Is There a Quantity Puzzle Within Countries? An Application Using US and Canadian Data

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Introduction

The "quantity puzzle" is a theoretical anomaly, because the presence of trade in financial assets should be motivated by insurance motives, which, in turn, should be reflected in quantities. Regions with particularly idiosyncratic local production patterns offer especially advantageous risk-sharing possibilities to each other and, therefore, have a particular incentive to consume out of an identical portfolio. In other words, consumption plans should be highly correlated between regions where fluctuations in local production are dissimilar, controlling for the possibility that capital flows be constrained institutionally. As is well known, the ranking in international data goes the opposite way.

In this paper, intranational data are brought to the issue. We ask whether fluctuations in disposable income (or retail sales) and output between US states and Canadian provinces give rise to a similar puzzle. They do not. We find that pairwise correlations in disposable income (or sales) are higher on

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average than output correlations in the universe formed by the states and provinces of the United States and Canada. Most importantly, within countries, bilateral risk sharing tends to be associated with low output correlations, as theory suggests it should. This stands in stark contrast with cross-country evidence, where fluctuations in GDP are high between financially integrated economies. This suggests that the positive association between financial linkages and co-movements in GDP is the reason for the quantity puzzle, as documented in Imbs (2004b), and it is specific to international data.

The jump from international to intranational data is not straightforward. First, states or provinces within a federation benefit from federal transfers, which may aim at achieving some form of income insurance. As a result, local disposable income should be measured net of (positive or negative) transfers if it is to be indicative of the local propensity to engage in risk sharing. Second, regional output fluctuations embed the influence of both aggregate and local shocks. But financial flows within countries can aim only at insuring against region-specific shocks, irrespective of aggregate developments. The extent of local risk sharing is measured equally well whether or not aggregate shocks are controlled for.

Third, the quantity puzzle pertains to aggregate consumption, but tests of risk sharing in local data are typically based on disposable income. Unlike Canada, the United States does not report consumption at the state level, but only net disposable income. This complicates the comparison between the two levels of aggregation. Indeed, disposable income embeds savings rates, which could push upwards pairwise correlations in local disposable income. But it does not affect the responsiveness of output correlations to measured risk sharing, which appears to be consistent with theory within the United States and Canada, but not, more generally, across countries. In addition, we use information on retail sales at the state level as an alternative to disposable income in the United States.

While there are several available measures of international financial integration, the equivalent within countries is virtually uncharted territory. This paper extends standard measures of local risk sharing to a bilateral context. In particular, Kalemli-Ozcan, Sørensen, and Yosha (2001, 2003) and Asdrubali, Sørensen, and Yosha (1996) measured local risk sharing by the extent of income insurance, captured by the responsiveness of local disposable income to local production. Here, we measure the responsiveness of pairwise differences in disposable income to pairwise output gaps. Between financially integrated regions, the difference in disposable income should be independent of the realization of shocks to output in both regions. Between autarkic regions, on the other hand, discrepancies in the income available for consumption are fully accounted for by differences in realized output. We implement this proxy to both international and intranational data, and show that the results across countries are similar to those implied by effectively observed measures of financial integration.

The paper follows the empirical methodology spelled out in Imbs (2004a, 2004b) and borrows from the extensive literature investigating the determinants of business cycle co-movements, on the one hand, while seeking to provide empirical estimates of the extent of risk sharing, on the other. It also leaves open the possibilities that trade, finance, and specialization are interrelated in ways that can affect the observed correlations in quantities. Section 1 details the data and measurement strategy. Estimations and results are presented in section 2, and the final section concludes.

1 Data and Measurement

1.1 International data

The international data used here are relatively standard and purport to reproduce the results documented in Imbs (2004b). Pairwise correlations in output and consumption are computed on the basis of yearly real GDP and aggregate consumption as reported in Version 6.1 of the Penn World Tables. Financial integration is measured as the pairwise sum of bilateral asset holdings reported in the Coordinated Portfolio Investment Survey (CPIS) conducted by the International Monetary Fund (IMF) in 2001. As an alternative, the paper also uses the restrictions indexes published in the IMF's Annual Report on Exchange Arrangements and Exchange Restrictions. They are summed pairwise, and the proxy reports the average number of countries with restrictions to financial flows, for each country pair.

The determinants of business cycle co-movements are standard. Bilateral trade flows are taken from the IMF's Direction of Trade Statistics database, and computed in an intensive form following Deardorff (1998), Clark and van Wincoop (2001), and Imbs (2004a, 2004b):

$$T_{ij} = \frac{1}{T}\sum_t \frac{(EX_{i,j,t} + IM_{i,j,t}) \ . \ NYW_t}{NY_{i,t} \ . \ NY_{j,t}}.$$

$EX_{i,j,t}(IM_{i,j,t})$ denotes total merchandise exports (imports) from country *i* to *j* in year *t*, NY_i denotes nominal GDP in country *i*, and NYW is world nominal output. Following Clark and van Wincoop (2001) and Imbs (2004a, 2004b), sectoral real value-added data are used to compute an index S of the similarity in sectoral structure:

$$S_{ij} = \frac{1}{T}\sum_{t}\sum_{n}^{N}\left|s_{n,i} - s_{n,j}\right|,$$

where $s_{n,i}$ denotes the GDP share of industry n in country i. S_{ij} is the time average of the discrepancies in economic structures of countries i and j, and reaches its minimal value zero for two identical countries.[1] The sectoral shares s are computed using one-digit value-added data covering all sectors of the economy, from the United Nations Statistical Yearbook. Combining all data sources and constraints gives rise to a database comprising a maximum of 63 countries, or 1,953 pairwise combinations. All details on the international data can be found in Imbs (2004b).

