

c. Instrument set C includes second lags of each of these log difference U.S. variables: seasonally-adjusted average hourly earnings in all sectors, total employment, labour force participation, and real oil prices. This instrument set also includes the third lag of the log difference in average hourly earnings in the manufacturing sector.

industries—contrary to the OLS case. In addition, even U.S. exports of heavy trucks is not significant at the 5 per cent level. A rationalization for this can be found in Knetter (1993) who explains that U.S. exports are only a very small share of total U.S. auto sales so that little attention is paid to their pricing strategies as opposed to other countries whose auto exports represent a greater share of production.

So far, the evidence presented is largely consistent with the existing empirical literature on PTM. However, as emphasized earlier, these results must be qualified to account for probable weak-instruments effects.

6. LR-based tests

Hypothesis tests in IV-regressions—as applied above—have recently been the subject of renewed attention. Indeed, several studies have documented serious size distortions in the presence of possibly weak instruments. The performance of the Wald test, on which the above evidence is based, was noted to be particularly poor.[13] In particular, Dufour (1997) has shown analytically that the size of the IV-based Wald tests may deviate arbitrarily from their nominal level because it is impossible to bound the Wald test criterion by a pivotal (exact nuisance-parameter-free) quantity. In a separate study, Staiger and Stock (1997) derive the asymptotic distribution of the Wald statistic allowing for weak instruments. From these, it is similarly evident that the χ^2 approximations are indeed poor, and more importantly, that the limiting null distributions are nuisance-parameter-dependent, which highly complicates statistical inference. Finally, Dufour and Khalaf (1998) show that the bootstrap method fails in the case of the Wald criterion. Their result seems to suggest that more involved simulation-based procedures based on the Wald criterion will not achieve size-control successfully.

In contrast, econometric theory suggests that LIML-based LR tests do not suffer from such problems. Indeed, the above cited studies, among others, have shown that bounds or bootstrap-type procedures may be applied to obtain valid (size-correct) tests based on the LR-LIML criterion. Dufour (1997) demonstrates that the test statistic admits, under general possibly non-linear hypotheses, an exact nuisance-parameter-free bound, which is derived in the paper. Dufour and Khalaf (1998) show how to implement the Dufour (1997) bound using simulation-based methods. Furthermore, they describe a procedure to obtain tighter exact bounds and illustrate the latter using

13. See, for example, Dufour (1997), Staiger and Stock (1997), Wang and Zivot (1998), Dufour and Khalaf (1998), and several others.

the linear structural restrictions case.[14] Wang and Zivot (1998), for their part, study the LR-LIML tests for the hypothesis that sets the full vector of endogenous variable coefficients to specific values and derive an symptotic bounds-based cut-off point that is valid whether the rank condition holds or not. They show that whereas the standard critical point (which holds only imposing identification) is χ^2 with degrees of freedom equal to the dimension of the endogenous variable coefficients vector, the bounding critical value is χ^2 with degrees of freedom equal to the number of instruments instead (and is valid regardless of identification considerations). In addition, Dufour and Khalaf (1998) show that the same cut-off may also be used if a subset of the endogenous variable coefficients is tested. As in Wang and Zivot, this result does not depend on the rank condition. For both hypotheses, the bounds test is conservative in the sense that rejections are conclusive. However, the simulation results in Dufour and Khalaf (1999) suggest that Monte Carlo (MC) tests, including parametric bootstrap-type procedures, based on the LR-LIML statistic, perform better than the corresponding bounds-test (in terms of size-controlled power).

In view of the above evidence, we adopt testing strategies that are known to perform well irrespective of the quality of available instruments. We apply both MC and the Wang-Zivot bounds LIML tests. The underlying formulae and test strategies are presented in the Appendix. The MC test uses 999 replications and may be summarized as follows. Drawing from the DGP with constrained LIML estimates, $N = 999$ simulated samples are generated which yield $N = 999$ simulated test statistics. Then the MC p-value is obtained from the rank of the observed value of the test statistic within the set [*observed* statistic, *simulated* statistics]. In other words, at level $\alpha = 0.05$, a *rank* exceeding $(1-\alpha)(N+1) = 950$ is interpreted as evidence against H_0. Refer to the Appendix for a more detailed exposition.

