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NON-LINEARITY AND CROSS-COUNTRY DEPENDENCE OF INCOME INEQUALITY

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We use top income data and the newly developed regime-switching Gaussian mixture vector autoregressive model to explain the dynamics of income inequality in developed economies within the past 100 years. Our results indicate that the process of income inequality consists of two equilibria identifiable by high inequality and high income fluctuations, and low inequality and low income fluctuations. Our results also imply that income inequality in the United States is the driver of income inequality in other developed economies. High wages and capital gains are found to be the likely channels for the U.S. influence.

JEL Codes: C32, C33, D30

Keywords: top 1 percent income share, GMAR, multiple equilibria, developed economies

1. Introduction

The history of the distribution of product, or income, inequality is embodied by large fluctuations in the share of income massing at the top.[1] According to Piketty 2014 (p. 274), in the history of inequality "there have been many twists and turns and certainly no irrepressible, regular tendency toward a natural equilibrium." In a similar vein, Roine and Waldenström (2011) found global and

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[1]Estimates on the level of global income inequality vary (see, among others, Sala-i-Martin, 2002; Anand and Segal, 2008; Milanovic, 2015), but the share of the total income going to the top income earners has not been this high in many developed economies since the 1920s (Alvaredo *et al.*, 2013).

country-specific break points from the top 1 percent income share series, which indicate that the process of income inequality could consist of different phases or equilibria. This study finds further evidence to support this by indicating that income inequality follows a regime-switching process where higher inequality leads to higher variance in income shares and vice versa. Our results also indicate that changes in income inequality in the United States (U.S.) have driven inequality in other developed economies during the past 100 years.

The structure of income inequality has varied quite heavily throughout the past century. In the period prior to World War II, high incomes consisted mostly of returns to capital, which was the main reason for high inequality during that era (Piketty and Saez, 2013; Piketty, 2014). However, the biggest driver of the resurgence of income inequality in developed economies after the 1970s has been the increasing share of high wages (Piketty, 2014). The variance of earnings has also been on the rise in developed economies during the same period (see, e.g., Daly Valletta 2008; Gottschalk and Moffitt, 2009; Beach *et al.*, 2010). Although increasing variance of earnings has occurred during a period marked by increasing income inequality, research on their relationship has been almost non-existent.[2] Moreover, to our knowledge, there are no studies looking at the possible dependence of income inequality of one individual country on that of others. In this study, we set out to fill these gaps.

As argued by Piketty (2014), income inequality seems not to have been following any kind of mean reversing process (see above). This has been confirmed in many econometric studies, which have been unable to reject the unit root hypothesis in the autoregressive models for different measures of income inequality (see, among others, Mocan, 1999; Parker, 2000; Jäntti and Jenkins, 2010; Malinen, 2012; Herzer and Vollmer, 2013). However, this is a problematic result, as the series of commonly used measures of income inequality, such as the Gini index and the top income share, are bounded between 0 and 1, while the unit root series has a time-increasing variance. The breaks in the top 1 percent income share series identified by Roine and Waldenström (2011) could be one reason for the non-rejection of unit root hypotheses. If breaks are actually shifts between different phases of income inequality identified by, for example, different levels of variance, there would be no tendency toward a single equilibrium, but shifts between multiple equilibria. A linear autoregressive model would be misspecified due to the observed jumps, whereas the so-called trend-break models would ignore the strong autocorrelation in the series.

We employ a newly developed Gaussian mixture autoregressive (GMAR) model studied in Kalliovirta *et al.* (2015) and its multivariate generalization, the Gaussian mixture vector autoregressive (GMVAR) model by Kalliovirta *et al.* (2016), to estimate the dynamic properties of income inequality. We use the GMAR and GMVAR models to identify the different regimes created by the breaks and autoregressive dynamics in the top income series, because they are able to model multiple equilibria. We analyze the top 1 percent income share data ranging from the end of the 19th century to the beginning of the 21st century for six countries: Australia, Canada, France, Finland, Japan, and the U.S.

[2]In the only study we could find, Beach *et al.* (2010) show that a rise in the total earnings variance in Canada after 1982 is mostly attributable to an increase in overall inequality.

According to our results, the process of income inequality has consisted of two or three different regimes. Two of these regimes are also found to be common to all the countries in the sample. Regimes are characterized by different means, or levels, and with different variances, or scales of variation. Moreover, our GMVAR results indicate that not only is the variance of income inequality highly dependent across countries, but income inequality in the U.S. is the driver of income inequality across our sample. The results of the impulse response analyses imply that role of the U.S. is stronger in the high-inequality, high-income-fluctuations regime, where all countries in our sample currently reside. Both institutional and economic changes in labor and capital markets emanating from the U.S., as well as globalization, are found to explain the influence of the U.S.

The rest of the paper is organized as follows. Section 2 presents the data and the GMAR and GMVAR models. Section 3 presents the univariate and panel estimations of the GMAR and GMVAR models. Section 10 discusses the economic implications of the estimation results and Section 11 concludes.

2. Data and Methods

The top 1 percent income share of population is used to proxy income inequality. It is the only aggregate measure of income inequality that currently contains enough observations for a meaningful testing of the long-run dynamics of income inequality.[3] The data on top income shares are obtained from the World Top Income Database (WTID; Alvaredo *et al*., 2013). At the time of writing, the WTID had long, continuous time series on six developing countries: Australia, Canada, Finland, France, Japan, and the U.S.[4] For these countries, the data on the top 1 percent income shares start at the end of the 19th or the beginning of the 20th century. For other countries, the data either start only after World War II and/or they have gaps extending over several years.

The dynamics in the 1 percent income share series can be described with traditional linear autoregressive models that explain the current value with past values in the series. However, earlier research indicates that those estimated models imply properties that are unreasonable for such bounded series (see Section 1). Further, to be able to combine both the structural breaks and the dynamics of the series, a non-linear time-series model needs to be employed. The univariate and multivariate non-linear time-series models that we have chosen (GMAR and GMVAR) combine linear autoregressive parts and structural break points. Thus, the number of breaks and their timing as well as the dynamic structure are determined from the observed data. The conclusions drawn about the regimes in the income share series

[3]Leigh (2007) has also demonstrated that the top 1 percent income share series have a high correlation with other measures of income inequality, such as the Gini index.

[4]For Japan, the observation from the year 1946 is missing, and it has been replaced with the average of the top 1 percent income share from the years 1945 and 1947. For Canada, the top 1 percent income share data is continued with the top 1 percent income share Longitudinal Administrative Databank (LAD) data after the year 2000. For Finland, the top 1 percent income share tax data are continued with the top 1 percent income share Income Distribution Survey (IDS) data after the year 1992. There was a large jump in the top 1% income series of Australia in 1951. Because it caused instability to multivariate estimations, it was phased out using the mean value of the series between 1950 and 1952.