1.2 Intranational data

This paper aims at asking from intranational data the same question Imbs (2004b) asked from international data, with focus on the impact of financial integration on output correlations. It is, therefore, important that the estimation in this paper be as close as possible to the standard international approach, with the implied data requirements. Most of the data on US states are taken from the Bureau of Economic Analysis (BEA). Yearly series on real gross state product (GSP) run from 1986 to 2000, while yearly data on state disposable income, net of taxes, run from 1969 to 2001. The data on disposable income are deflated using the same state-level deflator used to obtain real GSP series. The BEA also reports time series on total transfer receipts at the state level, from 1969 to 2001. These include retirement insurance; medical payments; family assistance, such as food stamps; and unemployment insurance. Local disposable income is measured net of taxes and all personal transfers. This residual is taken to reflect decentralized risk-sharing decisions.

There are issues of comparability. State-level information on consumption is not available for the United States, although it exists for the ten Canadian provinces and one territory. Unfortunately, such reduced coverage remains insufficient to draw general (or even significant) conclusions using intranational data. As a result, most existing studies have focused on disposable

1. Both the trade and specialization measures are based on time averages. Results do not change if initial values are used instead.

income data.[2] We do so here, as well, but we also use disaggregated measures of retail sales as an alternative measure of consumption behaviour.

None of this is an issue for the purpose of estimating the extent of local income insurance, but pairwise correlations in disposable income are not immediately comparable with pairwise correlations in aggregate consumption.[3] International correlations in net national income (NNI) would be the aggregate equivalent, but these, in turn, cannot easily relate to the quantity puzzle and are of less interest as a result. Furthermore, NNI is available for a relatively limited sample of countries, and the approach would truncate sizably the international dimension of the exercise. As a result, this paper compares the pairwise correlations in disposable income (or retails sales) and in local production as a test of the quantity puzzle in disaggregated data. How this affects the generality of the conclusions is discussed later.

Interstate trade flows are taken from the Commodity Flow Survey conducted by the US Bureau of Transportation in 1993 and 1997. The data are used to compute T_{ij} according to the definition used in the international context. The two observed years are averaged, but results are unchanged if using either one in isolation. State-level patterns of production are reported by the BEA from 1986 to 2000. Information is available for 62 sectors, and it is used to compute S_{ij} for each state. Population by state and other gravity variables used as instruments are also available from the BEA, or otherwise obtained from public sources.[4]

Most of the information on Canadian provinces comes from Statistics Canada's CANSIM database.[5] Series on real gross product per province run from 1984 to 2003. Real consumption is available from 1984 to 1996, and is deflated with the same index as real output. Interprovincial trade flows are available from 1981 to 2002, and are used to compute T_{ij} as an average over the whole period. S_{ij}, in turn, is computed using sectoral data specific to each province, available at the equivalent of a two-digit level of aggregation between 1984 and 2000. Results using initial values instead for both T_{ij} and S_{ij} are virtually identical. As for the US data, distance and other gravity variables are obtained from public sources.

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2. See, for instance, Kalemli-Ozcan, Sørensen, and Yosha (2003).
3. They should, however, be equal under perfect income insurance.
4. For instance, the bilateral distances between state capitals is readily available on the Internet.
5. The sample includes: Newfoundland, Prince Edward Island, Nova Scotia, New Brunswick, Quebec, Ontario, Manitoba, Saskatchewan, Alberta, British Columbia, and the Northwest Territories. Nunavut is not included because of insufficient time coverage.

Of prominent interest is a measure of financial integration between economic regions in the same country, which is not directly observable. Kalemli-Ozcan, Sørensen, and Yosha (2001, 2003) and Asdrubali, Sørensen, and Yosha (1996) introduced a proxy based on the responsiveness of local income to local production. In particular, for each US state, they estimated

$$y_{it} - d_{it} = \alpha + \beta y_{it} + \varepsilon_{it}, \tag{1}$$

where y_{it} denotes (the cyclical component of) output in region *i*, and d_{it} is (the cyclical component in) local net disposable income. β is an index of risk sharing in region *i*. As financial openness unhinges income from local production, local disposable income becomes idiosyncratic and unrelated to y_{it}, and β equals unity. At the other extreme, financial autarky means that local income is fully determined by local output, the dependent variable in equation (1) becomes white noise, and $\beta = 0$.

β provides an estimate of overall risk sharing in region *i*, but cannot be used directly in a bilateral context. The pairwise sum of β, for instance, is a noisy—and worse, potentially misleading—proxy for bilateral financial integration. Alaska and Texas could, for example, engage in financial transactions meant to limit income fluctuations in both states, whose production is specialized and thus potentially volatile. So, $\beta_i + \beta_j$ will tend to be high between regions that might be sharing risk with different third parties, rather than with each other. Indeed, effective risk sharing between the two states might well be virtually nonexistent, since they provide poor insurance possibilities for each other, as they both specialize in oil industries and may have highly correlated GSP as a result.

This paper follows the same intuition, using observed fluctuations in output and disposable income (or retail sales) to measure the extent of financial integration. But the estimation in equation (1) is extended to make it applicable to a bilateral context. In particular, we estimate

$$\Delta y_{ijt} - \Delta d_{ijt} = \alpha + \beta \ \Delta y_{ijt} + \varepsilon_{ijt}, \tag{2}$$

where Δx_{ijt} denotes differences in x_t between regions *i* and *j*. According to equation (2), capital flows from a region with high relative output into one with low relative output, with the purpose that fluctuations in disposable income (i.e., consumption) be equated across regions. Whether this is possible is measured by β, which captures the extent to which Δd_{ijt} is white noise in equation (2). Now, the possibility that Alaska and Texas choose to insure income with third parties rather than with each other is accounted for through the presence of the output-gap term, Δy_{ijt}. Estimates of β are now low, as they should be, between regions that offer little insurance motive to each other (y_{it} and y_{jt} move in lockstep), even when measured fluctuations in income are similar because of risk sharing with third parties (say, New York). The validity of the approach in equation (2) is verified in the international data set, comparing results implied by estimates of β to those implied using directly observable measures of bilateral financial integration.

