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Wiener Institut für Internationale Wirtschaftsvergleiche

The Vienna Institute for International Economic Studies

Working Papers | 68 | August 2010

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An Empirical Characterization of Redistribution Shocks and Output Dynamics

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An Empirical Characterization of Redistribution Shocks and Output Dynamics

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Abstract

What are the economic effects of redistributing one dollar from profits to labour income? We address this question for the post-World War II economies of the United States and Canada within a structural VECM procedure allowing for up to two breaks of unknown timing. In the United States the short-run spending effect on growth, set in motion by higher labour income, is strong enough to make such a redistribution an attractive, maybe provocative, policy alternative. Across the border in Canada, however, the negative medium-run capacity effect, brought about by diminished profits, dominates the picture more or less from the beginning and output slumps considerably, a result actually suggesting a – maybe even more provocative – redistribution toward profits. We discuss several possible explanations such as the formation of expectations and the different exposure to international trade. Methodologically, we provide a novel procedure to estimate cointegrating rank and break dates jointly.

Key words: *redistribution, structural VECM, joint estimation of cointegrating rank and multiple break dates*

JEL classification: *C32, C53, E12, E25*

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We are grateful to Helmut Lütkepohl, Massimiliano Marcellino, and Leo Michelis for constant advice.

1 Introduction

In the last two decades numerous authors and research groups have put in huge efforts and ingenuity to correctly disentangle and characterize the economic effects of shocks to technology and monetary policy. Issues on the economic effects of a "redistribution shock" or, say, transferring money between labor income and profits, although having a long tradition in macroeconomics, fell into disfavor in the main economic debate.[1] In light of the 2008 economic crisis and discussion on potential economic stimulus programs other than government spending and low interest rate policies, we address the question on how such a redistribution shock will affect output.

From a theoretical perspective the debate might be characterized by focusing on two channels which affect aggregate demand: a short-run spending effect on the one and a medium-run capacity effect on the other hand. The first spending argument rests on the post-Keynesian perspective of differing marginal propensities to spend out of labor income and profits.[2] Workers will spend more of an additional dollar than firms and, therefore, a redistribution toward labor income will increase output. The capacity argument positively links investment with realized and expected profits. As a consequence, a distributional shift toward profits (equivalently, away from labor income) might have a positive effect on investment, resulting in higher output growth over the medium-run.[3]

In view of the trade-off between these two channels, our purpose in this paper is to characterize the dynamic effects of a redistribution between labor income and profits on output in the United States and Canada during the post-World War II period. We do so using a small 3-dimensional structural vector time series model of quarterly labor income, profits and output. Identifying a redistribution shock is intricate, inasmuch as any output ef-

[1]See European Commission (2007, Chap. 5) for a recent overview of the debate.

[2]Though not going into detail, a view linked to economists related to the Keynesian and post-Keynesian tradition such as John Maynard Keynes himself, Richard Goodwin, Nicholas Kaldor, Michael Kalecki and Joan Robinson to name a few.

[3]From a more general perspective, this debate centers around the notion of wage-led versus profit-led economic expansion; see Bhaduri and Marglin (1990) for a model which can accommodate both views and Stockhammer et al. (2009) for an empirical application thereof.

fect generated by the redistribution has, in turn, an immediate effect back on profits and labor income. Profits may move with output under existing markup rates, and labor income changes as firms and workers try to adjust employment and wage rates. But exactly these "automatic" responses of labor income and profits render the identification problem to be non-recursive and standard procedures such as a simple Choleski decomposition are not applicable. Therefore, to achieve identification, we opt for a structural VECM approach and exploit its inherent long-run restrictions to obtain estimates of the automatic responses, and, by implication, identification of temporary and permanent shocks. The distinction between the types of shocks and our clear focus on redistribution as a possible aggregate demand instrument lead us to design a redistribution shock as a combination of transitory shocks to labor and profit income. We design it such that there is a one-for-one redistribution between labor income and profits. Having defined and identified the redistribution shock we can trace its dynamic effects on output. This is, in a nutshell, what we aim at in this paper.

Structural vector time series approaches, in general, have been widely used to assess the effects of technology shocks, monetary policy, and fiscal policy (see, in particular, King et al., 1991; Bernanke and Mihov, 1998; and Blanchard and Perotti, 2002). We argue that such an approach is equally well suited to study the effects of a redistribution between labor income and profits for two reasons. First, movements in labor income and profits can be treated as exogenous with respect to output because stabilization motives, in contrast to monetary policy, are rarely the predominant driving force behind fluctuations in the distribution. A redistribution shock is therefore exogenous with respect to output. Second, and similar to fiscal policy, decision and implementation lags rule out—at least within a quarter—most of the discretionary response of labor income and profits to unexpected contemporaneous changes in output. What is left then are the automatic responses of unexpected movements in output on labor income and profits for which we account for in our identification strategy to obtain a proper representation of a redistribution shock.

The long time span of the post-World War II period, however, most probably contains structural breaks. We therefore allow for up to two level breaks of unknown timing in our structural VECM. We do so by combining an "older" approach on how to estimate the break dates with recent advances in testing for the cointegrating rank when breaks are present. The older approach goes back to the unit root test of Zivot and Andrews (1992) and its numerous successors. The goal is to pick the two break dates such that most weight is given to the stationary alternative. Gregory and Hansen (1996) extend these unit root tests to residual-based cointegration analysis in which the weight is then always on the alternative hypothesis of the next higher cointegrating rank. We take up this testing strategy but perform the cointegration tests with two level breaks within the framework of Saikkonen and Lütkepohl (2000a,b), and Lütkepohl et al. (2004). This yields a two-step system-based framework. In the first step one removes the deterministic components by a feasible generalized least squares procedure. Then, in the second step, we apply the commonly used Johansen (1988, 1995) likelihood ratio test to the adjusted series. The prior adjustment is convenient because it results in an asymptotic null distribution of the cointegration rank test statistic which does not depend on the timing of the level breaks.

Besides the empirical question addressed and the issues related to identification, the "joint nature" of our estimation of the cointegrating rank and the two break dates, is the methodological contribution of this paper. Other approaches typically estimate the break dates before the cointegration analysis, either on the basis of an—with respect to the cointegrating rank—unrestricted model (see, e.g., Lütkepohl, Saikkonen and Trenkler, 2004; and Saikkonen, Lütkepohl and Trenkler, 2006) or on the basis of unit root tests (see, e.g., Koukouritakis and Michelis, 2009). Our joint estimation procedure removes an additional layer from the analysis, as it avoids the pre-test bias introduced by using different models for choosing cointegrating rank and break dates. The paper also provides Monte Carlo evidence showing the basic consistency of our joint estimation procedure.

Our main results underline the trade-off between the two transmission channels mentioned above. In the United States the short-run spending effect is strong enough to make a one-dollar redistribution from profits toward labor income successful in terms of output growth: output is above trend for two years, with a multiplier of 0.51 dollars, before the negative capacity effect eventually takes over. In Canada, however, this negative capacity effect dominates the picture more or less from the beginning. After initially increasing by 0.47 dollars, output slumps by notable 1.73 dollars within the first two years and reverts to trend only slowly thereafter. Thus, it is a redistribution toward labor income that has a positive short-run effect on output in the United States, whereas in Canada one will need to shift sources in the opposite direction to generate a positive effect. We discuss several issues in turn why the transmission of a redistribution shock differs in the two countries. Prominent candidates among them are differences in the formation of expectations maybe explained by growing gap in unionization and collective bargaining power and differences in the exposure to international trade.

2 Identification and Theoretical Background

Throughout the paper we will use a small 3-dimensional VECM with two level breaks including labor income, profits and output for the United States and Canada. The identification of more, possibly non-zero, contemporaneous relations—including the automatic responses—than the $k(k+1)/2$ distinct elements of the covariance matrix actually offer in a k-dimensional system is at the center of our methodology. Accordingly, we want to impose less than $k(k-1)/2$ restrictions on the contemporaneous relations. On the one hand, the small empirical model employed keeps these identification issues, arising in higher dimensional systems and simultaneous equation models, and uncertainties about variable definitions such as choosing among the various measures of the wage rate and employment, at a minimum. On the other hand, the parsimonious setup of the model ignores other potentially important macroeconomic variables hiding a detailed analysis of the transmission channel, e.g. through consumption or investment. Higher dimensional systems with possible cross-country linkages would require a different framework to study the effects of a redistribution in more detail. An option, for instance, could be the global error-correcting model of Pesaran et al. (2004)

or the global VAR of Dees et al. (2007). Because of difficulties with identification, cointegration and breaks we stick in this paper to our three variable model, which shall be sufficient to characterize the most relevant features of redistributing between labor income and profits.

Our strategy to achieve identification makes use of the endeavors of John Hicks and Paul Samuelson to absorb Keynes thoughts into neoclassical economics. The result, as Blanchard (1997) forcefully argues, is the core of usable macroeconomics. Two propositions build the basis for what we nowadays know as the neoclassical synthesis: first, in the short-run aggregate demand dominates movements in economic activity and, second, the economy tends to return to a steady-state growth path. A redistribution shock changes aggregate demand and, therefore, these two propositions characterize the long-run effects of such a shock, after all the complex short- and medium-run effects have worked out. Essentially, this characterization suggests a stable steady-state relation of labor income and profits with output, giving the notion of two cointegration relations in the data. Cointegration puts restrictions on the permanent impact matrix, thus reducing the number of restrictions we need to impose on the contemporaneous relations. From a methodological point of view this type of identification follows the VECM approach of King et al. (1991).

Suppose the observed sample is $\{y_t\}_{t=1}^T$ in which y_t is a 3-dimensional vector containing the logarithm of quarterly real macroeconomic time series on labor income (inc_t), profits (π_t), and output (gdp_t). Without loss of generality we can write a structural model linking the reduced-form residuals $\{u_t\}_{t=1}^T$ of the vector time series model with the mutually uncorrelated structural shocks $\{e_t\}_{t=1}^T$ as a so called B-model,

$$\begin{bmatrix} u_t^{inc} \\ u_t^{\pi} \\ u_t^{gdp} \end{bmatrix} = \begin{bmatrix} b_{11} & b_{12} & b_{13} \\ b_{21} & b_{22} & b_{23} \\ b_{31} & b_{32} & b_{33} \end{bmatrix} \begin{bmatrix} e_t^{inc} \\ e_t^{\pi} \\ e_t^{gdp} \end{bmatrix}, \tag{1}$$

or more compactly as $u_t = \mathsf{B}e_t$. Once we embed the structural model in a VECM with the two suggested cointegration relations the first two structural shocks will have transitory effects only. We interpret these shocks as labor income and profit shocks. The permanent shock, e_t^{gdp}, may be interpreted as a productivity or growth shock, but since that is not at issue, there is no point in exploring the permanent shock further. Admittedly, interpreting residuals in small dimensional systems as structural shocks is always perilous, and our interpretation of the two temporary shocks as labor income and profit shocks is no exception.

The structural model allows to formally define a linear combination of the two transitory shocks such that there is a one-for-one redistribution between labor income and profits.

Definition 1 (Redistribution shock) $e_s^{\Delta} = (1, \varepsilon, 0)'$ with $\varepsilon = -(b_{11} + b_{21})/(b_{12} + b_{22})$ such that $u_s^{inc} = -u_s^{\pi}$ at time s when the shock occurs.

The last equation of our structural model contains the effects of a redistribution shock on output, captured by $b_{31}e_t^{inc}$ and $\varepsilon b_{32}e_t^{\pi}$, in which the parameters b_{31} and b_{32} are the ones absorbing the marginal propensities to spend out of labor income and profits. These two parameters principally reflect two different channels through which the shock affects output within the quarter: the direct impact effect after all immediate feedback effects have unfolded, and any discretionary adjustment made to rules and laws that influence wage setting, employment, and profit opportunities. We follow the approach by Blanchard and Perotti (2002) in their study on the effects of fiscal policy and rule out the second channel because of decision and implementation lags. It usually takes policymakers more than a quarter to analyze, decide, and implement measures, if any, to respond to unexpected events.