TABLE 1
Estimation Results on the Top 1 Percent Income Share

	Australia	Canada	Finland	France	Japan	U.S.
Original Data						
First autocorrelation	0.94	0.96	0.95	0.96	0.98	0.95
Mean	8.2	11.4	8.8	10.8	12.2	12.8
Variance	4.5	8.8	9.5	14.9	24.4	14.7
GMAR Model						
Autocorrelation (φ_1)	0.90 (0.04)	0.95 (0.03)	0.94 (0.03)	1.11 (0.11)	1.33 (0.08)	1.14 (0.10)
Autocorrelation (φ_2)				−0.16 (0.12)	−0.42 (0.09)	−0.24 (0.10)
Mean 1 (μ_1)	4.8 (0.1)	9.6 (0.8)	4.9 (0.5)	8.4 (0.4)	8.1 (0.3)	8.2 (0.4)
Mean 2 (μ_2)	6.0 (0.4)	14.3 (4.4)	7.9 (1.6)	15.5 (2.2)	16.7 (1.5)	15.1 (1.2)
Mean 3 (μ_3)	9.1 (1.0)					
Variance 1 (γ_1)	0.01 (0.01)	1.4 (0.7)	0.5 (0.3)	0.6 (0.3)	0.6 (0.2)	0.2 (0.1)
Variance 2 (γ_2)	0.3 (0.1)	15.2 (9.6)	4.5 (2.1)	11.3 (5.8)	12.7 (5.2)	6.6 (2.5)
Variance 3 (γ_3)	2.9 (1.3)					
α_1	0.08 (0.1)	0.95 (0.08)	0.22 (0.2)	0.42 (0.3)	0.62 (1.4)	0.17 (0.3)
α_2	0.47 (0.2)					
Logl	−61.6	−74.6	−89.5	−48.9	−94.7	−84.0
AIC	141	161	191	112	201	182
BIC	164	176	206	130	218	200
N	0.84	0.53	0.93	0.03	0.09	0.19
A	0.40	0.15	0.37	0.70	0.85	0.04
H	0.13	0.32	0.23	0.70	0.05	0.80
Years	1921–2010	1920–2010	1920–2010	1915–2010	1886–2010	1914–2010

Notes: Standard errors (in parentheses) are calculated using the estimated Hessian. “Logl” is the log-likelihood value of the estimated model. The smaller value of the information criteria AIC and BIC indicates the preferred model. The linear AR(2) model is preferred for Finland by BIC, but it is rejected by the diagnostic tests. The p-values of the diagnostics tests of normality (N), remaining autocorrelation (A), and conditional heteroskedasticity (H) in the (quantile) residuals are given so that values below 0.001 are denoted with 0.

having different levels and levels of variation thus strictly depend on the particular choice of the non-linear model.

We assume that in each country, the observed top 1 percent income share series follows a regime-switching GMAR process. This assumption is reasonable, because regime switches are a natural way to adequately model jointly both the dynamic autoregressive (AR) structure of these series and the structural breaks found by Roine and Waldenström (2011). They especially allow for multiple equilibria, unlike the linear AR models. A similar regime-switching approach has been successfully used in, for example, Hamilton (1989) to model the U.S. business cycle. The Markov-switching AR model of Hamilton (1989) and the GMAR model are closely connected. They both contain two or more regimes and each regime consists of a linear AR model. These separate AR models can differ in their parameters and together they describe the dynamics of the system. The changes between the regimes are estimated from the observed data. However, the general flexibility of these regime-switching models comes with a price: one has to be careful how to interpret them, because instead of knowing the regime exactly at each point in time, only an estimate of the probability of the series being in a certain regime is available. These estimated probabilities are henceforth referred to as time-varying mixing weights. In the Hamilton model, the probability of a regime switch is constant and does not depend on the previous values of the series, whereas in the GMAR model the change in regime is varying in time and depends on the previous values of the series. If the regime-switching probability is constant, the regime switches occur independent of the current and previous levels of the series. Thus, in the Hamilton model, a switch is equally likely at high and low levels of the series. However, Roine and Waldenström (2011) find that the structural breaks tend to occur when the level of the top 1 percent income share series moves from high to low and low to high. The GMAR model allows for more flexibility and credibility in modeling the regime changes in the top 1 percent income share series.

The GMAR model has several advantageous properties compared to the Markov-switching AR model or other non-linear models, such as the popular STAR model. First, the GMAR model is more parsimonious, a considerable advantage when only yearly data for a hundred years or less are available. Second, the GMAR model is known to be stationary: it suffices that the usual stationarity condition of the conventional linear AR model is fulfilled in the regimes.[5] Third, as a direct consequence of the stationarity, the stationary distribution of the GMAR model is known exactly. Thus, we are able to make direct comparisons with the unconditional moments (means and variances) of the original observations (as in Table 1). The GMAR model with two regimes implies separate means and variances in each regime, and these can be interpreted to describe different equilibrium points. We utilize this property when we compare the differences between the regimes in the estimated GMAR models for the top 1 percent income share series.

[5]The top 1 percent income share series is bounded between 0 and 1. The unit root type non-stationarity in estimated linear models implies that the series is unbounded since the variance of the unit root series is unbounded (increases as a function of time). The properties of the data and the model do not match, although such models that employ the unit root assumption have been employed in modeling bounded series in the literature. Thus, we are able to make the stationarity assumption due to the particular model choice, and this assumption is in accordance with boundedness.

This would be unavailable if any other non-linear model had been used, because the conditions for making the transition from the conditional to the unconditional distribution would be unknown.[6] We give the model equations of the estimated GMAR models in Appendix A (in the online Supporting Information). To learn more about the GMAR model and its competing non-linear alternatives, see Kalliovirta *et al*. (2015).

To understand the joint behavior of the 1 percent income share series in all six countries, we employ the multivariate GMVAR model of Kalliovirta *et al*. (2016). In particular, this multivariate model is able to depict regime switches and dynamic structures *common* to all these six countries. The theoretical properties of the GMAR model explained above continue to hold in the multivariate model, where two or more regimes of separate linear VAR models describe the dynamics and the regime changes depend on the past values of the series. Again, the exact knowledge of the series being in a particular regime needs to be replaced by the probability of being in a particular regime described by the time-varying mixing weights. Appendix A contains more details on the estimated GMVAR model.

To conclude, the GMAR and GMVAR models are well suited to examine the non-linearity in the 1 percent income share series caused by structural breaks. However, our modeling approach is only approximative and thus similar to any other statistical model based on historical data.

3. Results

3.1. *The Univariate Model*

As a starting point for the analysis of each series, we estimated linear Gaussian AR models. Residual diagnostics (not reported) rejected these models due to non-normality and conditional heteroskedasticity, which is a clear indication of non-linearity in the modeled series. We also performed linearity tests on the top 1 percent income share series within a STAR model and the linearity hypotheses were rejected (see Table 2) in all countries. Table 1 presents the properties of the original series and the estimation results for GMAR models that pass the quantile residual diagnostics of Kalliovirta (2012).[7]

[6]This analytic transformation is explained in Kalliovirta *et al*. (2015).

[7]The accuracy of the mean, variance, and weight parameter estimates suffer from the lack of data. Testing the significance of the mixing weights is a theoretically highly demanding non-standard testing problem, common to all regime-switching models such as the STAR and Markov-switching models (for further explanation, see Kalliovirta *et al*., 2015), and it has not been solved yet for GMAR models. For the same reason, one cannot test the equality of the means or variances simply by comparing their estimates and standard errors, because these parameters are closely connected to the time-varying mixing weights. Further, testing the equality of means and variances jointly would again lead to the non-standard testing problem. However, we can test them separately. For example, in the income series for Canada, the likelihood ratio (LR) test for equality of means has a p-value of 0.31 and that for equality of variances has a p-value $<10^{-12}$. The quantile residual diagnostics indicate that the model with equal means describes the autocorrelation of the series inadequately. Thus, the model reported in Table 1 is chosen.

For this reason, we base the model specification on the theoretically appropriate quantile residual diagnostics, which support non-linearity over linearity in all six models. Further, information criteria such as AIC and BIC and residual diagnostics (Table 1) clearly indicate that the non-linear models are superior. More details on the estimated models and residual diagnostics are available upon request.

Clearly, the original series are persistent in all six countries, and the variances are also highly fluctuating, from around 24 in Japan to around 5 in Australia. The GMAR model finds two regimes in the top 1 percent income series in all countries except Australia, where three regimes are found. The series of France, Japan, and the U.S. require two lags in the GMAR model, whereas one lag is enough for the other three countries.

The regimes of the GMAR models seem to be marked with quite clear and similar characteristics in all countries. In one regime, the mean and the variance of the top 1 percent income series are clearly higher, whereas in the other regime both are considerably lower. So, in these countries, income inequality has consisted of two notably different regimes. The first one is a low-income-inequality, low-income-fluctuations regime and the second is a high-income-inequality, high-income-fluctuations regime. Even though the Australian series has three regimes, the same characteristics are found in all of them.