Estimates for β in equation (2) will tend to be high if Δd_{it} tends to remain constant over time and Δy_{it} does not, i.e., when risk sharing as measured by relative variations in disposable income is also warranted on grounds of measured variations in relative output. In other words, estimates for β tend to be high if consumption correlations are high and output correlations are low, which potentially creates a positive bias between β and pairwise correlations in consumption. That makes it impossible to use this paper's approach to investigate the effect of financial integration on consumption plans between regions within one country, where β is the only available proxy. But Imbs (2004b) showed that, across countries, the main theoretical anomaly pertains to the association between output correlations and financial integration, negative in theory yet positive in international data. As Lewis (1996) already showed, consumption plans are detached from available production to an extent that increases with financial flows, as predicted by theory, and their behaviour is therefore less of a puzzle.

This paper focuses on the correlation between estimates of β in equation (2) and bilateral output correlations. *Ceteris paribus*, β tends to be low between regions with correlated output (since they provide little opportunity for risk sharing), which potentially induces a negative bias in the relation this paper seeks to evaluate. But, as will become clearer, the paper stresses the difference between results at the international and intranational levels. That discrepancy cannot be ascribed to an attenuating bias that presumably prevails equally in both data sets.

2 Estimations and Results

This section introduces the paper's main estimations, inspired largely from Imbs (2004b), and presents the main results. We first report simple results pertaining to output correlations, then complicate the approach to allow for interrelations between the determinants of business cycle correlations. The section closes with extensions and robustness checks.

Table 1 summarizes the statistics of interest in both data sets. In particular, it reports the moments of pairwise correlations in output, consumption, disposable income, and retail sales, as well as the main characteristics of the estimates of β in international and intranational data. Several comments are in order. First, β is, on average, larger within than between countries,

Table 1
Summary statistics

	International data			Intranational data		
	Mean	**Min.-Max.**	**Obs.**	**Mean**	**Min.-Max.**	**Obs.**
Output correlations	0.0822 0.215	–0.566/0.782	1,770	0.2104 0.339	–0.816/0.952	1,341
Consumption correlations	0.0292 0.198	–0.539/0.847	1,770	0.4327 0.308	–0.720/0.938	1,330
Sales correlations				0.2733 0.355	–0.813/0.981	1,330
Risk-sharing index	0.1487 0.278	0/1.575	1,830	0.4418 0.423	0/2.029	1,330
Risk-sharing index (sales)				0.6421 0.973	0/6.582	1,330

Notes: All variables are computed using the cyclical components of output and consumption as implied by the Baxter and King (1999) filter, with adequate parameters for yearly data. Real series for intranational data are obtained using regional GSP deflators. Consumption is net of taxes and transfers. “Mean” reports both the cross-sectional average and the corresponding standard error.

regardless of whether it is estimated using income or sales data. This is consistent with findings in Kalemli-Ozcan, Sørensen, and Yosha (2003) that risk sharing tends to be larger within regions of the same country than between countries, although β here pertains to income insurance between pairs of regions, rather than multilateral risk sharing. Second, output correlations are larger on average than consumption correlations between countries, an illustration of the quantity puzzle. Third, within the same country, regional output co-movements tend to be much lower than co-movements in local disposable income. This indicates that things might be different within countries, at least in North America. The results in Imbs (2004b) suggested that financial integration pushed both international correlations in output and in consumption upwards. Therefore, the fact that there appears to be more risk sharing between the regions of a same country cannot account for the smaller gap between the two interregional correlations in Table 1, unless the results in Imbs (2004b) do not apply to intranational data. This paper determines whether this is the case.

Two caveats are in order. First, pairwise correlations in disposable income embed regional savings rates, which could push them upwards relative to bilateral correlation in consumption. The smaller discrepancy implied by retail sales data is consistent with this possibility. The gap between output and consumption correlations may be thinner for intranational data than Table 1 suggests. Second, the cross-sectional variation in all pairwise correlations is such that none of the claims just described is significant.

Such variation is what makes these cross-sections potentially informative for rigorous inference.

2.1 The correlations in output

We know little about the relationship between financial integration and cycle synchronization.[6] Most theoretical models predict that it should be negative, starting with Backus, Kehoe, and Kydland (1994). Under complete markets, capital flows into the economy hit by a positive technology shock, and away from the no-shock economy, and thus between economies that tend to be out of phase. Limited enforcement introduced in Kehoe and Perri (2002) endogenously limits the magnitude of international capital flows, but does not alter the general result that, if capital flows between two regions, they should tend to be asynchronized. Empirically, this negative link has proved to be elusive. Kose, Prasad, and Terrones (2003); Bordo and Helbling (2003); and Imbs (2004b) all conclude that if anything, financially integrated countries have positively correlated output fluctuations. An exception is Heathcote and Perri (2002), who explain the falling correlation between the US economy and an aggregate of the rest of the world with increased financial integration.

We know much more about other determinants of the international correlations in GDP fluctuations. Since Frankel and Rose (1998) found a large and significant effect of bilateral trade intensity on business cycle co-movements, many papers have concerned themselves with the relative importance of trade, finance, and other potentially important variables. To name a few, Imbs (1999); Clark and van Wincoop (2001); or Kalemli-Ozcan, Sørensen, and Yosha (2001) document a significant impact of specialization patterns. Alesina, Barro, and Tenreyro (2002) or Rose (2000) focus on currency unions. Imbs (2004a) and Baxter and Kouparitsas (2004) assess the relative magnitude of these channels.