Any effect on output through b_{31} and b_{32} may have an immediate effect back on profits and labor income. For any change in output, profits move under existing markup rates, and firms and workers try to adjust employment and wage rates accordingly. The parameters b_{13} and b_{23} implicitly take up these "automatic" responses though only implicitly as our structural model formulates relations for the shocks rather than the observable variables. The parameters, therefore, do not have an interpretation as automatic responses. This statement will become clear momentarily in Remark 1. In the same way,

b_{31} and b_{32} are not marginal propensities but are the direct impact effects.

Remark 1 The structural model in (1) nests all models which explicitly formulate relations between the observable variables and the shocks.[4] One simple, yet general, way to write such a model is

$$\begin{aligned} u_t^{inc} &= a_1 u_t^{gdp} + a_2 e_t^{\pi} + e_t^{inc} \\ u_t^{\pi} &= b_1 u_t^{gdp} + b_2 e_t^{inc} + e_t^{\pi} \\ u_t^{gdp} &= c_1 u_t^{inc} + c_2 u_t^{\pi} + e_t^{gdp}, \end{aligned}$$

in which the parameters c_1 and c_2 can be interpreted as marginal propensities to spend, and a_1 and b_1 are the automatic responses. Translated into a B-model we have

$$\begin{aligned} \begin{bmatrix} u_t^{inc} \\ u_t^{\pi} \\ u_t^{gdp} \end{bmatrix} &= \left(\frac{1}{1 - a_1 c_1 - b_1 c_2} \right) \\ &\times \begin{bmatrix} 1 - b_1 c_2 + b_2 a_1 c_2 & a_2 - a_2 b_1 c_2 + a_1 c_2 & a_1 \\ b_1 c_1 + b_2 - b_2 a_1 c_1 & 1 + a_2 b_1 c_1 - a_1 c_1 & b_1 \\ c_1 + b_2 c_2 & a_2 c_1 + c_2 & 1 \end{bmatrix} \begin{bmatrix} e_t^{inc} \\ e_t^{\pi} \\ e_t^{gdp} \end{bmatrix}, \quad (2) \end{aligned}$$

This simple model already already shows how complex the underlying structure of the model in (1) may be. By estimating directly the B-model we avoid imposing any specific structure on the relations between the observable variables and the shocks. Nevertheless, the model here is suggestive for one important reason: if we do not allow b_{13} and b_{23} to differ from zero in (1), all automatic response effects would disappear from the analysis.

Let us now describe the procedure how to just-identify our structural model. Proposition 6.1 in Lütkepohl (2005) shows that in a VECM y_t can be decomposed in the Beveridge-Nelson fashion into $I(1)$ and $I(0)$ components.

[4] See Amisano and Giannini (1997) and Lütkepohl (2005) for the different ways to set up a structural vector time series model.

Specifically,

$$y_t = \Xi \mathsf{B} \sum_{i=1}^{t} e_i + \sum_{j=0}^{\infty} \Xi_j^* \mathsf{B} e_{t-j} + y_0. \tag{3}$$

y_0 are the starting values of the process; $\sum_{j=0}^{\infty} \Xi_j^*$ is an infinite-order polynomial that contains only transitory effects with Ξ_j^* converging to zero as $j \to \infty$; and the common trends term $\Xi \mathsf{B} \sum_{i=1}^{t} e_i$ captures the permanent effects of shocks. Ξ has rank $k-r$ and can be written as

$$\Xi = \beta_\perp \left[\alpha_\perp' \left(I_k - \sum_{i=1}^{p-1} \Gamma_i \right) \beta_\perp \right]^{-1} \alpha_\perp', \tag{4}$$

in which $\alpha_\perp$ and $\beta_\perp$ are orthogonal complements of α and β such that $\alpha' \alpha_\perp = 0$ and $\beta' \beta_\perp = 0$.

We get the restricted maximum likelihood estimator of B by maximizing the concentrated log-likelihood function (omitting the constant),

$$\ln L_c(\mathsf{B}) = -\frac{T}{2} \ln |\mathsf{B}|^2 - \frac{T}{2} \operatorname{tr}\left(\mathsf{B}'^{-1} \mathsf{B}^{-1} \Sigma_u \right), \tag{5}$$

subject to the structural short- and long-run constraints,

$$C_{sr} \operatorname{vec}(\mathsf{B}) = c_{sr} \quad \text{and} \quad C_{lr} \operatorname{vec}(\Xi \mathsf{B}) = c_{lr}, \tag{6}$$

with the usual definitions for $\operatorname{tr}(\cdot)$ and $\operatorname{vec}(\cdot)$: $\operatorname{tr}(\cdot)$ denotes the trace of a matrix and the vec-operator transforms a matrix into a vector by stacking the columns. Using the rules of the vec-operator and a proper selection matrix $C_{\Xi\mathsf{B}}$ we can reformulate the long-run constraint, C_{lr}, as $C_{\Xi\mathsf{B}}(I_k \otimes \Xi)$, in which the operator $\otimes$ denotes the Kronecker product. The reformulation of the long-run constraint reveals its stochastic nature: C_{lr} includes the estimator for Ξ from (4). Finally, C_{sr} specifies short-run constraints by restricting elements of B directly, and Σ_u is the estimated covariance matrix from a reduced-form VECM specified later. We deliberately express the constraints in (6) in linear form in order to make the scoring algorithm of Amisano and Giannini (1997) applicable. The Amisano-Giannini scoring algorithm is numerically simpler and faster than maximizing (5) subject to nonlinear constraints. The scoring algorithm yields an asymptotically efficient and normally distributed maximum likelihood estimator of B (see, e.g., Lütkepohl, 2005, Chap. 9.3.2).

The permanent effects of the structural shocks are given by the matrix $\Xi\mathsf{B}$. As already noted, in a VECM some of the structural shocks have transitory effects only, depending on the cointegrating rank r. We can then restrict r columns in $\Xi\mathsf{B}$ to zero. To be more specific, because the matrix $\Xi\mathsf{B}$ has reduced rank $k-r$, each column of zeros stands for $k-r$ independent restrictions. As such, the r transitory shocks represent $r(k-r)$ independent restrictions only. In total we then have $k(k-1)/2 - r(k-r)$ missing restrictions which can be placed on B and $\Xi\mathsf{B}$ based on other statistical or theoretical considerations. From the statistical side, we get some guidance on how many restrictions we need to place on the contemporaneous impact matrix B. Because each of the r transitory shocks corresponds to a zero column in $\Xi\mathsf{B}$, there is no way to disentangle the transitory shocks with further long-run restrictions. The guideline is then to impose $r(r-1)/2$ restrictions on B directly (see, e.g., Lütkepohl, 2005, Chap. 9.2).

Applied to our 3-dimensional model these considerations imply the following strategy to get the three required restrictions. With the emphasized two cointegration relations we have two transitory shocks and one permanent shock. From that structure of the model we get two independent restrictions from the long-run properties. So there is one more restriction left which has to be imposed on the contemporaneous impact matrix B in order to disentangle the two transitory shocks. Our set of feasible options contains $b_{12} = 0$ or $b_{21} = 0$. The restriction $b_{12} = 0$ is theoretically more appealing: higher profits do not translate into additional labor income contemporaneously, while profits may react swiftly to labor income shocks. Implicitly we assume some rigidities on the labor market here.

3 Estimation Procedures

3.1 The Data

Our data sources are the NIPA and CANSIM tables from the Bureau of Economic Analysis and Statistics Canada. The set of data includes the logarithm of real labor income, corporate profits and output.[5] Specifically, we are using quarterly data from 1947:1 to 2008:4 for the United States and 1961:1 to 2008:4 for Canada. These are the longest possible time spans available for these two countries.

Labor income is the total compensation accruing to employees as remuneration for their work; it is the sum of wage and salary accruals and of supplements to wages and salaries before taxes. There are no transfer payments included. Corporate profits, or profits for short, are the current production incomes before taxes of organizations required to file corporate tax returns. With several differences profits simply consist of receipts less expenses as defined in the tax law. In particular, one such difference in both countries is the exclusion of capital gains and dividends received. Consequently, our two measures of labor income and profits do not add up to output. Multicollinearity is therefore not an issue.

Figure 1 plots the trend and cyclical characteristics of the data. All six series display a strong upward trend with profits being quite volatile, while the labor income and profit shares seem to be—with a few qualifications—relatively stable over the time.

In the United States labor income rose faster than productivity in the years after World War II and the labor income share increased steadily from 52 to 56 percent by the early 1960s. Later on in the 1960s President Johnson's Great Society social reforms mark the sharp increase in the labor income share to about 59 per cent. From then on the labor income share stays at this high level until the stagflation years of the 1970s. Strong output

[5] United States (http://www.bea.gov/National/ - retrieved on May 16, 2009): labor income (1.12, line 2), corporate profits (1.12, line 13), output (1.1.5, line 1), and the output deflator (1.1.4, line 1).

Canada (http://cansim2.statcan.gc.ca/ - retrieved on July 22, 2009): labor income (380-0001, item 2), corporate profits (380-0001, item 3), output (380-0001, item 1), and the output deflator (380-0003, item 1).

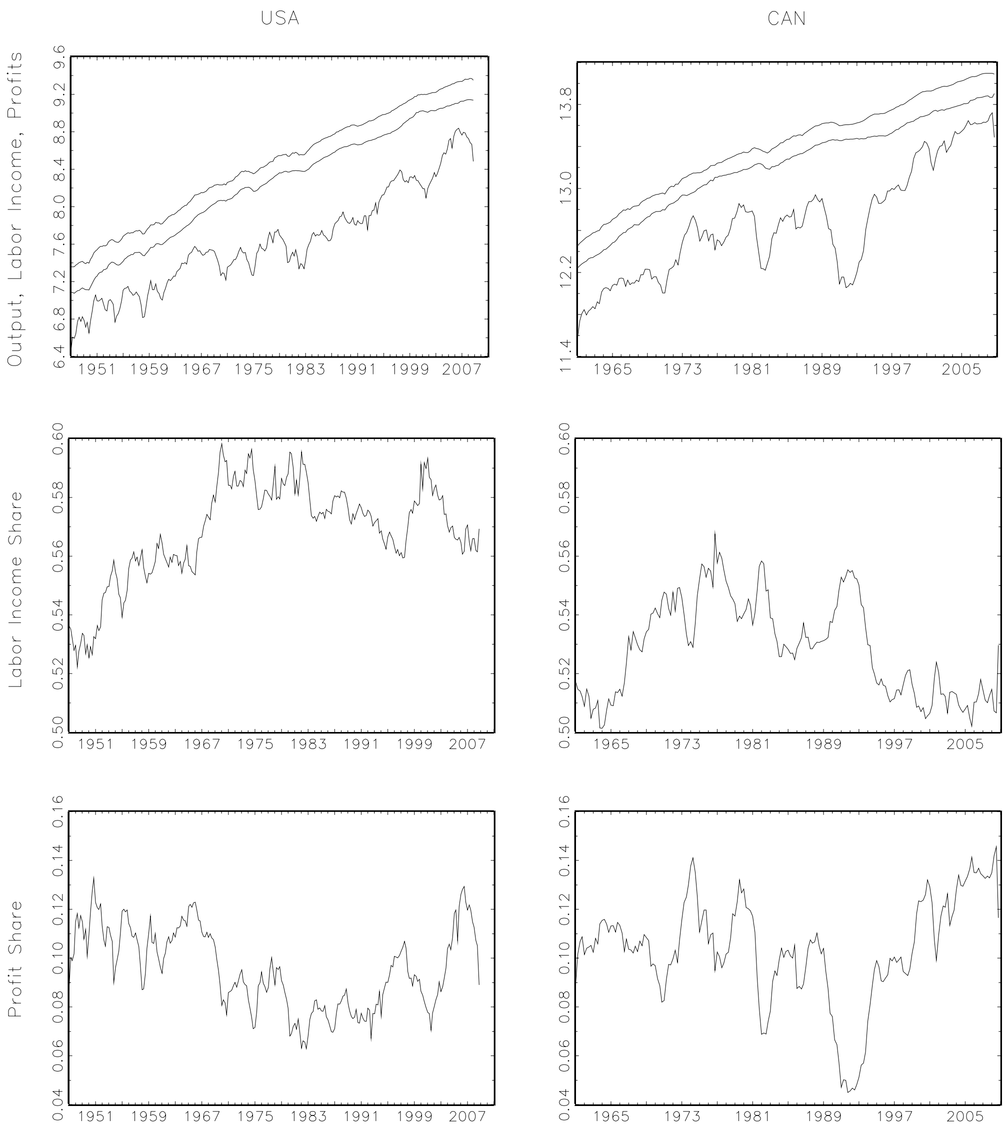

Figure 1 The Time Series with Labor Income-Output and Profit-Output Shares

Notes: Top panel: the order of the time series is output, labor income and profits; all variables are in logarithms and in real terms; to facilitate better graphing we add constants to these variables.