Further, our analysis in Table 2 points out that the evolution of the top 1 percent income series cannot be modeled adequately using a linear model. It is likely that the high autocorrelation observed in the original series may be a consequence of ignoring the non-linear structure—that is, the structural breaks—which leads to different constants and variances between the regimes. This indicates that although

TABLE 2
Estimation Results on the Top 1 Percent Income Share

	Australia	Canada	Finland	France	Japan	U.S.
AR(2) Model						
Logl	−80.8	−90.3	−93.1	−74.3	148.6	−119.8
AIC	170	189	194	157	305	248
BIC	180	199	204	167	316	258
N	0.11	0	0.35	0	0	0.66
A	0	0	0	0.80	0	0.80
H	0	0	0.04	0	0	0
STAR Model						
Non-linearity caused by y_{t-1}		**		***	***	
Non-linearity caused by y_{t-2}	**		*	***	***	*
Time t	***	*	*	*	*	***
Years	1921–2010	1920–2010	1920–2010	1915–2010	1886–2010	1914–2010

Notes: "Logl" is the log-likelihood value of the estimated model. The smaller value of the information criteria AIC and BIC indicates the preferred model. The linear AR(2) model is preferred for Finland by BIC, but it is rejected by the diagnostic tests. The p-values of the diagnostics tests of normality (N), remaining autocorrelation (A), and conditional heteroskedasticity (H) in the (quantile) residuals are given so that values below 0.001 are denoted with 0. The non-linearity tests are based on the STAR model, where only one variable at a time can depict the non-linearity in the model. Therefore, the STAR model is much simpler than the GMAR model, which allows several variables as well as the variances to depict the non-linearity in the model. Thus, the STAR model based tests may lack the power to reject the null hypothesis of linearity when the nature of the non-linearity is dependent on several variables. However, currently the non-linearity tests only exist for the STAR models. For more details on STAR model based non-linearity testing, see, for example, Teräsvirta (1998). The hypothesis of linearity rejected at 10 percent, 5 percent, and 1 percent are indicated by *, **, and ***, respectively.

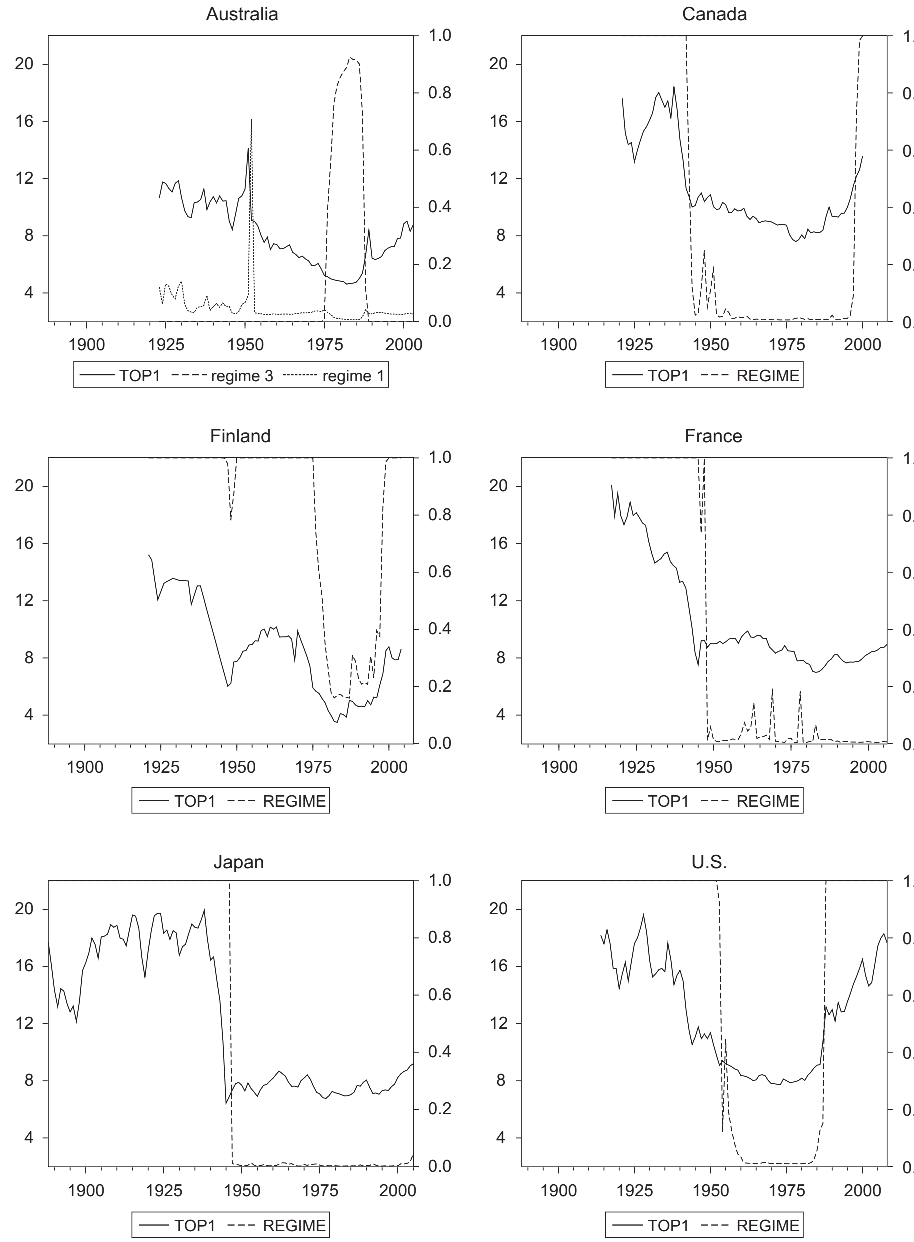

Figure 1. The Top 1 Percent Income Shares and Time-Dependent Mixing Weights for Australia, Canada, Finland, France, Japan, and the U.S., Based on the Univariate GMAR Model

the dynamics of income inequality can be *approximated* with a stochastic trend—that is, with a unit root process—this may not be its true form.[8]

[8]As already mentioned, the unit root type non-stationarity assumption in a model for bounded series is inconvenient. Further, it has been found in several studies on non-linear models that once the structural breaks are modeled, the autoregressive coefficients diminish in absolute value. This indicates that autocorrelation in the series is weaker than indicated by a linear model that cannot account for the structural breaks. This phenomenon is also visible in our estimation results.

Figure 1 presents the top 1 percent income shares and the estimated time-dependent mixing weights for the above-mentioned six countries. In all six subfigures, the mixing weights, $\hat{\alpha}_{1,t}$ (dashed line) or $\hat{\alpha}_{2,t}$ (dotted line), in the subfigure for Australia are given on the right-hand axis, while the share of total income earned by the top 1 percent of the income earners (solid line) is given on the left-hand axis.

In Australia, the probability that income inequality is in the third regime is above 90 percent until 1955. In 1955, the probability of the second regime begins to rise. The transition from the second regime into the first regime happens around 1975 and then moves back into the second regime in 1987. In 1999, the series moves back into the third regime. In Canada, France, and Japan, income inequality switches the regime right after World War II. The probability that the income inequality series is in the first regime increases to 99 percent in Canada in 1944, to 98 percent in France in 1948, and to 98 percent in Japan in 1948. In Finland, the probability of income inequality being in the first regime increases to 33 percent in 1976 and decreases to below 2 percent in 1998. In the U.S., the probability that income inequality is in the first regime increases to 61 percent in 1955. After 1988, the probability of the second regime is 100 percent.

The results based on GMAR models imply that many of the structural breaks found by Roine and Waldenström (2011) are points, where the series of income inequality changes regime and the characteristics of the series change in terms of means and variances. We find the following correspondences between the breaks of Roine and Waldenström (2011) and the regime switches: (1) in Australia, the regime change in 1987 corresponds to the structural break in the country-specific series in 1985; (2) in Canada, the country-specific break point in 1994 corresponds to the probability of the second regime beginning to increase in 1998; (3) in Finland, the probability of income inequality being in the first regime increases to 73 percent in 1981, which corresponds to the break in post-war data on Nordic countries, and the probability of the second regime rises to over 68 percent in 1997, which corresponds to the country-specific break in 1997; (4) in Canada, France, and Japan, the changes from the second regime into the first regime correspond to the global trend in the break point in 1946; and (5) in Australia and the U.S., the changes in regime around 1955 and 1987 correspond to the common structural break in 1953 and the common post-war break in Anglo-Saxon countries in 1987.