The results are summarized in Table 3. For simplicity, we have reparameterized the model so that the null hypothesis corresponds to an exclusion constraint on an endogenous variable coefficients subset; the reparametrized model and relevant hypotheses are included in Table 3.

It is striking to see how these test results differ from the Wald tests. In particular, two cases out of the six examined show that, while the Wald test p-value is 0.001 or less, both the corresponding ALR and the MC-LR test p-values are above 0.1, indicating that the sum of coefficients is not significantly different than zero even at the 10 per cent level. Inference based on the size-correct test now yields that there is no evidence for PTM in the Japanese exports of passenger cars and Canadian exports of heavy trucks to the United States. Furthermore, the MC-LR

14. As emphasized in Dufour and Khalaf (1998), linear restrictions on structural coefficients are actually non-linear, given the non-linear relations which map the structural coefficients onto the reduced-form ones.

reverses the IV conclusions drawn for U.S. exports of heavy trucks to Canada, yielding no PTM in this case either. Finally, as mentioned before, we use the above LR tests to test if Canadian automobile and truck export PTM elasticities equal one. We find that we cannot reject the null in either case, meaning that these two elasticities are not statistically different than one.

It is worth noting, at this stage, that Wald and LR tests may yield conflicting evidence even when the standard chi-square cut-off points are applied, whether instruments are poor or not (see, for example, Davidson and MacKinnon [1993, 456–458]). Furthermore, when there is more than one right-hand-side endogenous variable, the "quality" of instruments is not testable, since the underlying constraint involves a rank condition; see equation (A.6). Indeed, it is evident that unless $g=1$, the hypothesis that (A.6) fails to hold is not amenable to testing. Consequently, the fact that the Chi-square Wald and the bootstrap LR-LIML do not yield the same decision should not necessarily be interpreted as evidence of poor instruments. Nevertheless, what we are emphasizing here is that, given that the former test is known to be spurious and the latter has been shown to perform well whether instruments are "good" or "poor," we would rather rely on the decision of the LIML test.

To summarize, we can conclude that (1) pricing-to-market is supported in the small passenger car sectors of Japan and Canada,[15] as well as in the truck exports of these two countries; (2) there is no evidence for PTM in Japanese exports of passenger cars and Canadian exports of heavy trucks; (3) there is no evidence of PTM in any of the transportation export categories examined for the United States.

We can now see that these conclusions lend support to the Feenstra, Gagnon, and Knetter (1996) study where, in a Bertrand differentiated products model, it is suggested that countries with a very large share of total destination market sales should exhibit export prices with high pass-through.[16] Thus, there is evidence of PTM in the automobile and truck industries for Japan and Canada, both of which have a small market share in the U.S. market for these transportation categories. In addition, our results also support the viewpoint in Knetter (1993), Campa and Goldberg (1999), and Cheung and Chinn (1999) who suggest that the industry effect is a very

15. To be more specific, the automobile sector for Canada

16. This follows from the intuition that, if the demand function for a particular variety of a good is homogeneous of degree zero in all prices, and if all varieties of this good are supplied from the same country, a change in the exchange rate will lead to an equi-proportional change in all prices. In this case, since the demand elasticity of the particular variety is a function of all prices, it will not change. Indeed, it can be seen from the first-order condition of the maximization problem that all of the exchange rate change should be reflected in the import price. This intuition therefore implies that, if one country provides most of the varieties of a good, pass-through on these exports should be high. See Feenstra, Gagnon, and Knetter (1996) for more details.

Table 3: IV estimation and Wald & LR test results

	Industry group	α_1						Degrees of freedom
		IV & Wald results			LR results[a]			
		Estimate	Wald stat	p-value	ALR[b] p-value	W&Z p-value	MC-LR p-value	
JPUS	Passenger cars	0.8932	10.8403	0.0010	0.4414	0.9883	0.5450	82
	Trucks	0.6264	7.7243	0.0054	0.0628	0.6291	0.0250	83
CAUS	Automobiles	1.4195	20.7235	0.0000	0.0320	0.7086	0.0150	173
	Trucks	1.4252	19.2098	0.0000	0.0276	0.6782	0.0130	176
	Heavy trucks	1.3509	21.0055	0.0000	0.1799	0.9702	0.1150	174
USCA	Automobiles	0.8398	0.8514	0.3562	0.5916	0.9996	0.7580	197
	Trucks	0.5777	0.1903	0.6621	0.7368	0.9999	0.8040	200
	Heavy trucks	2.6326	3.2334	0.0721	0.0832	0.8086	0.1970	199