growth after the twin recession of the early 1980s led the labor income share to adjust downwards to a new level of 57 percent. Only at the end of the 1990s the labor income share surges again, mainly influenced by the general economic success of that decade. This development was, however, only short lived and came to a halt with the economic turbulence of the 2000s. Although naturally linked, the steady fall in the profit share in the United States by the mid-1980s had mostly other causes: American geopolitical power and thus the ability of the government to manipulate terms of trade in the interests of its large firms eroded over time; the rise in labor militancy brought on by low unemployment after 1964; and the intensification of competition in the 1960s—reflected both in the erosion of oligopoly pricing power within domestic industries and in increased trade competition from rivals such as Japan and Germany. Then, from the late 1980s onwards labor productivity rose faster than real wages, explaining the bouncing back of the profit share.

On the other side of the border a period known as the Great Canadian Slump dominates the picture. High interest rates, set to bring down inflation to a new target below two percent, played a key role in this deep economic and fiscal crisis of 1990-96 (see, e.g., Fortin 1996, 1999). Moreover, the real wage and employment adjusted such that the labor income share fell persistently back from about 54 percent to its pre-1967 level of 51 percent while the profit share recovered quickly and kept increasing thereafter. Since those turbulent years in the first half of the 1990s the Canadian economy has improved noticeably, in step with the neighbor's boom years. Moreover, Canada has become a role model of fiscal stability as the government has posted surpluses every fiscal year since 1996. Besides the Great Slump the Canadian time series experienced relatively large swings in the shares in the 1970s and early 1980s. During the 1973 oil crisis profits were soaring in oil rich Alberta, before the sharp negative effect of the global oil embargo on the industrial east, which suffered many of the same problems of the United States, swept away the effect of the boom in the west on nationwide profits. Then, from October 1975 to October 1978, the Canadian government installed wage and price controls in order to reduce the rate of inflation while, at the same time, suppressing the Phillips curve effect on unemployment that typically accompanies an anti-inflation policy. The program generally targeted wages by specific numerical guidelines and prices were controlled indirectly through control of profit margins. With these wage and price controls the Canadians followed the United States, which had a similar program in place already a few years early, but building on the experience of its neighbor the Canadians were able to establish a more successful program (see Barber and McCallum, 1982, Chap. 2). On top of that, Canada was hard hit by the recession of the early 1980s, with interest rates, unemployment, and inflation all being higher than in the United States.

Taken together, the visual inspection of the data verifies our hunch that output forms a stationary linear combination with both labor income and profits, at least when properly accounting for possible breaks in the comovement of the data. Our informal discussion of some historical facts in the United States and Canada will be a useful guide in the next section where the aim is to formally estimate and justify the timing of the breaks. The discussion of the historical events based on the shares is appealing, and without loss of generality in terms of the exact cointegration properties, since what matters are the breaks that show up in the comovements and may therefore disguise the "true" cointegrating rank.

3.2 Assumptions and Framework of the Reduced-Form Analysis

Following our empirical strategy we have to estimate the cointegration rank and the break dates. With respect to the latter we extend the setup of Lütkepohl et al. (2004) to allow for two level breaks of unknown timing. A structural break, or break for short, in our context is a rare event that disguises the otherwise stable stochastic comovements in the data, such as cointegrating relations. Ignoring a break may result in a misleading estimate of the cointegrating rank and the equilibrium relations through distorted size and power properties of conventional cointegration tests. Conversely, overestimating the number of structural breaks has the same negative side effects on rank and equilibrium relations.

We incorporate breaks into our analysis under the assumption that the k-dimensional vector of observable variables, $\{y_t\}_{t=1}^{T}$, is at most integrated of order one and has cointegrating rank r with a maximum of two structural breaks. Specifically, the vector process evolves according to

$$y_t = \mu_0 + \mu_1 t + \delta_1 d_{1t} + \delta_2 d_{2t} + x_t, \tag{7}$$

in which μ_0, μ_1, δ_1, and δ_2 are unknown $k \times 1$ parameter vectors; $d_{it} = 1$ for $t \geq T_i$, $i = 1, 2$, and zero otherwise with T_i denoting the time period when a structural break occurs; and x_t is an unobservable stochastic process which we assume to have a VAR(p) representation,

$$x_t = A_1 x_{t-1} + \cdots + A_p x_{t-p} + u_t. \tag{8}$$

The A_j's are the usual $k \times k$ parameter matrices and $u_t = (u_1, \ldots, u_k)'$ are the reduced-form residuals which are i.i.d. vectors with zero mean. This setup of the model can capture the dynamic interactions between the variables and their other properties discussed in Figure 1, such as the trending behavior and possible breaks. We therefore consider our model as a proper representation of the underlying data generating process.

The VECM($p-1$) form of x_t is

$$\Delta x_t = \Pi x_{t-1} + \Gamma_1 \Delta x_{t-1} + \cdots + \Gamma_{p-1} \Delta x_{t-p+1} + u_t, \tag{9}$$

in which Δ is the difference operator such that $\Delta x_t = x_t - x_{t-1}$, and with the obvious mapping $\Pi = -(I_k - A_1 - \cdots - A_p)$ and $\Gamma_j = -(A_{j+1} + \cdots + A_p)$ for $j = 1, \ldots, p-1$. The $k \times k$ matrix Π is of reduced rank, that is $\Pi = \alpha\beta'$ in which both α and β are $k \times r$ matrices of full column rank. We further define $\Psi = I_k - \Gamma_1 - \ldots - \Gamma_{p-1} = I_k + \sum_{j=1}^{p-1} j A_{j+1}$.

We can then write the data generating process (7) in VECM form, and in terms of observable variables only, as

$$\begin{aligned}\Delta y_t = \nu + \alpha\left(\beta' y_{t-1} - \theta_1 d_{1t-1} - \theta_1 d_{1t-1}\right)& \\ + \sum_{j=1}^{p-1} \Gamma_j \Delta y_{t-j} + \sum_{j=0}^{p-1} \left(\gamma_{1j} \Delta d_{1t-j} + \gamma_{2j} \Delta d_{2t-j}\right) + u_t,&\end{aligned} \tag{10}$$

in which Δd_{it-j} are impulse dummies with value one in period $t = T_i + j$ and zero elsewhere. The mapping with the parameters in (7) and (9) is now a bit cumbersome: for $i = 1, 2$ and $j = 1, \ldots, p-1$ we have $\nu = -\Pi\mu_0 + \Psi\mu_1$, $\beta'\mu_1 = 0$, $\mu_1 \neq 0$, $\theta_i = \beta'\delta_i$, $\gamma_{i0} = \delta_i$, and $\gamma_{ij} = -\Gamma_j\delta_i$. Appendix A.1 contains the level VAR version, without rank restrictions on Π, of the VECM.

Apparently, we have sneaked in a few non-trivial assumptions along the way which we now state and discuss more explicitly.

Assumption 1 *At most two structural breaks have occurred in the vector of observable variables,* $\{y_t\}_{t=1}^{T}$.

Admittedly the limit of two structural breaks is arbitrary and mostly determined by computational tractability. We believe, however, that our choice is appropriate and a good compromise to accommodate a sufficiently large number of breaks helping us to uncover the "true" cointegrating rank of Π. If someone still wants to estimate a higher number of structural breaks, the paper of Qu and Perron (2007) is a good point of departure. Strictly speaking, their approach deals with stationary systems. Based on the Bellman principle it offers a quite fast search algorithm over a prespecified maximum number of breaks. Although the Bellman principle provides a beautiful way to reduce the computational burden, it comes at a cost. In some preliminary research (neglecting non-stationarity issues), we find that the Qu-Perron search algorithm is extremely sensitive to the choice of the minimum regime length and the allowed maximum number of breaks.

Assumption 2 *A structural break is a shift in the level of* $\{y_t\}_{t=1}^{T}$*. Furthermore,* $\beta'\mu_1 = 0$ *and* $\mu_1 \neq 0$.

In other words, we purge all linear trends from the analysis. In line with the theoretical considerations about the economy's tendency to return to a steady-state growth path, the equilibrium relations between labor income, profits and output cannot linearly drift apart. Our setup of the data generating process (7) with two breaks is, however, flexible enough to allow the equilibrium relations to drift apart over some extended period. In the United States, for instance, the 1950s and 1960s may represent such a period (see

Figure 1). While we could capture this period with a broken linear trend in the equilibrium relations we choose to model it as an adjustment in the levels. A broken linear trend would "throw away" at least some information in the data (see, e.g., Enders, 2009). We will back up our argument here in the next sections.

We pay little attention to the linear trend in the data generating process (7) because it is the stochastic part which we are ultimately interested in. Generally, a unit root with drift approximates a trending variable well enough in a vector time series model (see, again, Enders, 2009). The linear trend in (7) is then implicitly generated by the intercept in (10) and the non-stationary behavior of the individual variables (see the unit root tests in Table 2). As such, the matrix Π in (9) must be of reduced rank, $r < k$. With full rank, $r = k$, the whole system would be stable and, in such a system, an intercept cannot generate the upward trend apparent in the series.

To bring our framework to life we need to develop procedures to estimate the cointegrating rank, r, and the timing of the two structural breaks, T_1 and T_2 simultaneously. Furthermore, we somehow have to accommodate the "at most two structural breaks" statement of Assumption 1, in order to asses the statistical significance of the breaks. These issues will be our main tasks in the remainder of this section.

3.3 Joint Estimation of Cointegrating Rank and Break Dates

Our joint estimation procedure combines the way how Gregory and Hansen (1996) determine the timing of a break with the procedure of Saikkonen and Lütkepohl (2000a,b) for testing the cointegrating rank of a vector process. The first ingredient, the Gregory-Hansen test, can be thought as the multivariate extension of unit root tests in the tradition of Zivot and Andrews (1992). These tests pick a break date such that the most weight is on the (trend) stationary alternative. The idea behind the second ingredient, the Saikkonen-Lütkepohl test, is to estimate the deterministic terms in (7) first and, in a second step, to apply a likelihood ratio (LR) type test, as in Johansen (1988, 1995), to the adjusted series. The prior adjustment for deterministic terms offers one crucial advantage: multiple breaks in the level leave the limiting distribution of the LR-statistic unaffected. Theorem 4.1 in Lütkepohl et al. (2004) and the remarks thereafter provide a formal discussion of this argument.[6] The Saikkonen-Lütkepohl test is therefore appealing for a grid search over all possible combinations of break dates. The Monte Carlo study in Section 3.5 shows evidence for the practical relevance of our joint estimation procedure.