3.2. *The Multivariate Panel Data Model*

Next, we combine the six individual series into a panel over the years 1921 and 2009 to find out whether the regime switches and other dynamics in these

TABLE 3
Diagnostics on the Multivariate Top 1 Percent Income Share Models

	Logl	AIC	BIC	N	A	H
VAR model	−396.7	991	1,234	0	0.91	0
GMVAR model	−366.9	830	948	0.11	0.05	0.28

Notes: "Logl" is the log-likelihood value of the estimated model. The smaller value of the information criteria AIC and BIC indicates the preferred (GMVAR) model. The p-values of the diagnostics tests of normality (N), remaining autocorrelation (A), and conditional heteroskedasticity (H) in the (quantile) residuals are given so that values below 0.001 are denoted with 0.

series move in tandem. The GMVAR model that passes quantile residual diagnostics has three regimes and a VAR structure with two lags common in all regimes. Thus, the regimes differ in their constant and covariance matrix parameters.[9]

We report the estimated GMVAR model component by component to make comparisons with the estimated univariate models easy, and we report the estimated Hessian-based standard errors in parentheses below. The estimated weight parameters for the first and second regimes in the GMVAR model are $\hat{\alpha}_1 = \underset{(0.28)}{0.14}$ and $\hat{\alpha}_2 = \underset{(0.28)}{0.85}$. Note that these estimates also yield the unconditional probabilities $P(s_{t,1} = 1) = 0.14$, $P(s_{t,2} = 1) = 0.85$, and $P(s_{t,3} = 1) = 0.01$. We denote the ith element of vector $\hat{\Omega}_2^{1/2}\boldsymbol{\varepsilon}_t$ with $u_{t,i}$ and report separately the estimated covariance matrix $\hat{\Omega}_2$, because it is not diagonal like $\hat{\Omega}_1$. The third regime is added to allow the constants of France and Japan to change within the second regime, so that there is no need for the third covariance matrix. Based on the likelihood ratio (LR) test statistics, we restrict the variance parameter for Finland to be the same in both regimes. According to the GMVAR model, the top 1 percent income share series of Australia, Canada, Finland, France, Japan, and the U.S. are as follows:

$$
\begin{aligned}
y_{t,\text{Aus}} = {} & \underset{(0.04)}{0.93}\, y_{t-1,\text{Aus}} + \underset{(0.04)}{0.02}\, y_{t-1,\text{U.S.}} \\
& + s_{t,1}\left(\underset{(0.27)}{0.23} + \sqrt{\underset{(0.02)}{0.08}}\,\varepsilon_{t,\text{Aus}}\right) + (1-s_{t,1})\left(\underset{(0.44)}{0.28} + u_{t,1}\right),
\end{aligned}
$$

$$
\begin{aligned}
y_{t,\text{Can}} = {} & \underset{(0.11)}{1.00}\, y_{t-1,\text{Can}} + \underset{(0.05)}{0.11}\, y_{t-1,\text{U.S.}} - \underset{(0.10)}{0.15}\, y_{t-2,\text{Can}} \\
& + s_{t,1}\left(\underset{(0.28)}{0.35} + \sqrt{\underset{(0.02)}{0.07}}\,\varepsilon_{t,\text{Can}}\right) + (1-s_{t,1})\left(\underset{(0.46)}{0.17} + u_{t,2}\right),
\end{aligned}
$$

$$
\begin{aligned}
y_{t,\text{Fin}} = {} & \underset{(0.03)}{0.91}\, y_{t-1,\text{Fin}} + \underset{(0.04)}{0.07}\, y_{t-1,\text{U.S.}} \\
& + s_{t,1}\left(\underset{(0.27)}{0.13} + \sqrt{\underset{(0.07)}{0.47}}\,\varepsilon_{t,\text{Fin}}\right) + (1-s_{t,1})\left(\underset{(0.41)}{-0.23} + u_{t,3}\right),
\end{aligned}
$$

$$
\begin{aligned}
y_{t,\text{Fra}} = {} & \underset{(0.03)}{0.88}\, y_{t-1,\text{Fra}} + \underset{(0.02)}{0.06}\, y_{t-1,\text{U.S.}} \\
& + s_{t,1}\left(\underset{(0.23)}{0.54} + \sqrt{\underset{(0.02)}{0.07}}\,\varepsilon_{t,\text{Fra}}\right) + s_{t,2}\left(\underset{(0.30)}{0.19} + u_{t,4}\right) + s_{t,3}\left(\underset{(0.49)}{0.83} + u_{t,4}\right),
\end{aligned}
$$

[9]Similar to the univariate models, the accuracy of the mean, variance, and mixing weights estimates suffers from the lack of data. One may suspect that there are several redundant mean and variance parameters based on their standard errors. However, the testing of their equivalence has to be based on the LR tests (e.g. a LR test for equality of means in regimes 1 and 2 for France has a p-value of 0.002), and hypotheses that contain unidentified nuisance parameters lead again to the non-standard testing problems common to all regime-switching models. Thus, we base our model selection on the information criteria and theoretically appropriate quantile residual diagnostics presented in Table 3, which strongly support the GMVAR model over the VAR model. Note also that compared to the univariate case, the joint modeling leads to more efficient parameter estimates for the regime variances.