a. The model is $\Delta x_t = c + \sum_{k=1}^{m} \alpha_{0k}(\Delta s_t - \Delta s_{t-k}) + \sum_{k=0}^{m} \alpha_{1k}\Delta r_{t-k} + \sum_{k=0}^{q} \alpha_{2k}\Delta\left(\frac{w_{t-k}}{p_{t-k}}\right) + \sum_{k=0}^{q} \alpha_{3k}\Delta y_{t-k} + \sum_{k=0}^{q} \alpha_{4k}\Delta\Omega_{t-k} + u_t$ where $\Delta\Omega_{t-k}$ is the vector containing the variables p^m_{t-k}/p_{t-k} and Δz_{t-k}, the endogenous variables are $\{\Delta r_t, \Delta w_t, \Delta y_t\}$, and the tested hypothesis is that $\sum_{k=0}^{m} \alpha_{1k} = 0$.

b. See notes on previous tables for definitions of the instrument sets and α_1. ALR denotes the asymptotic χ^2-based p-value. W&Z is the asymptotic bounds-based Wang & Zivot p-value. MC-LR is the Monte-Carlo LR p-value.

important factor in finding whether PTM is practised or not. Thus, we find that there is no PTM in the heavy truck industry, whether in Canada or the United States.

Finally, we should note two caveats in this study. One is that the model examined above considers all unanticipated changes in the exchange rate as temporary and all expected changes as permanent. However, firms' pricing strategies might be different if perceptions of temporary and permanent shocks to the exchange rate were different.[17] The other caveat is that, as we mentioned before, in the case of the transportation vehicle trade between Canada and the United States, we have abstracted from transfer pricing issues that may exist in the data because of the Canada-U.S. auto pact.[18] For both these reasons, we do not spend much time interpreting parameter estimate values, but instead concentrate on the testing strategies in line with our objectives as emphasized in the introduction.

7. Conclusion

Using tests that do not rely on the quality of instruments used, our study has shown that only four of the nine cases examined yielded evidence in favour of pricing-to-market in the transportation industry for three countries. Indeed, in three cases, our LR-based tests reversed the results associated with traditional Wald tests. Yet most of the published studies in this context use such tests, which casts doubts on the reliability of the partial-equilibrium-based evidence in favour of PTM.

The fact that we find, in the transportation industry, less evidence for PTM than thought previously may be due to many reasons. One is that, in its present form, the structural simultaneous equations model underlying the tested equations may contain insufficient information on pricing-to-market. In this case, a general-equilibrium model may be a more suitable framework within which to examine this question. Whether such an approach would prove more reliable, in general, is of course an open question. Another explanation, as stressed by Obstfeld and Rogoff, is that the assumption of local currency price stability in these models may not be entirely valid for countries other than the United States. Finally, the results might have been different in an error-correction-type setting where perceptions of permanent versus temporary exchange rate changes are treated in another way than in the present model.

17. This issue has already been addressed by a number of recent error-correction models of PTM.
18. In this context, the profit maximization problem and intra-firm pricing strategies are probably also influenced by differing taxation policies in the two countries. This is an aspect that we have abstracted from in this paper.

Notwithstanding the issues surrounding the choice of the proper model, we have demonstrated that relying solely on Wald tests indeed presents a danger of drawing the wrong conclusions from the hypothesis testing exercise and that the existing evidence on PTM needs to be revisited.