This paper focuses on the link between financial integration and business cycle co-movements within the United States and Canada. We know since Imbs (2004b) that financial integration tends to be associated with correlated GDP fluctuations. We now ask whether this is true between regions in North America, as well. To do this, we follow Imbs (2004b) in estimating jointly

$$\rho_{ij}^{Y} = \alpha_0 + \alpha_1 \ \phi_{ij} + \alpha_2 \ T_{ij} + \alpha_3 \ S_{ij} + \alpha_4 \ X_{ij} + \varepsilon_{ij}^{0}$$

6. This review summarizes the detailed discussion in Imbs (2004b).

$$T_{ij} = \gamma_0 + \gamma_1 \phi_{ij} + \gamma_2 I^1_{ij} + \varepsilon^2_{ij} \tag{S}$$

$$S_{ij} = \delta_0 + \delta_1 \phi_{ij} + \delta_2 I^2_{ij} + \varepsilon^3_{ij},$$

where ρ^Y_{ij} denotes the Pearson correlation between the cyclical components of GDP in regions i and j, and ϕ_{ij} is a measure of bilateral financial integration. X_{ij} is a vector of control variables that affect ρ^Y_{ij} directly. Any direct impact of finance on cycles is captured by estimates of α_1. In this general specification, it is possible that financial integration affects trade, via γ_1, or indeed specialization, via δ_1. The first channel reflects the possibility that trade and financial flows build on the same stock of local knowledge or infrastructures, as argued, for instance, in Lane and Milesi-Ferretti (2004). The second channel corresponds to the potential specialization effects of finance that Kalemli-Ozcan, Sørensen, and Yosha (2003) documented. We first estimate a version of the system (S), where γ_1 and δ_1 are set to zero, then generalize the estimation.

Identification of the system (S) requires distinct instrument sets for T_{ij} and S_{ij}, which are relatively standard. The gravity model of international trade justifies the use of geographic variables to instrument T, as they are both obviously exogenous and strong predictors of trade flows. Specialization is less easy to instrument. We follow Imbs (2004b) and instrument S with both the pairwise sum and difference of per capita output, which assumes that rich regions tend to be more diversified, and thus potentially more similar.

This choice for I^1, I^2 enables identification of the system (S) and also tackles issues of endogeneity, for instance of T to ρ^Y. But it leaves open the question of the endogeneity of ϕ_{ij}. In international data, we follow Imbs (2004b) and instrument observed financial integration from the CPIS data with institutional variables capturing the financial advancement of both countries. But the same cannot be achieved at the intranational level, for lack of comparable data between regions in the United States or Canada. It is important, however, to reiterate that the endogeneity of ϕ_{ij} creates an attenuating bias, as synchronized economies should have low values of ϕ_{ij}. This paper compares results in two data sets where the endogeneity of ϕ_{ij} is likely to prevail equally. The impossibility of coming up with instruments for ϕ_{ij} for US states and Canadian provinces cannot explain that these data generate different results from aggregate ones.

2.2 Results

Table 2 reports ordinary least squares regression of the system (S) where γ_1 and δ_1 are set to zero, which, de facto, shuts down all simultaneity

Table 2
The determinants of output correlations

	International data			Intranational data		
	(i) OLS	**(ii) IV**	**(iii) IV**	**(iv) OLS**	**(v) IV**	**(vi) IV**
Trade	0.0451	0.2295	0.0540	0.0022	0.0127	0.0157
	2.68	4.99	1.93	0.39	1.58	1.86
Structure	–0.3800	–0.4313	–1.2328	–0.8475	–0.4427	–0.5556
	–4.86	–2.65	–9.78	–9.98	–3.00	–3.63
Finance	0.0036	0.0136	0.1414	–0.1170	–0.1238	–0.0210
	2.98	6.13	4.70	–3.93	–4.25	–1.74
Obs.	465	274	778	643	643	643

Notes: The dependent variable is the bilateral correlation in the cyclical components of output. Specifications (i) and (ii) use directly observed measures of bilateral asset holdings collected by the IMF for 2001, and sums the values of all asset holdings in both directions. Specifications (iii)–(vi) use instead the proxy based on bilateral risk sharing, constructed on the basis of the cyclical component of real consumption (iii), real disposable income (iv) and (v), and real retail sales (vi). All series are detrended using the Baxter-King (1999) filter. The instruments for trade include distance, a binary variable capturing the presence of a common border, and the product of populations. The instruments for structure include the sum and pairwise differences in per capita output, and the instruments for finance (when it is directly observed) capture financial development in each economy, and include an index of accounting standards, a measure of the efficiency of the judicial system, whether votes at general meetings are cast cumulatively or proportionately, and the risks of expropriation and contract repudiation.
OLS: ordinary least squares.
IV: instrumental variable.

problems. In the international data, specifications (i) and (ii) make use of the CPIS data on bilateral asset holdings, instrumented in (ii). All other specifications use proxies for ϕ_{ij} given by the estimates of β in equation (2). Finally, estimations (iii) and (v) instrument trade and structure. In both data sets, the impact of trade and structure is consistent with results in the literature: intense bilateral trade linkages are associated with correlated business cycles, whereas economies with similar patterns of specialization are more correlated.

The striking result in Table 2 pertains to the measured role of finance. Across countries, we confirm the conclusion in Imbs (2004b), that financially integrated countries are significantly more synchronized. This happens whether ϕ_{ij} is measured using the IMF CPIS data, whether bilateral asset holdings are instrumented or not, and when ϕ_{ij} is measured using this paper's suggested proxy, based on aggregate consumption data. Estimates of α_1 are in all cases positive and significant. This is particularly remarkable in specifications (i) and (iii), where ϕ_{ij} is not instrumented and the estimates of α_1 (potentially) suffer from an attenuating endogeneity bias. Consistent with this, instrumentation in column (ii) does indeed increase the magnitude of α_1. Across regions within one country, on the other hand, α_1 is always negative and significant, whether T_{ij} and S_{ij} are instrumented or not, and whether β is estimated using disposable income or retail sales data. This suggests that the effect of financial integration on ρ_{ij}^{Y} is drastically different within the regions of a same country, and indeed potentially consistent with theory (though estimates for α_1 in columns (iv) to (vi) might not be as negative if we were able to instrument ϕ_{ij}).