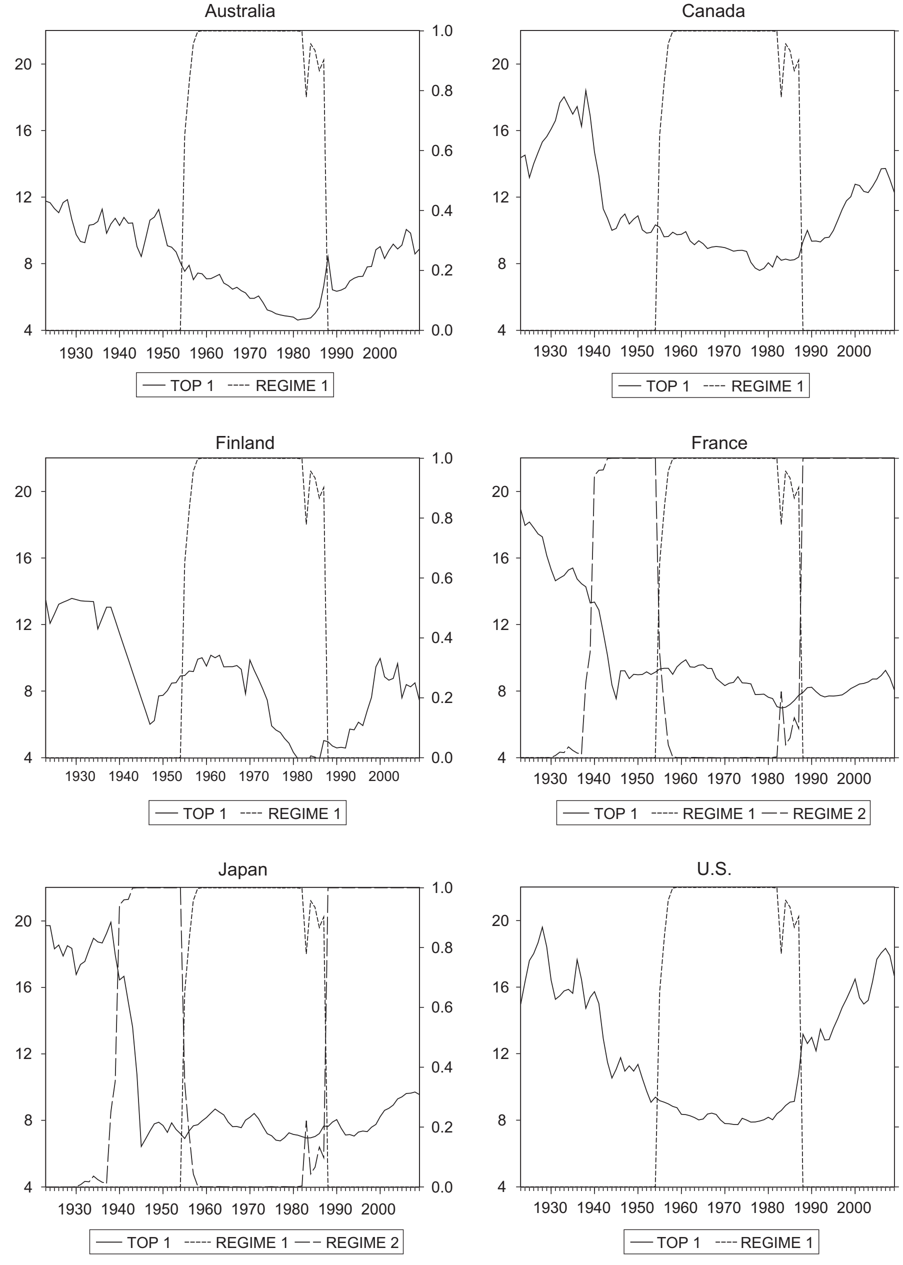

Figure 2. The Top 1 Percent Income Shares and Time-Dependent Mixing Weights for Australia, Canada, Finland, France, Japan, and the U.S., Based on the GMVAR Model

$$
\begin{aligned}
y_{t,\mathrm{Jpn}} = & \underset{(0.09)}{1.22}\, y_{t-1,\mathrm{Jpn}} + \underset{(0.03)}{0.13}\, y_{t-1,\mathrm{U.S.}} - \underset{(0.08)}{0.33}\, y_{t-2,\mathrm{Jpn}} \\
& + s_{t,1}\left(\underset{(0.28)}{-0.21} + \sqrt{\underset{(0.01)}{0.04}}\, \varepsilon_{t,\mathrm{Jpn}} \right) + s_{t,2}\left(\underset{(0.46)}{-0.75} + u_{t,5} \right) + s_{t,3}\left(\underset{(0.62)}{-0.16} + u_{t,5} \right).
\end{aligned}
$$

$$
\begin{aligned}
y_{t,\mathrm{U.S.}} &= \underset{(0.10)}{1.21}\, y_{t-1,\mathrm{U.S.}} - \underset{(0.10)}{0.28}\, y_{t-2,\mathrm{U.S.}} \\
&\quad + s_{t,1}\left(\underset{(0.37)}{0.55} + \sqrt{\underset{(0.01)}{0.03}}\,\varepsilon_{t,\mathrm{U.S.}}\right) + (1-s_{t,1})\left(\underset{(0.67)}{0.90} + u_{t,6}\right).
\end{aligned}
$$

The autoregressive dynamics within countries remain very similar to that found in the univariate models. However, the first lag of the top 1 percent income share of the U.S. affects the autoregressive dynamics of all countries, although the effect is not very large in Australia, where the lag coefficient is not statistically significant at the 5 percent level. The positive coefficients indicate that an increase (decrease) in income inequality in the U.S. leads to an increase (decrease) in income inequality in other countries the next year. Thus, the estimated model implies that changes in income inequality in the U.S. have been exported to countries across our sample. Because the autoregressive dynamics are the same in both regimes, this effect applies irrespective of the regime.

The mean vectors of the stationary distribution, solved using $\mu_m = A(1)^{-1}\phi_{m,0}$, are as follows:

$$
\boldsymbol{\mu}_1 = \begin{bmatrix} \mu_{1,\mathrm{Aus}} \\ \mu_{1,\mathrm{Can}} \\ \mu_{1,\mathrm{Fin}} \\ \mu_{1,\mathrm{Fra}} \\ \mu_{1,\mathrm{Jpn}} \\ \mu_{1,\mathrm{U.S.}} \end{bmatrix} = \begin{bmatrix} \underset{(0.6)}{6.0} \\ \underset{(0.5)}{9.0} \\ \underset{(1.2)}{7.5} \\ \underset{(0.4)}{8.6} \\ \underset{(0.6)}{7.8} \\ \underset{(0.5)}{8.5} \end{bmatrix}, \quad \boldsymbol{\mu}_2 = \begin{bmatrix} \underset{(1.8)}{8.6} \\ \underset{(1.9)}{12.0} \\ \underset{(1.8)}{7.6} \\ \underset{(1.2)}{8.5} \\ \underset{(2.5)}{9.2} \\ \underset{(2.1)}{14.0} \end{bmatrix}, \quad \text{and} \quad \boldsymbol{\mu}_3 = \begin{bmatrix} \mu_{2,\mathrm{Aus}} \\ \mu_{2,\mathrm{Can}} \\ \mu_{2,\mathrm{Fin}} \\ \underset{(2.0)}{13.7} \\ \underset{(3.2)}{14.3} \\ \mu_{2,\mathrm{U.S.}} \end{bmatrix}.
$$

The mean vectors of the stationary distribution of the GMVAR model have roughly the same values as those found in the univariate GMAR models. The differences are found in Australia, where the third, lowest mean regime found in the univariate GMAR model becomes redundant, and in Finland, where the mean value of the top 1 percent income share of low regime has increased significantly.

3.2.1. Time-Dependent Mixing Weights

Figure 2 depicts the top 1 percent income shares and the estimated time-dependent mixing weights for the above-mentioned six countries.

In all subfigures, the mixing weights, $\hat{\alpha}_{1,t}$ (dashed line), or $\hat{\alpha}_{2,t}$ (dotted line) in the subfigures for France and Japan, are given on the right-hand axis while the share of total income earned by the top 1 percent of the income earners (solid line) is given on the left-hand axis. At the beginning of the period, the series for France and Japan are in the third regime and the other series are in the second regime, with a probability of 100 percent. Both of these regimes have high mean and high variance. France and Japan change to the second, low-inequality regime around 1940.[10] Between 1955 and 1987, all the series are in the first regime with a probability of 99 percent.[11] This first regime has low mean and low variance. After 1988, the probability of the second regime is above 82 percent for all countries, indicating that income distribution has returned to the high-inequality, high-income-fluctuations regime. The regime changes common to all six countries in the multivariate model in 1953 and in 1987 are the same ones observable in the univariate model for the U.S. This further illustrates the significant effect that the U.S. series has on the dynamics for all the series in the multivariate model.

3.2.2. Details of the Regime Change

The striking similarity of the changes in regimes in the univariate GMAR model for the U.S. and in the multivariate GMVAR model raises the question of whether the U.S. series leads the change in the regime in our system. In Appendix B, we present contour plots of the estimated mixing weights of the GMVAR model. They show that a change in the conditional variance of the U.S. series leads the change in the regime in the multivariate model both in the 1950s and at the end of the 1980s. That is, the income inequality in the U.S. changes first and is followed by a similar change in inequality in the other countries in our sample. This implies that changes of income inequality in the U.S. have driven regime changes of inequality across our sample in the 1950s and in the 1980s.