Appendix 1

Description of data

For the case of Japanese exports to the United States, we use the same data as does Marston for the domestic and export prices of each of the three categories in the transportation industry,[19] that is, we use monthly domestic wholesale and export prices respectively from the Bank of Japan. The Canadian monthly automobile domestic and export prices data are from the Statistics Canada Industrial Product Price Indexes data base while those for the United States are taken from the BLS Production Price Indexes. From the IMF we obtain monthly wholesale or consumer price indexes for the different countries, which are used to deflate nominal variables. We also obtain monthly Japanese nominal wages from the same source. From the OECD, we obtain monthly industrial production indexes for all countries, which we use to represent real outputs. Canadian average hourly earnings for motor vehicles and trucks, bus bodies and trailers, and other variables of the labour market used as instruments are taken from Statistics Canada while, for the United States, wages in transportation and variables used as instruments are taken from the BLS. From the BIS, we obtain all the nominal exchange rates except for the monthly Canada-U.S. nominal exchange rate, which is taken from the Bank of Canada data base instead. The real price of raw materials for Canada is the monthly industrial commodity price index from Statistics Canada deflated by Canadian prices. For the United States and Japan, we use the U.S. commodity price index of metals and steel obtained from the BLS and deflated using U.S. prices. Finally, the price of oil is that of monthly West Texas intermediate crude oil in USD.

19. We are grateful to Richard Marston for kindly supplying this data.

Appendix 2

Description of the LIML estimator and the LR-based tests

Consider the IV regression using standard notation

$$y = Y\beta + X_1\gamma_1 + u = Z\delta + u \tag{A.1}$$

$$Y = X_1\Pi_1 + X_2\Pi_2 + V \tag{A.2}$$

$$Z = \begin{bmatrix} Y & X_1 \end{bmatrix}, \qquad \delta = \begin{bmatrix} \beta \\ \gamma_1 \end{bmatrix},$$

where y is $(Tx1)$, Y is (Txg) and denotes the right-hand-side included endogenous variable, X_1 is (Txk_1) and denotes the included exogeneous variables, and X_2 refers to the (Txk_2) matrix of available additional instruments. Equation (A.2) is the reduced form associated with the included right-hand-side endogenous variables. The associated *Limited Information* reduced form is

$$\begin{bmatrix} y & Y \end{bmatrix} = \begin{bmatrix} \pi_1 & \Pi_1 \\ \pi_2 & \Pi_2 \end{bmatrix} \begin{bmatrix} X_1 & X_2 \end{bmatrix} + \begin{bmatrix} v & V \end{bmatrix}, \tag{A.3}$$

where

$$\pi_1 = \Pi_1\beta + \gamma_1, \tag{A.4}$$

$$\pi_2 = \Pi_2\beta . \tag{A.5}$$

The necessary and sufficient condition for identification follows from the relation (A.5). Indeed, β is recoverable if and only if

$$rank(\Pi_2) = g . \tag{A.6}$$

In this context, the two-stage least squares (2SLS) structural coefficient estimator is

$$\begin{aligned} \hat{\delta} &= [Z'P(P'P)^{-1}P'Z]^{-1}Z'P(P'P)^{-1}P'y \\ P &= \left[X \;\; X(X'X)^{-1}X'Y\right] \\ X &= \left[X_1 \;\; X_2\right] \end{aligned} \quad . \tag{A.7}$$

Wald-type tests for linear restrictions on structural parameters are routinely associated with 2SLS estimation. For hypotheses of the form $R\delta = r$, where R is a known (qxg) matrix of rank q and r consists of known constants, the 2SLS-based Wald test statistic is

$$\begin{aligned} r_W &= \frac{1}{s^2}(r - R\hat{\delta})' - [R'(ZP(P'P)^{-1}P'Z)^{-1}R](r - R\hat{\delta}), \\ s^2 &= \frac{1}{T}(y - Z\hat{\delta})(y - Z\hat{\delta})' \end{aligned} \quad . \tag{A.8}$$

Imposing identification, the asymptotic null distribution of r_W is $\chi^2(q)$. In near-identified models, i.e., in poor-instruments situations, this distribution fails to hold, so that test procedures based on $\chi^2(q)$ cut-off points seriously over-reject. For an asymptotic theory conformable with under-identification, see, for example, Staiger and Stock (1997).