An explanation for the difference might be that ρ_{ij}^{Y} has other determinants in international data (but not in intranational data) that we fail to control for properly. Most prominently, aggregate business cycles might correlate positively because of international coordination in macroeconomic policies, in turn possibly correlated with our measures of ϕ_{ij}. If this were true, it would already foster our understanding of the quantity puzzle, since it would mean that business cycles are more correlated than consumption plans across countries because of convergence in policy. That would not leave much of a puzzle, and intranational data would offer a vindication for existing theories of capital flows and their effect on business cycle synchronization. It is, however, far from clear that positive estimates for α_1 in international data stem from policy coordination. For instance, Imbs (2004b) controls for measures of exchange rate volatility, differentials in inflation rates, or the presence of monetary or trade unions, and still finds positive measures for α_1, even in a sample excluding correlations between rich OECD (Organisation for Economic Co-operation and Development) economies, the most likely to coordinate their policies. Kose, Prasad, and Terrones (2003) conclude similarly.[7]

Table 3 now allows for γ_1 and δ_1 to differ from zero. The main result is unchanged: α_1 is positive in international data, but negative in intranational data. This is important, for it suggests that α_1 is not negative in intranational data because of the specialization effects of finance, which might be more acute within than between countries. The putative effects of finance on trade

7. Another difference between the two data sets pertains to the series utilized when estimating β. In international data, actual consumption is used, whereas local disposable income is used in intranational data. The latter embed local savings rates. But for this to explain the discrepancy in Table 2, we would need the pairwise difference in savings rate Δs_{ijt} to correlate very negatively with Δy_{ijt}. In other words, we would need $s_{it} < s_{jt}$ whenever $y_{it} > y_{jt}$. It is impossible to verify this possibility in intranational data, but in the literature, savings rates are usually found to increase with income, rather than the opposite. Furthermore, our results are by and large confirmed by state-level sales data, which are immune from this potential issue.

Table 3
Simultaneous estimations

	International data		Intranational data	
	(i) 3SLS	**(ii) 3SLS**	**(iii) 3SLS**	**(iv) 3SLS—Sales**
Correlation equation				
Trade	0.2226	0.0539	0.0125	0.0104
	4.82	1.93	1.55	1.26
Structure	–0.3980	–1.2371	–0.4426	–0.4627
	–2.43	–9.81	–3.00	–3.07
Finance	0.0131	0.1180	–0.1228	–0.0225
	5.70	3.41	–4.08	–1.88
Trade equation				
Finance	0.0103	0.0000	–0.2332	–0.0324
	1.82	0.00	–1.73	–0.64
Specialization equation				
Finance	–0.0027	–0.0269	0.0132	0.0235
	–2.02	–1.49	1.07	5.20
Obs.	274	778	643	643

Notes: Specification (i) uses directly observed measures of bilateral asset holdings collected by the IMF for 2001, and sums the values of all asset holdings in both directions. Specifications (ii)–(iv) use instead the proxy based on bilateral risk sharing, constructed on the basis of the cyclical component of real consumption (ii), real disposable income (iii), and real retail sales, (iv). All series are detrended using the Baxter-King (1999) filter. The instruments for trade include distance, a binary variable capturing the presence of a common border, and the product of populations. The instruments for structure include the sum and pairwise differences in per capita output, and the instruments for finance (in column (i)) capture financial development in each economy and include an index of accounting standards, a measure of the efficiency of the judicial system, whether votes at general meetings are cast cumulatively or proportionately, and the risks of expropriation and contract repudiation.
3SLS: three-stage least squares.

or specialization are accounted for in Table 3. Estimates of the residual effect may still be affected by the endogeneity of ϕ_{ij}, but not in a way that can explain the discrepancy between intra- and international data.

2.3 Extensions and robustness

In this section, we extend the estimations to include measures of the size of the financial sector. We then close with some sensitivity analysis.

Table 4 reproduces Table 3, adding a measure of the size of the financial sector in the correlation equation. This is meant to control for the possibility that actual risk sharing as measured by ϕ_{ij} (or β) be determined by local regulations affecting financial development, as banking regulations in the

Table 4
The determinants of output correlations—FIRE

	International data		Intranational data		
	(i) 3SLS	**(ii) 3SLS**	**(iii) 3SLS**	**(iv) 3SLS**	**(v) 3SLS—Sales**
Correlation equation					
Trade	0.2452	0.0581	0.0220	0.0246	0.0196
	5.60	2.08	2.66	3.11	2.33
Structure	–0.3761	–1.2529	–0.3938	–0.6774	–0.41147
	–2.24	–9.03	–2.91	–4.71	–2.96
Finance	0.0150	0.1179	–0.1223	–0.1197	–0.0182
	5.88	3.41	–4.10	–3.94	–1.55
FIRE	–0.2401	–0.0563	0.5292		0.5105
	–1.19	–0.60	2.82		2.66
Bank				–4.6038	
				–7.52	
Securities				5.5806	
				5.61	
Real estate				0.5384	
				1.88	
Holdings				–11.1211	
				–1.72	
Insurance				3.3603	
				2.17	
Trade equation					
Finance	0.0100	–0.0006	–0.2308	–0.2311	–0.0330
	1.76	–0.01	–1.71	–1.71	–0.65
Specialization equation					
Finance	–0.0029	–0.0271	0.0133	0.0133	0.0235
	–2.19	–1.45	1.07	1.07	5.20
Obs.	274	778	643	643	643

Notes: Specification (i) uses directly observed measures of bilateral asset holdings collected by the IMF for 2001, and sums the values of all asset holdings in both directions. Specifications (ii)–(v) use instead the proxy based on bilateral risk sharing, constructed on the basis of the cyclical component of real consumption (ii), real disposable income (iii) and (iv), and real retail sales (v). All series are detrended using the Baxter-King (1999) filter. FIRE denotes the pairwise sum of the share of the finance, insurance, and real estate sector in overall output. The instruments for trade include distance, a binary variable capturing the presence of a common border, and the product of populations. The instruments for structure include the sum and pairwise differences in per capita output, and the instruments for finance (in column (i)) capture financial development in each economy, and include an index of accounting standards, a measure of the efficiency of the judicial system, whether votes at general meetings are cast cumulatively or proportionately, and the risks of expropriation and contract repudiation.
3SLS: three-stage least squares.