3.2.3. Regime-Specific Covariances

Differing behavior of the series within regimes is also visible in the covariance matrices. In the first regime, where the means and variances are low, the covariance matrix $\hat{\Omega}_1$ is diagonal. Therefore, the shocks of the components are not contemporaneously connected, and in each country the variation is country specific. The estimated covariance matrix of the second (and the third) high-mean, high-variation regime,

$$\hat{\Omega}_2 = \begin{bmatrix} \underset{(0.09)}{0.50} & \underset{(0.07)}{0.13} & 0 & \underset{(0.05)}{0.12} & \underset{(0.07)}{0.16} & \underset{(0.09)}{0.24} \\ \underset{(0.07)}{0.13} & \underset{(0.10)}{0.54} & 0 & 0 & \underset{(0.07)}{0.14} & \underset{(0.10)}{0.17} \\ 0 & 0 & \underset{(0.07)}{0.47} & 0 & 0 & 0 \\ \underset{(0.05)}{0.12} & 0 & 0 & \underset{(0.06)}{0.30} & \underset{(0.06)}{0.22} & 0 \\ \underset{(0.07)}{0.16} & \underset{(0.07)}{0.14} & 0 & \underset{(0.06)}{0.22} & \underset{(0.11)}{0.57} & 0 \\ \underset{(0.09)}{0.24} & \underset{(0.10)}{0.17} & 0 & 0 & 0 & \underset{(0.18)}{0.91} \end{bmatrix},$$

[10]This is mostly due to fall in capital incomes caused by shocks; that is, depression and wars (see Piketty and Saez, 2006; Piketty, 2014). In Japan, there was also a major political regime shift in 1947, when the Empire of Japan was dissolved.

[11]The global break points found by Roine and Waldenström (2011) were in 1945 and in 1980.

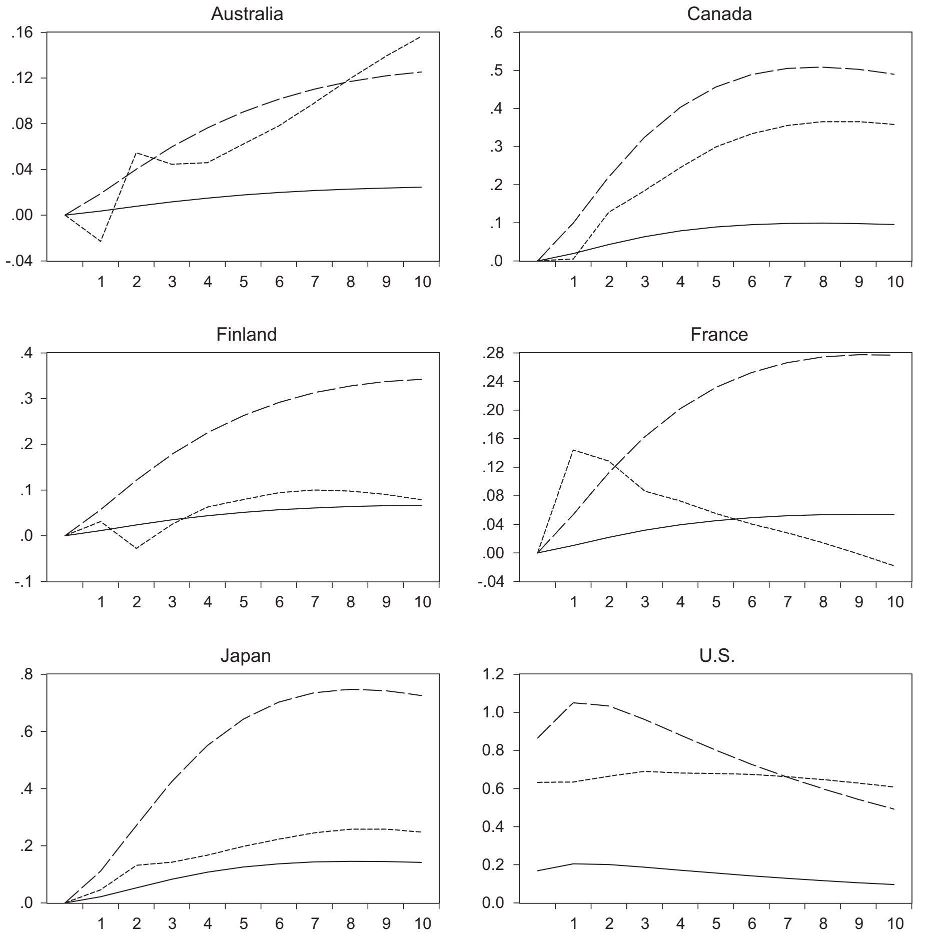

Figure 3. The Orthogonal Impulse Responses of All Countries to a Unit Change in the U.S. Top 1 Percent Income Share Series Based on Regime 1 in GMVAR Model (Solid Line), Regime 2 in the GMVAR Model (Dashed Line), and the VAR(2) Model (Dotted Line)

shows that, excepting Finland, the components affect each other through shocks, indicating a symmetric interdependence.[12] In the second (and the third) regimes, the means and variances are high and (excepting Finland) a shock in one country is contemporaneously connected to the shocks affecting the other countries. To make the strength of the dependence between countries easier to interpret, we also report the corresponding correlation matrix:

[12]The variance of the top 1 percent income share of the second regime in Finland is identical to that in the first regime (see the autoregressive dynamics in Section 3.2). However, the conditional variance of Finland changes due to the regime structure.

$$\begin{bmatrix} 1 & \underset{(0.13)}{0.26} & 0 & \underset{(0.12)}{0.32} & \underset{(0.12)}{0.30} & \underset{(0.12)}{0.36} \\ \underset{(0.13)}{0.26} & 1 & 0 & 0 & \underset{(0.12)}{0.24} & \underset{(0.13)}{0.24} \\ 0 & 0 & 1 & 0 & 0 & 0 \\ \underset{(0.12)}{0.32} & 0 & 0 & 1 & \underset{(0.10)}{0.53} & 0 \\ \underset{(0.12)}{0.30} & \underset{(0.12)}{0.24} & 0 & \underset{(0.10)}{0.53} & 1 & 0 \\ \underset{(0.12)}{0.36} & \underset{(0.13)}{0.24} & 0 & 0 & 0 & 1 \end{bmatrix}.$$

Although the effect of the U.S. on the top 1 percent income series of Australia is weak in terms of the autoregressive dynamics (see Section 3.2), the two-way contemporaneous effect is statistically significant (correlation = 0.36) in the high-mean, high-variance regime through the shock structure. The strongest dependence between shocks is observed for France and Japan, where the correlation is 0.53. To conclude, income inequality between France and Japan and between the Anglo-Saxon countries Australia, Canada, and the U.S. is contemporaneously connected through shocks in the high-mean, high-variance regime.

3.2.4. Impulse Response Analysis

To gain a better understanding of the dynamical system in the estimated GMVAR model, we compute the regime-specific orthogonal impulse responses of the top 1 percent income shares of all countries to a unit shock in the U.S. series. We also include a linear VAR model in the analysis to obtain more comparison.[13] We employ the orthogonal impulse responses, because the shocks in the high-level, high-income-fluctuations regime of the GMVAR model and in the VAR model are contemporaneously correlated. The results of the impulse response analysis are presented in Figure 3.

The impulse response functions in Figure 3 describe how a single one-unit impulse (change) in the top 1 percent income shares of the U.S. will affect the future values (levels) of top 1 percent income shares for the U.S. and for countries across our sample over the period of the next 10 years. A positive impulse response function indicates that an increase in the top 1 percent income share of the U.S. will increase its level in other countries in our sample, whereas a negative impulse response indicates that a unit increase in the U.S. causes a decrease in the level top 1 percent income shares in other countries.

The different dynamics between the regimes are clearly visible in Figure 3. In the (low-inequality and low-income-fluctuations) regime 1, a shock to the U.S. series has a negligible effect on the series of other countries. For the U.S., the effect is positive and decays slowly to zero. In the (high-inequality and high-income-fluctuations)

[13]Table 3 reports diagnostics on the estimated VAR(2) model. The log-likelihood in the GMVAR model is larger than in the VAR even though the VAR(2) model has 114 parameters and the GMVAR model has only 48. Table 3 shows that the information criteria support the GMVAR model. Further, the quantile residual diagnostics strongly support the GMVAR model over the VAR(2) model. More details are available upon request.

regime 2, a shock to the U.S. series has a strong non-negligible effect on all countries. Because the autocorrelation structure is the same in both regimes, the differences are explained by the different error covariance matrices in the regimes.