Imposing (A.4) and (A.5), LIML corresponds to maximizing the likelihood function associated with the equations (A.1) and (A.2)

$$\begin{aligned} l(y, Y|X_1, X_2) &= -\frac{T(g+1)}{2}\ln(2\pi) - \frac{T}{2}\ln|\Sigma| - \frac{1}{2}tr\Sigma^{-1}D'D, \\ D &= \left[y \;\; Y\right] - X\begin{bmatrix} \pi_1 & \Pi_1 \\ \pi_2 & \Pi_2 \end{bmatrix} \end{aligned} \tag{A.9}$$

where Σ denotes the error covariance. Numerical maximization may be considered, yet it is well known that an equivalent solution obtains through an eigenvalue/eigenvector problem based on the following determinental equation:

$$\left|\left[y \;\; Y\right]'M_1\left[y \;\; Y\right] - \lambda\left[y \;\; Y\right]'M\left[y \;\; Y\right]\right| = 0 \tag{A.10}$$

where

$$M = I - X(X'X)^{-1}X', \\ M_1 = I - X_1(X_1'X_1)^{-1}X_1', \tag{A.11}$$

and λ refers to the eigenvalue in question. Indeed, it can be shown by concentrating the likelihood function that the LIML problem corresponds to minimizing, with respect to β,

$$\lambda(\beta) = \frac{\begin{bmatrix} 1 \\ -\beta \end{bmatrix}' \begin{bmatrix} y & Y \end{bmatrix}' M_1 \begin{bmatrix} y & Y \end{bmatrix} \begin{bmatrix} 1 \\ -\beta \end{bmatrix}}{\begin{bmatrix} 1 \\ -\beta \end{bmatrix}' \begin{bmatrix} y & Y \end{bmatrix}' M \begin{bmatrix} y & Y \end{bmatrix} \begin{bmatrix} 1 \\ -\beta \end{bmatrix}}. \tag{A.12}$$

Formally, the LIML estimator of β and γ_1 is

$$\begin{bmatrix} \tilde{\beta} \\ \tilde{\gamma}_1 \end{bmatrix} = \begin{bmatrix} (Y'Y - \tilde{\lambda}Y'MY) & Y'X \\ X'Y & X'X \end{bmatrix}^{-1} \begin{bmatrix} Y' - \tilde{\lambda}Y'M \\ X' \end{bmatrix} \tag{A.13}$$

where $\tilde{\lambda}$ is the smallest root of (A.10), which corresponds to $\lambda(\tilde{\beta})$. Furthermore, the LIML error covariance estimate is

$$\tilde{\Sigma} = \frac{\begin{bmatrix} y & Y \end{bmatrix}' M \begin{bmatrix} y & Y \end{bmatrix}}{T} + \frac{(\tilde{\lambda} - 1)S}{T} \\ S = \frac{\begin{bmatrix} y & Y \end{bmatrix}' M \begin{bmatrix} y & Y \end{bmatrix} \begin{bmatrix} 1 \\ -\tilde{\beta} \end{bmatrix} \left(\begin{bmatrix} y & Y \end{bmatrix}' M \begin{bmatrix} y & Y \end{bmatrix} \begin{bmatrix} 1 \\ -\tilde{\beta} \end{bmatrix} \right)'}{\left(\begin{bmatrix} 1 \\ -\tilde{\beta} \end{bmatrix}' \begin{bmatrix} y & Y \end{bmatrix}' M \begin{bmatrix} y & Y \end{bmatrix} \begin{bmatrix} 1 \\ -\tilde{\beta} \end{bmatrix} \right)}. \tag{A.14}$$

For detailed proofs, see, for example, Davidson and MacKinnon (1993, chapter 18).

The LIML-based LR test statistic is

$$LR = T(\ln|\tilde{\Sigma}^0| - \ln|\tilde{\Sigma}|) = T(\ln|\tilde{\lambda}^0| - \ln|\tilde{\lambda}|) \tag{A.15}$$

where $\tilde{\Sigma}^0$ is the constrained covariance estimate and $\tilde{\lambda}^0$ is the constrained minimum of (A.10). In this paper, we test hypotheses of the form

$$\beta_{(1)} = \beta^0_{(1)} \tag{A.16}$$

which corresponds to the following partitioning of β,

$$\beta = \begin{bmatrix} \beta_{(1)} \\ \beta_{(2)} \end{bmatrix}, \tag{A.17}$$

and where $\beta^0_{(1)}$ is a known constant. Wang and Zivot (1998) studied the latter test given hypotheses of the form $\beta = \beta^0$, where β^0 is a known constant.[20]