Table 5
The determinants of output correlations—sensitivity

	International data			Intranational data		
	(i) HP	**(ii) T2**	**(iii) UNIDO**	**(iv) HP**	**(v) T2**	**(vi) T2—Sales**
Correlation equation						
Trade	0.0792	49.069	0.0834	0.0119	5.2782	5.0206
	2.44	5.70	3.71	1.45	4.14	3.83
Structure	–1.2618	–0.6967	–0.5582	–0.4509	–0.0403	–0.0611
	–8.83	–3.65	–12.21	–3.00	–0.24	–0.35
Finance	0.0786	0.0612	0.1002	–0.0659	–0.1260	–0.0310
	2.01	1.83	4.86	–1.67	–4.22	–2.61
Trade equation						
Finance	–0.0022	0.0012	–0.4779	0.2103	–0.0025	–0.0008
	–0.03	2.73	–0.81	1.21	–2.72	–2.30
Specialization equation						
Finance	–0.0329	–0.0337	–0.0183	–0.0196	0.0130	0.0235
	–1.76	–1.80	–0.98	–1.23	1.05	5.21
Obs.	778	778	1,491	643	643	643

Notes: All specifications use the proxy based on bilateral risk sharing, constructed on the basis of the cyclical component of real consumption (i) to (iii), real disposable income (iv) and (v), and real retail sales (vi). Specifications (i) and (iv) use the HP filter to isolate the cyclical component of time series when estimating β in equation (2). Specifications (ii), (v), and (vi) use an alternative measure for bilateral trade intensity, where exports and imports are normalized by both regions' total output. Column (iii), finally, uses a measure of *S* based on manufacturing data from the United Nations Industrial Development Organization (UNIDO). The instruments for trade include distance, a binary variable capturing the presence of a common border, and the product of populations. The instruments for structure include the sum and pairwise differences in per capita output.
T2 denotes the alternative measure of bilateral trade intensity used in Frankel and Rose (1998).

United States, for instance. Different outcomes at both levels of aggregation could then merely reflect different regulations of the financial sector within and between countries. Table 4 ensures they do not: estimates for α_1 continue to change signs with the level of aggregation, whether or not the (pairwise sum of the) shares of the finance, insurance, and real estate (FIRE) sector in overall output are controlled for. Interestingly, Table 4 also suggests that US states and Canadian provinces that have developed FIRE sectors tend to be more correlated, a result that does not hold significantly between countries. A decomposition of the FIRE sector suggests that this happens mostly because of the securities, insurance, and real estate sectors. In contrast, the prevalence of a large banking sector seems to be associated with low output correlations.

Table 5 conducts a sensitivity analysis. First, an alternative filter is used to isolate the cyclical component of output and to estimate β in equation (2). Second, an alternative measure of bilateral trade intensity is used, following

Frankel and Rose (1998), which consists in the sum of bilateral exports and imports simply normalized by both regions total output. Third, for international data, the UNIDO data are utilized to compute *S* on the basis of manufacturing production only. The results obtain in all cases. α_1 continues to change signs with the level of aggregation.

Conclusion

This paper uses intranational data on the United States and Canada to investigate the determinants of output correlations at the regional level. An important difference with the international evidence is uncovered. While financial integration tends to result in synchronized business cycles in the aggregate, the opposite is true within countries, at least in the United States and Canada. A novel proxy for bilateral risk sharing between regions is proposed to obtain this result. This specificity of intranational data may explain why, in the United States and in Canada, correlations in disposable income are larger than correlations in output, in contrast with the aggregate evidence. In regional data, the link between bilateral financial integration and output correlations is consistent with theory. There is no apparent quantity puzzle. This suggests that the key to the puzzle in aggregate data lies in understanding the determinants of aggregate capital flows and, in particular, why they tend to result in positively correlated business cycles.

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Discussion

Linda Goldberg

My discussion will focus on two areas—on the paper by Jean Imbs, and on understanding the institutions that may be contributing to the quantity puzzle. I will use as an example the role of banks in fostering real effects of financial integration. Banks are useful institutions for this purpose, since we can consider what their mechanisms for consumption smoothing would be, if they were involved in such smoothing. As a general comment on the quantity puzzle research agenda, I would like to see more discussion of which institutions are expected to participate in the type of insurance mechanism implicit in the quantity puzzle. More institutional and data details would be useful in this respect.

The paper by Jean Imbs is a useful study with carefully implemented empirics. It begins with a relationship that has surprised many economists. Specifically, while consumption is less correlated than business cycles across countries (the quantity puzzle), consumption is more correlated than business cycles across regions. Imbs offers an explanation whereby the specific impact of capital flows on international output correlations is behind the quantity puzzle. In other words, he posits that capital flows don't provide insurance internationally, but may do so intranationally (i.e., within a country's borders).

The paper is part of a large body of literature on intranational risk sharing. Within this literature, studies of US states emphasize the roles played by capital markets and/or fiscal federal systems: as in Atkeson and Bayoumi (1993); Asdrubali, Sørensen, and Yosha (1996); Athanasoulis and van Wincoop (2000); and Sala-i-Martin and Sachs (1992). There are also studies of Canadian provinces that emphasize the explicit mechanism for horizontal transfers among the provinces, i.e., the Canadian system of equalization (redistribution instead of insurance per se). Examples are Bayoumi and Masson (1995), Mélitz and Zumer (1999), and Obstfeld and Peri (1998). More recent papers focus on the importance of trade linkages.

Imbs argues that financial integration helps with risk sharing in Canadian provinces and in US states. But financial integration between economic regions in the same country is not directly observable. A proxy is introduced based on the responsiveness of local income to local production. While such a proxy has been used in other studies (mostly international), I find it less than compelling within a country, since the proxy can capture many things other than financial integration. It would be useful, therefore, to determine and more precisely document why financial institutions might behave differently within rather than across countries.

Do banks smooth consumption internationally? To explore this question, let us consider a real application. Specifically, how should we think about what real effects are associated with financial integration through banking? Consider two types of foreign exposures of banks: cross-border claims (claims on foreign debtors by US domestic banks) and local claims (claims on foreign debtors by branches of US banks located in the debtor's country). Figure 1 shows the total cross-border claims of US reporting banks starting in 2000 and continuing through mid-2004, and provides details on the components for Europe, Asia, and Canada.[1] Note that in total cross-border lending by the United States, Canada is the recipient of $26 billion (only 4 per cent of nearly $693 billion in 2004).