The joint estimation procedure starts by assigning the optimal lag length, p, to each possible combination of break dates, $\tau = (T_1, T_2)'$. Since, at this point, we do not have any information about the "exact" timing of the breaks and the proper rank specification the level VAR version of (10)—which is unrestricted with respect to the cointegrating rank—proves to be useful (see Appendix A.1). We estimate this version of (10) allowing for a maximum of four lags, $p_{max} = 4$, and under the following definition for T_1 and T_2. Let the breaks be at a fixed fraction, κ_1 and κ_2, of the sample size. Then,

$$T_1 = [\kappa_1 T] \quad \text{and} \quad T_2 = [\kappa_2 T] \quad \text{with} \quad 0.1 \leq \kappa_1 < \kappa_2 \leq 0.9, \qquad (11)$$

in which we impose a 10 percent trimming to eliminate endpoints and $[\cdot]$ denotes the integer part of the argument. Moreover, we set $\kappa_1 < \kappa_2$ such that $[\kappa_1 T] + p_{max} + k \leq [\kappa_2 T]$ in order to avoid singularity in the estimation and to sufficiently identify the break parameters in the cointegration space. For each of these possible pairs $\tau = (T_1, T_2)'$ we choose the optimal lag length according to the corrected Akaike information criterion of Hurvich and Tsai (1993) and denote it by τ_p.[7]

The maximum of the LR-statistics determines then the cointegrating rank, r, and the break dates, τ_p, endogenously through a grid search over

[6] There is, however, a consistency problem inherent in the Saikkonen-Lütkepohl procedure. The parameter μ_0 in (7) is not fully identified. It cannot be estimated consistently in the direction of $\beta_\perp$ and depends partly on the initial values in the procedure. This may be viewed as a drawback of our model setup. Still, we obtain cointegrating rank tests with desirable properties as there is a probability bound on the estimators (see, e.g., Saikkonen and Lütkepohl, 2000b; and Trenkler, Saikkonen and Lütkepohl, 2008).

[7] The Hurvich-Tsai criterion is a correction to Akaike's (1974) criterion and is especially designed for VARs: it makes use of a second order expansion of the Kulback-Leibler divergence and has better small sample properties in VARs than the Akaike criterion.

all possible values for r and τ_p as follows:

$$LR(\tau_p, r_0) = \sup_{\tau_p} -T \sum_{j=r_0+1}^{k} \ln\left(1 - \lambda_j(\tau_p)\right) \qquad r_0 = 0, 1, \ldots, k-1, \quad (12)$$

subject to

$$-T \sum_{j=\tilde{r}_0}^{k} \ln\left(1 - \lambda_j(\tau_p)\right) > LR^c(\tilde{r}_0, k) \qquad \tilde{r}_0 = 0, 1, \ldots, r_0 - 1, \quad (13)$$

in which the $\lambda_j(\tau_p)$'s are the ordered eigenvalues obtained by applying reduced rank regression techniques to the adjusted series, again, as in Johansen (1988, 1995). $LR^c(\cdot)$ denotes the critical values (see, e.g., Trenkler, 2003, Table 2), for which we obtain p-values by approximating the whole asymptotic distribution of the LR-statistic with a Gamma distribution using the response surface procedure of Trenkler (2008). For $r_0 > 0$ the constraint ensures a supremum such that all LR-statistics for ranks lower than r_0 are significant at, say, the 10 percent level.

Finally, from the series of LR-statistics we pick the maximum as

$$LR^* = \max_{r_0=0\ldots,k-2} LR(\tau_p, r_0) \quad \text{such that} \quad LR^* > LR^c(r_0, k), \quad (14)$$

and select the corresponding break dates, $\hat{\tau}_p$, and cointegrating rank, $\hat{r} = r_0 + 1$. If the inequality never holds there is no evidence for cointegration. According to Assumption 2 we exclude the alternative $\hat{r} = k$ (stationarity) and therefore the rank of Π in (9) can be at most $k-1$. With $r = 2$ and an optimal lag order $p = 2$ in the United States and $p = 4$ in Canada we find our VECMs to be adequate representations of the data generating process. Although a multivariate Breusch-Godfrey test indicates some leftover autocorrelation we continue our analysis with this lag specification: autocorrelation does not invalidate the cointegration tests (Lütkepohl and Saikkonen, 1999). Moreover, increasing the lag order does not fix the problem.

Table 1 confirms our hunch that a cointegrating rank of two is a proper choice in both countries. We can reject both hypotheses ($\mathrm{H}_0 : r_0 = 0$ and

Table 1 Joint Estimation of Cointegrating Rank and Break Dates

Country	H_0	Lags, p	Break dates, $\hat{\tau}$		Rank, $\hat{r}$	LR^*-stat.	p-value
USA	$r_0 = 0$	2	1958:1	1988:3		19.442	0.08
	$r_0 = 1$				2	15.370	0.00
CAN	$r_0 = 0$	4	1971:2	1976:2		20.363	0.06
	$r_0 = 1$				2	17.614	0.00

Notes: "Lags" is the order of the VECM in (10) which we assign to each possible pair of break dates prior to our joint estimation procedure. We select p based on the unrestricted model (26) and the corrected Akaike criterion of Hurvich and Tsai (1993). The corresponding p-values of the LR^*-statistics follow from the response surface procedure of Trenkler (2008).

Table 2 Lee-Strazicich LM Unit Root Test with Two Level Breaks

Country	Variable	Lags, p	Break dates, $\hat{\tau}$		LM-statistic	CV (90%)
USA	Labor income	8	1958:1	1988:3	−2.188	−2.763
	Profits	8			−2.531	
	Output	7			−2.122	
CAN	Labor income	8	1971:2	1976:2	−1.066	−2.763
	Profits	8			−2.380	
	Output	8			−0.664	

Notes: Lee and Strazicich (2003) LM unit root test with two exogenous level breaks (their model A). We determine the optimal number of augmented lags by the general-to-specific procedure, starting with a maximum number of $p_{max} = 8$ and using a 10 percent level of significance as the cut off for the last augmented lag. The 90 percent critical value is from Lee and Strazicich (2003), Table 1.

$H_0 : r_0 = 1$) at least at the 8 percent level of significance. By and large, the estimated break dates fit well into the discussion of historical events in Section 3.1. We defer a more in depth discussion of the break dates to the point where we have all statistical facts at hand.

Introducing breaks could potentially render the series trend stationary and Assumption 2 obsolete. We use the minimum Lagrange multiplier unit root test of Lee and Strazicich (2003) to bring clarity into this matter. The

Lee-Strazicich test has several desirable properties. Most importantly for our purpose is the inclusion of the breaks under both the null and alternative hypotheses. A rejection of the unit root null hypothesis therefore implies unambiguously trend stationarity. Table 2 shows the test results. Based on the model A version of the Lee-Strazicich test, the equivalent to our model setup in (7), we cannot reject the null for none of the six series at levels lower than the 10 percent level of significance. A result which supports the appropriateness of our Assumption 2.

The property of an unaffected limiting distribution is in stark contrast to a Saikkonen-Lütkepohl test with broken linear trends (see, e.g, Trenkler, Saikkonen and Lütkepohl, 2008) and to the other branch of cointegration tests with breaks, such as Johansen et al. (2000). These tests require a new set of critical values whenever the timing of a break changes. These critical values are larger the more balanced the various regimes are. Naturally, this property is unpractical for a grid search as it introduces a bias toward picking one relatively long regime: a reason why we refrain from using broken linear trends in our analysis.

To sum up, our joint estimation procedure provides a novel way to estimate cointegrating rank and break dates. It extends the two-step procedure of Lütkepohl et al. (2004) and Saikkonen et al. (2006) in which the break dates follow from minimizing the determinant of the residual covariance matrix of the level VAR version of (10) in a first step. Then, in a second step, the cointegrating rank is determined given the breaks. The use of different models at each step introduces a pre-test bias into their procedure. We instead use the level VAR version only to assign the optimal lag order to each possible pair of break dates. Therefore, our joint estimation procedure removes one layer of uncertainty.

3.4 The Reduced-Form VECM Estimator with Parameter Restrictions

There is by now one dominant method for estimating a reduced-form VECM: the reduced-rank maximum likelihood (ML) method of Johansen (1995). Although the ML estimation of VECMs is quite common under practitioners it may produce occasional outlying estimates of the cointegration parameters (see, e.g., Brüggemann and Lütkepohl, 2005; and Johansen, 1995, p. 184). A feature arising through the lack of finite-sample moments of the estimator (Phillips, 1994).

Thus, instead we use a two-stage feasible generalized least squares (GLS) estimator which does not share the unpleasant feature of the ML estimator. This alternative estimator for VECMs was first proposed by Ahn and Reinsel (1990). In addition, it offers a computationally simple and unproblematic way to place restrictions on the cointegrating matrix. Since the two-stage feasible GLS method has not attracted much attention under practitioners we briefly introduce the estimator here (see, e.g., Lütkepohl, 2005, Chap. 7.2.2 for a textbook treatment).

Consider a general k-dimensional sample $\{y_t\}_{t=1}^{T}$ with pre-sample values $\{y_t\}_{t=1-p}^{0}$, leaving the specific three variable case aside for the moment. To estimate the VECM specification in (10), let us define the variable matrices $\Delta Y = (\Delta y_1, \ldots, \Delta y_T)$, $Y_{-1} = (X_0^{lr}, \ldots, X_{T-1}^{lr})$, and $\Delta X = (\Delta X_0^{sr}, \ldots, \Delta X_{T-1}^{sr})$ in which the long-run ("lr") and short-run ("sr") matrices are $X_{t-1}^{lr} = (y_{t-1}, d_{1t-1}, d_{2t-1})'$ and $\Delta X_{t-1}^{sr} = (\Delta y_{t-1}, \ldots, \Delta y_{t-p+1}, 1_T, \Delta d_{1t}, \ldots, \Delta d_{1t-p+1}, \Delta d_{2t}, \ldots, \Delta d_{2t-p+1})'$. As in Section 3.2, we define the $n_d = 2$ break dummies $\{d_{it}\}_{t=1-p}^{T}$, $i = 1, 2$ such that the sequence is one for $t \geq T_i$ and zero otherwise. The corresponding parameter matrices are $\Pi = \alpha\beta^{*\prime}$ with $\beta^* = (\beta : \theta_1 : \theta_2)$, and $\Gamma = (\nu : \Gamma_1 : \cdots : \Gamma_{p-1} : \gamma_{10} : \cdots : \gamma_{1p-1} : \gamma_{20} : \cdots : \gamma_{2p-1})$.