It is easy to see from the above LIML formulae that constrained estimation under the latter hypothesis is trivial. To obtain the constrained estimates under (A.16), partition Y into $\begin{bmatrix} Y_{(1)} & Y_{(2)} \end{bmatrix}$ conformably with the null hypothesis and apply the unconstrained LIML technique to a reparameterized equation where the left-hand-side endogeneous variable is $(y - Y_{(1)}\beta^0_{(1)})$ and the right-hand-side included endogenous variable is $Y_{(2)}$. Alternatively, the likelihood may be maximized numerically under (A.4)-(A.5) and the tested constraint (A.16).

As is well known, the asymptotic null distribution of LR is $\chi^2(g)$, yet this limiting distribution depends on the rank condition for identification. Wang and Zivot show that under hypotheses of the form (A.13), the $\chi^2(k_2)$ provides a bounding limiting distribution for the LR criterion, whether (A.5) holds or not. Furthermore, Dufour and Khalaf (1998) show that the Wang-

20. Indeed, it is straightforward to see that the model at hand may be reparameterized so that the tested hypothesis corresponds to (A.16). For instance, for the model presented in the footnotes of Table 3, given $m = 2$ and $q = 1$, we reparameterize the model as:

$$\Delta x_t = \begin{array}{c} c + \sum_{k=1}^{2} \alpha_{0k}(\Delta s_t - \Delta s_{t-k}) + \alpha_{10}(\Delta r_t - \Delta r_{t-1}) + \Phi_1 \Delta r_{t-1} + \alpha_{12}(\Delta r_{t-2} - \Delta r_{t-1}) + \sum_{k=0}^{1} \alpha_{2k}\Delta\left(\frac{w_{t-k}}{p_{t-k}}\right) + \\ \sum_{k=0}^{1} \alpha_{3k}\Delta y_{t-k} + \sum_{k=0}^{q} \alpha_{4k}\Delta\Omega_{t-k} + u_t \end{array}$$

so that testing the aforementioned hypothesis amounts to testing the exclusion restriction that $\Phi_1 = 0$.

Zivot $\chi^2(k_2)$ bound also holds asymptotically under (A.16). We next show how to derive a bootstrap-type MC p-value that we will denote $\hat{p}_N$, where N refers to the number of MC replications. Our presentation here is only descriptive; see Dufour and Khalaf (1999) for further discussion on MC tests in simultaneous equations and related references.

Let LR_0 refer to the value of the test statistic obtained from the data.

- Draw N samples from the null DGP, imposing normality and using (A.3) and the constrained LIML sample-based estimates.
- From each simulated sample, compute the LR statistic $LR_j, j = 1, \ldots, N$.
- Given LR_0 and $LR_j, j = 1, \ldots, N$, obtain

$$\hat{G}_N = \frac{1}{N}\sum_{i=1}^{N} I_{A[0,\infty]}(LR_i - LR_0), \qquad I_A(x) = \begin{cases} 1, & x \in A \\ 0, & x \notin A \end{cases} . \tag{A.18}$$

In other words, $N\hat{G}_N$ is the number of simulated criteria $\geq LR_0$ and $\hat{R}_N = N - N\hat{G}_N + 1$ is the rank of LR_0 in the series $LR_0, LR_1, \ldots, LR_N$.

- Then, the MC p-value is

$$\hat{p}_N = \frac{N\hat{G}_N + 1}{N + 1}. \tag{A.19}$$

Dufour (1995) proves that, given general regularity conditions, the test based on $\hat{p}_N$ has the correct size asymptotically (as $T \to \infty$). Note that the number N of MC replications is not required to tend to infinity for the latter result to hold; this is the fundamental difference between the technique described here and the (closely-related) parametric bootstrap.