Data on foreign exposures of US banks, this time focusing on net local country claims by US banks, reveal a different pattern. As shown in Figure 2, Canada is the recipient of $17.6 billion of net local country claims by US reporting banks, approximately 20 per cent of $85.8 billion in 2004. These data suggest that in banking, US banks have more financial integration through local affiliates than through cross-border flows.

Figure 3 shows these two types of claims, normalized so that 2004Q1 = 100, plotted alongside normalized Canadian real GDP and US real GDP. It appears that both types of claims are procyclical with Canadian GDP, as is typical for bank lending. If this is the case, it is difficult, ex ante, to show that banks are playing a large role in international consumption smoothing. At the very least, Imbs and other researchers on the quantity puzzle should provide a much more nuanced explanation of which types of flows and to which types of parties would be the appropriate metrics for

1. Data are from the Federal Financial Institutions Examination Council (FFIEC) quarterly report: E.16 Country Exposure Lending Survey and Country Exposure Information Report, Table 1.

Figure 1
Total cross-border claims of US reporting banks

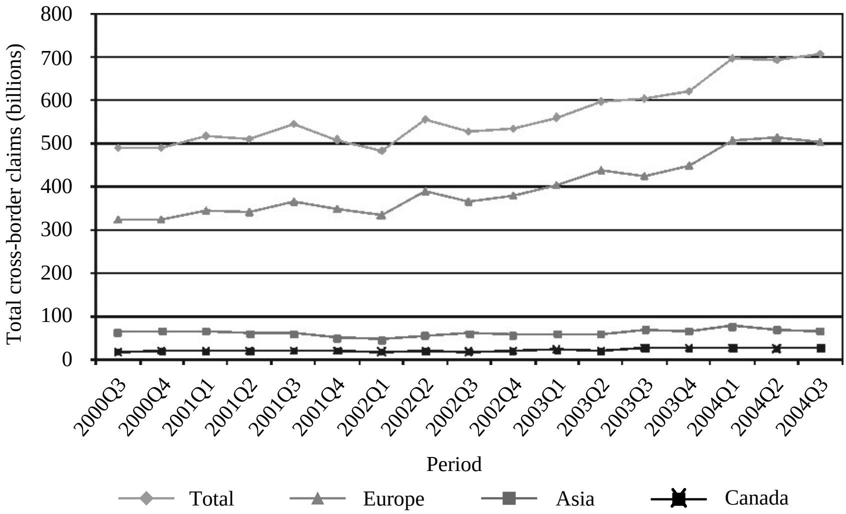


Figure 2
Net local country claims of US reporting banks

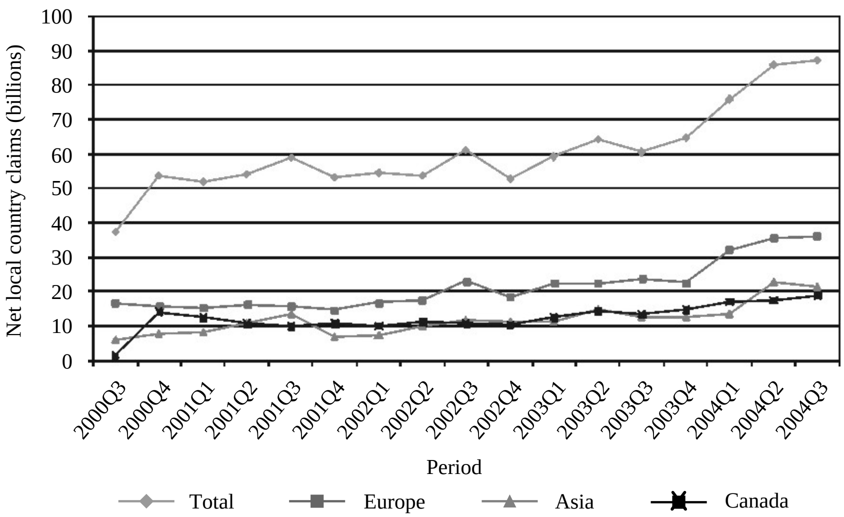
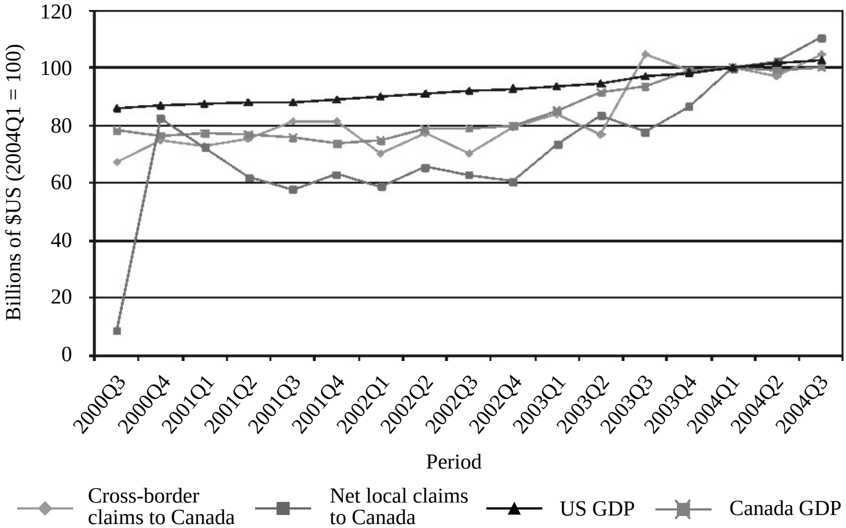