In the VAR model, the dynamic impact of the U.S. top 1 percent income shares on its future values begins on level lower than in regime 2, but it is more persistent than in the GMVAR model, because the impulse response function of the VAR model is larger for lags extending beyond 6 years. This might be explained by the fact that the largest root in the GMVAR model is 0.93, compared to 0.98 in the VAR model. Also, the impact of the U.S. on other countries, in general, is smaller in the VAR model than is observed in regime 2. One may interpret that the VAR model represents a weighted average model over the two regimes, so its impact is approximately a weighted average between regimes 1 and 2. Therefore, the role of the U.S. across our sample in regime 2 would be underestimated by the VAR model. We consider the impulse responses of the GMVAR model to be more reliable, because the GMVAR fits the data better according to diagnostic testing and allows for multiple equilibria, whereas the VAR model allows only a unique equilibrium. Moreover, the GMVAR is in line with the observed non-linear behavior of the top 1 percent income shares (Roine and Waldenström, 2011; Piketty, 2014).

To understand the overall effects of a shock in the U.S. series on the series of other countries, we compare the total accumulated impulse responses of the GMVAR model. In regime 1, the total accumulated effect is between 1 (in Australia) and 3 (in Japan). In regime 2, it is between 4 (in Australia) and 15 (in Japan). Accordingly, the impulse response analysis indicates that changes in income inequality of the U.S. have a persistent effect on the inequality of the other countries in our sample in the high-inequality, high-income-fluctuations regime. The effect is considerably smaller in the low-inequality, low-income-fluctuations regime.

To summarize, the U.S. income inequality series seems to affect the autoregressive dynamics of inequality (see Section 3.2) and lead regime changes (see Section 3.2.2) across our sample. In addition, a shock to the U.S. top 1 percent income share series seems to have a persistent effect on the level of top 1 percent income shares of other countries in the high-inequality, high-income-fluctuations regime. All these findings based on the estimated GMVAR model point to the conclusion that the income inequality of the U.S. has been the driver of changes in the income inequality of other developed economies over the past 100 years. This raises questions on the possible channels of the effect as well as on the policy implications, to which we turn next.

4. Discussion

Our results have three rather drastic implications. First, a shift from the first regime into the second regime indicates both a fall in the mean income and an increase in the uncertainty (variance) of future income for the bottom 99 percent.[14] Mean income will fall, because the increase in the mean income share of

[14]The top 1 percent income share has a high correlation with broader measures of income inequality (Atkinson *et al.*, 2011; Leigh, 2007), but it seems to have the biggest impact on the earnings of middle-income families (Thompson and Leight, 2013).

the top 1 percent is greater than any conceivable short- to medium-term GDP growth. Thus, the high-inequality regime is much more harmful for the bottom 99 percent of income earners. For the 1 percent, however, there is a tradeoff; in the second regime, they receive more income, but with a greater risk than in the first regime. The welfare implications may thus be beneficial for the 1 percent, if there is an overall improvement of their relative incomes, but also negative, if the increase in risk offsets the possible gains in the relative income.

Second, Malinen (2012) and Herzer and Vollmer (2013) have found that the stochastic parts of income inequality and the GDP per capita have a long-run *equilibrium* relationship. This indicates that larger stochastic fluctuations in the top 1 percent income share in the second regime translate to larger stochastic fluctuations in the GDP per capita, creating macroeconomic instability. This finding is supported by Berg and Ostry (2011), who find that higher inequality is associated with shorter growth spells and vice versa.

Third, the level of inequality in the U.S. was found to directly affect the future level of inequality in our sample of developed countries. This level effect is also visible in how the regime changes occur: the U.S. has been leading the regime change in our sample in the 1950s and especially at the end of the 1980s (see Appendix B). In addition, in the high-inequality, high-income-fluctuations regime, the changes in the level of inequality in the U.S. are transmitted to all other countries through the covariance structure of that regime. This observed dynamic dependence between the level and the changes of inequality in the U.S. is likely to diminish the control that individual countries have on their distribution of income.

These empirical findings naturally raise two important questions: what are the driving forces of regime switches and, more importantly, what is the role of the U.S. in these forces? Acemoglu and Robinson (2015) show that major changes in the top 1 percent income shares in Sweden and in South Africa have been associated with changes in economic and political institutions. In the U.S., institutional changes have also been associated with changes in the distribution of product. Campbell and Allen (2001) analyzed the average tax rate, the progressivity of taxation, and the population covered by taxes. They found four regimes describing the tax policies in the U.S. between 1916 and 1986:

1. 1916–17, 1923–33 (symbolic)
2. 1918–22, 1934–40 (fiscal crisis)
3. 1941–53 (war making)
4. 1954–85 (macroeconomic stability)

Regimes from 1916 to 1941 were characterized by a low average tax rate and a low coverage of population subjected to federal taxes. In the "war making" regime, the degree of progressivity of taxation, the average tax rate, and the tax coverage were increased dramatically. The "macroeconomic stability" regime was characterized by high progressivity, a high average tax rate, and a high coverage of taxation. This regime coincides with the low-inequality, low-income-fluctuations regime found in the top 1 percent income share series of the U.S. (see Section 3.1) and other developed economies (see Section 3.2). Did the U.S. drive the institutional change in other developed economies during the onset of this regime? This does not seem likely, because many countries adopted progressive income taxation and raised other taxes due to the war-related increase in government expenditures already in the 1940s (Galvin, 1981; Reinhardt and Steel, 2006; Kaneko, 2009). The U.S. had also a smaller effect on the regime change in countries in our sample in the 1950s (see Section 3.2.2). Liberal tax-lowering policies adopted under President Reagan's administration were likely to have a bigger influence on the resurgence of the high-inequality, high-income-fluctuations regime in developed economies at the end of the 1980s. During that time, almost all Organisation for Economic Co-operation and Development (OECD) countries limited the number of personal income tax brackets and lowered their top statutory tax rates (Torres *et al.*, 2012). Still, although the institutional change toward lower taxes originated from the U.S. in the 1980s, it cannot comprehensively explain the observed effect that the U.S. had on the income inequality in other developed economies of our sample during that period. Inequality in the U.S., for example, had the biggest effect on the levels of top 1 percent shares in Japan and in Canada (see Section 5), but they pursued rather different tax policies than the U.S. in the 1970s and 1980s. During the tax reform of 1986–8, Japan adopted value added tax (VAT) and raised several other taxes (Kaneko 2009). Canada had already adopted more liberal income taxation in the 1970s, before similar measures were introduced in the U.S. during President Reagan's administration (Galvin, 1981; Jacob, 1985). Most importantly, the adoption of changes in taxation and other institutions relevant to income inequality usually takes longer than the 1 year that was the estimated lag length for the effect of the U.S. on the inequality of other countries (see Sections 3.2 and 3.2.2). Therefore, the observed effect of the U.S. on the dynamics of income inequality in our sample also needs to arise due to factors beyond institutional change.

Galbraith and Rossi (2016) find that U.S. dollar (USD) exchange rates are a powerful predictor of industrial pay inequality in both developed and developing countries in the short run. This indicates, as the authors note, that inequality is driven more by global financial and macroeconomic factors than local policies controlling, for example, the adoption of technology and education.[15] According to Piketty (2014), high income inequality in developed economies before World War II was mostly due to the larger share of income obtained from concentrated capital. Fluctuations in dividends and stocks added volatility in the share of income going to the top income earners. After World War I, the global capital markets became highly integrated within a relatively short period of time (Obstfeld and Taylor, 1997), and by the 1920s the U.S. had accumulated the largest pool of private and public capital in the world (Bolt and van Zanden, 2013; Piketty, 2014). Between 1924 and 1931, the U.S. also provided some 60 percent of global private lending (Crafts and Fearon, 2010). In other words, the U.S. became the dominant power in capital markets after World War I. The effect of the U.S. on the global capital markets was multiplied during the Great Depression, which began in the U.S. and spread through the developed world. In the 1980s, the U.S. began to liberalize its financial sector, which led to a wave of financial liberalization in other developed economies (Jacob, 1985; Stiglitz, 2004). This increased the share of

[15]Furceri and Loungani (2015) also find that financial globalization leads to increasing inequality in developed economies.

private capital to income, but the renewed increase in income inequality in developed economies was mostly caused by the rise in high wages. Two thirds of the increase of income inequality that occurred in the U.S. after the mid-1970s is attributable to the increase in wages of the top 1 percent income earners, especially in the wages of top managers (Piketty, 2014). This aggravated income inequality in other developed economies, because the wages of top managers in Europe (and elsewhere) need to keep up with the wages in the U.S. (Petit, 2010; Gerakos *et al.*, 2013). Top managers in the U.S. also exported higher salaries to other countries when they took up job offers around the world. The high volatility of incentives, bonuses, and option prices (mostly through stock market fluctuations) of the top managers has added to the increase in fluctuations of the top incomes during past few decades (Gottschalk and Moffitt, 2009; Piketty, 2014).