Description of the test for exogeneity

We finally turn to exogeneity test. In this paper, we have focused on test methods that have been proven to be immune to the weak-instrument problem. In view of this, we have considered an exogeneity test statistic that is highly related to the well-known Hausman test, yet whose null distribution does not depend on the rank identification condition. The statistic is proposed in Dufour (1987) and may be derived as follows:

Obtain $\hat{V}$, the residuals from the regression of Y on X_1 and X_2. Then, the exogeneity test corresponds to the standard F-test for the exclusion of $\hat{V}$ in the (augmented) regression of y on Y, X_1, and $\hat{V}$. The test's null distribution is $F(g, T-k_1-k_2)$. Here we emphasize that this distributional result is exact and does not depend on the rank identification constraint. In other words, the F-based cut-off point is valid, irrespective of the "quality" of available instruments.

References

Ball, L. 1998. “Policy Rules for Open Economies.” NBER Working Paper No. 6760.

Bergin, P. and R.C. Feenstra. 1999. "Pricing-to-Market, Staggered Contracts and Real Exchange Rate Persistence." NBER Working Paper No. 7026.

Betts, C. and M. Devereux. 1996. "The Exchange Rate in a Model of Pricing-to-Market." *European Economic Review* 40: 1007–1021.

———. 1998. "Exchange Rate Dynamics in a Model of Pricing-to-Market." *Journal of International Economics,* forthcoming.

———. 2000. "The International Effects of Monetary and Fiscal Policy in a Two-Country Model." In *Money, Capital Mobility and Trade: Essays in Honor of Robert A. Mundell,* edited by G.A. Calvo, R. Dornbusch, and M. Obstfeld. Cambridge, MA: MIT Press.

Campa, J. M. and L. Goldberg. 1999. “Investment, Pass-Through and Exchange Rates: A Cross-Country Comparison.” *International Economic Review* 40: 287–314.

Cheung, Y.W. and M. Chinn. 1999. “Market Structure and the Persistence of Sectoral Real Exchange Rates.” NBER Working Paper No. 7408.

Davidson, R. and J.G. MacKinnon. 1993. *Estimation and Inference in Econometrics.* New York: Oxford University Press.

Dufour, J.-M. 1987. "Linear Wald Methods for Inference on Covariance Matrices and Weak Exogeneity Tests in Structural Equations." In *Time Series and Econometric Modelling,* edited by I.B. MacNeil and G.J. Umphrey, 317–338. D. Reidel Publishing.

———. 1995. "Monte Carlo tests in the presence of nuisance parameters, with econometric applications." C.R.D.E., Université de Montréal, Working Paper.

———. 1997. “Some Impossibility Theorems in Econometrics, with Applications to Structural and Dynamic Models.” *Econometrica* 65: 1365–1389.

Dufour J.-M. and L. Khalaf. 1998. "Simulation based finite and large sample inference methods in simultaneous equations." C.R.D.E., Université de Montréal, Working Paper.

———. 1999. "Monte Carlo tests methods in econometrics.” In *Companion to Theoretical Econometrics,* edited by B. Baltagi. Blackwell, forthcoming.

Faruqee, H. 1995. “Pricing to Market and the Real Exchange Rate.” *IMF Staff Papers* 42: 855–881.

Feenstra, R.C. 1989. "Symmetric Pass-Through of Tarriffs and Exchange Rates under Imperfect Competition: An Empirical Test." *Journal of International Economics* 27: 25–45.

Feenstra, R.C., J. Gagnon, and M. Knetter. 1996. "Market Share and Exchange Rate Pass-Through in World Automobile Trade." *Journal of International Economics* 40: 187–207.

Feenstra, R.C. and J.D. Kendall. 1997. “Pass-through of exchange rates and purchasing power parity.” *Journal of International Economics* 43: 237–261.

Gagnon, J. and M. Knetter. 1995. “Markup Adjustment and Exchange Rate Fluctuations: Evidence from Panel Data on Automobile Exports.” *Journal of International Money and Finance* 14: 289–310.

Giovannini, A. 1988. “Exchange Rates and Traded Goods Prices.” *Journal of International Economics* 24: 45–68.

Goldberg, P.K. and M. Knetter. 1997. "Goods Prices and Exchange Rates: What Have We Learned?" *Journal of Economic Literature* 35: 1243–1272.