Then, the VECM (10) rewritten in matrix notation is

$$\Delta Y = \Pi Y_{-1} + \Gamma \Delta X + U, \tag{15}$$

for which we can write the OLS estimator as

$$\left[\widehat{\Pi} : \widehat{\Gamma}\right] = \left[\Delta Y Y_{-1}' : \Delta Y \Delta X'\right] \begin{bmatrix} Y_{-1}Y_{-1}' : Y_{-1}\Delta X' \\ \Delta X Y_{-1}' : \Delta X \Delta X' \end{bmatrix}^{-1}. \tag{16}$$

Our focus is now to disentangle the cointegration matrix β from Π. For this purpose we work with the concentrated version of the VECM and replace the short-run parameters with their OLS estimators given Π, i.e.

$$R_0 = \Pi R_1 + U = \alpha\beta^{*\prime} R_1 + U, \tag{17}$$

in which R_0 and R_1 are the residual matrices from regressing ΔY and Y_{-1} on ΔX. We further split R_1 into its first r and last $k-r$ rows and denote the two sub-matrices by $R_1^{(1)}$ and $R_1^{(2)}$. Together with the common identifying restriction $\beta^* = (I_r : \beta_{k-r} : \theta_1 : \theta_2)$ we can then rewrite the concentrated model as

$$R_0 - \alpha R_1^{(1)} = \alpha\beta_{k-r}^{*\prime} R_1^{(2)} + U. \tag{18}$$

Further, we introduce possible over-identifying restrictions, of the form

$$\text{vec}\left(\beta_{k-r}^{*\prime}\right) = \mathbf{R}\varphi + \mathbf{r}, \tag{19}$$

on the cointegration matrix $\beta_{k-r}^* = (\beta_{k-r} : \theta_1 : \theta_2)$. φ is an unrestricted $m \times 1$ vector of unknown parameters, while $\mathbf{r}$ is the $r(k-r+n_d)$-dimensional vector of imposed constants; and $\mathbf{R}$ is an appropriately defined zero-one matrix of dimension $r(k-r+n_d) \times m$ such that (19) holds. A simple Wald test can be used to check the restrictions. Under the null hypothesis of statistical valid restrictions the Wald statistic has an asymptotic χ^2-distribution with the number of restrictions as degrees of freedom (see, e.g., Lütkepohl, 2004).

We confine ourselves to allowing restrictions on the cointegration matrix only. Our sole aim is to have a statistical tool for analyzing the structural breaks we have estimated with our joint estimation procedure. Essentially, we are testing the significance of the individual break parameters in the θ_1 and θ_2 vectors. Since in (10) the θ's are linear combinations of the δ's from our data generating process, a zero restriction here does not automatically imply a zero restriction there. As such, we estimate the parameter vectors of the impulse dummies—the γ's in (10)—unrestrictedly. According to Saikkonen and Lütkepohl (2000a), ignoring any form of restrictions on the impulse dummies will not do great damage to the other estimators.

We plug the restrictions (19) into the concentrated model (18) and solve for the GLS estimator of φ:

$$\breve{\varphi} = \left[\mathbf{R}'\left(R_1^{(2)}R_1^{(2)\prime}\right)\alpha'\Sigma_u^{-1}\alpha\right]^{-1} \times \mathbf{R}'\left(R_1^{(2)}\otimes\alpha'\Sigma_u^{-1}\right)\left[\operatorname{vec}\left(R_0-\alpha R_1^{(1)}\right)-\left(R_1^{(2)\prime}\otimes\alpha\right)\mathbf{r}\right]. \quad (20)$$

To make GLS operational (i.e. feasible) we need to replace the loading matrix, α, and the residual covariance matrix, Σ_u, with consistent estimators. For Σ_u such an estimator follows from (16) and (15) in the usual way and for α from the first r columns of $\widehat{\Pi}$. The estimator for α is a direct implication from the common identifying restriction we have imposed on β^*. We denote the resulting feasible GLS estimator by $\breve{\varphi}$. The unrestricted version of $\breve{\varphi}$ is available by defining $\mathbf{R}$ as an identity matrix of dimension $r(k-r+n_d)$ and $\mathbf{r}$ as a vector of zeros.

Proposition 7.6 in Lütkepohl (2005) shows the asymptotic properties of the estimator. $\breve{\varphi}$ goes in distribution to a Normal. Formally,

$$\left[\mathbf{R}\left(R_1^{(2)}R_1^{(2)\prime}\right)\otimes\left(\alpha'\Sigma_u^{-1}\alpha\right)\mathbf{R}\right]^{1/2}(\breve{\varphi}-\varphi)\rightsquigarrow N\left(0,I_k\right), \quad (21)$$

in which $\breve{\varphi}$ converges to its true value at the rate T, faster than the usual rate $\sqrt{T}$. It is therefore what we call a super-consistent estimator. Whether we impose a known cointegration matrix β^* or estimate it will not affect the asymptotic distribution of the OLS estimators of Π and Γ in (16).

Then, given the super-consistent estimator $\breve{\varphi}$, we can construct $\breve{\beta}^* = (I_r : \breve{\beta}^*_{k-r})'$ from (19) and consistently estimate all the other parameters in the VECM in a second stage. With $\breve{\beta}^*$ and possible zero restrictions of the form

$$\operatorname{vec}(\alpha:\Gamma) = \mathbf{S}\vartheta, \quad (22)$$

we can write (15) in vectorized form as

$$\operatorname{vec}(\Delta Y) = \left(\left[Y_{-1}'\breve{\beta}^* : \Delta X'\right]\otimes I_k\right)\mathbf{S}\vartheta + \operatorname{vec}(U), \quad (23)$$

in which $\mathbf{S}$ is a fixed zero-one matrix of dimension $k(r*k(p-1))\times n$ and ϑ

is a n-dimensional vector of free parameters. The GLS estimator of ϑ is now

$$\breve{\vartheta} = \left[\mathsf{S}' \left(\begin{bmatrix} \breve{\beta}^{*\prime} Y_{-1} Y_{-1}' \breve{\beta}^* : \breve{\beta}^{*\prime} Y_{-1} \Delta X' \\ \Delta X Y_{-1}' \breve{\beta}^* : \Delta X \Delta X' \end{bmatrix} \otimes \Sigma_u^{-1} \right) \mathsf{S} \right]^{-1} \times \mathsf{S}' \left(\begin{bmatrix} \breve{\beta}^{*\prime} Y_{-1} \\ \Delta X \end{bmatrix} \otimes \Sigma_u^{-1} \right) \operatorname{vec}(\Delta Y). \quad (24)$$

A consistent estimator for the residual covariance matrix, Σ_u, is readily available from the first stage and, as before, the feasible GLS estimator of ϑ goes in distribution to a Normal (see Lütkepohl, 2005, Proposition 7.7):

$$\sqrt{T}(\breve{\vartheta} - \vartheta) \rightsquigarrow N\left(0, \operatorname{plim} T \left[\mathsf{S}' \left(\begin{bmatrix} \breve{\beta}^{*\prime} Y_{-1} Y_{-1}' \breve{\beta}^* : \breve{\beta}^{*\prime} Y_{-1} \Delta X' \\ \Delta X Y_{-1}' \breve{\beta}^* : \Delta X \Delta X' \end{bmatrix} \otimes \Sigma_u^{-1} \right) \mathsf{S} \right]^{-1} \right), \quad (25)$$

in which $\breve{\vartheta}$ converges to its true values at the rate $\sqrt{T}$. A Wald test is again available to check the statistical validity of the restrictions.

3.5 Taking Stock

We now put the bits and pieces of the preceding sections together and present the results of the long-run parameters and the breaks. In addition, we provide evidence for the practical relevance of the two-stage feasible GLS estimator and a Monte Carlo study to assess the performance of our joint estimation procedure of cointegrating rank and break dates.

Two results stand out from Tables 3 and 4. First, there is no evidence for structural breaks in Canada. A joint test for the exclusion of the breaks cannot be rejected by a Wald test ($\chi^2(4) = 2.128$ [0.71]). For this result we get further support from a standard Saikkonen-Lütkepohl cointegration test without breaks: we still find two cointegration relations. Unlike Canada, the 1958:1 and 1988:2 breaks in the United States are partly necessary to reveal a cointegrating rank of two. These dates fit well into our discussion of the historical developments in Section 3.1. Specifically, the first break is only

Table 3 United States: Estimated Long-Run Parameters

Model		Labor income	Profits	Output	Break, T_1	Break, T_2
					1958:1	1988:3
M_0:	$\breve{\beta}^{*\prime}$	1.000	0.000	-1.024	-0.037	0.038
				(0.015)	(0.014)	(0.014)
		0.000	1.000	-0.801	0.035	-0.267
				(0.129)	(0.127)	(0.123)
	$\breve{\alpha}'$	-0.135	-0.848	-0.121		
		(0.032)	(0.252)	(0.041)		
		-0.003	-0.131	-0.009		
		(0.004)	(0.030)	(0.005)		
M_1:	$\breve{\beta}^{*\prime}$	1.000	0.000	-1.000	-0.045	0.000
					(0.008)	
		0.000	1.000	-0.865	0.000	-0.116
				(0.078)		(0.091)
				Wald statistic: $\chi^2(3) = 4.020$ [0.26]		
	$\breve{\alpha}'$	-0.092	-0.724	-0.090		
		(0.024)	(0.213)	(0.033)		
		0.000	-0.122	-0.008		
			(0.027)	(0.003)		
				Wald statistic: $\chi^2(1) = 0.173$ [0.68]		

Notes: Estimation by two-stage feasible GLS with $p = 2$ lags. The corresponding standard errors of the parameters are in parentheses; the p-values of the Wald statistics are in brackets. In the second stage (equation (22)), we estimate the loading matrix, $\breve{\alpha}$, either under M_0 or M_1 by taking the cointegration matrix, $\breve{\beta}^{*\prime}$, from the first stage (equation (19)) as given. The imposed restrictions represent the maximal possible that cannot be rejected by a Wald test.

significant in the labor income-output equation and adjusts for the long and steady increase of the labor income share after World War II while the 1988:3 break captures the U-shaped profile of the profit share. As we see from the empirical evidence here, our notion of a structural break as rare events that disguise the "true" cointegrating rank seems to be appropriate.

The second result is that a permanent increase in output sets in motion a redistribution from labor income toward profits in Canada, whereas in the

Table 4 Canada: Estimated Long-Run Parameters

Model		Labor income	Profits	Output	Break, T_1	Break, T_2
					1972:2	1976:2
M_0:	$\breve{\beta}^{*\prime}$	1.000	0.000	−0.911	0.012	0.009
				(0.038)	(0.039)	(0.041)
		0.000	1.000	−1.420	−0.030	0.260
				(0.297)	(0.306)	(0.321)
	$\breve{\alpha}'$	−0.056	−0.221	−0.051		
		(0.014)	(0.115)	(0.013)		
		0.004	−0.057	−0.000		
		(0.003)	(0.023)	(0.003)		
M_1:	$\breve{\beta}^{*\prime}$	1.000	0.000	−0.896	0.000	0.000
				(0.021)		
		0.000	1.000	−1.231	0.000	0.000
				(0.169)		
				Wald statistic: $\chi^2(4) = 2.128$ [0.71]		
	$\breve{\alpha}'$	−0.053	−0.232	−0.055		
		(0.015)	(0.119)	(0.013)		
		0.006	−0.048	0.000		
		(0.002)	(0.018)			
				Wald statistic: $\chi^2(1) = 0.601$ [0.44]		

Notes: Estimation by two-stage feasible GLS with $p = 4$ lags. The corresponding standard errors of the parameters are in parentheses; the p-values of the Wald statistics are in brackets. In the second stage (equation (22)), we estimate the loading matrix, $\breve{\alpha}$, either under M_0 or M_1 by taking the cointegration matrix, $\breve{\beta}^{*\prime}$, from the first stage (equation (19)) as given. The imposed restrictions represent the maximal possible that cannot be rejected by a Wald test.

United States we observe a redistribution away from (corporate) profits in favor of other non-labor incomes, such as capital gains and dividends. A Wald test cannot reject a one-by-one long-run movement of labor income and output (see the joint test with the breaks: $\chi^2(3) = 4.020$ [0.26]), while the long-run output elasticity of profits is significantly below one (0.87 percent). Since our measures of labor income and profits do not (and must not) add up to output, a permanent increase in output sets in motion a redistribution toward the "missing" part of overall profits, most importantly capital gains and dividends. This trend may reflect the relative dominant role of the stock markets in the United States. Across the border in Canada we estimate long-run output elasticities of labor income (0.90 percent) and profits (1.23 percent) significantly below and above unity. These long-run elasticities imply a redistribution from labor income to profits in response to a permanent increase in output: economic growth does not fully show up on Canadians' paychecks. This is a well-known fact in Canada and part of the, sometimes tempered, political discussion.