Figure 3
US reporting banks claims to Canada, with GDP data

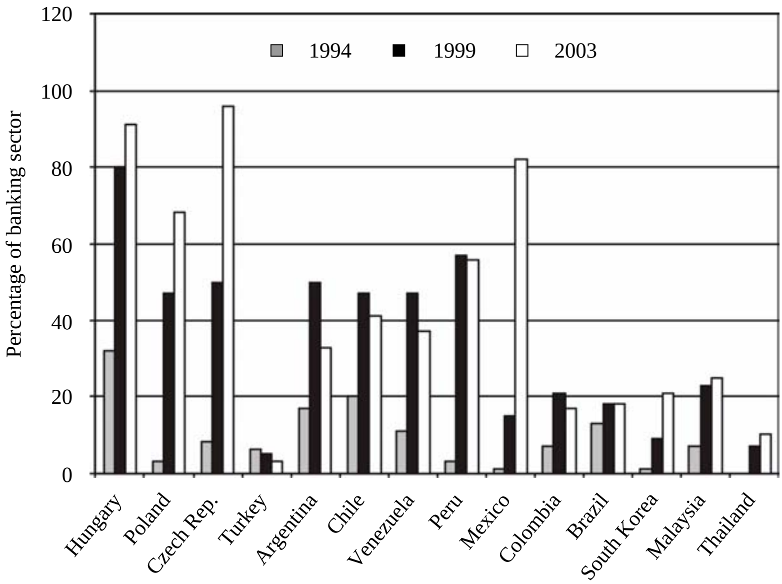


discussions on consumption smoothing. This type of observation has been made for Canada and other industrialized countries, but it also applies to emerging economies.

It is also interesting to think about whether the changing structure of the industrialized country banks within, for example, emerging financial markets would have any real effects on the economic determinants of claims across countries. The evolution of ownership has been impressive. As shown in Figure 4, by 1999, bank assets under foreign control were dramatically higher than five years before. By 2003, there were more nuanced developments. Still, procyclical lending occurs.

Why should there be a relationship between this type of financial intermediation and shock transmission? Theoretical arguments show us why banks may magnify or dampen local business cycles and transmit foreign shocks through the lending channel. The reasons offered for cyclicality in lending include: market risk changing cyclically, so that asset demands and supplies change (as in Berlin and Mester 1999); mismeasurement of difficulties in downturns and strengths in boom periods (as in Borio, Furfine, and Lowe 2001); and the type of intertemporal smoothing, leading to countercyclical loan demand, as in the literature behind the quantity puzzle discussions.

Figure 4
Bank assets under foreign control, 1994, 1999, and 2003
(Assets of banks in which foreigners own 40 per cent or more of bank)



Source: Mathieson and Roldos (2001), plus calculations using Fitch IBCA Bankscope.

Both foreign banks (multinationals) and domestic banks are procyclical suppliers of credit. Both contribute to international business cycle integration but not necessarily to smoothing. Domestic banks rely more on domestically generated sources of funds for lending activity, so their lending is highly procyclical. It is sensitive to domestic cycles and increases the amplitude of these cycles. Foreign banks rely more on source-country funds. They transmit slightly more of their own country shocks to markets in which they have a presence, but also reduce the amplitude of locally generated cycles.

Instead of leading to real consumption smoothing through lending channels, how else can real financial integration promote smoothing of local shocks? Some ideas to explore are: (i) through the development of local institutions; (ii) through bank integration with foreign head offices, which can lead to stronger risk-management systems/operational controls; and (iii) through product innovation and expansion of services (broader range of credit and deposit products, treasury, financial advisory services, etc.). There is also anecdotal evidence of spillovers to supervision.

Is there more of a story to tell for these alternative channels of effects in an intranational versus international setting? This point warrants exploration. The Jean Imbs paper takes us in the right direction and can usefully explore these themes further.

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General Discussion*

Andrew Rose responded to Eric Santor’s comments by noting that it was unlikely that exchange rate movements were the driving factor, as the analysis of daily and monthly data delivers similar results. Rose acknowledged that firms listed on the TSE may have different features, such as different ownership structure, but emphasized that it doesn’t change the fact that the expected marginal rates of substitution are not equal across markets. With respect to adequacy of the instruments used, Rose reported that the first-stage regressions explained about 80 per cent of the variance. He also responded that they had simply sorted the portfolios alphabetically. He welcomed further research by others on sorting strategies.

Jean Imbs responded to Linda Goldberg’s comments by noting that the paper’s findings were consistent with procyclical lending patterns that Goldberg had indicated. He agreed that it would be interesting to look at interregional and international sources of funding.

In the general discussion that followed, Charles Engel pointed out that Andrew Rose’s analysis assumes a representative investor rather than heterogeneous agents. Moreover, as purchasing-power parity (PPP) does not hold, prices will differ between the United States and Canada. Thus, the expected marginal rate of substitution will differ between countries even under perfect financial integration so long as markets are incomplete. Essentially what Rose estimates is a weighted average of marginal rates of substitution. Rose underscored that what was actually being estimated was the expected, not the actual, marginal rate of substitution. While

* Prepared by Robert Lafrance.

heterogeneity among investors is plausible, Rose doubted that this could account for the large differences in the expected marginal rate of substitution between markets as estimated in the paper.

In a follow-up comment on Imbs's paper, Engel pointed out that perfect financial markets do not imply perfect risk sharing as long as PPP doesn't hold. This raises the question of what we are trying to measure. If the real exchange rate is not a constant, then perfect risk sharing is not an optimal strategy. He asked whether the framework used in the paper could be extended to deal with that issue. Imbs replied that there are many reasons why risk sharing may not be occurring in the data. In practice, the ratio of marginal utilities is not equal to the real exchange rate. Including real exchange rates in the estimation was not likely to change his results. Gregor Smith sought to clarify the identification assumption in equation 8 in Rose's paper, since it seemed different from the conventional approach. Rose agreed. In effect, he had used an econometric trick to avoid explicit identification of factors. Lawrence Schembri asked how the difference in expected marginal rate of substitution between Canadian and US financial markets compared with other markets and how it might have changed over time. Rose replied that the only other comparison he had done was between the NYSE and NASDAQ, which showed, surprisingly, a similar result.

In reference to Imbs's paper, Graham Voss questioned pooling US and Canadian data, given that these two countries have very different banking systems. He suggested that one could look at the two countries separately. Imbs explained that pooling was required, since Canada had too few jurisdictions.