Hooper, P. and C. Mann. 1989. "Exchange Rate Pass-Through in the 1980s: The Case of U.S. Imports of Manufacturers." *Brookings Papers on Economic Activity* 0(1): 297-329.

Kasa, K. 1992. "Adjustment Costs and Pricing-to-Market: Theory and Evidence." *Journal of International Economics* 32: 1–30.

Klitgaard, T. 1999. "Exchange Rates and Profit Margins: The Case of Japanese Exporters." *Federal Reserve Bank of New York Economic Policy Review* 5: 41–54.

Knetter, M. 1993. “International Comparisons of Price-to-Market Behavior.” *American Economic Review* 83: 473–486.

Marston, R. 1990. "Pricing To Market in Japanese Manufacturing." *Journal of International Economics* 29: 217–236.

Obstfeld, M. and K. Rogoff. 1999. “New Directions for Stochastic Open Economy Models.” NBER Working Paper No. 7313.

Staiger, D. and J.H. Stock. 1997. "Instrumental Variables Regression with Weak Instruments." *Econometrica* 65: 557–586.

Wang, J. and E. Zivot. 1998. "Inference on Structural Parameters in Instrumental Variables Regression with Weak Instruments." *Econometrica* 66: 1389–1404.

Bank of Canada Working Papers
Documents de travail de la Banque du Canada

Working papers are generally published in the language of the author, with an abstract in both official languages. ***Les documents de travail sont publiés généralement dans la langue utilisée par les auteurs; ils sont cependant précédés d'un résumé bilingue.***

2000

2000-7	Non-Parametric and Neural Network Models of Inflation Changes	G. Tkacz
2000-6	Some Explorations, Using Canadian Data, of the *S*-Variable in Akerlof, Dickens, and Perry (1996)	S. Hogan and L. Pichette
2000-5	Estimating the Fractional Order of Integration of Interest Rates Using a Wavelet OLS Estimator	G. Tkacz
2000-4	Quelques résultats empiriques relatifs à l'évolution du taux de change Canada/États-Unis	R. Djoudad and D. Tessier
2000-3	Long-Term Determinants of the Personal Savings Rate: Literature Review and Some Empirical Results for Canada	G. Bérubé and D. Côté
2000-2	GAUSS™ Programs for the Estimation of State-Space Models with ARCH Errors: A User's Guide	M. Kichian
2000-1	The Employment Costs of Downward Nominal-Wage Rigidity	J. Farès and S. Hogan

1999

99-20	The Expectations Hypothesis for the Longer End of the Term Structure: Some Evidence for Canada	R. Lange
99-19	Pricing Interest Rate Derivatives in a Non-Parametric Two-Factor Term-Structure Model	J. Knight, F. Li, and M. Yuan
99-18	Estimating One-Factor Models of Short-Term Interest Rates	D. Mc Manus and D. Watt
99-17	Canada's Exchange Rate Regime and North American Economic Integration: The Role of Risk-Sharing Mechanisms	Z. Antia, R. Djoudad, and P. St-Amant
99-16	Optimal Currency Areas: A Review of the Recent Literature	R. Lafrance and P. St-Amant
99-15	The Information Content of Interest Rate Futures Options	D. Mc Manus
99-14	The U.S. Capacity Utilization Rate: A New Estimation Approach	R. Lalonde
99-13	Indicator Models of Core Inflation for Canada	R. Dion
99-12	Why Canada Needs a Flexible Exchange Rate	J. Murray
99-11	Liquidity of the Government of Canada Securities Market: Stylized Facts and Some Market Microstructure Comparisons to the United States	T. Gravelle
99-10	Real Effects of Collapsing Exchange Rate Regimes: An Application to Mexico	P. Osakwe and L. Schembri
99-9	Measuring Potential Output within a State-Space Framework	M. Kichian

Copies and a complete list of working papers are available from:
Pour obtenir des exemplaires et une liste complète des documents de travail, prière de s'adresser à:

Publications Distribution, Bank of Canada
234 Wellington Street Ottawa, Ontario K1A 0G9

Diffusion des publications, Banque du Canada
234, rue Wellington, Ottawa (Ontario) K1A 0G9

E-mail / Adresse électronique: publications@bank-banque-canada.ca
WWW: http://www.bank-banque-canada.ca/