The interpretation of cointegrating parameters as long-run elasticities can be problematic but works well in our case. The reason is the orthogonality property of the permanent change: it is orthogonal to both cointegrating vectors. To see why this property holds, let us consider a permanent one-percent output change that entails—everything else equal—a permanent change of labor income by $-\beta_{13}$ percent.[8] Having in mind the parallel (unknown) effect on profits, say ω, we model the permanent change, c, as $c(\omega) = (-\beta_{13}, \omega, 1, 0, 0)$. Then,

$$c'(\omega)\beta_1^* = (-\beta_{13}, \omega, 1, 0, 0)(1, 0, \beta_{13}, \theta_{11}, \theta_{12})' = 0$$
$$c'(\omega)\beta_2^* = (-\beta_{13}, \omega, 1, 0, 0)(0, 1, \beta_{23}, \theta_{21}, \theta_{22})' = \omega + \beta_{23},$$

in which we can set $\omega = -\beta_{23}$ to make $c(\omega)$ orthogonal to both cointegrating vectors. It is exactly this change in profits, the variable absent from the labor income-output relation, that keeps the system on the attractor set. Only in cases where this condition is met, the initial permanent change, c, produces the required effect and we can interpret $-\beta_{13}$ as the long-run output elasticity of labor income. Likewise, $-\beta_{23}$ is the long-run output elasticity of profits. This result is the essence of Propositions 1 and 2 in Johansen (2005).

As a cross-check for the robustness and practical relevance of the two-stage feasible GLS estimator we compare our results with the ones from the

[8] β_{13} is the output parameter in the labor income-output relation. Formally, in (10) we have $\beta^* = \begin{bmatrix} 1 & 0 & \beta_{13} & \theta_{11} & \theta_{12} \\ 0 & 1 & \beta_{23} & \theta_{21} & \theta_{22} \end{bmatrix}$, and we denote the first and second row of β^* by β_1^* and β_2^*.

reduced-rank ML method of Johansen (1995). While for the United States the long-run parameters are practically identical, the results for Canada are unreasonable and are quite likely outliers. For instance, the output elasticity of profits is -19.03 in the unrestricted model (M_0). We rule it out as a valid result, both sign and size are economically implausible. In the same way, including a linear trend in the cointegrating space causes troubles. In the United States, it suggests an output elasticity of profits of -3.03. The results for Canada are similar. Again, sign and size are hard to reconcile with common theoretical considerations on the long-run relation between profits and output.

Table 5 presents the results from our Monte Carlo study on the performance of the joint estimation procedure. We use the parameter estimates of the VECM in unrestricted form (model M_0) and simulate 1,000 series of the the original sample length using multivariate normal residuals. We impose structure on the residuals by pre-multiplying them with the contemporaneous impact matrix B, where B is from the estimation stage. We use the first p observations ($p = 2$ in the United States and $p = 4$ in Canada) to initialize all 1,000 runs. We then go through all the steps of Section 3.3, while keeping the lag order p fixed, and store the estimated break dates, $\hat{\tau}$, and cointegrating rank, $\hat{r}$, in each run. Table 5 shows the frequencies of $\hat{\tau}$ to lie in certain time intervals. Since the shift magnitudes of the breaks are relatively small (see Tables 3 and 4), our test performs modestly in finding the correct break dates. In an interval of plus-minus three quarters around the two break dates the hit rate is 5.7 percent in the United States and 1.9 percent in Canada. The somewhat larger shift magnitudes in the United States increase the frequency of finding the break dates closer to the true ones. While, at first glance, this performance seems to be rather disappointing, what matters most is the success of our test to identify the correct cointegrating rank. In most instances the hit rate is a full 100 percent. Even when the estimated break dates are outside the plus-minus three quarters intervals the cointegration test within our joint estimation procedure performs reasonably well. As a benchmark, using the standard Saikkonen-Lütkepohl cointegration test without breaks, the frequency of finding the correct cointegrating rank is 17.4 percent in the

United States and 73.9 percent in Canada. With one exception in Canada, these hit rates are significantly lower than the ones from our joint estimation procedure. Especially so in the United States where the shift magnitudes play a more important role than in Canada.

4 Empirical Results on Redistribution

After these technicalities let us summarize the results of our estimations regarding the effects of a redistribution from profits to labor income. We start by discussing the contemporaneous relations in Table 6 and move then on to the dynamic effects. Figures 2 and 3 present the effects of the two components of a redistribution shock—labor income and profits—individually. This intermediate step helps for a better intuition for the driving forces behind the changes set in motion by a combination of these two shocks in the redistribution experiment shown in Figure 4. Deriving an interpretation directly from Figure 4 is in fact a bit tricky since both effects overlap and the behavior of workers and firms and the respective aggregates becomes less clear.

Although the original estimates have the dimension of elasticities, it is more intuitive to discuss the results on a dollar for dollar basis. As such, we work with derivatives evaluated at the point of means. Moreover, when we talk about "dollars" we do not distinguish between United States an Canadian dollars for ease of reading. This would anyway be only a matter of labeling a unit.

Throughout the discussion of the results we mostly focus on the model with parameter restrictions (M_1) which is our preferred model based on the discussion above. However, to check the robustness and sensitivity we report the results from the unrestricted model M_0 as well.

4.1 Contemporaneous Effects

Table 6 reports the estimation results for the B matrix, based on the unrestricted model, M_0, and the restricted one, M_1, from above. The estimation procedure follows Section 2 which we extend to allow for over-identifying restrictions on the B matrix in the restricted model. We denote this model

Table 5 Joint Estimation Procedure: Monte Carlo Analysis

A. United States

	47:1–87:3	87:4–88:2	1988:3	88:4–89:2	89:3–08:4	Sum
47:1–57:1	0.211	0.007	0.001	0.010	0.030	0.259
	(0.41)	(1.00)	(1.00)	(1.00)	(1.00)	
57:2–57:4	0.044	0.001	0.002	0.005	0.019	0.071
	(1.00)	(1.00)	(1.00)	(1.00)	(1.00)	
1958:1	0.072	0.007	0.002	0.022	0.066	0.169
	(1.00)	(1.00)	(1.00)	(1.00)	(0.99)	
58:2–58:4	0.059	0.005	0.006	0.007	0.050	0.127
	(1.00)	(1.00)	(1.00)	(1.00)	(1.00)	
59:1–08:4	0.110	0.021	0.008	0.031	0.209	0.379
	(0.99)	(0.95)	(1.00)	(0.97)	(0.99)	
Sum	0.495	0.040	0.018	0.074	0.372	

B. Canada

	61:1–75:2	75:3–76:1	1976:2	76:3–77:1	77:2–08:4	Sum
61:1–71:2	0.095	0.009	0.004	0.043	0.343	0.494
	(0.88)	(1.00)	(1.00)	(1.00)	(0.99)	
71:3–72:1	0.020	0.002	0.000	0.013	0.067	0.102
	(1.00)	(1.00)		(1.00)	(0.99)	
1972:2	0.002	0.000	0.000	0.002	0.012	0.016
	(0.50)			(1.00)	(0.92)	
72:3–73:1	0.002	0.000	0.000	0.002	0.024	0.028
	(1.00)			(1.00)	(1.00)	
73:2–08:4	0.000	0.001	0.000	0.004	0.355	0.360
		(1.00)		(1.00)	(0.99)	
Sum	0.119	0.012	0.004	0.064	0.801	

Notes: Data generating process based on the unrestricted parameter estimates (model M_0, equation (10)) for the two countries. We present the relative frequencies (out of $1,000$ replications) of finding the break dates, $\hat{\tau}$, in a specific time interval and report in parentheses the ratio of how often the estimated cointegrating rank, $\hat{r}$, is $r = 2$ (at the 10 percent level of significance).

by M_1^*. Four results stand out. First, the contemporaneous effect from a labor income shock on output dominates the one from a profit shock, i.e. $b_{32} < b_{31}$. Specifically, a one-dollar shock to labor income increases output within the quarter by about 1.15 dollars in both countries. Thus, there is a modest impact multiplier effect at work creating these 15 cents in excess of the initial one-dollar input through the swift effect of the induced extra spending on aggregate demand. In the case of a one-dollar profit shock the retarded capacity effect implies a less than unity increase in output: 0.68 dollars in the United States and 0.82 dollars in Canada.

The second result is that the correlation between the reduced-form labor income and profit residuals is relatively high in the United States (about 0.43 in our sample) yielding a positive effect from a labor income shock on profits ($b_{21} = 0.47$). From a methodological point of view this effect implies that to achieve a one-dollar redistribution from profits to labor income we have to take away more than one dollar from firms, 1.47 dollars to be exact. In Canada the effect from a labor income shock on profits is, if at all, slightly negative which is a direct consequence of the low correlation between the reduced-form residuals (-0.05). In model M_0 this effect is a mere -0.03 and rather imprecisely estimated. Moreover, a Wald test cannot reject the over-identifying restriction $b_{21} = 0$. The parameter is therefore absent from M_1^*.

Third, the parameters b_{13} and b_{23} indicate the qualitative effect of the automatic responses. Indicative only because b_{13} and b_{23} subsume various other effects, for instance the marginal propensities to spend, as in (2). The sign of the parameters, however, may ultimately be driven by the sign of the respective automatic response (see Remark 1). Both automatic response channels seem to be positive in the United States, whereas in Canada the effect is negative and the other one is statistically not different from zero.

Fourth, the standard errors under M_0 are quite large which points toward a problematic identification and estimation of at least some of the parameters. These problems, however, seem to disappear in the model with parameter restrictions (M_1^*), where we observe a considerably reduced estimation uncertainty. Still, we might underestimate the "true" estimation uncertainty

Table 6 Contemporaneous Effects

Country	Model		b_{31}	b_{21}	b_{32}	b_{13}	b_{23}
USA	M_0	coeff.	1.199	0.473	0.663	0.190	0.194
		std.err.	(0.732)	(0.473)	(0.660)	(0.225)	(0.358)
	M_1^*	coeff.	1.156	0.471	0.682	0.205	0.186
		std.err.	(0.466)	(0.178)	(0.225)	(0.119)	(0.144)
CAN	M_0	coeff.	1.283	−0.026	0.592	−0.336	0.263
		std.err.	(0.360)	(0.074)	(0.354)	(0.332)	(0.744)
	M_1^*	coeff.	1.295	0.000	0.822	−0.360	0.000
		std.err.	(0.242)		(0.355)	(0.224)	
				Wald statistic: $\chi^2(2) = 5.388$ [0.068]			

Notes: The parameter estimates refer to the B matrix in equation (1) under the short-run identifying restriction $b_{12} = 0$. Model M_1^* extends M_1, the one with parameter restrictions in the long-run relations (see Tables 3 and 4), and tests over-identifying restrictions on the B matrix. We present the parameters as derivatives evaluated at the point of means and express them as dollar for dollar. b_{31} and b_{32} are the effects of labor income and profit shocks on output; b_{12} and b_{21} are the effects of a profit shock on labor income and vice versa ; and b_{13} and b_{23} indicate the direction (not the size) of the automatic responses. The imposed restrictions represent the maximal possible that cannot be rejected by a Wald test. Bootstrapped standard errors in parentheses (2,500 replications, see Appendix A.2).

as we take the restrictions as given in the bootstrap procedure.

What could be the economic explanations behind these results and the differences between the two countries? In the following we shortly discuss two arguments which might comprise important and reasonable explanations.

A first argument could be that the stronger exposure of Canada to international trade brings some unwelcome side effects of a positive labor income shock: an increase in unit labor costs and a loss in competitiveness. These effects will reduce exports and may undo any additional profit opportunities arising through the boost in domestic demand. The automatic response channel from output on profit is, therefore, absent in Canada ($b_{23} = 0$).