We tested these possible channels by replacing the top 1 percent income share of the U.S. in the GMVAR model (see Section 5) with an index that combines the wage share of the top 1 percent income earners in the U.S. and the nominal S&P 500 stock market prices as a share of GDP.[16] Figure 4 presents the top 1 percent income share of wages, the S&P 500 stock market prices as a share of GDP, their combined index, and the top 1 percent income share of the U.S. It shows that the index follows the U.S. top 1 percent income shares rather closely.

The value of the index rises steeply in the 1920s, falls in the 1940s, and starts to rise again in the 1980s.[17] While both capital gains and top wages seem to lead the index in the beginning of the sample, the surge in the index after the 1970s is mostly attributable to the rise in the wage share of the top 1 percent income earners.

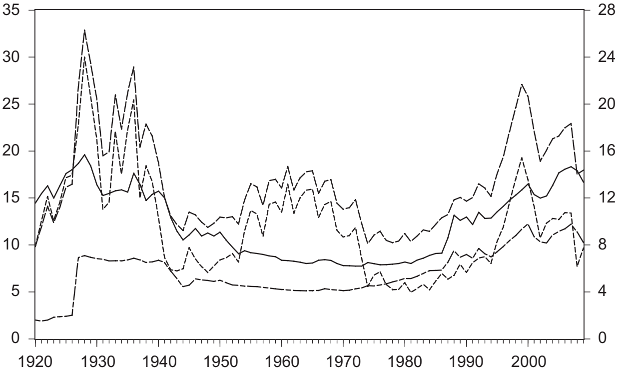

Figure 4. The Top 1 Percent Wage Share (Dashed/Dotted Line, Right-Hand Axis), the Share of the S&P 500 Stock Market Index to GDP (Dotted Line, Left-Hand Axis), Their Combined Index (Dashed Line, Right-Hand Axis), and the Top 1% Income Share (Solid Line, Right-Hand Axis) in the U.S. from 1920 to 2009

[16]Data on wages were obtained from Piketty (2014) and data on stock prices from Schularik and Taylor (2012).

[17]We also tested how well the combined index and the top 1 percent income share of the U.S. explain variations in each other. We found that the combined index has explanatory power on the changes in the top 1 percent income share of the U.S., but not the other way around, indicating causality running from the index to the incomes of the top 1 percent income earners in the U.S. The results are available upon request.

The results of the estimations, where the index replaces the top 1 percent income share of the U.S., are presented in Appendix C. According to the results, the index affects the autoregressive dynamics of all countries in a way that is strikingly similar to the top 1 percent income share of the U.S. (see Section 3.2). Although the magnitudes of the lag coefficients are somewhat smaller than with the top 1 percent income share, they are still statistically significant at the 5 percent level in all the same countries as with the top 1 percent income share of the U.S. For Australia and Canada, the dependence of their top 1 percent income share through the shock structure (covariance matrices) with the U.S. index is even stronger than with the U.S. top 1 percent income share series. These results indicate that developments in the capital and labor markets are the channels through which changes in the U.S. income inequality are transmitted to countries across our sample.

To summarize the discussion, it is likely that the economic and institutional change in labor and capital markets originating from the U.S. as well as globalization have been important factors behind the observed regime switches in the income inequality of developed economies. High income inequality, on the other hand, seems to have contributed to higher variance in the share of income of the top earners through two interlinked channels. First, periods of high income inequality have been associated with periods of concentrated capital (Piketty, 2014). Because financial capital has been an integral part of concentrated capital accumulation, the higher share of the volatile income from capital has increased the volatility of income of the top 1 percent. Second, during the latest era of globalization, the price fluctuations of incentives, bonuses, and options received by top managers have caused additional fluctuations in the top 1 percent income share series (Piketty, 2014). Many of these developments have originated from the U.S., which has enhanced the influence of the U.S. on the income inequality of other developed economies during the past 100 years. The U.S. has also been an influential advocate of globalization, possessing both the greatest incentives and capacity to advance it, especially after World War II (Moon, 1998). However, the U.S. was already in an elevated position as the dominant industrial power at the beginning of the 20th century (Baldwin and Martin, 1999). Therefore, the U.S. was a dominant global power during both waves of globalization in the 20th century, which were also marked by increasing income inequality (Baldwin and Martin, 1999; Milanovic, 2016).

5. Conclusions

In his path-breaking book, Piketty (2014) shows that income inequality has followed a U-shaped path in many developed economies, instead of the inverted-U-shaped path hypothesized by Kuznets (1955). The results presented in this paper add to this finding by showing that the level of inequality may determine the characteristics of income distribution in a similar manner to inflation: it can be either equal and stable or unequal and volatile. According to our results, all the countries in our sample currently reside in the high-inequality, high-income-fluctuations regime. Moreover, the results also indicate that changes in the dynamics of the income inequality of developed economies are affected by changes in the income inequality in the U.S.

As they stand, our results yield some unpleasant policy implications. Because an increase in the mean share of the top 1 percent income earners in the high-inequality, high-income-fluctuations regime is higher than any conceivable short- to medium-term growth of GDP, a shift to this regime is harmful for the bottom 99 percent of income earners. Larger fluctuations in the top 1 percent income share in the high-inequality, high-income-fluctuations regime also translate to larger stochastic fluctuation in the GDP per capita, because the stochastic parts of income inequality and GDP per capita have been found to have a joint *equilibrium* relationship (Malinen 2012; Herzer and Vollmer, 2013). This combination makes poor and middle-income households bearers of the costs of income inequality in more than one way: increasing income inequality lowers their share of the total income disproportionately and increases the uncertainty of their future income.[18] The attempts by sovereign nations to reduce the costs associated with income inequality may also be diminished by the dependence of their inequality on that in the U.S.[19]

The U.S. has been the leader of the capitalist world since the beginning of the 20th century. According to our results, this could also hold for the dynamics of income distribution, which seem to be more integrated across developed economies than previously thought. Our results indicate that changes in high wages and capital gains in the U.S. affect income inequality across our sample. The explanation for this could be that many of the major changes in the capital and labor markets have originated from the U.S. Moreover, globalization in the form of integration of capital and job markets has been likely to contribute to the convergent increases in income inequality in developed economies between the world wars and during the past few decades. Analysis of the possible causal channels of the interdependence of income inequality between developed countries should guide future research. The leading role of the U.S. may also be changing due to the rise of China and the developing countries. Nonetheless, because of the continuing integration of the world economy, it is likely that the dynamics of income inequality are destined to become even more interrelated between the developed economies, or globally, in the future.

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[18]In addition, the inequality may enforce itself in the high-inequality, high-income-fluctuations regime because stronger business cycle fluctuations can exacerbate income inequality (Ashley, 2007; Fawaz *et al*., 2012).

[19]See also Galbraith and Rossi (2016).

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Supporting Information

Additional Supporting Information may be found in the online version of this paper at the publishers website:

Appendix I: The Univariate and the Multivariate GMAR Models

The univariate GMAR model

The multivariate GMVAR model

Appendix II: Details of the regime change

Appendix III: Index series replaces US top 1% income share series