A second potential explanation could stem from the formation of expectations which might differ in the two countries. The prospects about the beneficial effects of a labor income shock for the future may heavily influence the decisions of American firms and workers. These positive expectations, then, reinforce the spending effects already within the quarter. Labor income reacts by more than the initial one-dollar shock ($b_{31} = 1.16$) and profits increase by 0.47 dollars. In Canada, however, the picture is less clear. Labor income still reacts by more than the initial input ($b_{31} = 1.30$), but it does not create the additional boost as in the United States. Firms just manage to maintain their profit levels ($b_{21} = 0$). Canadian firms might interpret a labor income shock in terms of a cost shock and hence do not increase capacities. This different interpretations of a labor income shock in the two countries might itself be explained by the fact that over the last 40 years the quantitative significance of unions and collective bargaining drifted apart in the two countries, with Canada experiencing an increasing unionization and higher bargaining power of workers (Riddell, 1993). On the other hand, the output effect of a profit shock is with 0.82 dollars relatively high. Taken together, these results give the hunch that spending drives the formation of expectations in the United States, whereas capacity considerations may be the driving factor in Canada.

4.2 Dynamic Effects of a Labor Income Shock

Figure 2 depicts the effects of a one-dollar labor income shock. One would expect a positive effect on the level of output mainly driven by additional spending. This effect can clearly be seen in the response of output. It is about one-for-one on impact with a multiplier effect at work thereafter, reaching a peak three quarters out at 1.92 dollars in the United States and 1.45 dollars in Canada after one quarter. Output starts then to decline in both countries but rather abrupt in Canada with the effect becoming statistically insignificant already after four to five quarters. The impulse responses from the unrestricted and restricted models lie practically on top of each other. Apparently, the exact specification of the restrictions on the breaks and the contemporaneous effects does not really matter here.

The formation of expectations after a labor income shock drives the dynamic pattern in the Unites States. Workers increase their spending and firms will produce more by using idle capacities or by investing in new capacities.

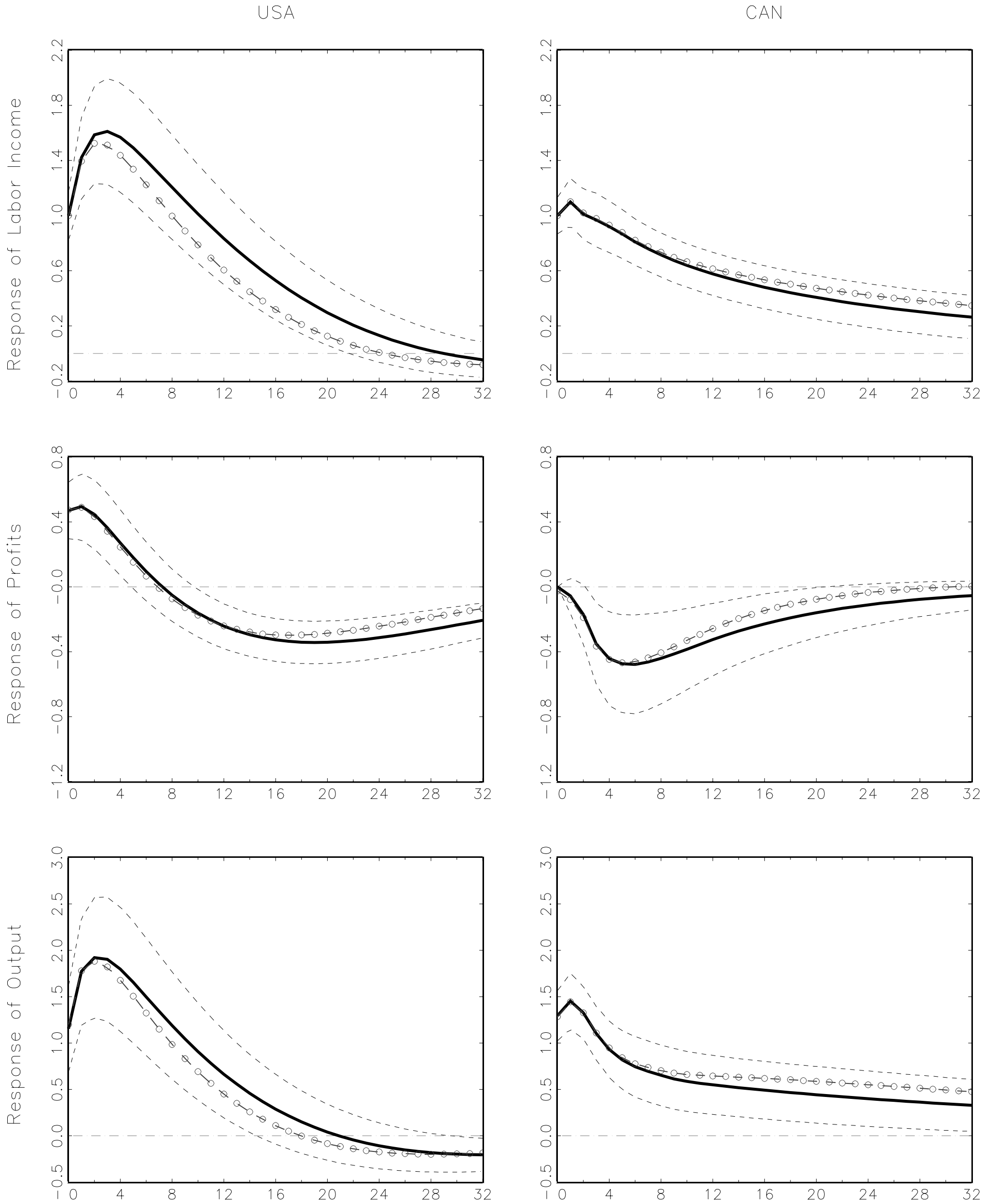

Figure 2 Response to a One-Dollar Labor Income Shock

Notes: The thick solid line depicts the impulse response from the model with parameter restrictions on the cointegration and the contemporaneous impact matrices (model M_1^*), surrounded by bootstrapped one-standard deviation bands (2,500 replications, see Appendix A.2). The dashed line with circles depicts the response based on the VECM with no parameter restrictions (model M_0).

This positive short-run effect is then phasing out over the medium-run as workers ask for higher wages in return for higher productivity and additional labor demand. As a consequence, profits eventually go below trend after six quarters. Together with the general upward adjustment of prices this "classic" channel explains the pronounced hump-shaped response of labor income and output.

In Canada where we have not found a significant contemporaneous reaction of profits, only higher labor income drives the multiplier process. The missing positive impact on firms arise, perhaps, through a different formation of expectations or a loss of competitiveness as discussed above. When hit by a labor income shock the spending stimulus does not lead Canadians to revise their expectations as much as their colleagues across the border. This difference induces a much faster decline of output in Canada and a relatively strong negative profit effect that reaches its trough seven quarters out at -0.48 dollars.

4.3 Dynamic Effects of a Profit Shock

Let us now turn to the effects of a one-dollar profit shock. We model a negative shock because of its later relevance to resemble a redistribution shock toward labor income according to Definition 1. Theoretically, the lower profits will have a negative effect on capacity decisions and will, perhaps, lead to a revision of expectations. In a standard AS-AD model (see, e.g., Blanchard, 2009) prices will then adjust downward over the medium-run and output reverts to its initial trend.

Figure 3 shows the empirical pattern. In the United States profits start recovering more or less right away, whereas in Canada profits respond in a hump-shaped manner, reaching a trough four quarters out at -1.51 dollars. By our short-run identifying assumption ($b_{12} = 0$), labor income does not change on impact, decreases then relatively smoothly before it reverts to trend. The higher influence of unions and collective bargaining makes the trend reversion particularly pronounced and slow in Canada. Output initially decreases by 0.68 dollars and reaches a trough at -0.99 dollars after four quarters in the United States; by 0.82 dollars and a trough at -1.91 dollars

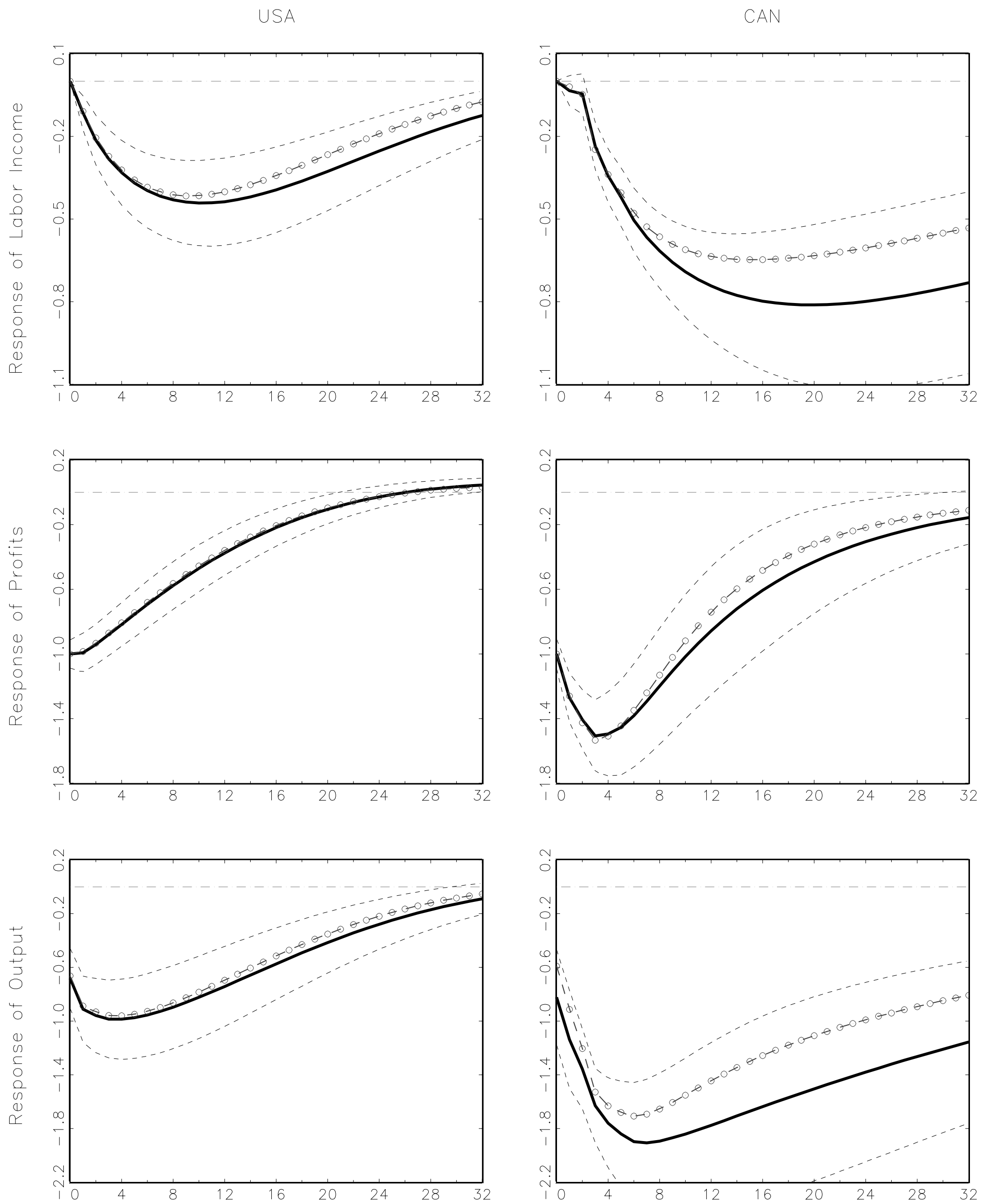

Figure 3 Response to a (Negative) One-Dollar Profit Shock

Notes: The thick solid line depicts the impulse response from the model with parameter restrictions on the cointegration and the contemporaneous impact matrices (model M_1^*), surrounded by bootstrapped one-standard deviation bands (2,500 replications, see Appendix A.2). The dashed line with circles depicts the response based on the VECM with no parameter restrictions (model M_0).

eight quarters out in Canada. Moreover, besides the much larger (negative) multiplier effect the output response in Canada is rather persistent.

Again, the formation of expectations plays a crucial role. After a profit shock, firms will reduce capacities or postpone investment projects and workers will cut down spending eventually. Especially in Canada, the decrease of output by 0.82 dollars on impact, and profits much below par, puts pressure on the labor market. A situation in which firms may become tougher on wage negotiations or adjust employment, thereby reinforcing the negative effect on labor income and output, making the recovery a long one.

The specification of the restrictions, on the breaks in particular, matter more here than in labor income shock scenario. While in the United States the impulse responses from the restricted and unrestricted model are virtually the same, the unrestricted model (M_0) would underestimate the effects on labor income and output.

4.4 Dynamic Effects of a Redistribution Shock

Definition 1 shows how to linearly combine the labor income and profit shocks to get an exact one-dollar redistribution on impact. Figure 4 shows the dynamic effects of a redistribution toward labor income.[9]

What may take one by surprise at first, turns out to be the central difference in the adjustment to a redistribution shock in the two countries. The profit response in Figure 3 is pretty much the mirror image of the labor income response in Figure 2 and vice versa. Following our discussions of the single shock scenarios, this cross-pattern verifies our hunch of a different formation of expectations in the two countries. Depending on whether expectations manifest through a stimulation of spending, as in the United States, or the adjustment of capacities, as in Canada, the transmission of a redistribution shock differs. Specifically, in the United States, output increases on impact by 0.15 dollars, reaches a peak two quarters out at 0.51 dollars and dips below trend after eight quarters before it steadily reverts to trend. Although in Canada output initially increases by 0.47 dollars, it takes

[9]The effects of the opposite experiment would be exactly symmetric.

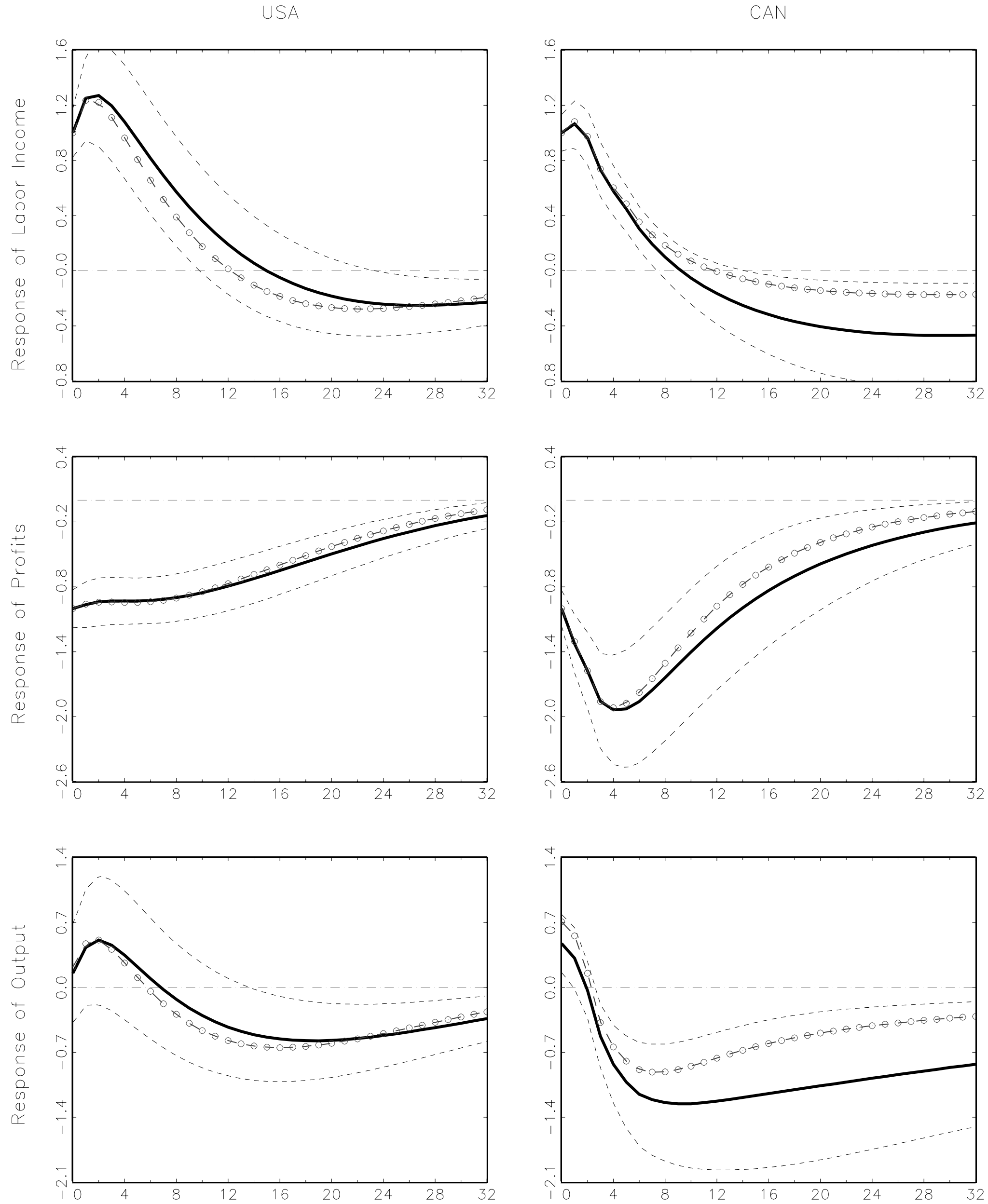

Figure 4 Response to a One-Dollar Redistribution from Profits to Labor Income

Notes: The thick solid line depicts the impulse response from the model with parameter restrictions on the cointegration and the contemporaneous impact matrices (model M_1^*), surrounded by bootstrapped one-standard deviation bands (2,500 replications, see Appendix A.2). The dashed line with circles depicts the response based on the VECM with no parameter restrictions (model M_0).

a rather persistent nosedive of 1.73 dollars within the first eight quarters. Put another way, Canadians would benefit form a redistribution toward profits.

When the transmission of a redistribution shock has a strong spending component, as in the United States, the response of output should be similar to the ones from other attempts to stimulate the economy through a spending stimulus. In fact, our output response is akin to the one Corsetti et al. (2009) get from a fiscal stimulus in which a debt-stabilizing policy systematically reduces spending below trend over the medium-run.

5 Conclusions

Our paper contributes to the debate about the output effects of a redistribution between labor income and profits. Fiscal stimulus packages or monetary policy measures, which might have triggered off the recovery after the 2008 crisis, inevitably imply a shift of income between labor and profits. In a more general view the discussion in our paper therefore relates to a long-standing debate on wage-led versus profit-led economic expansion. While most of the recent research focuses directly on the effects of government spending, we elaborate on the possible beneficial output effects of redistributing resources from profits toward labor income. A positive output effect, at least over the short-run, requires two main ingredients: a marginal propensity to spend out of labor income which exceeds the one of firms to spend an additional dollar of profits, and a medium-run capacity effect, brought about by lost profit opportunities, that does not crowd out the first effect too strong and too quickly.

We study how these opposing effects shape the output response in a 3-dimensional structural VECM with up to two breaks at unknown time using quarterly data on labor income, profits and output. Our analyzes focuses on the post-World War II economies of the United States and Canada. While in the United States, a one-dollar redistribution from profits to labor income, in fact, increases output long enough to call the experiment a success, the redistribution fails to produce the welcome output stimulus in Canada. After a short-lived output gain the Canadian economy plunges persistently below trend. As our VECM bestows the symmetry property, this results actually suggests a redistribution toward profits.

We discuss several related economic arguments in turn to provide explanations for our results. One argument concerns the formation of expectations generally. American firms and workers thrive on the spending stimulus triggered off by the labor income shock. This positive impulse on expectations overcompensates the unfortunate effects from the profit shock. The United States economy therefore expands after a one-dollar redistribution toward labor income, at least over the short-run. In Canada the expectation channel seems to be the other way round which might itself be explained by the differences in unionization in the two countries. Stated differently, growth is wage-led in the United States whereas it is profit-led across the border. Finally, because of the stronger exposure of Canada to international trade, a labor income shock brings some unwelcome side effects: an increase in unit labor costs and a loss in competitiveness. These effects will reduce exports and may undo any additional profit opportunities arising through the boost in domestic demand. The automatic response channel from output back on profits is, therefore, absent in Canada while it is positive in the United States.

Developing a proper econometric tool that jointly estimates break dates and cointegrating rank was, besides the economic question, the main objective of our paper. We provide Monte Carlo evidence showing the basic consistency of this joint estimation procedure.

The low-dimensionality of the VECM helps to keep econometric issues with break dates, cointegration, and identification at a minimum but, at the same time, limits the accuracy of our economic explanations. In ongoing work we draw on our conclusion from this paper and build a DSGE model which can explicitly shed more insights on the transmission mechanism of a redistribution shock.

Furthermore, allowing for country interdependencies, for instance, or comparing evidence across a larger set of countries strikes us as promising directions for future research. The global error-correcting framework of Pesaran et al. (2004) or the global VAR of Dees et al. (2007) may be good starting points for an extension along these lines. Since we focus on the econometric procedure, an extensive cross-country study was beyond the scope of the paper. Our results for the United States and Canada are, however, suggestive for the possible benefits and pitfalls of redistributing income between labor and profits.

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A Appendix

A.1 The Reduced-Form Level VAR

The level VAR version of the VECM in (10) proves useful whenever we have to make a statistical judgement, for instance about the lag order, before we have actually determined the cointegrating rank. We leave the cointegrating rank unrestricted in (10), that is $\beta = I_k$, and rewrite the model as

$$\begin{aligned} y_t = \mu_0^* + \mu_1^* t + \delta_1^* d_{1t} + \delta_2^* d_{2t} \\ + \sum_{j=1}^{p} A_j y_{t-j} + \sum_{j=0}^{p-1} \left(\gamma_{1j}^* \Delta d_{1t-j} + \gamma_{2j}^* \Delta d_{2t-j}\right) + u_t. \end{aligned} \quad (26)$$

After some rearranging we get the mapping with the parameters in (7), (9), and (10) as $\mu_0^* = \nu + \Pi\mu_1$, $\mu_1^* = -\Pi\mu_1$, $\delta_i^* = -\Pi\delta_i$, $\gamma_{i0}^* = \delta_i - \delta_i^*$, and $\gamma_{ij}^* = \gamma_{ij}$ for $i = 1, 2$ and $j = 1, \ldots, p-1$. Obviously the linear trend has found its way back into the model. With β being the identity matrix we can no longer maintain Assumption 2.

A.2 Bootstrapped Standard Errors

All the results in Section 4 come with bootstrapped standard errors. While the method is more general and applies to error bands for impulse responses and so on, we show it with the help of the contemporaneous impact matrix B. We derive the bootstrap covariance matrix of B as

$$\text{vec}(\hat{\Sigma}_{\mathsf{B}}) = N^{-1} \sum_{n=1}^{N} \Big(\text{vec}(\hat{\mathsf{B}}_n) - \text{vec}(\hat{\mathsf{B}})\Big)^2, \quad (27)$$

in which N is the number of bootstrap replications and $\hat{\mathsf{B}}_n$ is the estimate of the contemporaneous impact effect from the n-th replications. Using the estimate $\hat{\mathsf{B}}$ instead of the mean value of all $\hat{\mathsf{B}}_n$ ($n = 1, \ldots, N$) in the bootstrap, automatically accounts for the stochastic nature of the long-run constraint (see Section 2; Brüggemann, 2006; and Vlaar, 2004).

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